The Heterogeneous Effect of Diversity: Ascriptive Identity Cleavages and Redistributive Outcomes in Developed Democracies

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Abstract

The current consensus among comparative political scientists postulates that diverse democracies redistribute less than homogeneous ones. However, whereas homogeneous democracies do redistribute more on average, heterogeneous democracies exhibit high variation in redistributive outcomes. Why does ascriptive heterogeneity stifle redistribution in some cases but not in others? I argue that heterogeneity undermines redistribution when identity groups differ in their income levels and when democratic institutions allocate more policymaking power to dominant groups. Importantly, the policy outcomes are nonuniform: under these conditions, stronger groups undercut transfers for poorer, minority-heavy groups, but keep their own redistributive benefits protected and more exclusive. I find empirical support for these claims using cross-sectional time-series data from 19–22 developed democracies in the years 1980–2011. My paper thus outlines a more nuanced account of redistribution under diversity and, further, demonstrates that the policy implications of identity-related preferences are best analyzed together with class and political institutions.

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1 Introduction

How do ascriptive identity cleavages—ethnicity, race, religion, and language—shape redistributive policies in developed democracies? The growing attention to this question by comparative political economists, and particularly the influential work by Alesina and Glaeser (2004), cemented the notion that diverse countries redistribute less than homogeneous ones. Yet, although widely accepted, this straightforward theoretical prediction has mixed empirical support, particularly in developed democracies (Mau & Burkhardt, 2009; Pontusson, 2006; Steele, 2016; Taylor-Gooby, 2005). Moreover, newer research on individual-level preferences finds that the negative relationship between heterogeneity and support for redistribution depends upon additional demographic and socioeconomic factors (e.g., Alt & Iversen, 2017; Burgoon, 2014; Dahlberg, Edmark, & Lundqvist, 2012; Finseraas, 2012).

Recent data confirm this empirical tension. Figure 1 plots three different measures of redistribution levels against an index of ascriptive identity fractionalization in 19–22 OECD countries. Figure 1.A displays public social spending in 2011 as a share of gross domestic product (GDP), Figure 1.B presents the relative reduction in income inequality by tax and transfers in 2011, and Figure 1.C showcases the combined generosity of key social security programs in 2010 (Scruggs, 2014). The fitted lines indeed show negative correlations between ascriptive heterogeneity and all three aspects of redistribution. On average, as the literature expects, homogeneous countries spend more than heterogeneous ones on social programs, reduce a greater share of inequality, and offer more generous social security entitlements. However, the plots also reveal a second pattern: whereas homogeneous countries tend to cluster together more closely, heterogeneous ones scatter broadly. This tendency is illustrated with the matching box-and-whisker diagrams, which split the sample in half and plot the variation within each subgroup. In all three measures, the heterogeneous subsample stretches wide across the redistributive spectrum. Thus, the accepted notion that heterogeneity undermines redistribution seems justified but insufficient: higher heterogeneity in ascriptive identities influences redis-

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1 The Ascriptive Identity Fractionalization index integrates multiple data sources and identity dimensions. Its operationalization is expanded upon in later sections.
tributive outcomes, but not in all cases. Why is this so?

Figure 1: Ascriptive Heterogeneity and Redistribution

This paper seeks to address this gap and improve our understanding of the conditions under which ascriptive heterogeneity shapes redistributive outcomes. Building upon previous research on redistributive preferences, I argue that diverse democracies redistribute less when some ascriptive identity groups are richer than others, leading to lower willingness to share resources with the latter. Importantly, the redistributive implications are not uniform. As stronger identity groups seek to minimize intergroup redistribution, the negative implications center primarily on programs targeting poorer groups and on the latter’s access to public benefits. In contrast, stronger groups protect broader redistributive programs serving their members too and make them more exclusive. The extent to which these group interests translate into policy outcomes also depends on policymaking institutions. When governmental systems concentrate more policymaking power in the hands of dominant groups, the reinforcement of identity and class cleavages has a greater dampening effect on redistribution. Conversely, systems with more diffuse policymaking power and higher minority representation moderate the ability of rich identity groups to disregard the interests of the former. Thus, rather than considering as-
criptive diversity as an independent mechanism, I argue that it operates together with class and political institutions.

I find support in this theory using macrocomparative data on 22 developed democracies in the years 1980–2011. Specifically, I demonstrate that the negative effect of ascriptive heterogeneity on redistributive outcomes is mediated by starker income differences between identity groups and moderated by institutional power-sharing. Moreover, examination of multiple redistribution indicators reveals differential outcomes. An overlap of identity and class in heterogeneous countries implicate social assistance and unemployment programs, both covering risks characterizing economically-weaker populations, but not programs serving broader needs such as old-age and sickness. I also find that key social security programs are less inclusive but not less generous, implying higher differentiation between strong and weak recipients rather than crude cutbacks for everyone. Therein lies the answer to the empirical puzzle: heterogeneous countries tend to redistribute less than homogeneous ones on average, but this tendency varies by the overlap of identity and class, the target audiences of specific programs, and the level of policymaking power held by stronger groups. The paper concludes with several broader implications for the study of identity politics and social policy.

2 Existing Research on Identity and Redistribution

Social cleavages, broadly defined as salient division lines between different groups in society, have long been central in political science. Ascriptive identities—ethnicity, race, religion, and language—are a particularly potent type of cleavage given their inherent, indivisible, and relatively rigid nature (Lipset & Rokkan, 1967; Powell, Jr., 1982; Rae & Taylor, 1970), their central role in in-group coordination (Bates, 1983; Fearon & Laitin, 1996), and their mobilization by political actors (Chandra, 2004; Posner, 2004). In the past few decades, researchers demonstrated that ascriptive identity cleavages influence economic development and democratization (Easterly & Levine, 1997; Houle, 2015; Montalvo & Reynal-Querol, 2005), inter-group conflict (Fearon & Laitin, 2003; Wilkinson, 2008), public goods provision (Alesina,
Baqir, & Easterly, 1999; Habyarimana, Humphreys, Posner, & Weinstein, 2007), party systems (Clark & Golder, 2006; Ordeshook & Shvetsova, 1994), and voter behavior (Chandra, 2004; J. D. Huber, 2012). Identity politics were slower to penetrate the comparative study of redistributive politics, however. Mainstream studies of developed welfare states focus predominantly on class as the primary, and often only, relevant social cleavage. Redistributive outcomes are linked most commonly with the income distribution of otherwise-identical voters (Meltzer & Richard, 1981; Moene & Wallerstein, 2001; T. Romer, 1975), the political power of the left (Esping-Andersen, 1990; Korpi, 1983; Korpi & Palme, 2003; Stephens, 1979), and the institutional constrains on class-based actors (Hicks & Swank, 1992; E. Huber & Stephens, 2001; Iversen & Soskice, 2006). Ascriptive identity cleavages play a minor or no role in these theories.

In recent years, nonetheless, the redistributive implications of ascriptive diversity has drawn growing scholarly attention. The most common argument in this literature suggests that diverse societies redistribute less than homogeneous ones. Higher heterogeneity in ascriptive identities, the logic goes, fractures social ties, weakens interclass solidarity, and exacerbates collective action problems. Empirical support in this argument draws predominantly from the US, where the salient racial divide is linked repeatedly with reduced public services and social programs in all levels of government (Alesina et al., 1999; Quadagno, 1994; Ribar & Wilhelm, 1999; Skocpol, 1992). The most comprehensive articulation of this argument from a macrocomparative perspective is offered by Alesina and Glaeser (2004; see also Desmet, Ortuño-Ortín, & Weber, 2009; Sanderson, 2004), who maintain that higher ascriptive heterogeneity explains much of the historic gap between the American and European welfare states. They further argue that ascriptive cleavages precede and explain the emergence of class politics and electoral institutions, the primary culprits by previous research. Empirically, Alesina and Glaeser find

There is a small handful of references to ascriptive identity cleavages in this literature, yet mostly as a marginal factor operating indirectly. Stephens (1979), for instance, notes that ethnic and linguistic heterogeneity undermine the cohesion of labor organizations. Similarly, J. E. Romer, Lee, and Van der Straeten (2007) argue that anti-immigration sentiments strengthen right-wing parties and, therefore, lead to lower redistribution indirectly. Some attention was also given to the effect of religion on welfare states, primarily through Christian-Democratic parties (E. Huber, Ragn, & Stephens, 1993; Van Kersbergen & Manow, 2009) and religious networks (J. D. Huber & Stanig, 2011; Scheve & Stasavage, 2006). However, this discussion analyzes religion, namely Christianity, not as a contentious social cleavage but as a defined constituency and set of values.
that racial and ethno-linguistic diversity have a negative bivariate correlation with public social spending cross-nationally. The notion that ascriptive heterogeneity weakens redistribution has since become a common premise in comparative politics research.

Although highly influential, this argument was challenged both theoretically and empirically (for a comprehensive critique, see Pontusson, 2006). Taylor-Gooby (2005) and Mau and Burkhardt (2009), for example, argue that the negative correlation loses its significance when the sample is limited to Western countries and with a more careful statistical control for economic, political, and demographic differences. Furthermore, newer studies find that various individual and social factors condition personal preferences for redistribution in diverse societies (Steele, 2016). In particular, researches found that support in redistribution decreases most strongly when minorities are poorer or perceived as such, a possibility acknowledged in passing by Alesina and Glaeser themselves (2004, p. 134). These studies find that when identity groups differ more starkly in their income, intergroup affinity weakens (Lupu & Pontusson, 2011), social rivalry exacerbates (Corneo & Grüner, 2002; Shayo, 2009), and labor-market risks drift further apart (Alt & Iversen, 2017), all leading to lower public support in redistribution.

This type of explanation has been tested empirically in two often-overlapping ways, both of which paint only part of the picture. First, many studies focus on individual preferences as the primary outcome of interest (e.g., Alt & Iversen, 2017; Brady & Finnigan, 2014; Burgoon, 2014; Finseraas, 2012; Stichnoth, 2012). Public preferences are necessary but insufficient for inferences on actual policy outcomes, however, particularly given the importance of political institutions and the gradual pace in which social policies change over time (E. Huber & Stephens, 2001; Pierson, 1996). Second, there is increased focus on recent immigration and its mostly-negative effect on redistributive preferences and policies. Immigration is one of several causes for ascriptive heterogeneity, but as such is insufficient for a fuller understanding of the latter. New immigrants integrate slowly, do not automatically constitute a coherent political constituency (Michon & Vermeulen, 2013), and face unique barriers to full political and economic rights, parliamentary representation, and access to state services (Bird, Saalfeld, & Wüst, 2011; Dancygier, 2010; Dancygier, Lindgren, Oskarsson, & Vernby, 2015). Accordingly,
whereas there are some signs of a gradual decline in social spending due to increased immigration (Soroka, Johnston, Kevins, Banting, & Kymlicka, 2016), the implications found in the literature concentrate primarily on immigrant-specific policies such as tighter immigration rules, revised integration policies, and stricter immigrant access to welfare programs (Hemerijck, Palm, Entenmann, & Van Hooren, 2013; Koning & Banting, 2013; Sainsbury, 2012). Existing theoretical and empirical accounts of the relationship between ascriptive heterogeneity writ large and redistributive outcomes remain incomplete. I now turn to redress these shortcoming theoretically and empirically.

3 Theoretical Framework

I begin with three simplifying premises on the nature of ascriptive identity cleavages. Whereas these assumptions sacrifice certain nuances on group identities, they also help to focus the theoretical discussion on their redistributive implications. First, I assume that relevant ascriptive categories have already been formed. Thus, this paper does not explore long-term processes of identity formation. Second, contrary to the division imposed by many empirical studies, I assume that different types of ascriptive identities—ethnicity, race, religion, and language—have similar political implications once salient. All types share an alleged common genetic, historic, cultural, or spiritual decent, have relatively rigid and visible criteria, and improve social coordination (Chandra, 2006; Hale, 2004; Haller & Eder, 2015; Laitin, 2007). The type of ascriptive identities activated in specific cases is rather contingent upon historic processes (Wimmer, 2008) and can lead to comparable intergroup tensions, as many Americans (race), Irish (religion), Belgians (language), and Israelis (ethnicity) can attest. Third, for similar reasons, I also assume that ascriptive identities, as a unified category, do differ inherently from other politically-relevant classifications (e.g., gender, education, geography), particularly in

3 A comprehensive body of work demonstrates that ascriptive identities can be constructed endogenously by economic development, state borders, electoral institutions, and violent conflicts (e.g., Chandra, 2004; Laitin, 1986; Posner, 2005; Sambanis & Shayo, 2013). These insights, however, are of lower importance for contemporary developed democracies, where the primary social cleavages have largely been shaped decades and centuries ago. Constructivist arguments are “about the long-run formation, and the consequent stickiness, of identities” (Varshney, 2009, p. 288).
their cognitive availability and mobilizing power.\footnote{These inherent differences too, of course, depend in part on human perceptions (e.g., that one’s ethnicity implies shared fate) and may change over time, albeit very slowly. For example, in past centuries, class was largely seen as an inborn trait passed down through generations. Modernization, the rise of the middle class, and increased social mobility have gradually relaxed this perception in developed economies (Deutsch, 1961).}

3.1 Group Preferences and Policy

Ascriptively diverse democracies, I argue, do face stronger pressures against broad redistribution compared to homogeneous ones. However, these forces depend on each country’s class and institutional structures. The first step is the alignment of identity and class. Classic theories in both social psychology and political sociology argue that group identities grow more salient when multiple cleavages are reinforced, i.e., when the same individuals are grouped together by several social criteria (Brewer, 2000; Deschamps & Doise, 1978; Lipset & Rokkan, 1967). An overlap of ascriptive identities and class—that is, when some identity groups are also distinctly poorer than others—seems particularly relevant for redistribution, a process of cross-class income transfer. This expectation draws from several micro-level mechanisms, most of which explored in recent studies on redistributive preferences. First, higher inequality between identity groups increases their perceived social and cultural distance, leading in turn to decreased solidarity (Lupu & Pontusson, 2011) and to stronger stigmatization of poor minorities (Gilens, 1995; Kinder & Sears, 1981; Nelson, 1999). Second, an overlap of identity and class increases social rivalry: stronger identity groups oppose policies that may mobilize poorer identity groups into their communities and undermine their relative social status (Corneo & Grüner, 2002; Shayo, 2009). Third, given increasingly segmented labor markets, reinforcement of identity and class separates each group’s occupational risks and, accordingly, their redistributive interests (Alt & Iversen, 2017). Overlap of identity and class, then, both increases the salience of group identities and decreases intergroup solidarity. Indeed, the cited studies find lower general support in redistribution under such conditions.

The shift from simple preferences to policy implications, however, is more nuanced. Redistribution is not a unidimensional choice between more or less. To see why, it is useful
to consider redistributive policies as protection against various socioeconomic risks (Bonoli, 2005; Dryzek & Goodin, 1986; Rehm, 2016). The typical risks covered by social programs are distributed differently across classes. Some risks, like aging-related complications or sickness, have similar likelihood regardless of one’s income. Others, like sustained unemployment or material deprivation, are concentrated unequally among poorer classes with lower savings and skills. When some ascriptive identity groups are richer than others, their increased in-group bias should undermine only the latter type of redistribution. Shared risks, in contrast, benefit middle- and high-class groups too. Therefore, when identity and class overlap, I expect stronger identity groups to minimize redistribution to poorer ones, but also protect programs that cover broader risks relevant for their members.

This intuition is illustrated in Figure 2, which plots several hypothetical identity/class cleavage structures in a society comprising two ascriptive identity groups. Panel A plots perfectly cross-cutting cleavages, such that the two groups share an identical income distribution. Even if the majority cares only for the interests of its members, limiting any redistributive program will hurt them equally. This is not the case when identity and class overlap, as panel B demonstrates. Here, the richer identity group benefits nothing from redistributive policies targeting the poor. At the same time, its interest in programs covering interclass risks remains firm. As Panel C illustrates, the key mechanism requires a combination of high heterogeneity and class overlap. In relatively homogeneous societies with a significantly small minority, an overlap of identity and class does not crowd out the majority from the lower class and, accordingly, retains their interest in all aspects of redistribution.

For both normative and legal reasons, policies in developed democracies are unlikely to discriminate among recipients explicitly based on their ascriptive identities. When identity and class overlap, however, richer identity groups can minimize redistribution to poorer ones through two policy channels. The first, and straightforward, way is to cut down programs that

5 A perfect overlap of identity and income can also reflect the opposite case, whereby a minority is significantly richer than a poorer majority. For simplicity, I assume that strictly richer identity groups have significant economic influence even if they are smaller numerically. Importantly, it is hard to find many such instances in contemporary developed democracies. Examples of privileged minorities in developing settings, typically the result of colonialism, do indicate strong protection of their economic interests regardless of size (e.g., whites in Brazil or South Africa, ethnic Russians in soviet republics, ethnic Chinese in various Southeast Asian countries).
specifically target poorer, and hence minority-heavy, classes. At the same time, programs covering broader risks are prioritized instead. The second policy channel decreases the access of poorer minorities to broader programs. In particular, poorer classes can be excluded effectively in the labor market, where many developed economies experience growing segmentation between different tiers of workers. As studies emphasize, there is increasing divergence of interests between so-called “insiders,” workers with relatively secure jobs and strong political capital, and “outsiders,” low-skilled and vulnerable workers, often minorities and immigrants (Emmenegger, Häusermann, Palier, & Seeleib-Kaiser, 2012; Iversen & Soskice, 2015; Rueda, 2005). Thus, in addition to lower redistribution to the poor, an overlap of identity and income should also increase the exclusion of outsiders from labor-market protection.

Figure 2: Income Distribution under Different Cleavage Structures
3.2 Institutional Constraints

The discussion thus far assumes automatic translation of group preferences to policy. It is safe to assume that richer identity groups hold more political resources and influence than poorer ones. Nevertheless, their representatives are still constrained by democratic institutions, which vary significantly across different countries. In particular, the ability of stronger groups to realize their preferences depends on the level of institutional power-sharing with weaker groups. Quite intuitively, broader power-sharing in government amplifies the voice and influence of weaker groups and places more curbs on stronger ones. Therefore, higher power-sharing moderates the extent to which stronger groups can limit de facto redistribution to poorer ones. Conversely, higher concentration of power enables them to act on their group preferences more fully.

Two institutional nexuses affect power-sharing in government: the distribution of policymaking influence and the proportionality of representation (Lijphart, 1999; Powell, Jr., 2000; Tsebelis, 1995). First, policymaking is more constrained when more political actors have institutional veto power. A higher number of veto points forces policymakers to reach broader compromises that satisfy a wider set of preferences. Moreover, a higher number of institutional veto points have been specifically linked with lower retrenchment in redistributive programs (E. Huber & Stephens, 2001). Second, minority interests are expressed better when political representation is more proportional. Proportional electoral rules, as opposed to majoritarian systems, translate votes to seats more accurately, facilitate a higher number of political parties, and create more diverse and left-leaning coalitions in government (Bowler, Donovan, & Brockington, 2003; Clark & Golder, 2006; Iversen & Soskice, 2006; Shugart, 1994).

In sum, I expect that ascriptively diverse democracies redistribute less than homogeneous ones, but only as a function of certain structural conditions: (1) the overlap between ascriptive identities and class, (2) their target audiences, and (3) the level of institutional power-sharing in government. Figure 3 summarizes the argument’s implications.

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6 For simplicity, I assume that reelection-seeking politicians always promote the interests of identity groups within their electoral coalitions.
4 Data and Empirical Strategy

I test my theory using a series of models estimating how ascriptive heterogeneity, its overlap with class, and political power-sharing affect various aspects of redistribution in developed democracies. I use cross-sectional time-series data form the years 1980–2011 for 19–22 countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Israel, Italy, Japan, The Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, the United Kingdom, and the United States. The panel data are unbalanced, i.e., not all countries have available data for all years and models.

Figure 3: Summary of Theoretical Implications

Dampening effect:
1. Lower redist. to poorer classes
2. Lower redist. inclusiveness
4.1 Variables of Interest

4.1.1 Dependent Variables

The primary outcome of interest, redistribution, is examined from multiple perspectives. The first and simplest aspect is aggregate redistribution levels, measured by two complementary variables: (1) public social spending as a share of GDP, using data from the OECD Social Expenditure Database, and (2) reduction in income inequality before and after taxes and transfers, measured as the share of change in the Gini coefficient of household income, using Luxembourg Income Survey (LIS) data (Wang & Caminada, 2011). The use of two separate measures of redistribution adds robustness to the findings and harnesses each measure’s respective strengths. Social-spending data are available on an annual basis and for more countries, are better standardized, and include both cash transfers and in-kind services. The inequality reduction measure, meanwhile, captures de facto policy implications better and is not sensitive to changes in GDP levels or recipient numbers. All else equal, I expect that a combination of higher heterogeneity and overlap with class will decrease both measures similarly.

The second aspect unpacks redistribution to subcomponents, again using both public social spending and inequality reduction data. I group disaggregated data from both sources into four categories based on their covered risks: (1) age-related benefits, consisting of old-age and survivor transfers and services; (2) unemployment benefits, consisting of unemployment compensation and active-labor programs; (3) incapacity and sickness benefits, consisting of sick-pay compensation, occupational injury transfers, and disability benefits; and (4) assistance benefits, consisting of income maintenance, housing assistance, family and child allowances, and similar in-kind benefits. All else equal, I expect that a combination of higher heterogeneity

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7 The LIS data do not include Japan, New Zealand, and Portugal. In addition, the LIS data are not annual and spaced unequally over time by country. When analyzing these data, I therefore follow Persson, Roland, and Tabellini (2007, p. 19) and Lupu and Pontusson (2011, p. 324) in averaging the values of annual independent variables for the period between every two country-observations.

8 My disaggregation leaves out several components that are irrelevant for my theory as separate categories. In the OECD expenditure data, I exclude public spending on health as a disaggregated category, as it consists primarily of public healthcare services. Whereas public healthcare merits an examination in itself, it is closer to a public good than to interpersonal transfer and may thus reflect different mechanisms. I also do not include the “other” category in the OECD data, since, as the name implies, its content changes by country. In the LIS data, I exclude military service and veteran benefits.
and overlap with class will decrease only unemployment and assistance programs, both serve lower classes more distinctively, but not the other two types.

The third aspect of redistribution involves inclusiveness in labor-market programs. I use data from the Comparative Welfare Entitlement Project (CWED2; Scruggs, Jahn, & Kuitto, 2014) on two key social security programs: unemployment and sick-pay insurance. For each program, I compare two measures: (1) coverage, measured as the share of labor force insured under each program, and (2) wage replacement rate, calculated against the mean of an average single worker’s wage and an average four-person family’s wage. All else equal, I expect that a combination of higher heterogeneity and overlap with class will decrease only the level of program coverage, but not replacement rates.

4.1.2 Independent Variables

The theoretical framework marks three explanatory factors: heterogeneity in ascriptive identities, overlap of ascriptive identity and class, and institutional power-sharing. In line with the current literature, I measure ascriptive heterogeneity using one minus the Herfindahl index, which estimates the level of social fractionalization. There has been significant improvement in the quantity and quality of ascriptive fractionalization indices in recent years. Nonetheless, these indices pose two problems. First, the multitude of indices and sources raises the risk of cherry-picking. Second, contrary to my theoretical assumptions, available indices calculate separate scores for different ascriptive-identity types, typically ethnicity, religion, and/or language. This separation makes it difficult to examine comparable processes across equally-divided countries differing only in the type of salient identities.

To deal with both problems, I calculate an Ascriptive Identity Fractionalization (AIF) index that integrates multiple sources and identity types into a single country-score. The combination of several sources increases reliability and can gauge cleavage salience better by adding weight

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9 The CWED2 data do not include Israel.
10 The index is calculated as $F = 1 - \sum_{i=1}^{G} p_i^2$, where $p_i$ is the relative share of group $i$ in the general population and $G$ is the total number of groups.
to cleavages identified repeatedly. I use four databases, each relying on a different type of primary source: (1) three indices of ethnic, linguistic, and religious fractionalization based on encyclopedic sources (Alesina, Devleeschauwer, Easterly, Kurlat, & Wacziarg, 2003); (2) the Ethnic Power Relations’ index of ethnic fractionalization in politically-relevant groups based on expert surveys (Cederman, Wimmer, & Min, 2010); (3) the Cross-cutting Cleavages Dataset’s two indices of ethnic and religious fractionalization based on survey data (Selway, 2011); and (4) an index of linguistic fractionalization based on a genealogical linguistic tree analysis (Desmet, Ortúñ-Ortíñ, & Wacziarg, 2012). Per my theoretical premises, I assign equal weight to each identity type: I first average across all indices of a particular dimension, creating separate fractionalization scores for ethnicity, religion, and language, and then average again across all three identity types to produce a single AIF score per country. Figure 4 summarizes the index’s structure.

Due to data limitations, the AIF index is time-invariant, i.e., assigns a fixed score per country for the entire sample period. Although this is not ideal, ascriptive heterogeneity is considered quite stable in the literature, particularly over relatively short periods of 30 years or less (Alesina et al., 2003, p. 161). To corroborate this assumption, I recreated the AIF index using data from Patsiurko, Campbell, and Hall (2012), who calculate separate ethnic, religious, and linguistic fractionalization scores for 18 of the countries in my sample in 1985 and in 2000. The strong bivariate correlation between the two periods ($r = 0.93, p < 0.000$) indicates high

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*Figure 4: The Ascriptive Identity Fractionalization Index*

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11 Most indices count race under ethnicity, so it is not included separately. Although they rely on different primary sources, the different indices correlate strongly along their respective identity dimensions as expected.
stability over time.\footnote{There are, of course, growing immigration inflows to Western democracies in recent years. However, this should not destabilize the AIF scores significantly. First, immigrant populations arriving by the mid–2000s are included in the sources I use. Second, later immigration is relatively gradual: with few exceptions, annual immigration inflow rates leading to 2011 are less than 1% of the host population. Indeed, the bivariate correlation between the AIF index and foreign-born population shares in 2011 (OECD data) is relatively high ($r = 0.66$, $p < 0.005$). Third, newer immigrants are not absorbed immediately into the cleavage structure and the political system, and should thus get involved in policymaking quite slowly. Nevertheless, I reran all my models with an additional control for annual immigration inflows as a share of the population (OECD data). Despite a smaller sample size, my findings remain substantively unchanged.}

The second explanatory variable of interest is the overlap of ascriptive identities and class, measured using the Cross-cutting Cleavages Dataset created by Selway (2011). Selway aggregates data from various public opinion surveys to evaluate the distribution of group members in one cleavage across another. Two cleavages are cross-cutting perfectly if members of each group in the first cleavage distribute equally across all categories in the second cleavage. Conversely, two cleavages overlap perfectly if group composition replicates identically in both. Similar to the AIF measure, I average across two ascriptive identity scores in Selway’s data: cross-cuttingness of income and ethnicity and cross-cuttingness of income and religion.\footnote{Selway (2011) counts linguistic groups under ethnicity. Additionally, Greece has no data on ethnicity/income cross-cuttingness, but I nonetheless keep it in the sample using only its religion/income cross-cuttingness score. Ethnicity plays only a minor role in Greece: its ethnic fractionalization score is 0.076 compared to a sample mean of 0.22. As a robustness check, I omit Greece from all my models and find substantively unchanged results.} To align with my theoretical hypotheses, I invert the composite score to measure overlap.

Like the AIF index, the overlap scores are time-invariant due to data limitations. This constraint too should not pose a serious problem, as the level of overlap is expected to be both exogenous to redistribution and stable during the sample period. The surveys used to calculate the overlap score ask responders about their relative income bracket. Redistributive policies target inequality, i.e., decrease the gap between income levels and not their position relative to one another. It is, indeed, quite implausible that welfare systems make net contributors worse-off than net recipients. This premise does not imply that ascriptive identities and class are ossified indefinitely, only that socioeconomic structures change very slowly, if at all, due to prolonged processes involving investment in human capital, education, and infrastructure. Diagnostic and robustness tests, detailed below and in the paper’s supplemental material, corrobore both the exogeneity and stability of the overlap measure.\footnote{To test for exogeneity, I reran my models with an instrument for ethnic group inequality instead of overlap overlap scores.}
perceptions and stigmas, a central mechanism behind group preferences, change even slower.

The third and final explanatory variable is the level of institutional power-sharing, and specifically the number of veto points and level of electoral proportionality. Using data from Armingeon, Isler, Knöpfel, Weisstanner, and Engler (2016), the number of veto points are measured using an additive index with the following indicators: (1) degree of federalism, (2) presence of presidentialism, (3) degree of bicameralism, (4) use of referenda, (5) degree of majoritarianism, inverted to reflect proportionality, and (6) presence of judicial review. The resulting scale ranges from 0 to 8, where higher values indicate more veto points and thus lower concentration of power. Electoral proportionality is measured using Gallagher’s least-square index of electoral disproportionality, which captures the match between the share of party seats won in every election and the share of votes they received (Gallagher, 1991). To align with my theoretical framework, I multiply the scores by –1 to reflect proportionality. Descriptive values of all explanatory variables are presented in the supplemental material.

In addition to the primary variables of interest, I also control for several other explanatory factors associated in the literature with redistributive outcomes. My control variables include both political influences (cabinet partisanship and union power) and socioeconomic factors (unemployment rate, labor force participation, female participation in the labor force, the share of elderly population, logged trade, and logged GDP per capita). Table A1 summarizes the full list of control variables, their definitions, sources, and expected effects.

4.2 Model Specification

My empirical strategy consists of a series of single-equation error-correction models (ECM) using pooled regression analysis (Beck, 1991; Davidson, Hendry, Srba, & Yeo, 1978; De Boef & and found substantively similar results. To test for stability, I analyzed overlap scores for 97 comparable surveys—surveys conducted by the same project in the same country using the same questions, but in different years. I then calculated the annual difference rate between all comparable survey dyads. The results show random noise rather than a consistent pattern of temporal change. This analysis is explained in detail in the supplemental material.

I code Israel’s values using the same rules as Armingeon et al. (2016) as it is missing from the original dataset.
Due to indications of panel-specific heteroskedasticity, the estimation employs panel-corrected standard errors (PCSE, Beck & Katz, 1995, 2011). The ECM specification is particularly appropriate for redistribution data, known for their slow change over time and strong serial autocorrelation. This specification assumes that the outcome is in an equilibrium relationship with the explanatory variables, but that this relative stability can be disturbed by short-term shocks followed by a correction back to the long-term trend when the system adjusts. These dynamics are estimated by regressing changes in the dependent variable on the lagged values of all independent variables (long-term equilibrium relationship), the first difference of all dependent variables (short-term disturbances to the equilibrium), and the lagged value of the independent variable (the correction back to equilibrium). Since my primary explanatory variables are either time-invariant or slow-changing, I include them outside the error-correction dynamics. Their stable values are more appropriately interpreted as casting a constant, long-term influence on patterns of equilibrium and disturbances.

In formal terms, I estimate the following model structure:

$$\Delta R_t = \alpha + \beta_1 \Delta X_t + \gamma (R_{t-1} - \beta_2 X_{t-1}) + \beta_3 Z + \epsilon_t$$ (1)

where $\beta_1$ estimates the short-term effect of a vector of control variables $X$ on changes in redistribution level $R$, $\beta_2$ estimates the long-term effect of a one-unit increase in vector $X$, $\gamma$ is the error-correction term capturing the speed of adjustment back to equilibrium, and $\beta_3$ estimates the structural effect of vector $Z$ of time-invariant variables. My primary interest is in the latter effect. I interact the main explanatory variables to test their hypothesized conditionality.
(Brambor, Clark, & Golder, 2006; Kam & Franzese, 2007), first using a two-way interaction between AIF and overlap, and then a three-way interaction with each of the two institutional power-sharing measures.\(^{18}\)

## 5 Findings

### 5.1 Ascriptive Heterogeneity and Class

The first set of tests, presented in Table 1, estimate whether overlap with class mediates the effect of ascriptive heterogeneity on redistribution, at this point still in aggregate terms. As a point of reference, models 1 and 3 tests the direct effect of ascriptive heterogeneity on redistribution levels, as proposed by Alesina and Glaeser (2004). Contrary to their argument, I find no direct effect of ascriptive heterogeneity on neither social spending nor inequality reduction. Models 2 and 4, in contrast, support my conditional prediction. The negative interaction coefficients indicate that as the overlap of identity and class grows, ascriptive heterogeneity has an increasing negative effect on changes in both government spending and inequality reduction.

Figure 5 illustrates these patterns visually. The two graphs plot the estimated marginal effect of a change from complete homogeneity to complete heterogeneity given different levels of overlap with income. As expected, heterogeneity has a negative marginal effect on redistribution that grows as the two cleavages overlap more closely. Furthermore, the negative effect takes effect only past some minimal threshold of overlap. Interestingly, when overlap levels are very low, the spending model estimates that higher heterogeneity will slightly expand redistribution levels. This outcome may be explained by a race to the top between rival identity groups. The micro-mechanisms described earlier imply that deep identity cleavages should increase in-group bias. When all groups have strong identities but similar shares of poor members, increased group loyalties may thus create simultaneous motivations for higher

---

\(^{18}\) As a robustness check, I reran my models with the alternative bin-estimator approach for interactions suggested by Hainmueller, Mummolo, and Xu (2016). Where their code executes, their procedure supports my findings. Additional details are reported in the supplemental material.
<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>AIF</td>
<td>-0.241 (0.214)</td>
<td>4.389*** (1.144)</td>
<td>-0.062 (0.045)</td>
<td>0.389*** (0.126)</td>
</tr>
<tr>
<td>Overlap</td>
<td>18.495*** (4.712)</td>
<td>1.158** (0.567)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>AIF × Overlap</td>
<td>-45.767*** (11.898)</td>
<td>-4.066*** (1.339)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Veto</td>
<td>0.001 (0.018)</td>
<td>-0.001 (0.018)</td>
<td>-0.001 (0.004)</td>
<td>-0.003 (0.004)</td>
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</table>

**Short-Term Relationships**

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coef. (PCSE)</th>
<th>Coef. (PCSE)</th>
<th>Coef. (PCSE)</th>
<th>Coef. (PCSE)</th>
</tr>
</thead>
<tbody>
<tr>
<td>ΔLeft</td>
<td>0.001 (0.001)</td>
<td>0.000 (0.001)</td>
<td>-0.000 (0.000)</td>
<td>-0.000 (0.000)</td>
</tr>
<tr>
<td>ΔChristDem</td>
<td>0.008** (0.004)</td>
<td>0.009** (0.004)</td>
<td>0.001*** (0.000)</td>
<td>0.001*** (0.000)</td>
</tr>
<tr>
<td>ΔCWB</td>
<td>0.032 (0.050)</td>
<td>0.025 (0.050)</td>
<td>0.036 (0.008)</td>
<td>0.036 (0.008)</td>
</tr>
<tr>
<td>ΔUnemployment</td>
<td>0.135*** (0.037)</td>
<td>0.127*** (0.037)</td>
<td>0.044** (0.002)</td>
<td>0.044** (0.002)</td>
</tr>
<tr>
<td>ΔLabForce</td>
<td>-0.167*** (0.042)</td>
<td>-0.175*** (0.042)</td>
<td>-0.004 (0.005)</td>
<td>-0.004 (0.005)</td>
</tr>
<tr>
<td>ΔFemLabForce</td>
<td>0.329 (0.089)</td>
<td>0.309*** (0.087)</td>
<td>0.010 (0.010)</td>
<td>0.010 (0.010)</td>
</tr>
<tr>
<td>ΔElderly</td>
<td>0.188 (0.202)</td>
<td>0.443*** (0.218)</td>
<td>0.029*** (0.011)</td>
<td>0.033*** (0.012)</td>
</tr>
<tr>
<td>ΔLogTrade</td>
<td>-1.728*** (0.558)</td>
<td>-1.615*** (0.547)</td>
<td>0.043 (0.067)</td>
<td>0.027 (0.062)</td>
</tr>
<tr>
<td>ΔLogGDP</td>
<td>-17.400*** (2.013)</td>
<td>-17.854*** (1.983)</td>
<td>0.041 (0.092)</td>
<td>0.019 (0.089)</td>
</tr>
</tbody>
</table>

**Long-Term Relationships**

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coef. (PCSE)</th>
<th>Coef. (PCSE)</th>
<th>Coef. (PCSE)</th>
<th>Coef. (PCSE)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Left_{t-1}</td>
<td>0.000 (0.001)</td>
<td>0.000 (0.001)</td>
<td>0.000 (0.000)</td>
<td>0.000 (0.000)</td>
</tr>
<tr>
<td>ChristDem_{t-1}</td>
<td>0.001 (0.001)</td>
<td>0.003** (0.001)</td>
<td>0.001*** (0.000)</td>
<td>0.001*** (0.000)</td>
</tr>
<tr>
<td>CWB_{t-1}</td>
<td>0.052* (0.029)</td>
<td>0.045 (0.029)</td>
<td>0.007 (0.007)</td>
<td>0.007 (0.007)</td>
</tr>
<tr>
<td>Unemployment_{t-1}</td>
<td>-0.011 (0.010)</td>
<td>-0.024** (0.011)</td>
<td>0.002 (0.002)</td>
<td>0.002 (0.002)</td>
</tr>
<tr>
<td>LabForce_{t-1}</td>
<td>-0.004 (0.007)</td>
<td>-0.015* (0.008)</td>
<td>0.001 (0.001)</td>
<td>0.001 (0.001)</td>
</tr>
<tr>
<td>FemLabForce_{t-1}</td>
<td>0.026*** (0.010)</td>
<td>0.031*** (0.010)</td>
<td>0.008** (0.004)</td>
<td>0.008** (0.004)</td>
</tr>
<tr>
<td>Elderly_{t-1}</td>
<td>0.034** (0.016)</td>
<td>0.054*** (0.017)</td>
<td>0.001 (0.002)</td>
<td>-0.001 (0.002)</td>
</tr>
<tr>
<td>LogTrade_{t-1}</td>
<td>-0.012 (0.067)</td>
<td>-0.045 (0.066)</td>
<td>0.033** (0.016)</td>
<td>0.036** (0.016)</td>
</tr>
<tr>
<td>LogGDP_{t-1}</td>
<td>-0.099 (0.133)</td>
<td>-0.164 (0.131)</td>
<td>0.003 (0.025)</td>
<td>0.007 (0.024)</td>
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**Error-Correction Term**

<table>
<thead>
<tr>
<th>Variable</th>
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<th>Coef. (PCSE)</th>
<th>Coef. (PCSE)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Spending_{t-1}</td>
<td>-0.038*** (0.010)</td>
<td>-0.055*** (0.011)</td>
<td>-0.493*** (0.152)</td>
</tr>
<tr>
<td>Gini Reduction_{t-1}</td>
<td>-0.038*** (0.010)</td>
<td>-0.055*** (0.011)</td>
<td>-0.493*** (0.152)</td>
</tr>
</tbody>
</table>

**Lagged First-Difference**

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coef. (PCSE)</th>
<th>Coef. (PCSE)</th>
<th>Coef. (PCSE)</th>
</tr>
</thead>
<tbody>
<tr>
<td>ΔSpending_{t-1}</td>
<td>0.172*** (0.055)</td>
<td>0.157*** (0.054)</td>
<td></td>
</tr>
<tr>
<td>ΔLogGDP_{t-1}</td>
<td>9.078*** (1.840)</td>
<td>8.075*** (1.831)</td>
<td></td>
</tr>
<tr>
<td>Constant</td>
<td>0.861 (1.357)</td>
<td>0.447 (1.339)</td>
<td>-0.465** (0.200)</td>
</tr>
</tbody>
</table>

Observations: 627 627 188 188
Countries: 22 22 19 19
R²: 0.544 0.558 0.487 0.528

*p < 0.1, **p < 0.05, ***p < 0.01, panel-corrected standard errors in parentheses.

Table 1: The Interactive Effect of Identity and Class Overlap on Redistribution

redistribution for poor in-group peers, increasing overall redistribution in turn. A process in this spirit has been occurring in Belgium since the late-1990s, as both the federal and the regional Flemish governments expanded overlapping redistributive programs simultaneously—a process fueled to a large degree by Flemish motivation to establish a separate redistributive sys-
tem and by federal fortification of the national system in response (Béland & Lecours, 2008; Cantillon, 2011). This empirical pattern should not be overstated, however, as most heterogeneous countries have overlap levels near or above the threshold beyond which heterogeneity has a dampening effect on redistribution.

![Graphs](attachment:graphs.png)

**Figure 5:** The Marginal Effect of Ascriptive Heterogeneity on Overall Redistribution as a Function of Overlap

Most control variables perform as expected in Models 2 and 4. Christian-Democratic cabinets tend to increase redistribution both in the short and the long term, although left-leaning cabinets do not produce the expected positive effect. Interestingly, central wage bargaining loses its statistically significant effect once the overlap measure is added, implying that its direct effect on redistribution is eclipsed not by heterogeneity in itself (Stephens, 1979) but by the latter’s interaction with class. The index of institutional veto points, considered independently for now, does not have a significant effect on changes in redistribution.

Moving to socioeconomic factors, higher unemployment increases redistribution in the short term, reflecting higher immediate demand, but also decreases social spending over the long run, reflecting a shrinking tax base. Higher labor force participation decreases social spending both immediately and over the long run, although its effect on inequality reduction is insignificant. Higher female participation in the labor force increases both measures of redistribution, as expected. An older population increases social spending in the long run, although it improves
inequality reduction only in the short term. More trade, associated with opposing pressures on the welfare state, shows a mixed pattern: it correlates with short-term reduction in government social spending, but also with long-term increase in inequality reduction. GDP growth has the expected positive effect on changes in social spending, for which it acts as the denominator, but not on inequality reduction. Finally, the negative and statistically significant error-correction terms in all models corroborate the sense of a stable long-term relationship that corrects itself following short-term disturbances.

5.2 Program Type, Coverage, and Replacement Rates

My theory, nonetheless, expects negative pressures only on social policies serving poorer identity groups. Table 2 presents the same model as before but with disaggregated redistribution measures as the outcome. The results, reported in truncated form for ease of presentation, support my hypotheses: the negative interactive effect of heterogeneity and overlap is statistically significant only for unemployment and social assistance, the two program types related most closely to lower socioeconomic position. In contrast, redistribution for old-age and for incapacity, two risks that extend beyond class, remain unaffected by heterogeneity irrespective of overlap levels. This outcome recurs in both government spending and inequality reduction.

The negative effect on unemployment benefits is particularly interesting. Poorer workers face a larger risk of unemployment, yet it still partially applies for stronger employees. How do better-off groups defend themselves against labor market risks while minimizing intergroup redistribution? Earlier, I suggested that they can do so by making labor market programs more selective rather than reducing benefit levels for everyone. This expectation is supported in models 13–16 in Table 3, which estimate coverage and replacement rates.\textsuperscript{19} Social insurance for both unemployment and sick-pay shows the same expected pattern: the interaction of het-

\textsuperscript{19} Models 13–16 add union density as an additional control variable (collinearity with central wage bargaining is ruled out) because some countries employ a Ghent System, in which social security services are distributed by labor unions rather than a state agency. Therefore, union membership is expected to have a direct mechanical effect on labor market coverage. Central wage bargaining is preferred elsewhere due to better data availability. As a robustness check, I reran all other models with union density instead of central wage bargaining and found substantively unchanged results.
erogeneity and overlap has a negative effect only on the share of covered workers, not on the compensation granted to those who are included.

The positive interaction effects for replacement rate display a curious trade-off, illustrated by the marginal effect plots in Figure 6. Whereas higher ascriptive heterogeneity does not affect benefit generosity under high levels of overlap with class, benefit generosity decreases when groups are more equal. The reason may be simple budget constraints: to avoid overspending, wider coverage (under cross-cutting cleavages) may force governments to transfer less per recipient. This interpretation implies that ascriptive heterogeneity and class affect coverage directly and replacement rates indirectly. When identity and class overlap more closely,
Table 3: The Interactive Effect of Identity and Class Overlap on Social Security Entitlements

The full set of control variables, with the addition of union density, is not reported for ease of presentation. Complete estimations are detailed in the supplemental material.

Figure 6: The Marginal Effect of Ascriptive Heterogeneity on Social Security Security Coverage and Replacement Rates as a Function of Overlap
5.3 Institutional Power-Sharing

The final part of my theory concerns the level of institutional power-sharing in policymaking. I test this proposition using a three-way interaction: the first conditional relationship between fractionalization and overlap is now further conditioned upon the veto point index and electoral proportionality. I use each measure separately, since proportionality is also reflected in the veto point index. Because three-way interactions are more demanding statistically, I estimate these models using only the larger sample of social-spending data. With the previous findings in mind, I focus only on overall spending, unemployment, and social assistance. I expect a negative sign for the two-way interaction as before (implying the same negative effect), but a positive sign for the three-way interactions (implying that more power-sharing counteracts the original negative effect).

The results, presented in Table 4, largely support these expectations. The clearest corroboration is found in social assistance spending, which targets lower classes most distinctly: as expected, both a higher number of veto points and higher electoral proportionality moderate the negative effect of heterogeneity and overlap. This finding is illustrated visually in Figure 7, which plots the marginal effect of heterogeneity as a function of overlap given different levels of institutional power-sharing. For ease of interpretation, the boldfaced dashed lines indicate where the slopes are different from zero with statistical confidence of 90% or above. The plots demonstrate that an increase in the number of veto points and in electoral proportionality do not change the negative direction of the slope, but they do gradually flatten it. In other words, power-sharing institutions have the power to moderate the negative relationship, but not reverse it. Nevertheless, the findings also show that a relatively high number of veto points, but not proportionality, can eliminate this effect.

The moderating effect of power-sharing institutions is weaker for overall spending and unemployment. The interactive effect of heterogeneity and overlap on overall social spending is moderated by a larger number of veto points, as expected, but electoral proportionality does not have a similar influence. This weaker result is consistent with my previous findings, which
<table>
<thead>
<tr>
<th>Veto Points Index</th>
<th>(17) Overall Spending</th>
<th>(18) Unemployment</th>
<th>(19) Assistance</th>
</tr>
</thead>
<tbody>
<tr>
<td>AIF</td>
<td>16.120***</td>
<td>0.869</td>
<td>4.176***</td>
</tr>
<tr>
<td></td>
<td>(5.562)</td>
<td>(1.661)</td>
<td>(1.341)</td>
</tr>
<tr>
<td>Overlap</td>
<td>36.924**</td>
<td>0.335</td>
<td>15.584***</td>
</tr>
<tr>
<td></td>
<td>(16.563)</td>
<td>(5.104)</td>
<td>(3.968)</td>
</tr>
<tr>
<td>Veto</td>
<td>0.742</td>
<td>-0.105</td>
<td>0.294***</td>
</tr>
<tr>
<td></td>
<td>(0.478)</td>
<td>(0.173)</td>
<td>(0.107)</td>
</tr>
<tr>
<td>AIF × Overlap</td>
<td>-137.950***</td>
<td>-2.627</td>
<td>-39.654***</td>
</tr>
<tr>
<td></td>
<td>(50.687)</td>
<td>(14.955)</td>
<td>(12.043)</td>
</tr>
<tr>
<td>AIF × Veto</td>
<td>-3.442**</td>
<td>0.197</td>
<td>-0.801**</td>
</tr>
<tr>
<td></td>
<td>(1.512)</td>
<td>(0.527)</td>
<td>(0.376)</td>
</tr>
<tr>
<td>Overlap × Veto</td>
<td>-4.874</td>
<td>1.741</td>
<td>-2.838***</td>
</tr>
<tr>
<td></td>
<td>(4.499)</td>
<td>(1.717)</td>
<td>(1.016)</td>
</tr>
<tr>
<td>AIF × Overlap × Veto</td>
<td>26.804*</td>
<td>-3.469</td>
<td>7.308**</td>
</tr>
<tr>
<td></td>
<td>(13.726)</td>
<td>(4.876)</td>
<td>(3.344)</td>
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<td>620</td>
<td>627</td>
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<tr>
<td>Countries</td>
<td>22</td>
<td>22</td>
<td>22</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.563</td>
<td>0.482</td>
<td>0.217</td>
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<table>
<thead>
<tr>
<th>Electoral Proportionality</th>
<th>(20) Overall Spending</th>
<th>(21) Unemployment</th>
<th>(22) Assistance</th>
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</thead>
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<tr>
<td>AIF</td>
<td>5.184***</td>
<td>1.737**</td>
<td>0.718</td>
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<td></td>
<td>(1.786)</td>
<td>(0.849)</td>
<td>(0.465)</td>
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<td>Overlap</td>
<td>22.050***</td>
<td>7.906***</td>
<td>5.772***</td>
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<td>(5.805)</td>
<td>(2.947)</td>
<td>(1.700)</td>
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<td>Prop$_{t-1}$</td>
<td>-0.023</td>
<td>0.000</td>
<td>0.045*</td>
</tr>
<tr>
<td></td>
<td>(0.110)</td>
<td>(0.040)</td>
<td>(0.026)</td>
</tr>
<tr>
<td>AIF × Overlap</td>
<td>-55.819***</td>
<td>-17.542**</td>
<td>-9.885**</td>
</tr>
<tr>
<td></td>
<td>(16.746)</td>
<td>(7.909)</td>
<td>(4.346)</td>
</tr>
<tr>
<td>AIF × Prop$_{t-1}$</td>
<td>0.203</td>
<td>-0.026</td>
<td>-0.312**</td>
</tr>
<tr>
<td></td>
<td>(0.503)</td>
<td>(0.204)</td>
<td>(0.130)</td>
</tr>
<tr>
<td>Overlap × Prop$_{t-1}$</td>
<td>0.355</td>
<td>0.105</td>
<td>-0.382*</td>
</tr>
<tr>
<td></td>
<td>(0.942)</td>
<td>(0.354)</td>
<td>(0.228)</td>
</tr>
<tr>
<td>AIF × Overlap × Prop$_{t-1}$</td>
<td>-2.212</td>
<td>-0.010</td>
<td>2.647**</td>
</tr>
<tr>
<td></td>
<td>(4.350)</td>
<td>(1.803)</td>
<td>(1.130)</td>
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<td>Observations</td>
<td>627</td>
<td>620</td>
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<tr>
<td>Countries</td>
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<tr>
<td>$R^2$</td>
<td>0.559</td>
<td>0.480</td>
<td>0.213</td>
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</table>

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$, panel-corrected standard errors in parentheses. The full set of control variables is not reported for ease of presentation. Complete estimations are detailed in the supplemental material.

Table 4: The Three-way Interactive Effect of Identity, Class Overlap, and Power-Sharing Institutions on Social Spending
The plotted values for electoral proportionality are percentiles 25, 50, 75 and 95 in its sample distribution.  

**Figure 7:** The Marginal Effect of Ascriptive Heterogeneity on Social Assistance Spending as a Function of Overlap and Power-Sharing Institutions

Indicate that most of the action occurs in class-related programs and not across the board. The negative effect on unemployment, meanwhile, is not moderated by either measure of intergroup power-sharing.

I also run the same three-way interaction models with coverage and replacement rates as the dependent variables, but they mostly produce null results. The only exception is moderation of the negative effect on sick-pay coverage by more veto points. These results may imply that institutional power-sharing assists weaker groups better in highly-visible government actions, such as passing a budget or welfare reforms, but less so in the gray areas of policy rule adjustments. Indeed, decisions on coverage criteria and benefit levels are often made in the bureaucracy and are accordingly subject to fewer veto points and lower public and legislative attention. This possibility justifies additional research.

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20 The full estimations are reported in the supplemental material.
5.4 Robustness Checks

Several diagnostic tests and alternative model specifications, all elaborated further in the supplemental material, validate the robustness of my findings. First, I cross-validate the estimations by dropping each country at a time to verify that the outcomes are not driven by influential cases (Beck & Katz, 2011). I then rerun all models while dropping in turn cases that may stand out for theoretical reasons: Norway, due to its irregular rise in GDP; Greece, due to its partial overlap score (see footnote 13); The US, due to the common critique that its unique racial history and welfare state drive previous findings; and Israel, due to its high overlap score. The findings remain substantively unchanged in all cases.

Second, for models using annual data, I verify that the findings are not driven by the yearly lags and differences. I do so by replacing the values of all (time-variant) independent variables with their moving averages for the previous three years (i.e., the mean value of lags 1 to 3) and rerun all models. The results remain substantively unchanged.

Third, I test my hypotheses with several alternative model specifications (see footnote 16). I use a simple cross-sectional between-effects estimator for the entire sample period, a standard lagged dependent variable (LDV) model, and a minimal LDV model with fewer control variables. Despite their lesser fit to the data structure, all models support my conclusions.

Fourth, I rule out the possibility that the results are driven by reverse or simultaneous causality between redistribution and overlap. As noted before, this concern is assuaged both theoretically and operationally. Nonetheless, to dispel remaining doubts, I rerun all models with an exogenous instrument of ethnic inequality created recently by Alesina, Michalopoulos, and Papaioannou (2016). The ethnic inequality instrument maps nighttime satellite imagery of light density, which reflect economic development, onto historic homelands of different ethnic groups within each country. It then calculates a Gini coefficient of inequality in light density between said subnational regions.\(^{21}\) The instrument is strictly independent of redistributive

\(^{21}\) I use scores from the earliest available year (1992) based on data from Geo-referencing of Ethnic Groups (GREG), excluding groups smaller than 1% of the population.
policies, but has two significant weaknesses: first, it refers only to ethnicity, and, second, its focus on historic homelands excludes cleavages stemming from non-geographic factors as migration, slave trade, or religious conversion. Even so, it correlates reasonably well with my overlap measure \((r = 0.67, p < 0.001)\). The results, reported in more detail in the supplemental material, corroborate my conclusions.

Fifth, I confirm that the findings are not driven by a single ascriptive identity type nested within the aggregate measures. Using my fractionalization subindices and Selway’s cross-cuttingness scores, I rerun all my models using separate interaction terms for ethnicity and for religion. The results find that both types have independent effects in some models but not others. This inconsistent pattern verifies that each type of cleavage is relevant in itself, but at the same time supports my assumption that separating ascriptive identities by type is arbitrary and risks an omitted variable bias.

Finally, I verify that the narrower program coverage found in Table 3 can indeed be interpreted as exclusion of weaker workers. To do so, I estimate whether my explanatory variables also predict non-redistributive labor-market policies benefiting strictly stronger workers. Specifically, I use the OECD’s Employment Protection Legislation (EPL) index for individual and collective dismissals of regular contracts, considered a measure of pro-insiders policy in segmented labor markets (Rueda, 2005). Since the EPL scores hardly change over time, I run a cross-sectional model with a between-effects estimator which includes all the dependent variables from models 13–16. As expected, the estimation finds a positive and statistically significant interactive effect of heterogeneity and overlap on EPL. In other words, heterogeneous democracies with a tighter identity/class overlap also tend to legislate stronger employment protection policies that benefit better-off workers.
6 Conclusion

This paper explored when and how ascriptive identity cleavages shape redistribution in developed democracies. The theoretical and empirical analysis suggests that higher ascriptive heterogeneity matters for redistributive outcomes, but, contrary to common assumptions, it does not act independently or homogeneously. Instead, diversity operates together with class and political institutions: deeper ascriptive identity cleavages undermine redistribution most strongly when identity groups differ in their income levels and when stronger groups concentrate more policymaking power in their hands. Under these conditions, richer groups can more easily and selectively cut benefits for poorer, minority-heavy groups while keeping their own redistributive interests protected. Redistributive policies are therefore driven by a combination of bottom-up factors (multidimensional societal structures) and top-down factors (political institutions and policy design). These findings portray a more nuanced relationship between heterogeneity and redistribution than commonly assumed and shed new light on the unexplained variation in redistribution among heterogeneous countries, the puzzle with which the paper started.

My analysis has several broader implications for the study of identity politics and social policy. First, the growing body of work on individual preferences should not ignore macro-level societal and institutional structures. My findings show that recently-theorized individual-level mechanisms of social distance, social rivalry, and risk differences do imply policy outcomes, but insufficiently so. Including structural factors such as cleavage configurations and political institutions enables us to explain cross-country variation in policy outcomes better. Hence, even as group tensions arise at the individual level, this paper underscores their dependency on intergroup power-distribution and on policy design, two factors that remain underexplored in the current discussion on identity politics and redistribution.

Second, social policies should be examined more closely, particularly in the context of identity politics. Redistributive outcomes are typically analyzed aggregatively, yet different components of social policy vary significantly in their underlying goals, target audiences, inclusiveness, and implementation. My findings thus point at the importance of additional research
on policy design and concrete instruments by which social groups may be differentiated in the
distribution of resources. Furthermore, this research avenue can shift the theoretical discussion
from repeated theories of collective action and social solidarity to more nuanced frameworks
of competing group interests and unequal allocation of political power.

Third, long-standing cleavage structures and power-sharing institutions are important for
the rapidly growing debate on recent immigration and social policy. My analysis implies that
identity politics did not begin with current immigration inflows but have long shaped social
policy. As such, recent changes in the ascriptive makeup of developed democracies build upon
previous intergroup dynamics and policy equilibria. More attention should be given to the
interaction of new and old identity divisions, intergroup coalitions and rivalries, and intergroup
power balance. Heterogeneous countries whose social policies are already shaped in light of
old intergroup tensions may deal differently with incoming immigration compared to more
homogeneous societies, and, moreover, new immigrants may alter the mutual affinities and
interests of old ascriptive identity groups. Recent immigration, in other words, is the most
recent development in the ongoing dynamics of identity politics in the developed world.
### Appendix

<table>
<thead>
<tr>
<th>Variable</th>
<th>Definition</th>
<th>Data Source</th>
<th>Expected Effect</th>
</tr>
</thead>
<tbody>
<tr>
<td>Left</td>
<td>The share of cabinet portfolios held by left-wing parties</td>
<td>Swank (2013), EJPR Political Data Yearbook&lt;sup&gt;a&lt;/sup&gt;</td>
<td>+</td>
</tr>
<tr>
<td>ChristDem</td>
<td>The share of cabinet portfolios held by Christian-Democratic parties</td>
<td>Swank (2013)</td>
<td>+</td>
</tr>
<tr>
<td>CWB</td>
<td>Centralization of wage bargaining</td>
<td>Visser (2015)</td>
<td>+</td>
</tr>
<tr>
<td>Union Density</td>
<td>Net union density (models 13–16)</td>
<td>Visser (2015)</td>
<td>+</td>
</tr>
<tr>
<td>Unemployment</td>
<td>Unemployment rate as a share of the total labor force</td>
<td>World Economic Outlook Database</td>
<td>+/-</td>
</tr>
<tr>
<td>LabForce</td>
<td>Civilian labor force participation as a share of population aged 15 or above</td>
<td>OECD Labor Statistics</td>
<td>-</td>
</tr>
<tr>
<td>FemLabForce</td>
<td>Female participation as a share of the civilian labor force</td>
<td>OECD Labor Statistics</td>
<td>+</td>
</tr>
<tr>
<td>Elderly</td>
<td>The share of population aged 65 or above</td>
<td>World Development Indicators</td>
<td>+</td>
</tr>
<tr>
<td>LogTrade</td>
<td>The log of trade (the sum of exports and imports of goods and services) as a share of GDP</td>
<td>World Development Indicators</td>
<td>+/-</td>
</tr>
<tr>
<td>LogGDP</td>
<td>The log of gross domestic product based on purchasing-power-parity (PPP) per capita</td>
<td>World Economic Outlook Database</td>
<td>+/-</td>
</tr>
</tbody>
</table>

<sup>a</sup> Swank’s database does not include Israel. I use Swank’s coding rules to calculate the cabinet portfolio allocation for left parties in Israel using data from the European Journal of Political Research (EJPR) Political Data Yearbook. Israel does not have Christian-Democratic parties.

**Table A1: Control Variables: Definitions, Sources, and Expected Effect on Redistribution**
References


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Cambridge: Harvard University Press.


