Abstract

The current consensus among comparative political scientists postulates that diverse democracies redistribute less than do homogeneous ones. However, whereas homogeneous democracies do redistribute more on average, diverse democracies exhibit high variation in redistributive outcomes. Why does ascriptive heterogeneity stifle redistribution in some cases but not in others? In this paper, I argue that diversity undermines redistributive outcomes when identity groups differ more starkly in their income levels. More importantly, under these conditions, the policy outcomes are not uniform: rather than general cutbacks, richer groups selectively underprioritize benefits and access for poorer, minority-heavy groups while keeping their own redistributive interests protected. The result is not just less redistribution aggregately, but a more exclusionary and regressive welfare state that serves the social needs of better-off identity groups. I find empirical support in these hypotheses using macrocomparative panel data on multiple redistributive aspects in 22 developed democracies in the years 1980–2011. My paper thus outlines a conditional and more nuanced relationship between diversity and redistributive outcomes than commonly assumed, as well as several broader lessons for research of identity politics and social policy.
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 social scientists, 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 and Burkhardt, 2009; Pontusson, 2006; Steele, 2016; Taylor-Gooby, 2005). Moreover, newer research on individual-level preferences finds that the negative relationship between diversity and popular support for redistribution depends upon additional demographic and socioeconomic factors (e.g., Alt and Iversen, 2017; Burgoon, 2014; Dahlberg, Edmark and Lundqvist, 2012; Finseraas, 2012).

Figure 1: Ascriptive Diversity and Three Measures of Redistribution. CV stands for Coefficient of Variation.

Recent data confirm this empirical tension. Figure 1 plots three different measures of redistribution levels against an index of ascriptive identity fractionalization\(^1\) in 19–22 OECD coun-

\(^1\)The Ascriptive Identity Fractionalization (AIF) index integrates multiple data sources and identity dimen-
tries. 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 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 heteroskedastic pattern: whereas homogeneous countries tend to cluster together more closely, heterogeneous ones vary more broadly by all three redistributive measures. This tendency is illustrated with the matching box-and-whisker diagrams, which split the sample in half and plot the variation within each subgroup. Thus, the accepted notion that diversity undermines redistribution seems correct but insufficient: higher heterogeneity in ascriptive identities influences redistributive outcomes, but not in all cases. Why is this so?

This paper addresses this gap by discussing the combined role played by ascriptive identities and class in shaping de facto redistributive outcomes. Building upon previous research on redistributive preferences, I argue that diverse democracies redistribute less in practice when some ascriptive identity groups are richer than others and have lower willingness to support the latter. Importantly, switching from preferences to policy outcomes, I hypothesize that the redistributive implications are not uniform. As stronger identity groups seek to minimize intergroup redistribution, the negative effect concentrates primarily in programs targeting poorer groups and on the latter’s access to social benefits. At the same time, richer groups protect broader redistributive programs that serve their members and make them more exclusionary. Diversity with higher intergroup inequality, therefore, leads to more regressive and less inclusive welfare state with more differentiation between the needs it addresses.

I find support in these hypotheses using macrocomparative panel data on 22 developed democracies in the years 1980–2011. Specifically, I show that the negative relationship between ascriptive diversity and redistributive outcomes is mediated by the level of income differences, including ethnicity, religion, and language. I elaborate upon its operationalization in later sections.
ferences between identity groups. More importantly, I find differential outcomes on various redistributive aspects. When identity and class cleavages reinforce one another, higher diversity curbs primarily welfare programs that target the needs of poorer groups (unemployment, social assistance, and public healthcare), but does not affect programs addressing cross-class risks such as old-age and incapacity. Under these conditions, furthermore, key social security programs cover fewer recipients, even as their generosity per (fully covered) recipient remains high. These combined findings imply higher differentiation between strong and weak recipient groups rather than crude cutbacks across the board. 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 reinforcement of identity and class divisions and by the target audiences of specific programs. The paper concludes with several broader implications for the study of identity politics and social policy.

2 Existing Research on Diversity and Redistribution

Social scientists have long established that ascriptive identities—ethnicity, race, religion, and language—are particularly potent politically given their inherent, indivisible, and relatively rigid nature (Lipset and Rokkan, 1967; Rae and Taylor, 1970), their central role in in-group coordination (Bates, 1983; Fearon and Laitin, 1996), and their mobilization by political actors (Chandra, 2004; Posner, 2004). Different studies have demonstrated that ascriptive identity cleavages influence economic development and democratization (Easterly and Levine, 1997; Houle, 2015; Montalvo and Reynal-Querol, 2005), intergroup conflict (Fearon and Laitin, 2003; Wilkinson, 2008), public goods provision (Alesina, Baqir and Easterly, 1999; Habyarimana, Humphreys, Posner and Weinstein, 2007), party systems (Clark and Golder, 2006; Ordeshook and Shvetsova, 1994), and voter behavior (Chandra, 2004; Huber, 2012). Identity politics were slower to penetrate the comparative study of redistributive politics, however. Mainstream research of developed welfare states has focused predominantly on class as the primary, and often only, relevant social cleavage. Redistributive outcomes are often explained
by the income distribution of otherwise-identical voters (Meltzer and Richard, 1981; Moene and Wallerstein, 2001), the political power of the left (Esping-Andersen, 1990; Korpi, 1983; Korpi and Palme, 2003; Stephens, 1979), and the institutional constrains on class-based actors (Hicks and Swank, 1992; Huber and Stephens, 2001; Iversen and Soskice, 2006). Ascriptive identity cleavages play a minor or no role in these theories.²

In recent years, nonetheless, more attention is given to the redistributive implications of ascriptive diversity. The accepted wisdom arising from this literature suggests that diverse societies redistribute less than do homogeneous ones, as higher ascriptive heterogeneity fractures 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, Baqir and Easterly, 1999; Quadagno, 1994; Ribar and Wilhelm, 1999; Skocpol, 1992). Alesina and Glaeser (2004) offer the most comprehensive comparative presentation of this argument, maintaining that higher diversity explains much of the historic gap between the American and European welfare states (see also Desmet, Ortuño-Ortín and Weber, 2009; Sanderson, 2004). They further argue that ascriptive cleavages precede and explain the emergence of class politics and electoral institutions, the primary drivers of welfare policy according to previous research. Empirically, Alesina and Glaeser find 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 statistical power

²There is a small handful of references to ascriptive identity cleavages in this literature, yet mostly as an indirect factor. Stephens (1979), for instance, notes that ethnic and linguistic heterogeneity undermine the cohesion of labor organizations that promote welfare state expansion. Using a similar logic, Romer, Lee and Van der Straeten (2007) argue that anti-immigration sentiments strengthen right-wing parties and thus decrease redistribution indirectly. Some attention was also given to the effect of religion on welfare states, primarily through Christian-Democratic parties (Huber, Ragin and Stephens, 1993; Van Kersbergen and Manow, 2009) and religious networks (Huber and Stanig, 2011; Scheve and Stasavage, 2006). However, this discussion analyzes religion not as a contentious social cleavage but as an organized constituency and a set of values.
when the sample is limited to Western countries and with a more careful consideration of other 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, researchers found that support in redistribution decreases most strongly when minorities are poorer or perceived as such. These studies argue that higher inequality between identity groups increases their perceived social and cultural distance, leading in turn to decreased solidarity (Lupu and Pontusson, 2011) and to stronger stigmatization of poor minorities (Gilens, 1995; Kinder and Sears, 1981; Nelson, 1999). It further exacerbates social threat, as richer identity groups oppose policies that may mobilize poorer ones into their communities and undermine their relative social status (Corneo and Grüner, 2002; Shayo, 2009). Finally, intergroup income inequality sets apart each group’s occupational risks and, accordingly, their redistributive interests (Alt and Iversen, 2017).

This conditional prediction has been tested empirically in two often-overlapping ways, both of which paint only part of the picture (e.g., Alt and Iversen, 2017; Brady and Finnigan, 2014; Burgoon, 2014; Finseraas, 2012; Stichnoth, 2012). First, many studies focus on individual preferences as the primary outcome of interest. Yet, public preferences on redistribution are typically measured broadly and are insufficient to infer concrete policy outcomes. Second, there is increased focus on recent immigration and its mostly-negative effect on redistributive preferences and policies. Recent immigrant inflows are a growingly important factor in Western democracies. Nevertheless, immigration is one of several causes for ascriptive heterogeneity, and as such is insufficient for a fuller understanding of the latter. New immigrants integrate slowly, do not automatically constitute a coherent political constituency (Michon and Vermeulen, 2013), and face unique barriers to full political and economic rights, parliamentary representation, and access to state services (Bird, Saalfeld and Wüst, 2011; Dancygier, 2010; Dancygier, Lindgren, Oskarsson and Vernby, 2015). Accordingly, while there are signs of a gradual decline in social spending due to sustained immigration (Soroka, Johnston, Kevins, Banting and 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 and Van Hooren,
Existing theoretical and empirical accounts of the relationship between ascriptive diversity writ large and redistributive outcomes remain incomplete. I turn next to address this gap.

### 3 Theoretical Propositions: The Interplay of Identity and Class

Following previous work, I expect ascriptively diverse democracies to face stronger pressures against broad redistribution than do homogeneous ones. Yet, as the recent literature on redistributive preferences establishes, these forces should vary by each country’s alignment of identity and class. The shift from simple preferences to policy implications, however, is not straightforward. Redistributive policymaking is not a dichotomous decision between more or less government involvement. To see why, it is useful to consider the two primary roles of the welfare state. First, welfare states reduce the inequalities created by market forces through income transfer from rich to poor. Second, welfare states provide social insurance against various socioeconomic risks. The latter, nevertheless, are distributed differentially across classes. Some risks, like age-related complications or enduring sickness, are not a function of class and are therefore similarly probable across all groups. Moreover, aging and incapacity pose a particular threat to the occupational advantage held by high-level workers. Other risks, like sustained unemployment, material deprivation, or sudden healthcare costs, do correlate with class-based factors such as labor market skills and savings. Thus, they are inherently more likely among poorer groups.

When some ascriptive identity groups are richer than others, their increased in-group bias should thus undermine only programs focused on income redistribution and on lower-class

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3 The discussion in this paper assumes that relevant ascriptive categories have already been formed and does not explore long-term processes of identity formation. An influential body of work shows that ascriptive identities can be endogenous to economic development, state borders, electoral institutions, and violent conflicts (e.g., Chandra, 2004; Laitin, 1986; Posner, 2005; Sambanis and Shayo, 2013). These insights, however, are of lower importance for contemporary developed democracies, where the primary social cleavages have largely been shaped decades or even centuries ago (Lipset and Rokkan, 1967). Indeed, constructivist arguments are “about the long-run formation, and the consequent stickiness, of identities” (Varshney, 2009, p. 288).

4 For simplicity, I assume that reelection-seeking politicians always promote the interests of identity groups within their electoral coalitions. In the discussion that follows, therefore, I refer directly to identity groups as the key political actors.
risks. Protection against shared risks, in contrast, equally benefiting middle- and high-class
groups, should remain protected.

Underprioritization of programs serving the poor is not the only channel through which
richer identity groups can minimize intergroup redistribution. Another path is to decrease the
latter’s access to programs that address shared risks. In particular, poorer classes can be ex-
cluded effectively in the labor market, where many developed economies experience growing
segmentation between different tiers of workers. As the literature emphasizes, there is increas-
ing 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 and Seeleib-Kaiser, 2012; Iversen and Sos-
kice, 2015; Rueda, 2005). Richer groups, therefore, can also entrench harsher access criteria
that exclude poorer workers from full labor-market protection.

The theoretical intuition is illustrated in Figure 2, which plots several hypothetical income
distributions in a society comprising two ascriptive identity groups. In all cases, income as-
sumes a common left-skewed distribution, with a higher concentration of citizens at the middle
and lower classes. Panel A plots a perfectly cross-cutting cleavage structure, such that the two
identity groups distribute similarly across all classes. Even if each group cares only for the in-
terests of its in-group members, limiting any redistributive program will hurt everyone equally.
This is not the case when identity and class divisions reinforce one another, as panel B demon-
strates. Here, members of the rich identity group benefit nothing from welfare policies targeting
the needs of the poor. At the same time, their interest in programs covering cross-class risks
remains firm. Only in case B, but not in case A, we should expect more limited redistributive
policies, specifically ones serving the poor.

Two nontrivial scenarios, portrayed in panels C and D, reinforce this logic. Panel C con-
siders high intergroup income differences in a relatively homogeneous society. If the poor
minority is very small in size, intergroup inequality does not crowd out the dominant group
from the lower class. Instead, the majority benefits from all types of welfare policies regard-
less of the small minority’s position. This case, therefore, underscores that the juxtaposition of
identity and class is insufficient in itself. The key theoretical mechanism, then, requires both high heterogeneity and reinforcement of identity and class.

Panel D examines a scenario where identity and class cleavages reinforce one another, but now the minority are strictly rich rather than poor. Prima facie, we should expect an opposite outcome compared to panel B, as the poorer majority group may establish a lower/middle class coalition that will force the rich minority to redistribute income broadly. This argument, however, is weak for two reasons. First, because of the left-skewed nature of income distributions, for a minority to be strictly richer than the lower and middle classes it must be small in size. Panel D, for example, has the same majority-to-minority proportion as does panel C, even as
the spread is different. Here too, then, society is actually quite homogeneous. As it will grow
more diverse, the minority’s members will fill the ranks of the middle class and fracture the
interclass coalition against the rich. Second, even if the rich are a relative minority, a growing
literature finds that policymaking tends to react more strongly to the interests of the rich (Bart-
tels, 2015; Gilens, 2012; Peters and Ensink, 2015), implying that a (sufficiently sizable) rich
minority would be politically strong even without an absolute majority.\footnote{Empirically, it is hard to find many current instances of small but dominant minorities in contemporary de-
veloped democracies. Examples of privileged minorities in developing settings, typically the result of colonial
legacies, do indicate strong protection of their economic interests (e.g., whites in South Africa, ethnic Russians in
soviet republics, ethnic Chinese in various Southeast Asian countries).} These points, then,
underpin the initial intuition: high diversity with larger intergroup class differences should lead
to less redistribution for the poor.

In sum, I expect that higher ascriptive diversity, when it is reinforced by broader intergroup
class differences, would dampen redistributive outcomes nonuniformly, concentrating primar-
ily on (1) welfare programs targeting the needs of the poor, and (2) universal access to programs
providing labor-market protection.

4 Data and Empirical Strategy

I test these theoretical hypotheses using a series of models estimating how ascriptive hetero-
geney and its juxtaposition with class differences correlate with multiple aspects of welfare
policy in developed democracies. I use cross-sectional time-series data form the years 1980–
2011 for 22 countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Ger-
many, 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.
4.1 Variables of Interest

4.1.1 Dependent Variables

I examine the primary outcome of interest, redistribution, 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 (SOCX), 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 data from the Luxembourg Income Survey (LIS; Wang and 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. Inequality reduction, meanwhile, captures de facto policy implications and is not as sensitive to changes in GDP levels or recipient numbers. All else equal, I expect that a combination of higher heterogeneity and cleavage reinforcement 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) incapacity and sickness benefits, consisting of sick-pay compensation, occupational injury transfers, and disability benefits; (3) unemployment benefits, consisting of unemployment compensation and active-labor programs; and (4) assistance benefits, consisting of income maintenance, housing assistance, family and child allowances, and similar in-kind benefits. Additionally, I examine (5) public social spending on universal healthcare services. All else equal, I expect that a combination of higher heterogeneity and

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6The 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.

7There is no LIS data on public healthcare, which typically involves public services rather than cash transfers. My disaggregation leaves out several components that are unhelpful as separate categories. In the OECD expen-
higher class differences will decrease only unemployment, social assistance, and public health-care programs, which benefit lower classes disproportionately, whereas old-age and incapacity programs should not be affected.

The third aspect of redistribution involves inclusiveness in labor-market protection programs. I use data from the Comparative Welfare Entitlement Project (CWED2; Scruggs, Jahn and 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.\(^8\) All else equal, I expect that a combination of higher heterogeneity and higher class differences will decrease only the level of program coverage (i.e., access to programs), not replacement rates (i.e., generosity for those who remain included).

4.1.2 Independent Variables

My hypotheses mark two explanatory factors: heterogeneity in ascriptive identities and the reinforcement of identity and class cleavages. In line with the current literature, I measure ascriptive heterogeneity using one minus the Herfindahl index, which estimates the level of social fractionalization.\(^9\) There has been significant improvement in the quantity and quality of ascriptive fractionalization indices in recent years. Nevertheless, these indices have two notable problems. First, the multitude of indices raises the risk of post hoc cherry-picking. Second, available indices calculate independent scores for different ascriptive identity dimensions, typically separating race, ethnicity, religion, and/or language. These different types of identities, however, all share an alleged common genetic, historic, or spiritual decent, have relatively rigid and visible criteria, and improve social coordination (Chandra, 2006; Hale, 2004; Haller and Eder, 2015; Laitin, 2007). Therefore, an imposed separation makes it difficult to compare sim-

diture data, I exclude the “other” category, since, as the name implies, its content varies by country. In the LIS data, I exclude military service and veteran transfers.\(^8\) The CWED2 data do not include Israel.\(^9\) 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.

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ilar processes across equally-divided countries differing only in the type of salient identities that developed there historically (Wimmer, 2008). We thus need a common measure to compare the implications of ascriptive intergroup tensions in such cases as the US (race), Belgium (language), Ireland (Religion), or Israel (ethnicity).

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 to divisions identified repeatedly. I draw from 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 and Wacziarg, 2003); (2) the Ethnic Power Relations’ index of ethnic fractionalization in politically-relevant groups based on expert surveys (EPR; Cederman, Wimmer and 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, Ortuño-Ortín and Wacziarg, 2012). 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.\textsuperscript{10} Figure 3 summarizes the index’s structure.

\begin{figure}[h]
\centering
\includegraphics[width=\textwidth]{figure3.png}
\caption{The Ascriptive Identity Fractionalization (AIF) Index}
\end{figure}

\textsuperscript{10}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.
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 then in 2000. The strong bivariate correlation between the two periods ($r = 0.93, p < 0.000$) implies high stability over time.11

The second explanatory variable of interest is the reinforcement of ascriptive identity and class cleavages, 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 if the distribution of members in the first cleavage does not predict their distribution in the second. Conversely, two cleavages reinforce one another if their membership distribution replicates 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.12 To align with my theoretical hypotheses, I invert the composite score to measure cleavage reinforcement.

Like the AIF index, the cleavage reinforcement scores are time-invariant due to data limitations. This constraint too should not pose a serious problem, however, as reinforcement levels are expected to be both stable during the sample period and exogenous to redistribution. The

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11There are, of course, growing immigration inflows to Western democracies in recent years. However, this should not destabilize the AIF scores significantly during my sample period. 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 influence 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.

12Selway (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 combined ethnic fractionalization score is 0.076 compared to a sample mean of 0.22. As a robustness check, I omitted Greece from all my models and found substantively unchanged results.
surveys used to calculate the reinforcement scores ask responders about their *relative* income bracket. Redistributive policies lessen inequality, i.e., 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. For my purposes, the measure’s reliance on membership distribution is preferable to measures of intergroup income inequality (e.g., Baldwin and Huber, 2010; Houle, 2015), which rely on mean incomes that may be endogenous to redistributive policies. 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. Furthermore, intergroup perceptions and stigmas, a central mechanism behind group preferences, change even slower. Several diagnostic and robustness tests corroborate both the exogeneity and stability of the cleavage reinforcement measure.\(^{13}\)

In addition to the two primary variables of interest, I also control for several other explanatory factors associated in the literature with redistributive outcomes. My control variables include institutional features (a combined index of institutional veto points), political power balance (cabinet partisanship and union centrality), 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 definitions, sources, and expected effects of all control variables. The descriptive statistics of the dependent, independent, and control variables are presented in the supplemental material.

### 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 and Yeo, 1978; De

\(^{13}\)To test for exogeneity, I reran my models with an instrument for group inequality instead of cleavage reinforcement and found substantively similar results. This test is elaborated upon later in this paper and in the supplemental material. To test for stability, I analyzed the cleavage reinforcement scores in 97 comparable surveys—surveys conducted by the same data-collecting project, in the same country, using the same question wording, only 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 in the juxtaposition of identity and class. This procedure is explained in more detail in the supplemental material.
Due to indications of panel-specific heteroskedasticity, the estimation employs panel-corrected standard errors (PCSE; Beck and 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 its explanatory variables, but that this relative stability can be disturbed by short-term shocks followed by a correction back to the long-term trend as 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 time-invariant, I include them outside the error-correction dynamics. Their stable values are interpreted as casting a constant, long-term influence on patterns of equilibrium and disturbances.

Formally, I estimate the following model:

\[ \Delta R_{i,t} = \alpha + \beta_1 \Delta X_{i,t} + \gamma (R_{i,t-1} - \beta_2 X_{i,t-1}) + \beta_3 Z_i + \epsilon_{i,t} \]

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 explanatory variables. My primary interest is in the latter effect. I interact the two main explanatory variables to test their hypothesized conditionality (B Brambor, Clark and Golder, 2006; Kam and Franzese, 2007).^{15}

\(^{14}\)Lagrange multiplier tests indicate that the ECM structure decreases, but not eliminates, serial autocorrelation when estimating annual spending data. To solve this problem, I add the lagged first difference of the dependent variable (\( \Delta R_{i,t-1} \)) and of its denominator (\( \Delta GDP_{i,t-1} \)). The first difference of the dependent variable (but not of GDP) is also added when estimating unemployment coverage, diagnosed with a similar problem. This fix eliminates the remaining serial autocorrelation.

\(^{15}\)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.
Several alternative model specifications, including a simpler cross-sectional model with a between-effects estimator, support the same theoretical and substantive conclusions. I use the ECM as the baseline specification because it both controls for additional variables (unlike a simple cross-sectional model with fewer degrees of freedom) and captures intricate political dynamics better by separating stable long-term effects from short-term disturbances (unlike a simpler Lagged Dependent Variable model). I elaborate more on these alternative specifications later and in the supplemental material.

5 Findings

5.1 Diversity, Class, and Aggregate Redistribution

The first set of tests, presented in Table 1, evaluate the prediction that cleavage reinforcement mediates the relationship between ascriptive diversity on redistributive outcomes, at this point still in aggregate terms. As a point of reference, models 1 and 3 test the direct effect of ascriptive heterogeneity on these outcomes, 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 cleavage reinforcement of identity and class grows, ascriptive diversity has an increasing negative effect on changes in both government spending and inequality reduction.

Figure 4 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 cleavage reinforcement with income. As expected, higher diversity has a negative marginal effect on redistribution, which grows as the two cleavages overlap more closely. Furthermore, the negative influence is significant only past some minimal threshold of cleavage reinforcement.

\[16\]

To verify that this null result is not an artifact of my AIF measure, I reran models 1 and 3 using Alesina and Glaeser’s original ethnic, linguistic, and religious fractionalization measures. None of their indices produce a statistically significance result.
### Table 1: The Interactive Effect of Identity and Class Reinforcement on Social Spending and on Inequality Reduction

Interestingly, when income and class cross-cut one another, 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. When all groups have strong identities and in-group bias, but similar shares of poor members, increased group loyalties may thus create

<table>
<thead>
<tr>
<th>Explanatory Variables</th>
<th>Overall Spending</th>
<th>Inequality Reduction</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Coef.</td>
<td>PCSE</td>
</tr>
<tr>
<td>AIF</td>
<td>-0.241</td>
<td>(0.214)</td>
</tr>
<tr>
<td>Reinforcement</td>
<td>18.495 **</td>
<td>(4.712)</td>
</tr>
<tr>
<td>AIF × Reinforcement</td>
<td>-45.767 ***</td>
<td>(11.898)</td>
</tr>
<tr>
<td>Veto</td>
<td>0.001</td>
<td>(0.018)</td>
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</tbody>
</table>

### Short-Term Relationships

<p>| | | | |</p>
<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>∆Left</td>
<td>0.001</td>
<td>(0.001)</td>
<td>0.000</td>
</tr>
<tr>
<td>∆ChristDem</td>
<td>0.008 **</td>
<td>(0.004)</td>
<td>0.009 **</td>
</tr>
<tr>
<td>∆CWb</td>
<td>0.032</td>
<td>(0.050)</td>
<td>0.025</td>
</tr>
<tr>
<td>∆Unemployment</td>
<td>0.135 ***</td>
<td>(0.037)</td>
<td>0.127 ***</td>
</tr>
<tr>
<td>∆LabForce</td>
<td>-0.167 ***</td>
<td>(0.042)</td>
<td>-0.175 ***</td>
</tr>
<tr>
<td>∆FemLabForce</td>
<td>0.329 ***</td>
<td>(0.089)</td>
<td>0.309 ***</td>
</tr>
<tr>
<td>∆Elderly</td>
<td>0.188</td>
<td>(0.202)</td>
<td>0.443 **</td>
</tr>
<tr>
<td>∆LogTrade</td>
<td>-1.728 ***</td>
<td>(0.558)</td>
<td>-1.615 ***</td>
</tr>
<tr>
<td>∆LogGDP</td>
<td>-17.400 ***</td>
<td>(2.013)</td>
<td>-17.854 ***</td>
</tr>
</tbody>
</table>

### Long-Term Relationships

<p>| | | | |</p>
<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Left_{t-1}</td>
<td>0.000</td>
<td>(0.001)</td>
<td>0.000</td>
</tr>
<tr>
<td>ChristDem_{t-1}</td>
<td>0.001</td>
<td>(0.001)</td>
<td>0.003 **</td>
</tr>
<tr>
<td>CWB_{t-1}</td>
<td>0.052 *</td>
<td>(0.029)</td>
<td>0.045</td>
</tr>
<tr>
<td>Unemployment_{t-1}</td>
<td>-0.011</td>
<td>(0.010)</td>
<td>-0.024 **</td>
</tr>
<tr>
<td>LabForce_{t-1}</td>
<td>-0.004</td>
<td>(0.007)</td>
<td>-0.015 *</td>
</tr>
<tr>
<td>FemLabForce_{t-1}</td>
<td>0.026 ***</td>
<td>(0.010)</td>
<td>0.031 ***</td>
</tr>
<tr>
<td>Elderly_{t-1}</td>
<td>0.034 **</td>
<td>(0.016)</td>
<td>0.054 ***</td>
</tr>
<tr>
<td>LogTrade_{t-1}</td>
<td>-0.012</td>
<td>(0.067)</td>
<td>-0.045</td>
</tr>
<tr>
<td>LogGDP_{t-1}</td>
<td>-0.099</td>
<td>(0.133)</td>
<td>-0.164</td>
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</tbody>
</table>

### Error-Correction Term

<p>| | | |</p>
<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Spending_{t-1}</td>
<td>-0.038 ***</td>
<td>(0.010)</td>
</tr>
<tr>
<td>Gini Reduction_{t-1}</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

### Lagged First-Difference

<p>| | | | |</p>
<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>∆Spending_{t-1}</td>
<td>0.172 ***</td>
<td>(0.055)</td>
<td>0.157 ***</td>
</tr>
<tr>
<td>∆LogGDP_{t-1}</td>
<td>9.078 ***</td>
<td>(1.840)</td>
<td>8.075 ***</td>
</tr>
</tbody>
</table>

<p>| | | | |</p>
<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>0.861</td>
<td>(1.357)</td>
<td>0.447</td>
</tr>
<tr>
<td>Observations</td>
<td>627</td>
<td></td>
<td>627</td>
</tr>
<tr>
<td>Countries</td>
<td>22</td>
<td></td>
<td>22</td>
</tr>
<tr>
<td>R²</td>
<td>0.544</td>
<td></td>
<td>0.558</td>
</tr>
</tbody>
</table>

* p < 0.1, ** p < 0.05, *** p < 0.01, panel-corrected standard errors in parentheses.
simultaneous motivations for higher redistribution to poor in-group peers, increasing overall redistribution as a result. 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. This process is fueled to a large degree by Flemish motivation to establish a separate redistributive system and by federal fortification of the national system in response (Béland and Lecours, 2008; Cantillon, 2011). This empirical pattern should not be overstated, however, as it is both small in size and does not repeat in inequality reduction.

Most, although not all, 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, but left-leaning cabinets do not produce the expected positive effect. Interestingly, central wage bargaining loses its statistically significant effect once the interaction is added, implying that its direct effect on redistribution is eclipsed not by diversity in itself but by the latter’s juxtaposition with class. The index of institutional veto points does not have a notable effect on changes in redistribution.

Among socioeconomic factors, higher unemployment shows a complex pattern: it increases
redistribution in the short term, reflecting higher immediate demand, but also decreases social spending over the long run, likely 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, displays a mixed influence: it correlates with short-term reduction in 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 shocks.

5.2 Program Type, Coverage, and Replacement Rates

My hypotheses, nonetheless, expects negative pressures only on social policies serving poorer identity groups. Table 2 presents the same models but with disaggregated measures of spending and inequality reduction as the outcome. The results, reported in truncated form for ease of presentation, support my hypotheses: the negative interactive effect of heterogeneity and cleavage reinforcement is statistically significant only for unemployment, social assistance, and public healthcare spending, the three program types related most closely to universal access and lower-class needs. By contrast, redistribution for old-age and for incapacity, two risks shared across classes, remain unaffected by diversity irrespective of its reinforcement levels with class. Notably, this outcome repeats in both government spending and inequality reduction.

The negative effect on unemployment benefits is particularly interesting. Poorer workers face a larger risk of sustained unemployment, yet middle-class employees are not fully immune to it. Earlier, I suggested that better-off groups can protect themselves against labor-market risks, while minimizing income transfer to poorer groups, by restricting the latter's access to
### Table 2: The Interactive Effect of Identity and Class Reinforcement by Program Type

<table>
<thead>
<tr>
<th>Outcome: Public Social Spending</th>
<th>(5) Age</th>
<th>(6) Incapacity</th>
<th>(7) Unemployment</th>
<th>(8) Assistance</th>
<th>(9) Healthcare</th>
</tr>
</thead>
<tbody>
<tr>
<td>AIF</td>
<td>-0.687**</td>
<td>0.058</td>
<td>1.252***</td>
<td>1.502***</td>
<td>1.763***</td>
</tr>
<tr>
<td>(0.333)</td>
<td></td>
<td></td>
<td></td>
<td>(0.37)</td>
<td>(0.416)</td>
</tr>
<tr>
<td>Reinforcement</td>
<td>-1.343</td>
<td>1.485</td>
<td>5.029***</td>
<td>5.875***</td>
<td>5.262***</td>
</tr>
<tr>
<td>(1.306)</td>
<td></td>
<td></td>
<td></td>
<td>(1.513)</td>
<td>(1.705)</td>
</tr>
<tr>
<td>AIF × Reinforcement</td>
<td>4.395</td>
<td>-1.268</td>
<td>-11.779***</td>
<td>-15.208***</td>
<td>-15.06***</td>
</tr>
<tr>
<td>(3.062)</td>
<td></td>
<td></td>
<td></td>
<td>(3.625)</td>
<td>(3.934)</td>
</tr>
<tr>
<td>Observations</td>
<td>627</td>
<td>627</td>
<td>620</td>
<td>627</td>
<td>637</td>
</tr>
<tr>
<td>Countries</td>
<td>22</td>
<td>22</td>
<td>22</td>
<td>22</td>
<td>22</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.464</td>
<td>0.218</td>
<td>0.476</td>
<td>0.209</td>
<td>0.313</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Outcome: Inequality Reduction</th>
<th>(10) Age</th>
<th>(11) Incapacity</th>
<th>(12) Unemployment</th>
<th>(13) Assistance</th>
</tr>
</thead>
<tbody>
<tr>
<td>AIF</td>
<td>-0.202</td>
<td>-0.053</td>
<td>0.073</td>
<td>0.424**</td>
</tr>
<tr>
<td>(0.15)</td>
<td></td>
<td></td>
<td></td>
<td>(0.179)</td>
</tr>
<tr>
<td>Reinforcement</td>
<td>-1.168</td>
<td>0.167</td>
<td>0.196</td>
<td>1.964***</td>
</tr>
<tr>
<td>(0.916)</td>
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<td></td>
<td></td>
<td>(0.686)</td>
</tr>
<tr>
<td>AIF × Reinforcement</td>
<td>1.939</td>
<td>0.19</td>
<td>-0.803*</td>
<td>-4.971***</td>
</tr>
<tr>
<td>(1.845)</td>
<td></td>
<td></td>
<td></td>
<td>(1.8)</td>
</tr>
<tr>
<td>Observations</td>
<td>87</td>
<td>79</td>
<td>80</td>
<td>88</td>
</tr>
<tr>
<td>Countries</td>
<td>19</td>
<td>19</td>
<td>19</td>
<td>19</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.492</td>
<td>0.702</td>
<td>0.572</td>
<td>0.49</td>
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</tbody>
</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. Full estimations are detailed in the supplemental material.

The positive interaction effects for replacement rate display a curious trade-off, illustrated by Table 3, which estimates changes in levels of coverage and replacement rates. Social insurance for both unemployment and sick-pay shows the same pattern: the interaction of heterogeneity and cleavage reinforcement has a negative effect on the share of covered workers, but not on the compensation granted to those who are included.

The full set of control variables is not reported for ease of presentation. Full estimations are detailed in the supplemental material.

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17Models 14–17 include union density as an additional control variable (collinearity with central wage bargaining is ruled out) since some countries employ a Ghent System, where 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, despite a smaller sample.

---

21
<table>
<thead>
<tr>
<th></th>
<th>Unemployment</th>
<th>Sick-pay</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(14)</td>
<td>(15)</td>
</tr>
<tr>
<td>AIF</td>
<td>-0.273***</td>
<td>0.199**</td>
</tr>
<tr>
<td></td>
<td>(0.093)</td>
<td>(0.084)</td>
</tr>
<tr>
<td>Reinforcement</td>
<td>-0.531**</td>
<td>0.501**</td>
</tr>
<tr>
<td></td>
<td>(0.209)</td>
<td>(0.207)</td>
</tr>
<tr>
<td>AIF × Reinforcement</td>
<td>2.475***</td>
<td>-1.865**</td>
</tr>
<tr>
<td></td>
<td>(0.815)</td>
<td>(0.768)</td>
</tr>
<tr>
<td>Observations</td>
<td>628</td>
<td>556</td>
</tr>
<tr>
<td>Countries</td>
<td>21</td>
<td>20</td>
</tr>
<tr>
<td>(R^2)</td>
<td>0.098</td>
<td>0.188</td>
</tr>
</tbody>
</table>

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

Table 3: The Interactive Effect of Identity and Class Reinforcement on Social Security Entitlements

the marginal effect plots in Figure 5: higher ascriptive heterogeneity decreases benefit generosity when identity and class cross-cut one another. The reason may be simple budget constraints: to avoid over-spending, wider coverage (under cross-cutting cleavages) may force governments to transfer less per recipient. This interpretation implies that ascriptive diversity and class affect coverage directly and replacement rates indirectly. When identity and class reinforce one another more closely, higher heterogeneity leads to narrower coverage and, therefore, to weaker budgetary pressures to cut benefit generosity. When identity and class cut across one another, however, higher heterogeneity motivates all groups to increase inclusion of their members, and hence forces them to reduce benefit generosity to keep the system sustainable.

I support this explanation with two additional tests, both reported in more detail in the supplemental material. First, to corroborate the interpretation that diversity and higher cleavage reinforcement increase the exclusion of weaker workers, I examine whether they predict other labor-market policies benefiting strictly stronger workers. Specifically, I estimate whether the same factors correlate with 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 ran a cross-sectional model with a between-effects estimator including all the dependent
variables from models 14–17. As I expect, I find a positive and statistically significant interaction coefficient. In other words, supporting my findings, diverse democracies with higher intergroup class differences also tend to protect better-off workers more strongly.

Second, to support the interpretation of a trade-off between coverage and generosity, I reran the same models with the multiplication of coverage and replacement rates as the outcome of interest. In both unemployment and sick-pay programs, this multiplication remains statistically unaffected by the interaction of diversity and reinforcement. This null finding indicates that the combined package of coverage and generosity per recipient remains stable across different cases regardless of their diversity and cleavage structure.

5.3 Robustness Checks

Several diagnostic tests and alternative specifications, all elaborated upon in the supplemental material, validate the robustness of my findings. First, I cross-validate all models by dropping each country at a time to verify that their fit is not driven by influential cases (Beck and 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, the denominator of social spending; Greece, due to its partial cleavage reinforcement score (see footnote 12); The US, due to the common critique that its unique racial history and welfare policies drive previous findings; and Israel, due to its extreme cleavage reinforcement 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 structure of the data. 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 using several alternative model specifications. Specifically, I use a simple cross-sectional between-effects estimator that averages variable values 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, all models corroborate my conclusions.

Finally, I rule out the possibility that the results are driven by an endogenous relationship between redistribution and cleavage reinforcement. As discussed previously, there are both theoretical and empirical reasons that alleviate this concern. Nonetheless, to dispel remaining doubts, I rerun all models with an exogenous instrument of ethnic inequality created by Alesina, Michalopoulos and Papaioannou (2016). The ethnic inequality instrument maps nighttime satellite imagery of light density, reflecting 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. The instrument is strictly independent of redistributive policies, but has two significant weaknesses: first, it refers only to ethnicity, and, second, the focus on historic homelands excludes cleavages based in non-geographic factors such as migration, slave trade, or religious conversion. Even so, it correlates reasonably well with my cleavage reinforcement measure ($r = 0.67, p < 0.001$). The results, reported in more

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18 I use scores from the earliest available year (1992) based on data from Geo-referencing of Ethnic Groups (GREG). Ethnic groups smaller than 1% of the population are excluded.
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 diversity matters for redistributive outcomes, but, contrary to the common assumption, it does not act independently or homogeneously. Instead, in line with recent findings on redistributive preferences, I demonstrate that welfare outcomes too are affected by the combination of diversity and class. Specifically, I find that deeper ascriptive diversity dampens redistributive outcomes when income differences between identity groups are sufficiently large and increasing. More importantly, I show that redistributive outcomes are affected nonuniformly. When diversity combines with broad intergroup income differences, politically dominant and richer identity groups selectively cut benefits and access for poorer, minority-heavy groups while keeping their own redistributive interests protected. The result is not fewer social services for everyone, but a more regressive and exclusionary welfare state that prioritizes the social needs of better-off identity groups. These findings portray a more nuanced relationship between diversity 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 contributions for the study of identity politics and social policy. First, the growing body of work, and important insights, on identity politics and individual preferences tell only part of the story. My findings indicate that individual-level mechanisms such as social distance, social rivalry, and skill differences do imply negative policy outcomes, but insufficiently so. Individual-level theories, then, should be complemented by macro-level research of actual policy outcomes, the types of available policy tools that can promote them, and the conflicting interests that each serves or undermines.

Second, as we turn to consider social policy outcomes, my analysis emphasizes the often-
overlooked variation between different types of social policies. Redistributive outcomes are typically analyzed aggregately, yet different social policies differ 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 resource distribution. Furthermore, this avenue of research can shift the theoretical discussion from often-discussed problems of collective action and solidarity to more nuanced frameworks of competing group interests and unequal allocation of political power.

Third, long-standing cleavage structures are important for the rapidly growing debate on recent immigration and its influence on social policy. My analysis implies that identity politics did not begin with current migrant inflows but have long influenced redistributive patterns. As such, it is fair to assume that recent changes in the ascriptive makeup of developed democracies build upon previous intergroup dynamics and policy equilibria. More attention should thus 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 ascriptive tensions may deal differently with incoming immigration compared to more homogeneous welfare states, 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>VP</td>
<td>Additive veto points index: (1) degree of federalism, (2) presence of presidentialism, (3) degree of bicameralism, (4) use of referenda, (5) degree of proportionality, and (6) presence of judicial review</td>
<td>Armingeron, Isler, Knöpfel, Weisstanner and Engler (2016)(^a)</td>
<td>+</td>
</tr>
<tr>
<td>Left</td>
<td>The share of cabinet portfolios held by left-wing parties</td>
<td>Swank (2013)(^b)</td>
<td>+</td>
</tr>
<tr>
<td>ChristDem</td>
<td>The share of cabinet portfolios held by Christian-Democratic parties</td>
<td>Swank (2013)(^b)</td>
<td>+</td>
</tr>
<tr>
<td>CWB</td>
<td>Centralization of wage bargaining</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>

\(^a\) I code Israel’s values using the same rules as Armingeron et al., as it is missing from the original dataset.

\(^b\) Since Israel is missing from Swank’s database, I use the same 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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URL: 10.1016/j.jdeveco.2004.01.002


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**URL**: www.uva-aias.net/208
