The Structure of Inequality and Demand for Redistribution*

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The recent growth of empirical literature on the politics of inequality in advanced capitalist societies is impressive. We now have a large number of studies that investigate the relationship between the distribution of market income and the degree of redistribution through taxes and transfers across countries and over time (e.g., Kenworthy and Pontusson 2005; Mahler and Jesuit 2006; Milanovic 2000). More recently, comparative political economists have begun to tackle the question of how inequality affects the policy preferences of voters and partisan politics (e.g., Finseeras 2009; Pontusson and Rueda 2008). Informed by the theoretical model typically attributed to Meltzer and Richard (1981) and by the alternative model proposed bv Moene and Wallerstein (2001), this literature essentially seeks to establish whether or how (or under what conditions) the *level of inequality* matters to redistributive politics. Simply put, the literature asks, does rising inequality generate more (or less) redistribution?

The basic purpose of this paper is to make the case that we ought to pay more attention to the *structure of inequality*. As suggested some time ago by Kristov, Lindert, and McClelland (1992), there are good reasons to suppose that dispersion in the upper half of the income distribution has different political implications than dispersion in the lower half of the distribution. The alternative perspective we put forth here boils down to the following proposition: voters in the middle of the income distribution will be inclined to ally with low-income voters and support redistributive policies when the distance between the middle and the poor is small and when the distance between the middle and the rich is large.

We test this proposition using an error-correction model of the kind specified by Iversen and Soskice (2006), with data from thirteen OECD countries. Based on data from the Luxembourg Income Study (LIS), our measure of redistribution is the percentage change in the Gini coefficient for household income brought about by taxes and transfers. For our measures of the structure of inequality, we use OECD data on wages (gross earnings) among full-time employees. Our dataset includes a few observations from the 1970s, but primarily draws on observations from 1980-2004. Controlling for government partisanship, electoral rules, and a number of other variables that the existing literature identifies as determinants of redistribution, we find that dispersion of wages in the upper half of the distribution (measured by the 90-50 wage ratio) is associated with more redistribution, while dispersion of wages in the lower half (measured by the 50-10 ratio) is associated with less redistribution. In supplementary analyses, we show that these results hold up when we control for national skill profiles, ethnic fractionalization, and immigration. Probing the causal mechanisms behind our main results, we also present descriptive survey evidence on preferences for redistribution and speculate about the role of government partisanship in the process linking the preferences of middle-income voters to government policy.

The Dependent Variable

Like much of the recent comparative literature on the politics of redistribution – notably Bradley et al. (2003) and Iversen and Soskice (2006) – we use LIS data to measure redistribution and operationalize this variable as the percentage change in Gini coefficients that we observe as we move from household income before taxes and transfers ("gross market income" in LIS terminology) to household income after taxes and transfers ("disposable income").[1] Also in keeping with existing literature, our analysis is restricted to redistribution of income among working-age households (see also Kenworthy and Pontusson 2005; Milanovic 2000). Specifically, our dependent variable is the effect of taxation and income transfers on the distribution of income among households headed by someone between the ages of 25 and 59.[2] As noted already, the statistical model that we estimate is an error-correction model, which means that the dependent variable is the change in redistribution since the preceding observation of redistribution.

The justification for restricting our analysis to working-age households is that generous public pension systems reduce the incentive for individuals to accumulate savings. In countries with very generous public pension systems, a majority of the retired population have virtually no market income at all. Studies of redistribution that include the retired population thus yield very high levels of market inequality and, in a sense, exaggerate the redistributive effects of public spending in countries with generous public pension systems.[3]

Our dataset draws primarily on LIS-based estimates of income inequality among working-age households generated by Lane Kenworthy (see Kenworthy and Pontusson 2005). Kenworthy's dataset includes at least two observations of inequality measured in terms of both gross market income and disposable income for twelve OECD countries, and one observation for Belgium. While Kenworthy only reports market income inequality on a post-tax basis (i.e., net market income) for France, a different LIS-based dataset assembled by Mahler and Jesuit (2006) includes four observations of gross market income for France and one additional observation for Belgium. Since the Mahler-Jesuit data are virtually identical to the Kenworthy data when they overlap, we have simply added these observations to our dataset. Altogether, we have 75 observations of redistribution in our dataset. In estimating an error-correction model, we lose the first observation for each country, so the *N* of our analyses is actually 62. For Belgium, our analysis only includes a single observation. At the other end of the spectrum, our analysis includes 8 observations for Canada. On average, our sample includes 4.8 observations per country.[4]

Our Theoretical Framework

Our starting assumption is that the support of voters in the middle of the income distribution is usually critical to winning elections in liberal democracies and that most parties are sophisticated enough to recognize this and opportunistic enough to adjust their policies accordingly. These assumptions are similar to those of the canonical median-voter model associated with Romer (1975) and Meltzer and Richard (1981). While we posit that parties make strategic adjustments in order to win elections, we do not rule out that parties also have either core constituencies they seek to please or ideologically-motivated policy commitments. Towards the end of the paper, we will elaborate on how government partisanship might be incorporated into our theoretical framework. For the time being, we set aside the issue of how voter preferences are translated into government policy and focus on the conditions that determine the preferences for redistribution of middle-income voters or, in other words, the "median voter."

In the model proposed by Romer (1975) and Meltzer-Richard (1981), the median voter is purely self-interested and seeks to maximize her income in the short term. Redistribution takes the form of a universal flat-rate benefit that is financed by a linear income tax. Holding the deadweight costs of taxation constant, the demand for redistribution by the median voter in this setup becomes a function of the distance between her income and the mean income. Assuming that all income earners are citizens and exercise their right to vote, a mean-preserving increase of inequality makes the median voter more supportive of redistribution. Crucially, the median voter does not have to know anything about the distribution of income in order for this result to hold. Indeed, the model assumes that the median voter does not care about her relative position in the income distribution.

Building on Kristov, Lindert, and McClelland (1992), our framework attributes a more complex set of motivations to voters. First, we posit that preferences for redistribution are partly informed by other-regarding motivations, specifically by sympathies (or antipathies) for fellow citizens. Secondly, we question the notion that self-interest boils down to short-term income maximization. To the extent that voters form their preferences based on some form of calculation of their own interests, we think it plausible to assume that their calculations are informed by some assessment, however imperfect, of their prospects of income mobility (downward as well as upward). Thirdly, and closely related to the second point, our framework assumes that voters care about relative as well as absolute gains and losses.

Much like Iversen and Soskice's (2006) model, the model of redistributive politics developed by Kristov, Lindert, and McClelland (1992) conceives the electorate as divided into three income-based groups of more or less the same size: the poor, the middle, and the affluent. The critical question is whether the middle-income group will form a redistributive coalition with the poor or will join with the affluent in defending the existing income distribution. For Iversen and Soskice (2006), the answer to this question depends on electoral rules (see below). For Kristov, Lindert, and McClelland (1992), by contrast, it depends on the structure of the income distribution. Specifically, these authors hypothesize that if the distance between the middle and the (near) bottom of the income distribution is small, the median voter will sympathizes with the poor or, alternatively, include the possibility of becoming poor into her cost-benefit calculus. If the distance between the middle and the (near) top of the distribution is small, the median voter will lean against redistribution for similar reasons. In other words, the Kristov, Lindert, and McClelland (1992) model suggests that the combination of relatively small income differences in the lower half of the distribution and relatively large income differences in the upper half provides the most favorable conditions for redistributive politics.

The implication of this argument is that the recent empirical literature on the relationship between inequality and redistribution is misguided insofar as it relies on a single measure of inequality such as the Gini coefficient or the 90-10 ratio. To appreciate the political consequences of inequality, we need incorporate separate measures of the spread of incomes in the lower and upper halves of the income distribution.[5] In what follows, we rely on the OECD's database on relative earnings among full-time employees as the source of such measures. Specifically, we estimate separate effects for the 90-50 wage ratio and the 50-10 wage ratio, expecting the former to be positively associated with redistribution and the latter to be negatively associated with redistribution.

Our main justification for using wages (rather than gross household income from LIS) as the basis for our measures of market inequality is that the OECD database provides continuous series of annual observations over much of the time period covered by our data on redistribution. This is a requirement for estimating the type of error-correction model that Iversen and Soskice (2006) specify.[6] Also, it should be noted again that our measure of redistribution is restricted to working-age households, which renders wage inequality a more reasonable measure of market inequality than it would be if the dependent variable pertained to all households. Finally, it is at least plausible to argue that wage inequality constitutes a particularly visible form of market inequality (cf. Pontusson and Rueda 2008). To the extent that voters know anything about the distribution before taxes and transfers in the society at large, they are likely to know more about the distribution of wages than the distribution of other types of market income.

For illustrative purposes, Table 1 ranks the thirteen countries included in our analysis by level of redistribution in 2000, or in years as close to 2000 as possible.[7] The table then reports 90-50 and 50-10 wage ratios for the same year and, in the last column, a summary measure that we think of as the "pro-redistribution skew" of the wage structure. This variable is simply the result of subtracting the 50-10 ratio from the 90-50 ratio. The correspondence between country rankings on redistribution and skew is far from perfect, but the correlation between redistribution and skew is 0.703. It is noteworthy that Denmark, Finland, and Sweden are among the four countries with the highest levels of redistribution and also among the four countries with the greatest skew in the wage distribution. In the top rankings, Belgium has more redistribution than we would predict based on the structure of wage inequality while France has less. At the other end of the spectrum, Switzerland and Canada stand out as the two countries in which the lower half of the wage distribution is more dispersed than the upper half. Along with the US, these two countries also figure in the bottom rung of the ranking based on the level of redistribution. As the next iteration of our paper will illustrate, there is a lot of over-time variation as well as cross-sectional variation in redistribution in our dataset. This justifies the use of an error-correction framework for estimating the determinants of redistribution.

[Table 1 about here]

The micro-foundations of the hypotheses put forth by Kristov, Lindert, and McClelland (1992) clearly need to be developed further. For now, let us just make a few points to clarify where we think this effort will take us. To begin with, we deliberately equivocate on the extent to which voter preferences are motivated by social norms and affinities or by calculations of self-interest. In our view, it is not only difficult for analysts to disentangle these motivations, it also difficult for individual social agents to do so. Solidarity becomes an operative behavioral norm, we think, when individuals have some rational reason to suppose that it might serve their own interests over the long run.

With regard to the self-interested part of the equation, the premise of our approach is that voter preferences for redistribution are informed not only by calculations of short-term gains and losses, but also by expectations of income mobility. According to Milanovic's (2000) analysis of LIS data, taxes and transfers typically reduce the income of the share of households between the 50th and 60th percentiles of pre-fisc income distribution, but these losses are quite small by comparison to those of households above the 60th percentile. On the other hand, households below the 40th percentile typically benefit from the redistributive schemes in existence in most OECD countries. It seems plausible to suppose that someone in the 50th percentile of the pre-fisc income distribution would be inclined to support further redistribution if she thinks there is a realistic possibility that she would end up in, say, the 30th percentile in the foreseeable future. Conversely, we might expect the median income earner to support cuts in redistribution to the extent that she believes that she might end up in the 70th percentile.

Suppose further that mobility between any two positions in the income distribution tends to increase as the distance between them declines.[8] By this logic, a smaller 50-10 wage ratio signifies a greater probability of downward mobility for the median income-earner while a smaller 90-50 ration signifies a greater probability of upward mobility.

One might object that the greater probability of upward or downward mobility when wages are compressed is offset by fact that the gains associated with moving from the 50th to the 70th percentile or the losses associated with moving from the 50th to the 30th percentile will be correspondingly smaller. However, this objection assumes that voters are only concerned with absolute gains or losses from redistribution. To the extent that voters care about their relative position in the income distribution, an increase in the (perceived) probability of falling from the 50th to the 30th percentile should always be associated with a stronger preference for redistribution among voters in the 50th percentile of the income distribution.

We readily recognize that other factors will also affect perceptions of the prospects of income mobility. In particular, it seems clear that perceptions of mobility vary across countries for reasons that have very little to do with objective contemporary conditions (cf. Alesina and Glaeser 2004; Piketty 1995). It seems likely that ideological constructions and historical experiences of income mobility would swamp any effects of the information that middle-income voters might derive from observing the current wage structure in a simple cross-sectional analysis. Again, however, our analysis focuses on change over time, and country-specific ideological constructions do not appear to vary significantly over the time period covered by our analysis.

Turning to the issue of social affinity, we want to emphasize that what we have in mind here is different from altruism. If middle-income voters are motivated by altruism in any strict sense, their sympathy for the poor should increase with the distance between the income of the poor and the median income. As we understand it, the notion of social affinity implies a form of "altruism" or, more appropriately, "solidarity" that is bounded by perceptions of common group membership or shared experience. We suppose that middle-income voters sympathize with the poor (or affluent) when they perceive the poor (or affluent) as living lives that are similar to their own. We suppose further that in income differentials are a relatively good proxy for variation in living conditions between the poor and the middle, on the one hand, and between the middle and the affluent, on the other.

At the same time, it seems quite obvious, especially in light of the American experience, that the degree to which middle-income voters feel affinity with the poor partly depends on whether or not the poor belong to the same racial or ethnic group. Our emphasis on social affinity has much in common with Alesina and Glaeser's (2004) discussion of race and ethnic-religious fractionalization as obstacles to redistributive politics (cf. Gilens 2000; Luttmer 2001). We do not wish juxtapose our argument about the structure of income inequality to the literature that emphasizes race and ethnicity as the basis of social affinities. However, it deserves be noted that there appears to be no association whatsoever between Alesina and Glaeser's fractionalization measures and social spending if the analysis is restricted to OECD countries (see Alesina and Glaeser 2004: 141-43). From our perspective, what matters to the politics of redistribution is not racial or ethnic fragmentation per se, but rather the way that racial or ethnic cleavages map onto the income distribution. Alesina and Glaeser agree: in their words, the crucial variable is whether "there are significant numbers of minorities among the poor," in which case "the majority population can be roused against transferring money to people who are different from themselves" (134).

Unfortunately, we are not aware of any cross-national dataset that captures the distribution of racial and ethnic minorities across the income/wage structure. For lack of appropriate data, our main results ignore the influence of race and ethnicity, but we will in due course present the results of supplementary analyses that attempt to take into account such considerations.

Like much of the recent empirical literature on the determinants of redistribution, we include voter turnout in our baseline model and conceive this variable essentially as an inverse proxy for income skew in voting.[9] As commonly noted, the variable that matters in the Romer-Meltzer-Richard model is the distance between the income of the median voter and the mean income, not the distance between the median income (in the population at large) and the mean income (e.g., Nelson 1999). If a mean-preserving increase of inequality is associated with a decline in (relative) turnout among low-income voters, the distance between the income of the median voter and the mean income may well remain unchanged (or decline). In such a scenario, we would not expect rising inequality to produce an increase in demand for redistribution. By our slightly different logic, we also expect voter turnout to be associated with more redistribution. In our framework, higher voter turnout means that the median voter is closer to the poor and farther from the affluent, and hence more likely to join a redistributive coalition with the poor.

Other Theories and Variables

We do not wish to argue that the structure of inequality alone provides an adequate account of temporal or cross-sectional variation in redistribution. In testing the core predictions generated by our theoretical framework, we need (and want) to take into account the effects of other potentially relevant variables. To identify these variables, we draw on existing literature.

To begin with, government partisanship features prominently in the existing literature on the politics of redistribution, notably in work by Bradley et al. (2003) and Iversen and Soskice (2006). The standard assumption in the literature is that parties of the Left and the Right draw their core support from different ends of the income distribution and that Left parties are more inclined to engage in redistribution than Right parties. Regarding the putative need for parties to gain the support of the median voter to win elections, proponents of partisanship as a major determinant of redistribution adopt one or several of the following positions: (1) convergence on the median voter is only necessary under majoritarian electoral rules; (2) who the median voter is depends on voter mobilization, and this in turn depends on the efforts of ideologically committed party activists; (3) even if turnout remains constant, the policy preferences of the median voter are sufficiently uncertain as to allow parties to pursue their own preferences; and (4) parties deviate from the median-voter platform between elections (and especially when they win elections by large margins).

To take into account partisan effects, our baseline model includes Cusack's "cabinet center of gravity" (CABCOG) index as a right-hand-side variable. This index relies on the average of three expert surveys to classify parties on the Left-Right continuum and weights party scores by the share of cabinet portfolios held by different parties (Cusack and Engelhardt 2002). The index is here standardized to vary between 0 and 1, with higher values representing more Right-leaning governments. In keeping with conventional wisdom, we expect partisanship measured in this fashion to be negatively associated with redistribution. That is, we expect a period of Right-leaning government to result in a reduction of redistribution.

It is important to keep in mind that the following analysis pertains to contemporaneous partisan effects. By contrast, Bradley et al. (2003) estimate the effects of a cumulative measure of cabinet shares held by Left parties over the entire postwar era. They do so in order to capture the notion that continued electoral success and participation in government by Left parties will induce other parties to engage in strategic repositioning or, in other words, to become more favorable to redistribution (see also Huber and Stephens 2001). We are sympathetic to this basic idea and plan to explore it for the next iteration of this paper (see below).

Our empirical models also include union density as a right-hand-side variable. Proponents of power resources theory have increasingly come to emphasize the effects of government partisanship, but unionization remains, we think, the most obvious measure of working-class mobilization, which is the key theoretical variable in this tradition.[10] Generalizing across OECD countries, unions do not typically organize workers at the very bottom of the wage distribution, but their members are drawn disproportionately from the lower half of the distribution. Controlling for income and other relevant demographics, union members are more likely to vote than non-union members and also more likely to support redistribution than non-union members (see Pontusson and Kwon 2006). Arguably, unionization strengthens the representation of low-income interests in policy-making as well as electoral politics. For these reasons, we expect union density to be associated with greater redistribution.

Another strand of the recent literature focuses on the question of why countries with proportional representation (PR) tend to have more redistributive governments than countries with majoritarian electoral rules. Alesina and Glaeser (2004) bring this question to the fore, but the causal mechanisms of their account are ambiguous. Persson and Tabellini (2000, 2003) argue that electoral rules affect the types of spending incumbent politicians choose. While majoritarian electoral rules favor geographically targeted spending, PR favors more broad-based or, in other words, more universalistic spending programs. In a somewhat different vein, Persson, Roland, and Tabellini (2007) propose a model in which the effect of PR on government spending hinges on the greater probability of a coalition government under PR rules. In this model, each party in the governing coalition has a strong incentive to reward its core constituency with spending financed by taxing all voters since the electoral losses associated with higher taxes are shared by all members of the coalition. While the original Persson-Tabellini model pertains to the allocation of government spending and does not explain why spending levels tend to be higher under PR, the more recent Persson-Roland-Tabellini model focuses on spending levels and does not make any strong predictions about the extent to which government spending is redistributive under different electoral rules.

Iversen and Soskice's (2006) alternative take on the role of electoral rules proceeds from the observation that government participation by Left parties has been much more common in PR countries than in majoritarian countries since 1945. In their formal model, uncertainty about party commitments leads the median voter to favor Center-Right parties over Center-Left parties under majoritarian rules while centrist parties will prefer to ally with Left parties over Right parties under PR. By a very similar logic, Ticchi and Vindigni (Forthcoming) argue that “consensual democracies should be expected to be ruled relatively often by center-left coalitions, more willing to tax and redistribute income, while the more fiscally conservative right should have an advantage in majoritarian countries” (3). For Ticchi and Vindigni as well as Iversen and Soskice, the redistributive effects of electoral rules operate through government partisanship. It would seem to follow that the direct effect of electoral rules will disappears once we control for the effect of government partisanship.

To adjudicate among different explanations of the association between PR and redistribution is not a primary goal of the analysis that follows. In the first instance, we simply want to control for the elector-systems effect identified by the literature. To this end, we include Gallagher’s (1991) measure of proportionality in our empirical models. This index begins with the sum of squared absolute deviations of individual party seat shares from their respective shares of the vote. This sum is then divided by two and the square root taken. Thus the index ranges from 0, meaning pure proportionality between vote and seat shares, to infinity as disproportionality increases. For ease of interpretation, we standardize this measure to vary between 0 and 1, and invert it so that larger values refer to higher levels of proportionality rather than of disproportionality. Unlike a dichotomous classification of countries as having either majoritarian or proportional electoral systems, the Gallagher index captures variations within each broad type (among other things, due primarily to variation in the average magnitude of electoral districts) and varies over time.

Going back to a somewhat earlier theoretical discussion, we also control for constitutional veto points, as measured by Huber, Ragin, and Stephens (1993). This composite and time-invariant index measures the multiplication of policy decision nodes as a result of federalism, bicameralism, and referenda. Following the existing literature, we expect constitutional veto points to be associated with less redistribution.

Finally, all the results we report below are based on estimating models that include the rate of unemployment. More readily than any other variable that we have been able to identify, the unemployment rate serves as a way to control for changes in the share of the working-age population that is eligible for redistributive social transfers. So long as unemployment insurance coverage and generosity remain unchanged, an increase in unemployment translates more or less automatically into greater redistribution (cf. Kenworthy and Pontusson 2005).[11]

Statistical Model

Our main results are based on estimating an error-correction model of the kind specified by Iversen and Soskice (2006) and modified by Vernby and Lindgren (2009). This model treats the level of redistribution today ($R_{i,t}$) as a function of previous levels of redistribution ($R_{i,t-1}$) and policies ($P_{i,t}$) that cause redistribution to deviate from the status quo. The model incorporates a scalar ρ, which captures the speed with which levels of redistribution respond to changes in government policy:

$$R_{i,t} = \rho\left[\alpha + \beta P_{i,t} - R_{i,t-1}\right] + R_{i,t-1} + u_{i,t} \quad (1)$$

Given that available data on redistribution are unequally spaced, while values for the independent variables are annual, Iversen and Soskice (2006) modify this basic model by replacing the lagged values of redistribution until reaching the previous observation of redistribution. By this procedure, we get:

$$R_{i,t} = \rho\alpha \sum_{s=0}^{N}(1-\rho)^{s} + \rho\beta \sum_{s=0}^{N}(1-\rho)^{s}\rho P_{i,t-s} + (1-\rho)^{N+1}R_{i,t-N+1} + \sum_{s=0}^{N}(1-\rho)^{s}\rho u_{i,t-s} \quad (2)$$

$$R_{i,t} - (1-\rho)^{N+1}R_{i,t-N+1} = \rho\alpha \sum_{s=0}^{N}(1-\rho)^{s} + \rho\beta \sum_{s=0}^{N}(1-\rho)^{s}P_{i,t-s} + \sum_{s=0}^{N}(1-\rho)^{s}u_{i,t-s} \quad (3)$$

The right-hand side of the model thus includes a constant followed by the cumulative effect of partisan government policies over the period N between the current observation of redistribution and the previous one (s represents the lags in years). The other independent variables are similarly calculated in terms of cumulative effects. Thus the estimated model takes into account the complete time series of annual data even though observations of the dependent variable are not available annually. In other words, each observation of redistribution is thought to be affected by the cumulative effect of independent variables in all the years since the previous observation of redistribution. Like Iversen and Soskice, we estimate this model with different values of ρ and present the results for the version that explains the most variance.

If we assume that the error term in equation (1) is not serially correlated, it follows that the errors in equations (2) and (3) are also not serially correlated. However, since the error terms in (2) and (3) depend on N, the errors must be heteroskedastic. Iversen and Soskice use feasible generalized least squares to correct for this heteroskedasticity. However, as Vernby and Lindgren (2009) note, this procedure only corrects for heteroskedasticity within countries, but not between them. We therefore follow Vernby and Lindgren's (2009) recommendation of using weighted least squares to account for both types of heteroskedasticity. [12]

Main Results

Table 2 reports our primary set of results. For comparative purposes, model 1 includes only a single, all-encompassing measure of inequality – the 90-10 ratio – as is often done by analysts seeking to control for the level of inequality. Model 2 shows our main results once we replace the 90-10 ratio with the 90-50 and the 50-10 ratios, thereby accounting for the structure of inequality. While the coefficient for the 90-10 ratio is tiny and insignificant, the coefficients for the 90-50 and 50-10 ratios are both sizeable and clear the 99% confidence threshold. Consistent with our theoretical expectations, the coefficient for the 90-50 ratio is positive and the coefficient for the 50-10 ratio is negative. In other words, dispersion of the upper half of the wage distribution is associated with more redistribution while dispersion of the lower half is associated with less redistribution.

[Table 2 about here]

Turning to the other variables included in our main model, neither voter turnout nor government partisanship appear to be associated with changes in redistribution. The absence of significant partisan effects poses a challenge to power resource theory as well as Iversen and Soskice's (2006) argument about the association between PR and redistribution. Our results are particularly troublesome for Iversen and Soskice's theory because we also obtain a strong electoral-system effect. On average, redistribution increased more in countries with more proportional electoral arrangements over the observations of redistribution included in our dataset, but this effect does not appear to be attributable to the greater incidence of government participation by Left parties in countries with PR systems. As noted earlier, our model is not designed to capture the effects of long-term electoral dominance by parties of the Left or the Right, and we shall return to this question in due course.[13]

While the absence of partisan effects also poses a challenge to power resource theory, our results provide support for the other prediction that emerges from this theoretical tradition: as measured by union density, working-class (or low-income) mobilization is strongly associated with greater redistribution. We also obtain significant coefficients for constitutional veto points and rates of unemployment, and the signs of these coefficients are consistent with our expectations. Again, it should be noted that our measure of constitutional veto points is time invariant. In the error-correction framework adopted here, the negative coefficient for the variable means that, on average, redistribution increased less (or decreased more) between observations of redistribution in countries with more constitutional veto points.

Regarding the effects of the structure of inequality on redistribution, our results stand in marked contrast to those reported by Schwabish, Smeeding, and Osberg (2006). In their analysis of social spending, measured in percent of GDP, the 50-10 ratio for market income among working-age households is weakly associated with *more* social spending, while the 90-50 ratio is strongly associated with *more* social spending. Quite reasonably, Schwabish, Smeeding, and Osberg (2006) interpret the former result to mean that eligibility for means-tested social spending increases with the 50-10 ratio and that government responsiveness to the policy preferences of the affluent increases (or, in other words, that government responsiveness to the preferences of the median voter declines) with the 90-50 ratio.

There are many possible reasons for the discrepancies between our results and theirs: their dependent variable is different (spending rather than redistribution, levels rather than change) as are their measures of inequality, and they include different control variables in their model. We intend to probe these differences further. For now, suffice it to note that Schwabish, Smeeding, and Osberg (2006) include public support for social spending and redistribution in their empirical model. In our theoretical framework, 90-50 and 50-10 ratios matter to the politics of inequality through their effects on the policy preferences of middle-income voters. Controlling for such preferences, we have no expectations about the effects of the structure of inequality on redistribution. We shall return to the question of public opinion below.

National Skill Profiles

Readers familiar with recent work in the varieties-of-capitalism tradition might suspect that the 50-10 wage ratio is an inverse proxy for skill specificity and that the causal mechanism underlying our finding is different from the one we have proposed. Iversen and Soskice (2001) demonstrate convincingly that individuals with more specific skills are more likely to support social spending (controlling for income, which is negatively associated with skills). On a cross-national basis, Estevez-Abe, Iversen, and Soskice (2001) measure skill specificity by the share of an age cohort engaged in secondary and tertiary vocational training, and show that this measure is associated with compression of wage differentials as measured by the 90-10 ratio (169-78). Iversen and Soskice (2001: 888) show that vocational training share (VTS) is also correlated with government spending on income transfers on a cross-national basis. Finally, Iversen (2005: 148-54) presents a modified version of the analysis of redistribution in Iversen and Soskice (2006) that includes VTS as a right-hand-side variable. With redistribution data for 14 countries over the period 1967-97, Iversen finds that VTS has a strong positive effect on redistribution and suggests that the introduction of this variable eliminates the negative effect of the 90-10 wage ratio obtained in a reduced-form model.

In theory, the Iversen-Soskice argument about skills pertains to demand for social insurance, not demand for redistribution. It is not entirely obvious why individuals with more specific skills should prefer a more equal distribution of income. These studies appear to treat redistribution as a more or less unintended by-product of the public provision of social insurance. In our view, this treatment of redistribution as a secondary issue is rather puzzling, seemingly at odds with the oft-repeated adage that politics is all adopt "who gets what." Why should we not think of insurance and redistribution as separate (overlapping) dimensions of the welfare state, each of which is politically contested? For our purposes, it should also be noted that even if controlling for national skill profiles were to eliminate the negative association between lower-half wage inequality and redistribution, this would still leave us with the need to account for the positive association between upper-half wage inequality and redistribution. One of the strengths of our theoretical framework is that it predicts the positive 90-50 ratio as well as the negative effect of the 50-10 ratio.

Iversen's (2005) analysis uses annual observations of VTS for the period 1980-1995. For most countries, this measure changes very little over time. Table 3 reports two models that include VTS as an independent variable in our analysis. The first set of results (model 3) are based on restricting our analysis to observations of redistribution over the period 1981-2000, thus minimizing the extrapolation of missing data on vocational training. For the second set of results (model 4), we extrapolated missing data so as to include our entire set of observations of redistribution. Contrary to the results that Iversen (2005) reports, we do not find any significant effects of vocational training in either model.[14] Including the VTS variable does seem to somewhat weaken the effect of the 50-10 wage ratio. In model 3, the size of the coefficient is noticeably smaller than in model 2 (Table 2), and the coefficient is only significant at the 90% level. However, it is seems likely that much of this discrepancy is due to differences in the composition of the sample. When we include our full sample (in model 4), the size of the 50-10 coefficient is nearly the same as in model 2, though somewhat less precisely estimated (95% confidence). The effects of the other variables in our model, including the 90-50 ratio, are largely unaltered by including the VTS variable. In short, we are fairly confident that skill specificity is not an omitted variable that confounds the inferences we draw from our main results.

[Table 3 about here]

Ethnicity and Immigration

Another potential problem with the results presented in Table 2 concerns the absence of controls for the racial or ethnic composition of the population. It seems likely that countries with large 50-10 ratios are also characterized by racial and ethnic cleavages and, in particular, that immigration has been an important cause both of rising differentials between the low end and the middle of the wage distribution and of increasing racial and ethnic fractionalization across OECD countries since the 1960s. The 50-10 ratio might thus be conceived as a proxy for the representation of minorities at the bottom of the wage distribution, and racist or anti-immigrant attitudes may be the causal mechanism behind the negative association between the 50-10 ratio and redistribution. We do not conceive this interpretation of the effect of lower-half wage distribution as a fundamental challenge to our theoretical framework, for the notion of social affinity/hostility – whether based on class or ethnicity – is an integral part of this framework. Of course, we would prefer to know the extent to which the effect of lower-half wage inequality operates through identities (and conflicts) that are racial or ethnic.

The first column of Table 4 (model 5) reports the results we obtain when we include a time-invariant measure of ethnic fractionalization in our model. The particular fractionalization index used here is the one developed by Alesina and collaborators. Based on the data from the late 1990s, this index is one minus the Herfindahl index of ethnolinguistic group shares in the population. It thus measures the probability that two people, drawn at random from the population, will be from different ethnic groups (Alesina et al. 2003). The time-invariant character of this index is, of course, a major drawback for our purposes. Perhaps more importantly, the Alesina index does not speak directly to the critical question of where ethnic minorities fall along the income/wage distribution. To illustrate this point, Switzerland (0.53) receives a higher score than the US (0.49) on the Alesina index. Though it seems quite plausible that ethnic fractionalization helps explain the limited extent of redistributive social spending in Switzerland, the Swiss experience clearly does not fit the mold of middle-income voters turning against redistribution because ethnic minorities are over-represented among the poor. To the best of our knowledge, however, there is no other index that is obviously better than the Alesina index on either of these counts.

[Table 4 about here]

As measured by the Alesina index, ethnic fractionalization actually turns out to have a positive coefficient in the error-correction framework adopted here, but the coefficient does not come close to being statistically significant. For our purposes, what is most noteworthy about these results is that the effects of 90-50 and 50-10 wage ratios actually become stronger when we control for ethnic fractionalization. This is particularly true for the 50-10 wage ratio. It is also noteworthy that we obtain a positive and significant coefficient for voter turnout when we include ethnic fractionalization in our model. The effects of the other variables are largely unchanged (and government partisanship remains insignificant).

In model 6 (also presented in Table 4), we substitute a measure of immigration for the ethnic fractionalization variable. Based on the data collected by Rafaela Dancygier, this variable combines two separate immigration-related measures. For Australia, Canada, and the US, the figures refer to the proportion of the population that is foreign-born, while for the other ten countries in our sample they refer to the non-citizen proportion of the population. The Dancygier dataset provides annual observations on a fairly regular basis from 1980 onwards and includes earlier observations for some countries. The results presented in Table 4 are based on linear interpolation of missing data between observations and extrapolation of other data based on a no-change assumption. (Again, we intend to revisit the issue of missing data for the next iteration of this paper).

Surprisingly, the immigration variable turns out to have a positive and significant coefficient. There are at least two possible explanations for this puzzle. First, the observed effect might be spurious if EU member states have experienced more immigration over the last one or two decades relative to the other countries in our sample (Australia, Canada, and the US) and have also (for different reasons, not adequately specified by our model) engaged in more redistribution. This would seem particularly plausible to the extent that relatively skilled labor has migrated between EU member states. The second –

and, we think, more plausible – explanation is that immigration increases the share of the population that qualifies for social benefits (on account of higher rates of unemployment and lower pay among immigrants) and that this effect has not (yet?) been offset by the kind of political backlash that we might have expected. We cannot resolve the immigration puzzle here. Again, the important point for our present purposes is that controlling for the share of the population made up of non-citizens or foreign-born does not alter our finding that dispersion of the upper half of the wage distribution is associated with more redistribution while dispersion of the lower half is associated with less redistribution.

Public Opinion

Our theory postulates that the structure of inequality shapes the preferences of middle-income voters for redistribution and that this accounts for the 90-50 and 50-10 effects that we observe in the macro-level analysis presented above. To the extent that existing survey evidences allows us to do so, we are, of course, keen to test empirically the claims that upper-half inequality makes middle-income voters more supportive of redistribution while lower-half inequality makes them less supportive. In this section, we present some preliminary evidence based on ISSP surveys.

Over the period 1985-99, there were six different rounds of ISSP surveys that asked respondents whether or not they agreed with the statement, "it is the responsibility of the government to reduce the differences in income between people with high incomes and those with low incomes." Respondents were given five options: "strongly agree," "agree," "neither agree nor disagree," "disagree" and "strongly disagree."[15] At one point or another, these surveys were administered in ten of the thirteen countries included in the preceding analysis (with Belgium, Denmark, and Finland not represented), but the frequency of the surveys varies by country. Every one of the six surveys was conducted in Australia, the UK, and the US, but for the Netherlands and Switzerland we only have one survey (from 1987). Altogether, the dataset we have assembled from the ISSP consists of 38 country-year observations of responses to this question.

To create a summary measure of public support for redistribution, we coded survey responses on a scale from 1 for "strongly disagree" to 5 for "strongly agree" and calculated an average score for each country-year. Though it should be possible to restrict the analysis of ISSP data to respondents in the middle of the income distribution – which, after all, are those of interest to our theory – we have yet to explore this possibility. For now, we assume that average public support for redistribution is a reasonable proxy for support among middle-income voters.

To begin, Figure 1 addresses the association between public preferences for redistribution and actual redistribution, i.e., the dependent variable in the preceding analysis. This scatterplot is based on matching ISSP-based observations of public support with LIS-based observations of redistribution within two years of each other. In our view, the figure provides fairly compelling support for the proposition that public preferences matter to policy outputs in the realm of redistribution. This is particularly true if we disregard the two Swedish data points from the 1990s that stand out as outliers in Figure 1. (In response to a deep employment crisis and sharp increases in recipients of means-tested transfers as well as unemployment insurance, Swedish governments redistributed a good deal more in the 1990s than we would have predicted based on public opinion surveys.)

[Figure 1 about here]

Our claim that public preferences matter to redistributive policy outputs is consistent with the central thesis of Brooks and Manza (2007) and supported by their empirical analysis (cf. Alesina and Glaeser 2004: Ch. 7). However, Brooks and Manza have surprisingly little to say about why public support for social protection and redistribution varies across countries or over time. Again, our point here is that the structure of inequality helps explain public preferences.

Figures 2 and 3 plot average support for redistribution against 50-10 and 90-50 wage ratios. The pattern in Figure 2 is clear and strongly supports our theory: average support for redistribution is inversely correlated with the distance between the 50th and 10th percentile of the wage distribution. However, Figure 3 directly contradicts our theory, showing that average support for redistribution is also inversely correlated with the distance between the 90th and the 50th percentile. The five data points in the lower right-hand corner of Figure 3 all come from the US, but even if we were to exclude these data points we still do not see the predicted positive association between upper-half inequality and support for redistribution. The problem here has to do with the fact that 90-50 and 50-10 ratios are not independent of each other: higher 90-50 ratios tend to be associated with higher 50-10 ratios (and the US stands out as a particularly unequal country on both counts).

[Figures 2 and 3 about here]

To see the effect of the 90-50 ratio on public support, we must control for the effect of the 50-10 ratio, as we do in our macro-level analysis. Figure 4 illustrates this point by plotting average public support against the measure of pro-redistribution skew introduced in Table 1. Again, we simply subtract the 90-50 ratio from the 50-10 ratio to obtain this measure of skew. Though it is not terribly strong, there is clearly a positive association between skew and public support for redistribution in Figure 4.

[Figure 4 about here]

We recognize the need to engage in a more systematic analysis of the ISSP data. First, we would like to focus the analysis specifically on the preferences of middle-income voters. Secondly, we would like to take into account other potential determinants of individual-level preferences in a multivariate framework.

Government Partisanship

The absence of partisan effects is a striking and, from our point of view, troublesome feature of the results presented in Tables 2-4. We wish to emphasize again that government partisanship may matter in a number of ways that are not captured by our empirical models. For the next iteration of this paper, we plan to explore the idea that electoral dominance by either parties of the Left or the Right leads to strategic adaptation by parties on the opposite end of the Left-Right divide and hence diminished partisan effects. The most obvious way to test this hypothesis with our data would be to construct some measure of cumulative government partisanship over a more extended period of time (say, two election cycles).

Observing that cumulative Left government is associated with more redistribution while contemporaneously measured Left government is not would constitute fairly compelling evidence that Right parties adapt to Left dominance by pursuing more redistributive policies. In that case, the obvious next question becomes whether the structure of wage inequality helps to explain the incidence of Left government (or, more precisely, Left participation). We imagine that this investigation would take the form of estimating a model with government partisanship as the dependent variable and 90-50 and 50-10 wage ratios as right-hand-side variables.

Ultimately, we would like to develop and test the following set of propositions linking the structure of inequality to redistributive policy:

1. Compression of the lower half of the distribution and dispersion of the upper half render middle-income voters more supportive of redistribution.
2. When middle-income voters support redistribution, Left parties enjoy a strategic advantage over Right parties.
3. When Left parties enjoy a strategic advantage, Left-leaning and Right-leaning governments alike implement redistributive policies.

Final Note

The research agenda sketched above treats the structure of wage inequality as exogenous. This strikes us as fairly legitimate. At the same time, the question of why the structure of inequality varies across countries and over time is obviously an important one. We cannot possible tackle this complicated question here. For now, suffice it to note that Pontusson, Rueda and Way's (2002) analysis suggests that unionization and centralized wage bargaining separately and jointly tend to produce the kind of wage structure we have identified as uniquely favorable to redistributive politics. In their analysis, unionization and bargaining centralization are associated with more compressed 90-50 and 50-10 wage ratios, but the effects of these institutional variables on the 50-10 ratio is about three times as large as their effects on the

90-50 ratio. For fairly obvious reasons, solidaristic unions have been much more able to impose their distributive preferences among workers in the lower half of the wage distribution than among workers in the upper half of the distribution. (One of the obvious reasons is that workers in the lower half are more likely to be union members.)

Notes

[1] See Appendix 1 for a complete list of our variables and data sources.

[2] Note also that the measure adjusts for household size in the conventional LIS fashion (household income divided by the square root of the number of household members).

[3] One might object that this is but one instance of "second-order effects" that call into question the premise that the distribution of "market income" is unaffected by government policies (cf. Esping-Andersen and Myles 2007). Suffice it to say that pensions are undoubtedly the most significant instance of "second-order effects" and also the one that is easiest to set aside.

[4] See Appendix 2 for further details. Covering the period 1967-97, the dataset used by Bradley et al. (2003) and Iversen and Soskice (2006) mixes observations of redistribution through *transfers only* for Belgium, France, and Italy with observations of redistribution through *taxes and transfers* for eleven countries. (John Stephens has subsequently produced a corrected version). Our dataset differs from that used by these authors on three counts: (1) we do not include Italy (for which LIS data do not allow us to calculate gross household income); (2) our observations for Belgium and France are noticeably different; and (3) we include sixteen additional observations from the period 1998-2004.

[5] To our knowledge, only two contributions to the recent literature have sought to test the hypotheses proposed by Kristov, Lindert, and McClelland: Moene and Wallerstein (2003) and Schwabish, Smeeding, and Osberg (2006). In both cases, the dependent variable is social spending rather than redistribution. The models estimated differ from ours in many other respects as well. Moene and Wallerstein (2003) fail to find a significant difference between the effects of upper-half and lower-half wage inequality; Schwabish, Smeeding, and Osberg (2006) report results that are diametrically opposed to ours. We shall return to this discrepancy below.

[6] There are still quite a few missing years in the OECD wage data. We deal with this problem in the same manner as Iversen and Soskice (2006): for missing observations between two observations, we interpolate data based on the assumption of linear change; for missing observations at the beginning or end of a time series, we extrapolate the earliest or most recent observation available. For the next iteration of this paper, we will to verify that our results are robust to the exclusion of interpolated and extrapolated data.

[7] For France, the figures refer to 1994. For all the other countries they refer to either 1999 or 2000.

[8] See Fields and Ok (1999) for a review of relevant literature on income mobility.

[9] Mahler (2008) demonstrates that aggregate voter turnout and income skew in voting are indeed closely correlated on a cross-national basis.

[10] See Korpi (2006) for a recent restatement of the core tenets of power resources theory.

[11] In earlier analyses (Lupu and Pontusson 2009), we also included GDP per capita and female labor force participation as control variables (following Iversen and Soskice 2006). Neither of these variables proved to be statistically significant, and including them does not alter any of the findings reported below.

[12] This procedure is now automated using Vernby and Lindgren's (2009) `dvgreg` command in Stata.

[13] It deserves to be noted that we do obtain a significant effect of government partisanship when we restrict our analysis to the same time period as Iversen and Soskice's analysis, i.e., 1967-97 (see Lupu and Pontusson 2009). It is also noteworthy that the electoral-systems effect diminishes when we introduce separate measures for upper-half and lower-half wage inequality. As Lupu and Pontusson (2009) document, countries with PR tend to be characterized by more compressed 50-10 wage ratios relative to 90-50 wage ratios (as a result of centralized wage bargaining and strong unions).

[14] Again, our statistical model is the same as that used by Iversen. See endnote 4 regarding differences

between our datasets. Culpepper (2007) argues that VTS is not a good measure of skill specificity because it includes tertiary vocation training with a strong general skills component. This is a debate that we cannot pursue here.

[15] More recent ISSP surveys have reworded the question and eliminated the middle option ("Neither agree not disagree").

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Table 1. *Redistribution and wage structure, ca. 2000*

	Redistribution	90-50 ratio	50-10 ratio	Pro-redistribution skew
Denmark	38.6 (1)	1.74	1.44	0.30 (3)
Belgium	36.8 (2)	1.47	1.33	0.14 (8)
Finland	35.8 (3)	1.72	1.41	0.31 (2)
Sweden	35.6 (4)	1.69	1.39	0.30 (3)
Norway	28.7 (5)	1.44	1.39	0.05 (10)
Netherlands	28.6 (6)	1.76	1.66	0.10 (9)
Australia	27.3 (7)	1.80	1.65	0.15 (7)
France	27.1 (8)	1.93	1.59	0.34 (1)
Germany	25.6 (9)	1.84	1.59	0.25 (5)
UK	23.9 (10)	1.88	1.84	0.04 (11)
Canada	20.3 (11)	1.81	2.00	-0.19 (12)
USA	16.3 (12)	2.24	2.05	0.21 (6)
Switzerland	9.2 (13)	1.69	2.00	-0.31 (13)

Sources: See Appendix 1.

Table 2. *Determinants of redistribution, main results*

	(1)	(2)
90-10 ratio	-0.327	
	(1.205)	
90-50 ratio		14.40***
		(3.907)
50-10 ratio		-8.845***
		(2.657)
Turnout	0.0199	0.0520
	(0.0441)	(0.0383)
Partisanship (Right)	-0.787	-1.228
	(2.630)	(2.235)
Unionization	0.156***	0.184***
	(0.0411)	(0.0363)
Proportionality	11.87***	8.635***
	(2.946)	(2.676)
Veto points	-1.757***	-1.468***
	(0.444)	(0.393)
Unemployment	1.328***	1.209***
	(0.2000)	(0.175)
Constant	5.357	-6.681
	(7.164)	(9.727)
Observations	62	62
Number of countries	13	13
R^2	0.822	0.870
ρ	0.73	0.79

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1 (two-tailed tests)

Table 3. *Determinants of redistribution, controlling for national skill profiles*

	(3)	(4)
90-50 ratio	14.46***	16.34***
	(5.227)	(4.486)
50-10 ratio	-5.490*	-7.712**
	(3.169)	(2.962)
Turnout	0.0572	0.0709
	(0.0479)	(0.0437)
Partisanship (Right)	-3.084	-1.415
	(2.413)	(2.265)
Unionization	0.207***	0.203***
	(0.0437)	(0.0421)
Proportionality	8.961***	7.650**
	(2.975)	(2.914)
Veto points	-1.579***	-1.264***
	(0.501)	(0.455)
Unemployment	1.278***	1.179***
	(0.197)	(0.179)
Vocational training	0.0314	0.0426
	(0.0541)	(0.0483)
Constant	-14.77	-14.57
	(14.73)	(13.15)
Observations	47	62
Number of countries	13	13
R^2	0.896	0.872
ρ	0.77	0.88

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1 (two-tailed tests)

Table 4. *Determinants of redistribution, controlling for ethnic fractionalization and immigration*

	(5)	**(6)**
90-50 ratio	17.45***	15.51***
	(4.370)	(3.550)
50-10 ratio	-13.67***	-11.00***
	(4.133)	(2.525)
Turnout	0.0827*	-0.00314
	(0.0428)	(0.0383)
Partisanship (Right)	-0.527	0.0165
	(2.290)	(2.009)
Unionization	0.171***	0.160***
	(0.0370)	(0.0339)
Proportionality	8.184***	12.18***
	(2.675)	(2.654)
Veto points	-1.593***	-2.407***
	(0.398)	(0.453)
Unemployment	1.101***	1.220***
	(0.188)	(0.161)
Ethnic fractionalization	6.106	
	(4.057)	
Immigration		0.320***
		(0.0944)
Constant	-6.087	-4.033
	(9.639)	(8.947)
Observations	62	62
Number of countries	13	13
R^2	0.875	0.892
ρ	0.74	0.91

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1 (two-tailed tests)

Figure 1. *Redistribution and public support for redistribution*

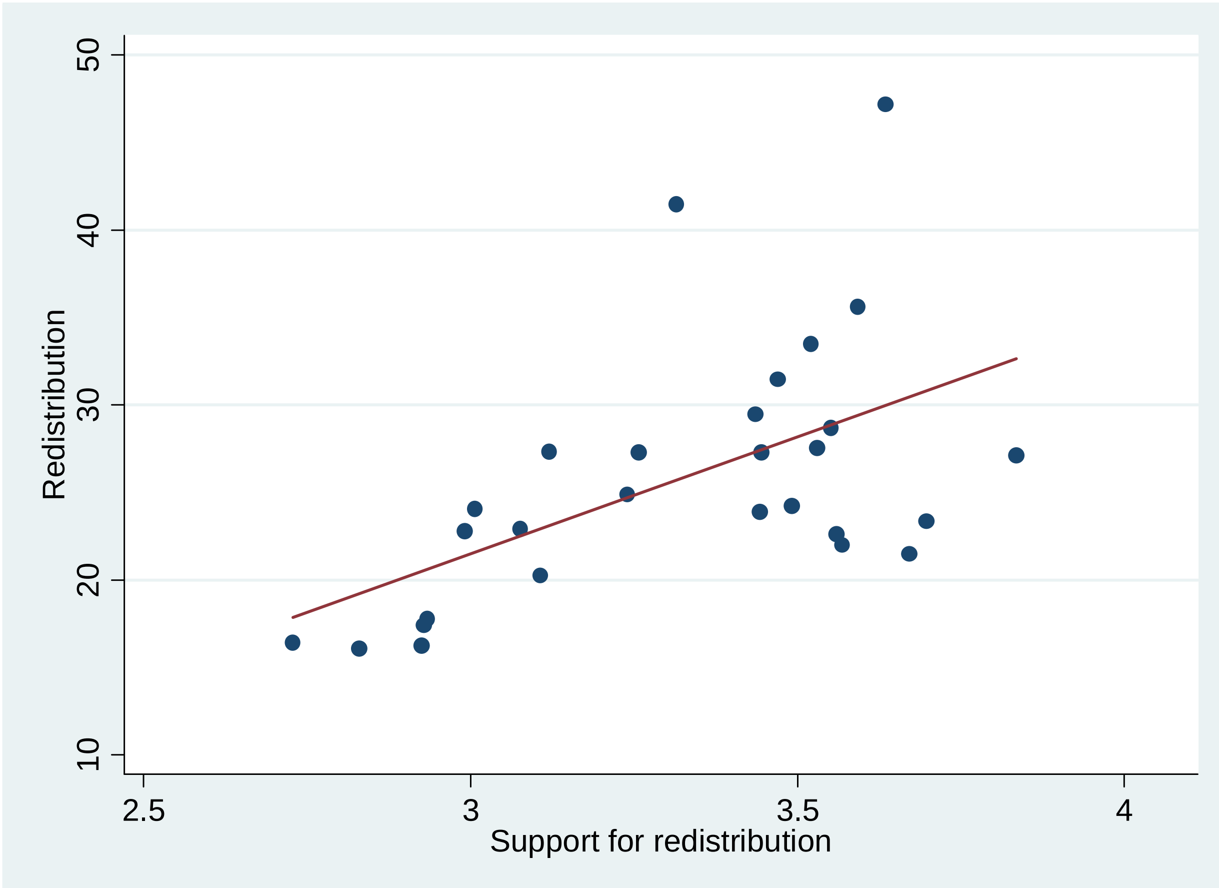


Figure 2. *Public support for redistribution and 50-10 wage ratios*

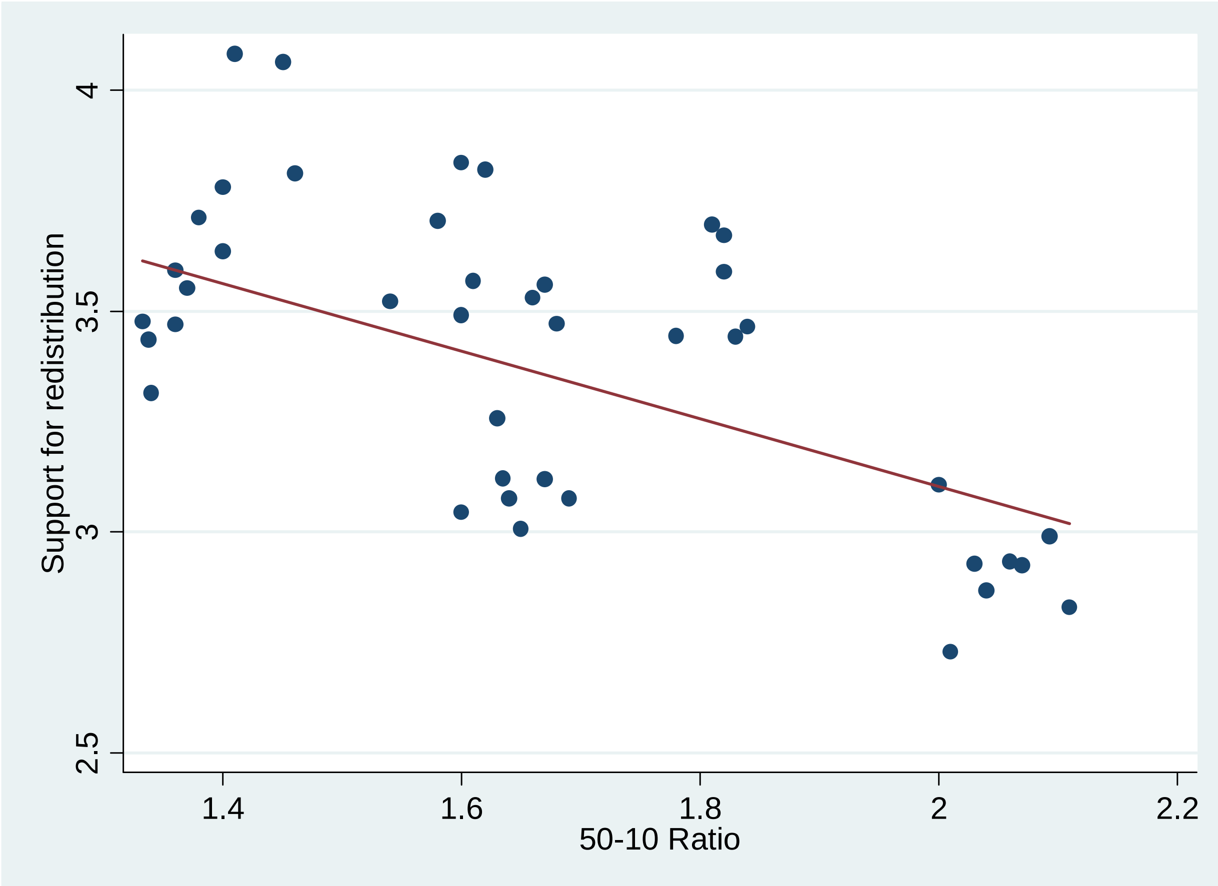


Figure 3. *Public support for redistribution and 90-50 wage ratios*

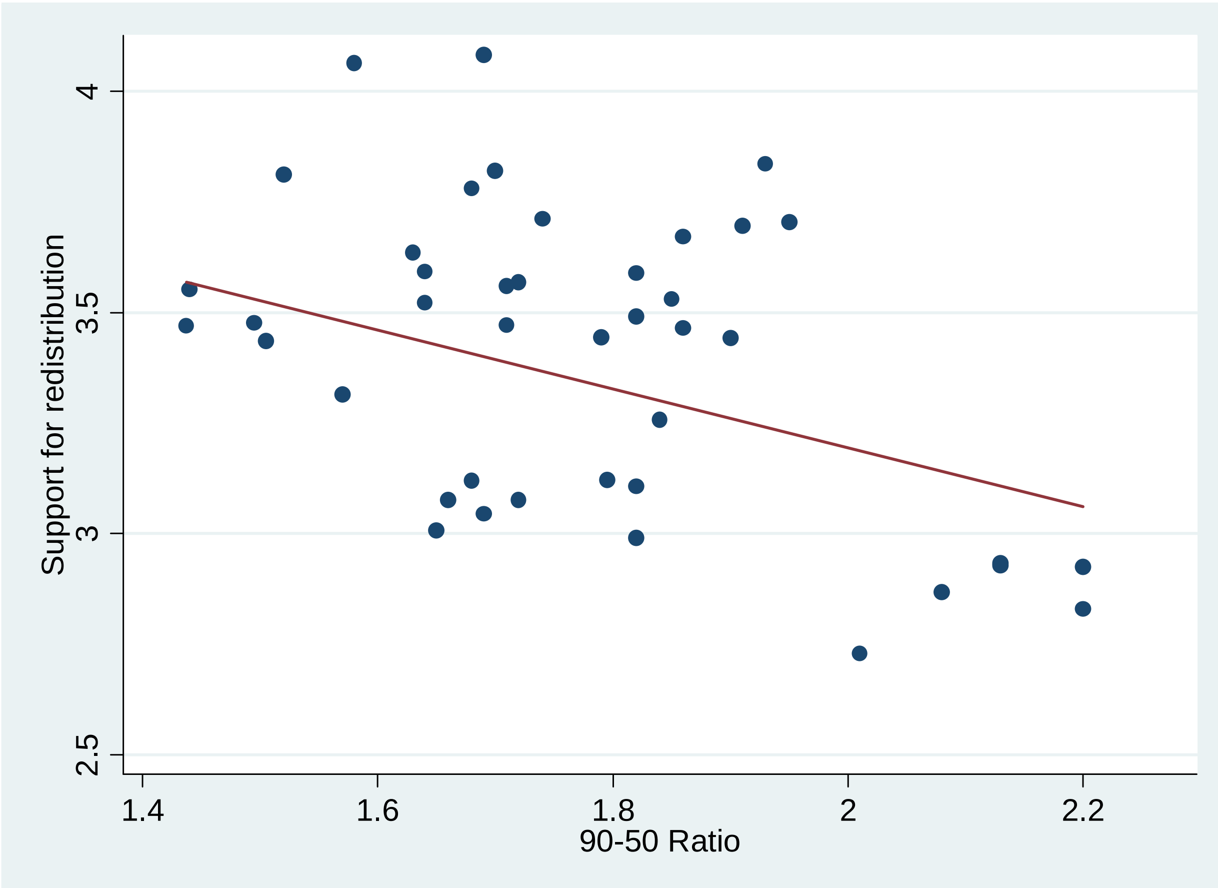
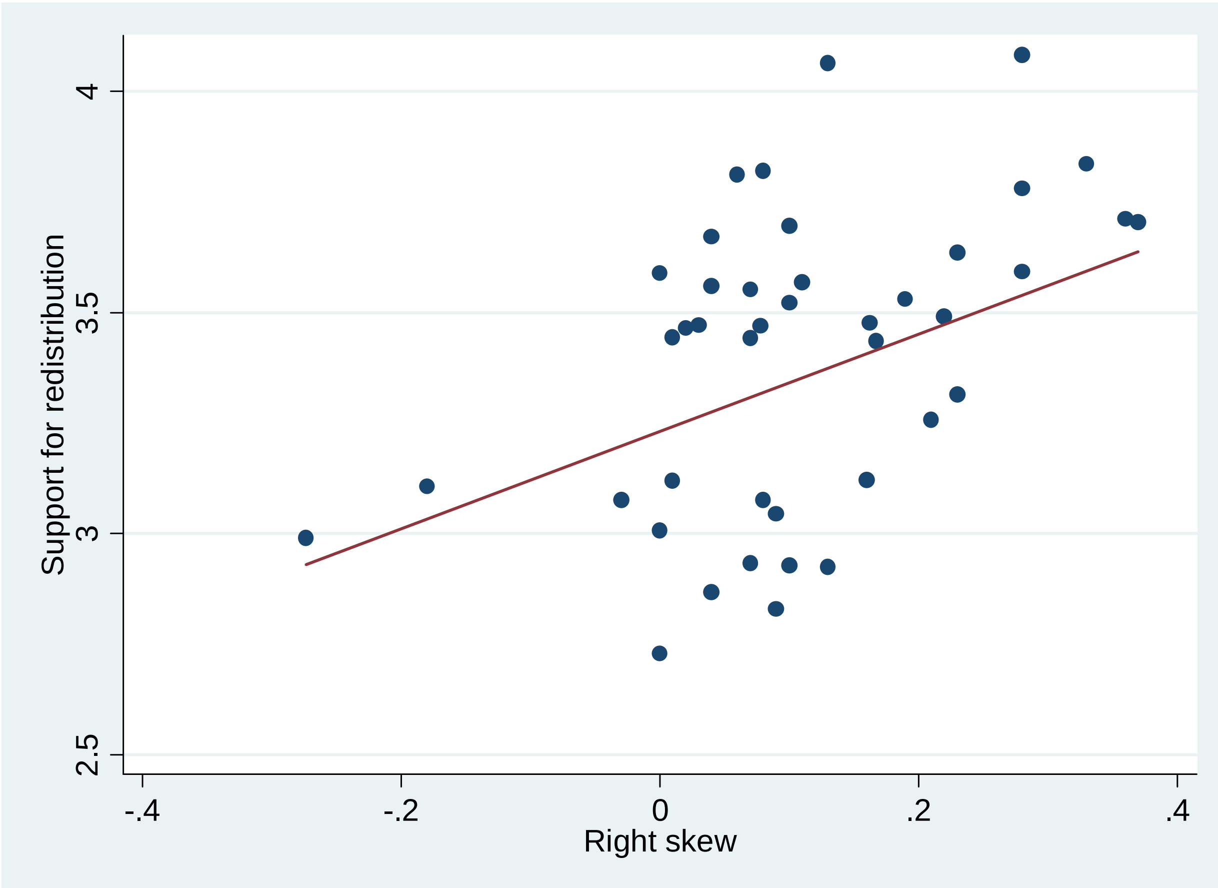


Figure 4. *Public support for redistribution and pro-redistribution skew*



Appendix 1: *Definitions and sources of variables*

Variable	Definition	Source
Redistribution	Percentage change in Gini coefficients as we move from gross market income (i.e., household income before taxes and transfers) to disposable income (i.e., income after taxes and transfers)	Lane Kenworthy, complemented with data from www.lisproject.org/publications/fiscalredistdata/fiscred.htm
90-50 Ratio	The earnings of a worker in the 90th percentile of the earnings distribution as a share of the earnings of the worker with a median income	OECD Database on Relative Earnings
50-10 Ratio	The earnings of the worker with a median income as a share of the earnings of a worker in the 10th percentile of the earnings distribution	OECD Database on Relative Earnings
Partisanship	An index of the partisan left-right "center of gravity" of the cabinet based on the average of three expert classifications of government parties' placement on a left-right scale, and weighted by their decimal share of cabinet portfolios (the index goes from left to right and is standardized here to vary between 0 and 1)	Cusack and Engelhardt (2002)
Proportionality	An index of proportionality measured as the square root of one-half the sum of squared absolute deviations of individual party seat shares from their respective shares of the vote (we standardize the index to vary between 0 and 1 and invert it such that larger values refer to higher levels of proportionality)	Gallagher (1991)
Unemployment	Annual rate of unemployment	Armingeon et al. (2006)
Unionization	Annual union density measure	Golden, Wallerstein, and Lange (2006)
Turnout	Turnout (as a percentage of eligible voters) in the most recent national election for each year	Armingeon et al. (2006)
Veto points	A composite measure of the number of constitutional veto points (that is, important decision nodes resulting from bicameralism, federalism, referenda, etc.) that exist in a political system	Huber, Ragin, and Stephens (1993)
Vocational training	Share of an age cohort engaged in secondary and tertiary vocational training	Iversen (2005)
Ethnic fractionalization	One minus the Herfindahl index of ethnolinguistic group shares	Alesina et al. (2003)
Immigration	Australia, Canada, US: proportion of the population that is foreign-born Other countries: non-citizen proportion of the population	Dancygier dataset

Appendix 2. *Observations of redistribution included in our dataset*

Country	Years
Australia	1981, 1985, 1989, 1995, 2000, 2003
Belgium	1992, 1997
Canada	1971, 1975, 1981, 1987, 1991, 1994, 1997, 1998, 2000
Denmark	1987, 1992, 1995, 2000, 2004
Finland	1987, 1991, 1995, 2000
France	1979, 1984, 1989, 1994
Germany	1973, 1978, 1981, 1983, 1984, 1989, 1994, 2000
Netherlands	1983, 1987, 1991, 1994, 1999
Norway	1979, 1986, 1991, 1995, 2000
Sweden	1967, 1975, 1981, 1987, 1992, 1995, 2000
Switzerland	1982, 1992, 2000, 2002
UK	1969, 1974, 1979, 1986, 1991, 1994, 1995, 1999
USA	1974, 1979, 1986, 1991, 1994, 1997, 2000, 2004