

# **Inequality, Mobility, and Social Affinity in the Politics Redistribution\***

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## **Abstract**

Against the current consensus among comparative political economists, we argue that inequality matters for redistributive politics in advanced capitalist societies, but it is the structure of inequality, not the level of inequality, that matters. Our theory posits that middle-income voters will be inclined to ally with low-income voters and support redistributive policies when the distance between the middle and the poor is small relative to the distance between the middle and the rich. We test this proposition with data from thirteen OECD countries and find that redistribution increases as the dispersion of wages in the upper half of the distribution increases relative to the dispersion of wages in the lower half of the distribution. We also present descriptive survey evidence on preferences for redistribution among middle-income voters that is consistent with our account of the causal mechanisms behind these results.

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 by 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?

While some scholars (most notably Milanovic 2000) have found significant effects of the level of inequality on redistribution, the current consensus seems to be that inequality does not matter directly for the politics of redistribution, at least not in a major way. The existing macro-comparative literature instead seeks to explain redistributive outcomes in terms of the effects electoral rules (Persson, Roland, and Tabellini 2007; Persson and Tabellini 2003), government partisanship (Bradley et al. 2003; Iversen and Soskice 2006), national skill profiles (Estevez-Abe, Iversen, and Soskice 2001; Iversen 2005; Iversen and Soskice 2001), unionization (Korpi 2006; Pontusson and Kwon 2006), linguistic-ethnic diversity (Alesina and Glaeser 2004) and religiosity (Scheve and Stasavage 2006).

We argue in this paper that inequality does matter for redistributive politics in advanced capitalist societies, but it is the *structure* of inequality, not the *level* of inequality, that matters. 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. These reasons have to do with the prospects of income mobility as well as social affinities between individuals occupying different places in the income distribution. Premised on the idea that the support of middle-income voters is critical to the introduction and successful implementation of redistributive policies, the theory we develop in the following pages boils down to the following

proposition: in the absence of overlapping ethnic cleavages, middle-income voters will be inclined to ally with low-income voters and support redistributive policies when the distance between the middle and the poor is small relative to the distance between the middle and the rich.

We test this proposition estimating an error-correction model of the kind specified by Iversen and Soskice (2006) with data from thirteen OECD countries over the period 1974-2004.<sup>1</sup> With change in redistribution as the dependent variable, this model specification provides a stronger basis for testing causal claims than the standard panel specification with the dependent and independent variables measured in levels. Controlling for government partisanship, electoral rules, and other variables identified as determinants of redistribution in the existing literature, we find that redistribution increases as the dispersion of wages in the upper half of the distribution increases relative to the dispersion of wages in the lower half of the distribution. We also present descriptive survey evidence on preferences for redistribution among middle-income voters that is consistent with our account of the causal mechanisms behind these results.

## **1. Theory and hypotheses**

The canonical median-voter model proposed by Romer (1975) and further developed by Meltzer and Richard (1981) clearly leaves something to be desired as far as the challenge of explaining cross-national variation in the extent of government redistribution is concerned. There are two possible sources of the model's limitations in this regard. One is that the model mistakenly assumes that the preferences of the median voter determine government policy. The second is that the model fails to adequately specify the determinants of the preferences of the median voter. Our discussion in this paper focuses on the latter question. Akin to standard median-voter models, we assume that the support of voters in the middle of the income distribution is usually critical to winning elections in liberal democracies. We also assume that most parties are sophisticated enough to recognize this and opportunistic enough to adjust their policies accordingly.<sup>2</sup>

In the Romer-Meltzer-Richard (RMR) model, 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, any increase of inequality that involves an increase in the between the median income and the mean income makes the median voter more supportive of redistribution.<sup>3</sup> 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, henceforth KLM), our theory attributes a more complex set of motivations to voters than the RMR model. First, we posit that preferences for redistribution are partly informed by other-regarding motivations, specifically by sympathies or antipathies for fellow citizens (or, more precisely, fellow residents). 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). Finally, our framework assumes that voters care about relative as well as absolute gains and losses.

Like the more recent model proposed by Iversen and Soskice (2006), the model of redistributive politics developed by KLM (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 hinges on electoral rules and partisan politics. For KLM, 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 sympathize with the poor or, alternatively, include the possibility of becoming poor in her cost-benefit calculus, and in turn support

redistribution. 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.

KLM's core idea calls into question the reliance on a single measure of inequality, such as the Gini coefficient or the 90-10 ratio, in recent empirical literature on the relationship between inequality and redistribution. To appreciate the political consequences of inequality, we need take account of the dispersion within each of the two halves of the income distribution. Following KLM's empirical strategy, we estimate separate effects for dispersion in the top half of the income distribution (measured as the 90-50 wage ratio) and the bottom half of the income distribution (measured as the 50-10 wage ratio).<sup>4</sup> We expect the former be positively associated with redistribution and the latter negatively associated with redistribution. Going beyond KLM, we also estimate a model with the ratio of the 90-50 ratio to 50-10 ratio as the measure of the structure of inequality. This measure, which we refer to as the *pro-redistribution skew* of the wage structure, captures the combinatorial logic of our core argument by relating dispersion in the upper half of the distribution to dispersion in the lower half of the distribution.

Our theory posits that middle-income voters must decide whether to ally with the poor or the affluent and that such decisions are based, at least in part, on perceptions of social distance informed by income differentials. In this context, knowing the distance to the poor is not very meaningful without also knowing the distance to the affluent. More specifically, an increase of the 90-50 ratio will only have the predicted effect of moving middle-income voters towards greater support for redistribution insofar as it is not offset by a corresponding increase of the 50-10 ratio. Put differently, our theory implies 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.

To clarify, our theory does *not* stipulate that dispersion of the bottom half of the distribution conditions the marginal effects of increasing the dispersion of the top half of the distribution or vice-versa. Rather, our theory is combinatorial in the sense that the distance to the poor and the distance to the affluent jointly determine the coalitional proclivities and policy preferences of middle-income voters.

Hence an empirical model interacting 90-50 and 50-10 ratios does not represent an appropriate test of our theory. (As a practical matter, multicollinearity among the variables of interests renders the estimation of such a model problematic.)

For illustrative purposes, Table 1 ranks the thirteen countries included in our analysis by the extent of redistribution in 2000 (or in years as close to 2000 as possible). The measure of redistribution used here, and in the analysis that follows, is the percentage change in the Gini coefficient (for working-age households only) brought about by taxes and transfers. In the second and third columns of Table 1, we report 90-50 and 50-10 wage ratios for the same year as the observation of redistribution and, in the last column, we report our measure of pro-redistribution skew. As illustrated by Figure 1, redistribution and skew turn out to be closely correlated on a cross-national basis. At one end of the spectrum, 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. 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 extent of redistribution. The US stands out as something of an outlier in Figure 1 in that the skew of its wage structure is quite pronounced, comparable to that of Australia and the Netherlands, yet the US government redistributes less than any other government except that of Switzerland.

[Table 1 and Figure 1 about here]

It should be noted that our theoretical framework deliberately equivocates 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 (cf. Converse 1964).

With regard to the self-interest part of the equation, our premise 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, taxes and transfers typically reduce the income of the share of households between the 50<sup>th</sup> and 60<sup>th</sup> percentiles of pre-fisc income distribution, but these losses are quite small by comparison to those of households above the 60<sup>th</sup> percentile. On the other hand, households below the 40<sup>th</sup> percentile typically benefit from the redistributive schemes in existence in most OECD countries. We therefore argue that someone in the 50<sup>th</sup> 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 30<sup>th</sup> percentile in the foreseeable future. Conversely, we expect the median income-earner to support cuts in redistribution to the extent that she believes that she might end up in the 70<sup>th</sup> percentile.

It seems very plausible to suppose that mobility between any two positions in the income distribution tends to increase as the distance between them declines. It follows that we should observe more mobility across the income distribution in countries with more compressed income distributions. Figure 2 presents some descriptive evidence suggesting that this is indeed case. Based on panel-data collected by the OECD (1996), the vertical axis of this figure shows the percentage in 1986 of low-paid workers – defined as workers earnings less than 65% of the median wage – whose earnings exceed 95% of the median wage in 1991. On the horizontal axis, we report inequality measured as the 90-10 wage ratio for 1991. At least for the subset of OECD countries for which this data is available, there is a strong negative correlation between mobility and inequality.<sup>5</sup>

[Figure 2 about here]

The salient point for us is that a smaller 50-10 ratio signifies a greater probability of downward mobility for the median income-earner while a smaller 90-50 ratio signifies a greater probability of upward mobility.<sup>6</sup> 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 50<sup>th</sup> to the 70<sup>th</sup> percentile or the losses associated with moving from the 50<sup>th</sup> to the 30<sup>th</sup> percentile will be correspondingly smaller. However, this objection assumes that voters are only concerned with *absolute* gains or losses

from redistribution. Studies have shown that voters care about their *relative* position in the income distribution (e.g., Luttmer 2005), such that an increase in the (perceived) probability of falling from the 50<sup>th</sup> to the 30<sup>th</sup> percentile should always be associated with a stronger preference for redistribution among voters in the 50<sup>th</sup> percentile of the income distribution.

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 involves *parochial altruism* or, in other words, altruism bounded by perceptions of common group membership or shared experience (cf. Fowler and Kam 2007). Akin to Shayo's (2009) emphasis on *perceived* distance between groups as a crucial parameter of social identification, 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 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.

Needless to say, income is not the only relevant source of social distance. Other obvious sources include race, ethnicity, and religious practices. Our emphasis on social affinity has much in common with Alesina and Glaeser's (2004) discussion of race and linguistic-ethnic fractionalization as obstacles to redistributive politics (cf. Gilens 2000; Luttmer 2001). From our perspective, however, what matters to the politics of redistribution is not racial or ethnic fractionalization per se, but rather the way that racial or ethnic cleavages map onto the income distribution.<sup>7</sup> All else equal, we would expect middle-income voters to feel more affinity with the poor at lower 50-10 ratios, but the effect will surely be attenuated to the extent that ethnic minorities are overrepresented among the poor. In our view, this is why the US turns out to have so little redistribution even though the wage distribution is, from a comparative perspective, characterized by a fairly pro-redistribution skew. (By the same logic, our theory predicts middle-income citizens will feel more affinity for the poor and prefer greater redistribution when ethnic minorities are concentrated among the affluent, a salient characteristic of at least some colonial or post-

colonial settings).

Our theory also implies that income skew in voting should matter for redistribution. As commonly noted, the variable that matters in the RMR 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 our framework, less income skew in voting 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.

Table 2 summarizes the macro-level hypotheses generated by the preceding discussion. Like many other theories in comparative political economy (e.g., Iversen 2005; Moene and Wallerstein 2001, 2003), the one sketched above has implications for individual preferences for redistribution as well as redistributive outcomes. In future work, we intend to derive the individual-level implications of our theory and test them empirically. For now, we focus on the macro implications.<sup>8</sup>

[Table 2 about here]

## **2. Variables, data, and methods**

### *The dependent variable*

Drawing on the Luxemburg Income Study (LIS), we measure redistribution 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”).<sup>9</sup> In keeping with much of the existing literature using LIS data (e.g., Bradley et al. 2003; Iversen and Soskice 2006; Kenworthy and Pontusson 2005), our analysis is restricted to working-age households (i.e., households headed by someone between the ages of 25 and 59). This is because generous public pension systems reduce the incentive for individuals to accumulate savings. In countries with very generous public pension systems, many retirees 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.<sup>10</sup>

Our dependent variable is constructed using 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. This produces a total of 76 country-year observations of redistribution in our dataset, but we have dropped the four earliest (pre-1974) observations in order to (a) achieve more balance across countries in the time periods covered by our analysis and (b) minimize the need to extrapolate values for some independent variables backwards in time.<sup>11</sup> We thus end up with 72 country-year observations over the period 1974-2004 in our analysis. In estimating an error-correction model, we lose the first observation for each country, so the  $N$  of our analysis is 59. For Belgium, we have only one observation, but for all other countries we have at least four observations and for Canada and the US we have as many as eight observations.

As shown in Appendix B, we observe substantial over-time variation in redistribution within the countries in our dataset. In nine out of thirteen countries, the difference between the minimum and maximum observations exceeds five percentage points and in four countries the difference exceeds ten percentage points. This type of variation is, of course, a necessary condition for estimating our empirical model, which treats change in redistribution as the dependent variable (and thus allows us to identify causal relationships with greater confidence). An analysis of variance indicates that within-country variation accounted for 82% of the total variation in our dependent variable.

### *Independent variables*

The test our argument about the political consequences of the structure of inequality, we use separate measures of upper-half and lower-half wage inequality derived from the OECD's database on relative earnings among full-time employees. Our main reason for using wages (rather than, for example, gross household income from LIS) as the basis for our measures of inequality is that the OECD database provides reasonably long, continuous series of annual observations.<sup>12</sup> Such time series are necessary for estimating error-correction models of the type we employ. Aside from data availability, however, wage inequality seems particularly relevant because we believe that voters care about their position in the income redistribution. As Pontusson and Rueda (2008) suggest, wage inequality constitutes a particularly visible form of market inequality: 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. Since our core hypothesis is combinatorial, our primary variable of interest is pro-redistribution skew, which captures dispersion in the top half of the wage distribution relative to that of the bottom half. We construct this measure simply by dividing the 90-50 wage ratio by the 50-10 wage ratio.

As indicated above, our theory stipulates that we should control for the distribution of ethnic minorities across the income distribution in estimating the effects of the structure of inequality. Unfortunately, existing indices linguistic-ethnic and religious fractionalization are not only time-variant – making their inclusion in dynamic models a rather dubious proposition – but also fail to speak to the critical question of where minorities are situated along the income distribution.<sup>13</sup> Instead, we include a cumulative measure of immigration in our empirical models as a proxy for the presence of poor minorities.<sup>14</sup> Consistent with Alesina and Glaeser's (2004: 175-77) discussion of immigration as a threat to European welfare states, the assumption here is that immigrants are overrepresented at the lower end of the income distribution such that their presence weakens support for redistribution among middle-income voters.

Like much of the recent empirical literature on the determinants of redistribution, we also include voter turnout in all of our empirical models and conceive this variable as an inverse proxy for income skew in voting. Our theory predicts that redistributive alliances between middle-income and poor voters will be more likely to prevail at higher levels of turnout (or, in other words, at lower levels of income skew in voting).<sup>15</sup>

We do not wish to argue that the social affinities shaped by the structure of inequality and ethnic-religious diversity alone provide an adequate explanation of temporal or cross-sectional variation in redistribution. In testing the hypotheses identified above, we want control for other potentially relevant variables. 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 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. To take into account partisan effects, the models we estimate include Cusack's cabinet-center-of-gravity 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). We standardize Cusack's index to vary between 0 and 1, with higher values representing more Right-leaning governments, and expect partisanship measured in this fashion to be negatively associated with redistribution.

While our main results pertain to short-term partisan effects, we discuss below the results of supplementary analyses with a cumulative measure of partisanship. Following Huber and Stephens (2001) and Bradley et al. (2003), the motivation behind this alternative specification is the idea that parties adapt strategically to social and political conditions that favor their opponents. For instance, we might expect sustained electoral success and government participation Left parties to induce Center-Right parties to adopt positions to embrace redistributive policies, in which case reversals of government and opposition might not be associated with significant policy shifts.

In addition to government partisanship, all of our empirical models 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 (see Korpi 2006). Unions do not typically organize workers at the very bottom of the wage distribution, but in most OECD countries their members are drawn disproportionately from the lower half of the distribution. Moreover, 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). For these reasons, we expect union density to be associated with greater redistribution. Controlling for the effects of unionization is for our purposes because unionization tends to be associated with compression of 50-10 wage ratios. In other words, union density must be included in the model to avoid spurious associations between wage compression and redistribution.

Another strand of the recent literature on the political economy of redistribution focuses on the question of why countries with proportional representation (PR) tend to have more redistributive governments than countries with majoritarian electoral rules. 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 the latter model, each party in the governing coalition has a strong incentive to reward its core constituency with spending financed by taxing all voters, but it is not altogether obvious why this should lead to more redistribution in PR countries.

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 disappear once we control for the effect of government partisanship.

In order to control for the electoral-systems effect identified by these authors, we include Gallagher’s (1991) measure of proportionality in our empirical models. This index ranges from 0 (pure proportionality between vote and seat shares) to infinity as disproportionality increases.<sup>16</sup> 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.

The existing literature suggests that we must also control for national skill profiles. Iversen and Soskice (2001) argue convincingly that individuals with more specific skills are more likely to support social spending and show that “vocational training share” (VTS), measured by the share of an age cohort engaged in secondary and tertiary vocational training, is correlated with government spending on income transfers on a cross-national basis. In addition, Estevez-Abe, Iversen, and Soskice (2001) show that VTS is associated with compression of wage differentials on a cross-national basis (169-78) and Iversen (2005: 148-54) reports a strong positive effect of VTS on redistribution among working-age households. In short, skill specificity could be the lurking variable behind any positive bivariate association between compression of the lower half of the wage distribution and redistribution.

In the aforementioned works, Iversen and his collaborators use VTS data from the period 1980-95. One of the two components of Iversen’s VTS index, graduates of tertiary vocational programs in percent of the tertiary age cohort, cannot be updated because UNESCO no longer publishes the requisite data on graduates of tertiary vocational programs. However, it is possible to construct continuous and apparently consistent time series on the percentage of secondary-school students enrolled in vocational programs for the thirteen countries included in our analysis over the entire period 1980-2005. This will

serve as our measure of vocational-training intensity in the analysis that follows. Aside from data availability, it is arguably a better measure than Iversen’s in that it sidesteps the complicated question of the kind of skills that are acquired through tertiary vocational training programs.<sup>17</sup> It is certainly an appropriate control variable for our purposes, given that our primary concern here is to ensure that national skill profiles do not confound the relationship between 50-10 wage ratios and redistribution.

All of our models include two additional variables not directly associated with any competing theory of redistributive politics. The first of these is the unemployment rate. More readily than any other variable, 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). The second control variable of this kind is the female labor-force participation rate. Iversen and Soskice (2006) include this variable on the grounds that “women’s participation in the labor market is likely to affect redistributive spending because it entitles some women to benefits (...) for which they would otherwise not be eligible” (174). A second motivation, equally important from our point of view, is the fact that women are more likely to be part-time workers than men, meaning that their hourly wages are generally lower. As the OECD data on relative wages pertain to full-time employees, female labor-force participation is likely to pick up effects of inequality, particularly low-end inequality, that are not captured by wage ratios.

### *The statistical model*

The regression results we report below 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 ( $P_{i,t-1}$ ) that cause redistribution to deviate from the status quo. The model incorporates a scalar  $\rho$ , which captures the speed with which levels of redistribution respond to changes in government policy:

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

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:

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

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

The right-hand side of the model thus includes a constant followed by the cumulative effect of the independent variables over the period  $N$  between the current observation of redistribution and the previous one ( $s$  represents the lags in years). 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 (2006), we estimate this model with different values of  $\rho$  and present the results for the version that explains the most variance.<sup>18</sup>

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, not between them. We therefore follow Vernby and Lindgren's (2009) recommendation of using weighted least squares to account for both sources of heteroskedasticity.<sup>19</sup>

The key point to keep in mind is that the coefficient estimates we present below pertain to the cumulative effects of *levels* of our independent variables (appropriately discounted) on *changes* in the dependent variable. In estimating these effects, our statistical model draws on variation both across countries and within countries over time.<sup>20</sup>

### 3. Empirical results

Table 3 presents our estimation results. Models 1-3, our main results, are identical except for the way that wage inequality is measured. Including the 90-10 ratio as a measure of the level of inequality, model 1 provides a benchmark for the other two models. In model 2, we take account of the structure of inequality by replacing the 90-10 ratio with the 90-50 and 50-10 ratios. In model 3, we in turn replace the 90-50 and 50-10 ratios with our combinatorial measure of pro-redistribution skew.

[Table 3 about here]

Regardless of the particular measure of wage inequality we use, the effects of the other variables in these models are by and large consistent. In all three models, we fail to observe any significant effects of either government partisanship or societal investment in specific skills. On the other hand, we consistently obtain significant positive effects of unemployment and female labor-force participation. Once we control for the structure of wage inequality, we also observe a statistically significant positive association between unionization and redistribution.

The effects of electoral rules are less consistent across the models three models. As the existing literature leads us to expect, the coefficient for proportionality is always positive, but this coefficient fails to reach statistical significance in model 2. It is noteworthy that the size of the coefficient for proportionality drops by more than 30 percent once we control for the structure of wage inequality. This suggests that one of the reasons why countries with PR tend to be characterized by more redistributive government than countries with majoritarian electoral rules is that these countries are also characterized by more pro-redistribution skew of their wage structure (see Lupu and Pontusson 2009).

Turning to the variables of direct relevance to our theory, let us first consider the effects of inequality. In model 1, the coefficient for the 90-10 wage ratio has a negative sign (contrary to expectations that might be derived from the RMR model), but its size is only about half of the standard error. Consistent with our expectations, and the empirical results in Kristov, Lindert and McLelland (1992), model 2 shows that the effects of wage dispersion in the upper and lower halves of the wage distribution run in opposite directions: while dispersion of the upper half is associated with

more redistribution, dispersion of the lower half is associated with less redistribution. These effects are only significant with 90% confidence, but it is important to note that the 90-50 and 50-10 ratios are quite closely correlated with each other.<sup>21</sup> What is remarkable is that we are able to estimate separate effects of these variables at all.

Combining the 90-50 ratio and the 50-10 ratio into a single measure of the structure of inequality in model 3 eliminates the problem of multicollinearity. Again, our measure of pro-redistribution skew also captures the combinatorial logic of our core argument by relating 90-50 and 50-10 ratios to each other. Consistent with our expectations, the coefficient for skew is positive and highly significant in a statistical sense (better than 99% confidence). We hasten to point out that this does not prove that the effects of upper-half and lower-half wage inequalities matter jointly. It is possible that the effect of skew is simply an additive effect of the two component variables, with the improvement in statistical significance being a consequence of eliminating the multicollinearity between them. In the end, we prefer model 3 over model 2 for theoretical reasons.<sup>22</sup>

With regard to voter turnout, our results show a consistent positive effect on redistribution. This result confirms our hypothesis that the median voter at high levels of turnout tends to be closer to the poor than the median voter at low levels of turnout and therefore more inclined to support redistributive policies. As hypothesized, we also find that immigration is consistently associated with less redistribution, though this result is just shy of statistical significance in model 2. It is noteworthy that the effect of immigration diminishes considerably when we control for the structure of wage inequality. In model 1, the relative size of the immigrant population and the proportionality of the electoral system seem to absorb the effects of the structure of wage inequality. Finally, it deserves to be reiterated that models 2 and 3 both identify effects of the structure of inequality while controlling for two potential confounding variables, in addition to immigration: unionization (which is itself associated with redistribution), skill specificity (which is not associated with redistribution).

Based on model 3, the following simulations illustrate the substantive significance of the key variables identified by our theory. While holding the 90-50 wage ratio constant, reducing the 50-10 wage

ratio (i.e., compressing the lower half of the wage distribution) by one standard deviation increases pro-redistribution skew by 0.19. With all other variables at their means, this in turn increases redistribution by 4.48 according to our results. Conversely, holding the 50-10 ratio constant and reducing the 90-50 ratio (i.e., compressing the upper half) by one standard deviation decreases pro-redistribution skew by 0.10. With all other variables at their means, this translates in turn into a reduction of redistribution by 2.35. With all other variables again at their means, increasing turnout by one standard deviation increases redistribution by 4.34 while increasing immigration by one standard deviation reduces redistribution by 2.62. Keeping in mind that the mean level of redistribution in our dataset is 25.98, the effects that we have identified must surely be considered substantively as well as statistically significant.

Models 4 and 5 present analyses designed to probe the robustness of our main results. First, model 4 replicates model 3 while dropping six country-year observations that constitute outliers in the sense that their standardized residuals are more than 1.75 standard deviations away from the mean.<sup>23</sup> Second, model 5 replicates model 3 with de-trended variables. A methodological concern that the comparative political economy literature rarely addresses is the potential for non-stationarity, which could induce spurious correlations. We use an augmented Dickey-Fuller test to check for a trending non-stationary process in our data (Davidson and MacKinnon 1993). The results of these tests are reported in Appendix C. Four of our variables in model 3 (female labor-force participation, immigration, unionization, and vocational training) do appear to trend over time. We correct for this by de-trending these variables using a moving-average process with one lag.

Dropping the outliers does not alter any of our findings except the effects of electoral rules. In model 4, we no longer observe any systematic association between proportionality and redistribution. The positive effects of union density, unemployment, female labor-force participation, voter turnout, and pro-redistribution skew remain strong and highly significant, as does the negative effect of immigration. And we still do not observe any significant effects of government partisanship or vocational training. The results with de-trended variables again confirm these main results.

Our models reveal two other noteworthy findings. Though some of our results confirm that more

proportional electoral rules promote redistribution, this electoral-system effect does not appear to be nearly as robust, or substantively significant, as previous studies lead us to expect. The complete absence of significant partisan effects is also noteworthy, calling into question recent applications of power resource theory as well as Iversen and Soskice's (2006) argument about the association between PR and redistribution. Recall that our approach focuses entirely on more or less short-term partisan effects: typically, the effects of different government constellations over a period of 3-6 years. In supplementary analyses, we have instead included a cumulative measure of government since 1960 and still do not obtain any significant partisan effects (results available upon request). On the other hand, it deserves to be noted that we do obtain a significant (short-term) effect of government partisanship when we restrict our analysis to the same time period as the analyses of Bradley et al. (2003) and Iversen and Soskice (2006), i.e., the period prior to 1998, and that we also obtain significant partisan effects in majoritarian countries for the entire period covered by our analysis (see Lupu and Pontusson 2009). The results presented above should not be taken to mean that government partisanship is of no consequence to redistributive politics. Nevertheless, the role of partisanship in the formulation and implementation of redistributive policies appears to be highly contingent and ought not to be taken for granted.

Our findings about the political consequences of the structure of wage inequality are remarkably robust. Our results clearly indicate that government policy tends to become more redistributive when earnings in the upper half of the wage distribution are more dispersed and that it tends to become less redistributive when earnings in the lower half are dispersed.<sup>24</sup> There is at least some evidence that these effects are combinatorial, i.e., that it matters how the dispersion in one half of the distribution compares to dispersion in the other half, but on this score our analysis is by no means definitive. Confirming other predictions of our theory, the results presented above also show that government policy tends to be more redistributive when voter turnout is high and, somewhat less decisively, that immigration is associated with less redistribution.

#### 4. The structure of inequality and middle-income preferences

Our theory postulates that the structure of inequality shapes the preferences of middle-income voters for redistribution and that this is why dispersion of the upper of the wage distribution and compression of the lower half are associated with more redistributive government. We are unable to think of any other, equally plausible account of the empirical regularities uncovered by the preceding analysis. At the same time, the preceding analysis does not speak directly to the causal mechanisms identified by our theory. Drawing on data from the International Social Survey Programme (ISSP), this section presents further evidence in support of these underlying mechanisms.

Over the period 1985-2000, eight rounds of ISSP surveys 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” and provided respondents with the following five response options: “strongly agree,” “agree,” “neither agree nor disagree,” “disagree” and “strongly disagree.”<sup>25</sup> At one point or another, the ISSP administered this survey question in all but one of the countries included in the preceding analysis of the determinants of redistribution, but the frequency of surveys varies a lot by country (on the one hand, zero for Belgium and one for Denmark and Finland and, on the other hand, eight for the UK and the US). Altogether, the dataset that we have assembled from the ISSP consists of 55 country-year observations of responses to the aforementioned survey question in twelve countries.

In keeping with our theoretical framework, the data presented in Figures 3 and 4 is restricted to survey respondents in the middle third of the distribution of household income (based on self-reported income).<sup>26</sup> As Goodrich and Rehm (2008) argue, creating a summary measure of support for redistribution by arbitrarily numerical scores to categorical responses and averaging these scores on a national basis is a rather dubious proposition.<sup>27</sup> To avoid the pitfalls identified by Goodrich and Rehm (2008), we measure support for redistribution instead as the percentage of middle-income respondents who either “strongly agree” or “agree” with the statement that it is the responsibility of the government to reduce income differences.

From the point of view of our theory, there are two separate questions to be asked of the ISSP data. First, do the preferences of middle-income voters matter to the redistributive policies pursued by governments? And, secondly, does the structure of inequality shape the policy preferences of middle-income voters? Figure 3 addresses the former question. To generate this figure, we have interpolated and, in a few instances, extrapolated LIS-based observations of redistribution to provide a match with each of the 55 ISSP-based observations of middle-income preferences for redistribution. Not surprisingly perhaps, we find a reasonably strong correlation between these two variables.

Three data points stand out in Figure 3. The only two Swiss data points stand out at the bottom end on account of the exceptionally low level of redistribution in Switzerland. Evidently, the Swiss case is distinguished less by the lack of public support for redistribution than by a distinctive disconnect between public opinion and policy outputs. Perhaps more importantly, the data point for the Netherlands in 2000 constitutes an influential outlier at the top end. According to our data, support for redistribution among Dutch middle-income voters fell sharply in the 1990s, but government policy did not fully register this shift.<sup>28</sup> Disregarding these three data points, the association between middle-income preferences and redistribution shown in Figure 3 is quite strong.

[Figure 3 about here]

Several recent contributions to the comparative study of welfare states and redistribution, most notably Brooks and Manza (2007), stress that public opinion matters (see also Schwabish, Smeeding, and Osberg 2008). Relative to this literature, two things distinguish our approach: first, we consider the preferences of middle-income voters to be critical from a comparative perspective and, secondly, we aspire to explain why the preferences of middle-income voters vary across countries and over time. (These authors have surprisingly little to say as to why public support for social protection and redistribution varies across countries.)

Figure 4 addresses the consequences of the structure of inequality for the preferences of middle-income voters. Previously on the horizontal axis, support for redistribution among middle-income ISSP respondents appears on the vertical axis of this figure, being plotted against our measure of pro-

redistribution skew of the wage structure. Though the fit is far from perfect, there is clearly a positive association between skew and middle-income support for redistribution in Figure 4. It is noteworthy that the seven observations at the bottom of the vertical axis and in the middle of the horizontal are all US observations. Again, we attribute the uniquely low level of middle-income support for redistribution in the US to the role that race plays in separating middle-income voters from the poor (cf. Gilens 2000; Alesina and Glaeser 2004).

[Figure 4 about here]

As the US case illustrates most vividly, the structure of wage inequality is not the only thing that matters to the preferences for redistribution among middle-income voters. In future work, we plan to estimate a multivariate model with support for redistribution, as measured above, as the dependent variable. We also plan to explore individual-level implications of our theory using ISSP data. Our theory implies that changes in the structure of inequality should have different consequences for the preferences of poor, middle-income, and affluent respondents.

For our present purposes, we would very much like to be able to add middle-income support for redistribution into the regression models presented in the previous section. Our theory predicts that this variable should be strongly associated with redistribution and that its inclusion would make the effects of immigration as well as the structure of inequality (and immigration) go away. Unfortunately, incorporating ISSP data into our main dataset would require a huge amount of extrapolation, for we do not have any ISSP observations of public opinion preceding (or coinciding with) 23 of out the 59 observations of redistribution that serve as the dependent variable in the preceding macro-comparative analysis. Given the significant over-time variation in support for redistribution that we observe in the ISSP data, extrapolating the data that far backwards in time – sometimes over a decade – seems entirely inappropriate.<sup>29</sup> For the time being, then, Figures 3 and 4 must suffice as evidence that the causal mechanism identified by our theory is plausible.

## 5. Conclusion

While the existing literature on the comparative political economy of redistribution has paid some attention to the effects of voter turnout, it has paid very little attention to the effects immigration and has entirely missed the effects of the structure of inequality. The theory we propose posits that compression of income differentials is associated with income mobility and also generates social affinities between individuals occupying different positions in the income distribution. Our macro-comparative analysis indeed shows that government policy tends to become more redistributive when earnings in the upper half of the wage distribution are more dispersed and that it tends to become less redistributive when earnings in the lower half are dispersed. It also shows that voter turnout is associated with more redistribution and that immigration is associated with less redistribution.

We do not wish to juxtapose our argument about the structure of inequality to the emphasis on racial and ethnic divisions in some of the recent literature on redistribution, but we want to amend the thrust of this literature by insisting that racial and ethnic divisions matter to the extent that racial and ethnic minorities occupy distinctive positions in the income distribution. Thus conceived, race and ethnicity become an integral part of the structure of inequality.

Our theory focuses on how the structure of inequality shapes the coalitional proclivities and policy preferences of middle-income voters, but the claim that proximity matters to self-interested calculations as well as solidaristic norms obviously applies to voters across the entire income spectrum. We do not argue that governments are exclusively or even primarily responsive to the policy preferences of middle-income voters. Our argumentation focuses on middle-income voters because we believe that their preferences for redistribution are more likely to vary across countries and over time than the preferences of the poor or the affluent. This is, of course, itself a proposition that can and should be empirically verified. The descriptive analysis of survey data presented here provides preliminary evidence in support for our claim that the structure of inequality shapes the preferences of middle-income voters and that these preferences in turn have an impact on government policy, but more work must be done on this score.

Quite legitimately, macro-comparative analyses of the type presented above invite questions about endogeneity. Our theory and empirical analysis treat the structure of inequality as an exogenous variable that causes changes in redistribution, but is it not equally plausible to treat redistribution as a cause of the structure of inequality? Specifically, it seems quite plausible to suggest that the redistributive government policies promote compression of the 50-10 wage ratio by setting a floor for competition in the labor market, i.e., by providing citizens with a “social wage” that employers must exceed in order to attract any workers at all. It is important to recognize in this context that the dependent variable in our analysis is not the level of redistribution, but rather change in the level of redistribution, and that our measures of the structure of wage inequality (as well as the other independent variables) temporarily precede our measures of the dependent variable. In addition, there is very little reason to suppose that redistributive government policies might cause dispersion of the upper half of the wage distribution. Quite the contrary: Hibbs (1987) argues persuasively that high marginal tax rates serve to compress market wages in the upper half of the distribution by reducing the net gains associated with any given wage increase, meaning that reverse causality should work in the opposite direction. In short, we do not believe that endogeneity represents a serious challenge to our interpretation of the relationship between wage-structure skew and redistribution.

What, then, explains cross-national and temporal variation in the structure of inequality? From a comparative perspective, labor-market institutions provide the most obvious point of departure for addressing this question. According to Pontusson, Rueda, and Way’s (2002) analysis, unionization and bargaining centralization are both associated with more compressed 90-50 as well as 50-10 wage ratios, but the effect of these institutional variables on the 50-10 ratio is about three times as large as their effect on the 90-50 ratio. Even operating within highly centralized bargaining systems, solidaristic unions appear to have been far better 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. The reasons for this have to do partly with the distribution of union members across the wage hierarchy and partly with the willingness of employers to accommodate union preferences. In any case, Pontusson, Rueda and

Way's (2002) analysis suggests that strong unions and centralized bargaining tend to generate the kind of wage structure that we have identified as uniquely favorable to redistributive politics.

Labor-market institutions alone clearly do not provide a sufficient explanation of the structure of wage inequality, let alone the structure of inequality in a more expansive sense. In the American case, the last twenty-five or thirty years have witnessed not only a marked increase of overall inequality, but also a major shift in the structure of inequality as a small group of very affluent people at the top of the income distribution have gained relative to everyone else (Piketty and Saez 2003). Based on our data on relative wages among the full-time employed, pro-redistribution skew increased from 0.99 in 1973 to 1.12 in 2001. While the decline of unions constitutes an important backdrop to the growth and reconfiguration of inequality, other developments – most obviously increased corporate reliance on equity markets and the emergence of new forms of executive compensation – have played a major, arguably more proximate causal role.

An implication of our theory is that we should observe a leftward shift of American politics as a result of increasing skew in the distribution of income. As we have noted repeatedly, the concentration of racial-ethnic minorities in the bottom of the income distribution represents an obvious obstacle to the formation of redistributive coalition of poor and middle-income voters in the US, but there is very little evidence to suggest that the overlap of race and income became more pronounced or that racism became a more salient feature of American politics in this period. While the Bush era clearly contradicts the implication of our theory, Obama's victory in 2008 might conceivably be interpreted as a political manifestation of long-term changes in the structure of inequality.

## Notes

<sup>1</sup> The thirteen countries are Australia, Belgium, Canada, Denmark, Finland, France, Germany, the Netherlands, Norway, Sweden, Switzerland, the UK, and the US. These are all the countries that satisfy both of the following criteria: (a) the Luxembourg Income Study allows us to measure redistribution in the manner specified below, and (b) the OECD database on relative earnings provides multiple observations of earnings inequality prior to the last LIS-based observation of redistribution.

<sup>2</sup> We do not rule out the possibility that parties also have core constituencies they seek to please or, alternatively, that they also have ideologically-motivated policy commitments. All of our empirical models include government partisanship as an independent variable, although we never find any statistically significant effects of this variable. This finding certainly does not mean that government policy can simply be “read off” the preferences of the median voter, as caricatures of median-voter models would have it. For our purposes, however, it seems both possible and necessary to bracket the question of how political systems aggregate and translate voter preferences into policy outputs.

<sup>3</sup> It is commonplace to argue that any mean-preserving increase of inequality will make the median voter more supportive of redistribution, but this is technically not correct. It is possible to imagine scenarios in which falling relative incomes at the bottom of the distribution are offset by increasing relative incomes at the top, such that median and mean incomes remain constant.

<sup>4</sup> We discuss the measurement of these variables in more detail below.

<sup>5</sup> Based on a more systematic analysis, Aaberge et al. (2002) show that relative earnings mobility in Denmark, Norway, and Sweden exceeded relative earnings mobility in the US over the 1980s. Blau and Kahn’s (1996) comparative analysis of gender pay differentials captures the underlying logic here very well. Using data from the 1980s, they show that the average female wage falls in a lower percentile of the overall wage distribution in Sweden than the US (the 28<sup>th</sup> as compared to the 32<sup>nd</sup> percentile). Because the Swedish wage distribution is much more compressed, however, the gender gap (average female wage in percent of average male wage) is much smaller in Sweden.

<sup>6</sup> Other factors will surely affect perceptions of the prospects of income mobility as well. 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). The fact that our empirical model is dynamic, focusing on change over time, enables us to sidestep the question of how to incorporate ideological constructions of income mobility.

<sup>7</sup> Though their fractionalization measures fail to capture it, Alesina and Glaeser (2004) clearly agree with this point. 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).

<sup>8</sup> To our knowledge, only two previous contributions to the comparative literature have explored separate effects of inequality in the upper and lower halves of the distribution. While Moene and Wallerstein (2003) fail to find any significant difference between the effects of low-end and high-end inequality, Schwabish, Smeeding, and Osberg (2006) report results that are very different from ours (more on this below). Consistent with our theory, Corcoran and Evans’ (2008) analysis of local spending on public education in the US finds that that low-end inequality is associated with less spending while high-end inequality is associated with more spending.

<sup>9</sup> Our measure adjusts for household size in the conventional LIS fashion (household income divided by the square root of the number of household members).

<sup>10</sup> 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.

<sup>11</sup> See Appendix B for further details. Note that including the four early observations does not significantly alter our results.

<sup>12</sup> There are still annual observations of wage ratios missing from our dataset. We have dealt 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 two other variables – immigration and vocational training – we have very little or no data for the 1970s. In both cases, we have extrapolated 1980s observations backwards in time as needed.

<sup>13</sup> To illustrate the latter problem, Switzerland receives a higher score than the US (0.53 compared to 0.49) on the index of ethno-linguistic fractionalization proposed by Alesina et al. (2003). Ethnic fractionalization might have something to do with the limited extent of redistributive social spending in Switzerland, but 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.

<sup>14</sup> Based on the data collected by Rafaela Dancygier, our immigration variable refers to the proportion of the population that is foreign-born for Australia, Canada, and the US, while it refers to the non-citizen proportion of the population for the other countries in our sample.

<sup>15</sup> Mahler (2008) demonstrates that aggregate voter turnout and income skew in voting are indeed closely correlated on a cross-national basis.

<sup>16</sup> Gallagher's index is the square root of the sum of squared absolute deviations of individual party seat shares from their respective shares of the vote divided by two.

<sup>17</sup> Culpepper (2007) argues persuasively that tertiary vocational training typically involves a strong general-skills component and does not have the same political-economic meaning as secondary vocational training.

<sup>18</sup> Note that our model (and equations) differs slightly from Iversen and Soskice's in that our independent variables are never contemporaneous with our dependent variable ( $s$  is never equal to zero). In our specification, each observation of redistribution is determined by temporally prior observations of the independent variables. This makes theoretical sense since policies rarely have immediate effects and also minimizes potential endogeneity problems. (Our results are substantively the same when we instead utilize the contemporaneous model proposed by Iversen and Soskice [2006].)

<sup>19</sup> This procedure is now automated using Vernby and Lindgren's (2009) `dvgreg` command in Stata. To

use this command without contemporaneous observations of our independent variables (see previous note), we used one-year lags of all our variables.

<sup>20</sup> Given the dynamic structure of the model, and the fact that the time dimension of our data is smaller than the cross-sectional dimension ( $T < N$ ), it would be inappropriate to include fixed effects (Hsiao 2005; Kristensen and Wawro 2007). The same holds for time-invariant variables, such as the commonly-used index of constitutional veto points (Huber, Ragin, and Stephens 1993) or the Alesina-Easterly index of linguistic-ethnic fractionalization (Alesina et al. 2003).

<sup>21</sup> The correlation between the 90-50 and 50-10 ratios is 0.71 for our entire sample and as high as 0.89 for some countries.

<sup>22</sup> Estimating a model with the difference between the 90-50 and the 50-10 ratio as the measure of skew (rather than the ratio of the 90-50 to the 50-10), we obtain results that are very similar to those of model 3: the statistical significance of the difference measure only clears the 95% level, but the R-square of this is only marginally lower than that of model 3. It should also be noted that if we add an interaction term to model 2, the coefficients for the 90-50 ratio, 50-10 ratio, and interaction term all fail to reach statistical significance. (This is not surprising given that the correlation between the 90-50 ratio and the interaction term is 0.88 while the correlation between the 50-10 ratio and the interaction term is 0.94). Finally, we have also estimated models that include the 90-10 wage ratio along with skew: the skew effect is stronger than in model 3 and the effect of the 90-10 ratio remains insignificant, though the sign of the coefficient becomes positive. These results are available upon request.

<sup>23</sup> The six outliers are Germany 1983; Netherlands 1987, 1994, and 1999; Norway 1986; and Sweden 1992.

<sup>24</sup> Our results stand in marked contrast to those reported by Schwabish, Smeeding, and Osberg (2006), who find that 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 less social spending. Quite reasonably, these authors interpret their results to mean that eligibility for means-tested social spending increases with the 50-10 ratio and that the political influence of the affluent increases with the 90-50 ratio.

There are many possible reasons for the discrepancies between our results and theirs. Among other things, their models include a variable designed to capture public support for social spending and redistribution. In our 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.

<sup>25</sup> Surveys including this question were fielded as part of the Role of Government module (1985, 1990, 1996), the Social Inequality module (1987, 1992, 1999), and the Environment module (1993, 2000). More recent ISSP surveys still ask the same question, but the middle option (“Neither agree nor disagree”) has been eliminated and, as a result, the responses are not comparable.

<sup>26</sup> Where household income was coded using a limited set of income bands, we include respondents from the income bands that get us as close as possible to the middle third of respondents.

<sup>27</sup> Goodrich and Rehm’s critique is directed against Brooks and Manza (2007), but applies equally to Schwabish, Smeeding, and Osberg (2006).

<sup>28</sup> The Netherlands is the only one of these countries in which the overall redistributive effect of taxes and transfers actually declined from the early 1980s to the late 1990s (see Kenworthy and Pontusson 2005: 455).

<sup>29</sup> Schwabish, Smeeding, and Osberg (2006) apparently circumvent this problem by using an item on interpersonal trust from the World Values Survey as a proxy for support for redistribution.

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**Table 1. Redistribution and wage structure (ca. 2000)**

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|             | <b>Redistribution</b> | <b>90-50 ratio</b> | <b>50-10 ratio</b> | <b>Pro-redistribution skew</b> |
|-------------|-----------------------|--------------------|--------------------|--------------------------------|
| Denmark     | 38.6 (1)              | 1.74               | 1.44               | 1.21 (3)                       |
| Belgium     | 36.8 (2)              | 1.47               | 1.33               | 1.11 (6)                       |
| Finland     | 35.8 (3)              | 1.72               | 1.41               | 1.22 (1)                       |
| Sweden      | 35.6 (4)              | 1.69               | 1.39               | 1.22 (1)                       |
| Norway      | 28.7 (5)              | 1.44               | 1.39               | 1.04 (10)                      |
| Netherlands | 28.6 (6)              | 1.76               | 1.66               | 1.06 (9)                       |
| Australia   | 27.3 (7)              | 1.80               | 1.65               | 1.09 (7)                       |
| France      | 27.1 (8)              | 1.93               | 1.59               | 1.21 (3)                       |
| Germany     | 25.6 (9)              | 1.84               | 1.59               | 1.16 (5)                       |
| UK          | 23.9 (10)             | 1.88               | 1.84               | 1.02 (11)                      |
| Canada      | 20.3 (11)             | 1.81               | 2.00               | 0.91 (12)                      |
| USA         | 16.3 (12)             | 2.24               | 2.05               | 1.09 (7)                       |
| Switzerland | 9.2 (13)              | 1.69               | 2.00               | 0.85 (13)                      |

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*Note:* For France, the figures refer to 1994; for all the other countries they refer to either 1999 or 2000.

See Appendix A for variable definitions and sources.

**Table 2.** *Hypothesized effects on redistribution*

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| <b>Variable</b>         | <b>Hypothesized effect</b> |
|-------------------------|----------------------------|
| 90-50 ratio             | +                          |
| 50-10 ratio             | -                          |
| Pro-redistribution skew | +                          |
| Poor ethnic minorities  | -                          |
| Income skew in voting   | -                          |

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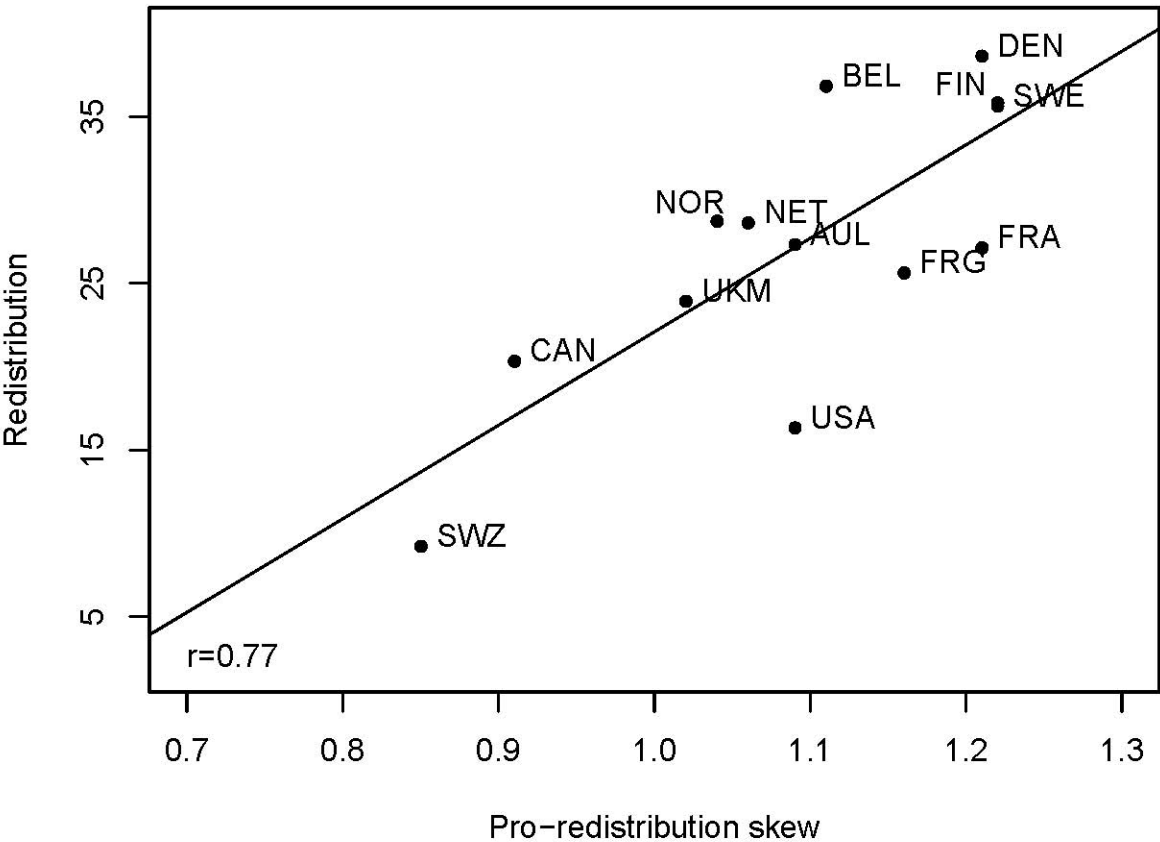
**Table 3. Determinants of redistribution**

|                         | (1)                  | (2)                  | (3)                  | (4)                   | (5)                  |
|-------------------------|----------------------|----------------------|----------------------|-----------------------|----------------------|
| 90-10 ratio             | -0.866<br>(1.708)    |                      |                      |                       |                      |
| 90-50 ratio             |                      | 11.44*<br>(6.253)    |                      |                       |                      |
| 50-10 ratio             |                      | -5.774*<br>(3.437)   |                      |                       |                      |
| Pro-redistribution skew |                      |                      | 15.29**<br>(5.950)   | 17.78***<br>(4.683)   | 14.10**<br>(6.011)   |
| Voter turnout           | 0.206***<br>(0.0569) | 0.203***<br>(0.0544) | 0.192***<br>(0.0522) | 0.180***<br>(0.0420)  | 0.194***<br>(0.0515) |
| Immigration             | -0.327***<br>(0.116) | -0.204<br>(0.122)    | -0.252**<br>(0.107)  | -0.235***<br>(0.0856) | -0.248**<br>(0.106)  |
| Partisanship            | -3.064<br>(2.876)    | -2.460<br>(2.736)    | -2.390<br>(2.682)    | -1.501<br>(2.160)     | -2.405<br>(2.691)    |
| Unionization            | 0.0927<br>(0.0584)   | 0.162**<br>(0.0629)  | 0.127**<br>(0.0497)  | 0.153***<br>(0.0405)  | 0.120**<br>(0.0498)  |
| Proportionality         | 7.599**<br>(2.937)   | 4.687<br>(2.981)     | 5.220*<br>(2.799)    | 3.347<br>(2.216)      | 4.954*<br>(2.862)    |
| Vocational training     | -0.0320<br>(0.0785)  | 0.0246<br>(0.0772)   | -0.0354<br>(0.0507)  | -0.0619<br>(0.0390)   | -0.0189<br>(0.0527)  |
| Unemployment            | 1.367***<br>(0.216)  | 1.233***<br>(0.214)  | 1.251***<br>(0.202)  | 1.143***<br>(0.163)   | 1.213***<br>(0.199)  |
| Female labor            | 0.335***<br>(0.104)  | 0.236**<br>(0.108)   | 0.234**<br>(0.105)   | 0.223**<br>(0.0882)   | 0.244**<br>(0.0988)  |
| Constant                | -22.57*<br>(11.64)   | -30.90*<br>(15.88)   | -33.34***<br>(8.776) | -33.48***<br>(7.053)  | -32.57***<br>(8.385) |
| Observations            | 59                   | 59                   | 59                   | 53                    | 59                   |
| R-squared               | 0.829                | 0.846                | 0.848                | 0.913                 | 0.845                |
| Rho                     | 0.590                | 0.640                | 0.650                | 0.700                 | 0.620                |

Standard errors in parentheses

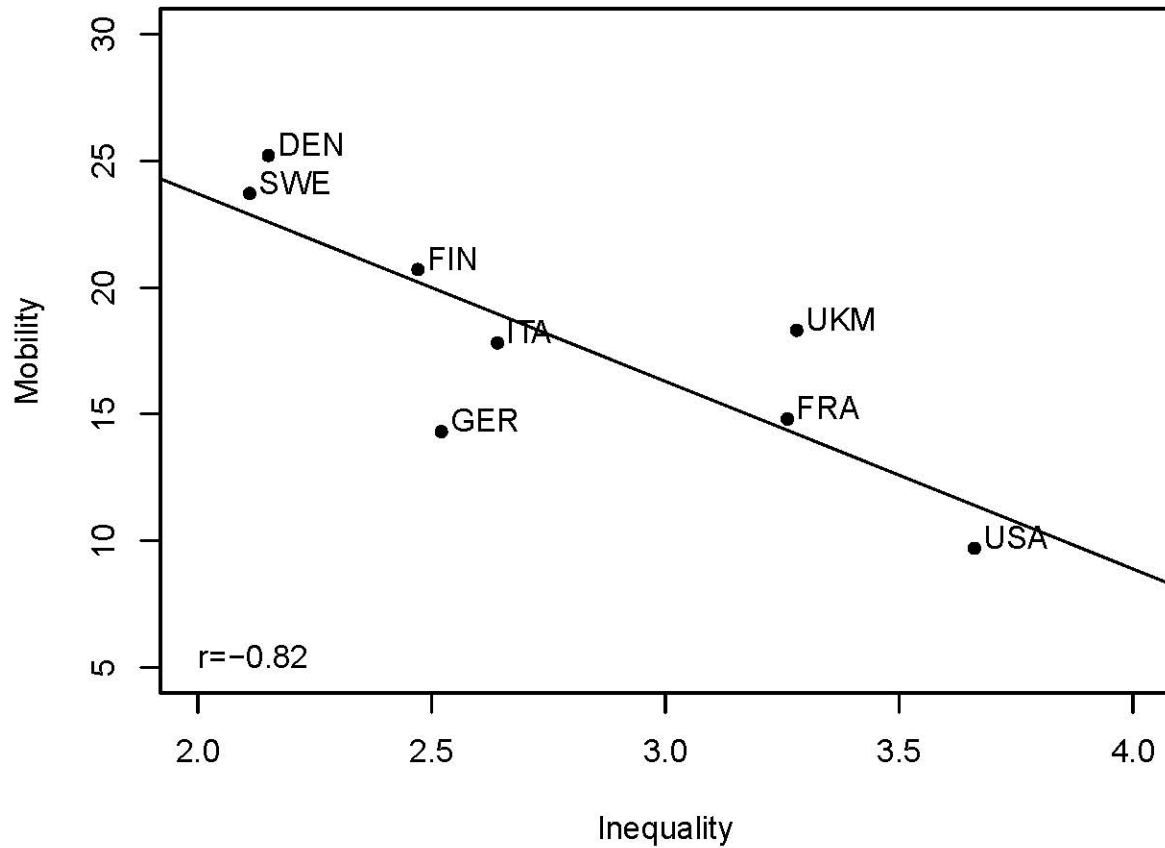
\*\*\* p&lt;0.01, \*\* p&lt;0.05, \* p&lt;0.1

Figure 1. Redistribution and pro-redistribution skew (ca. 2000)



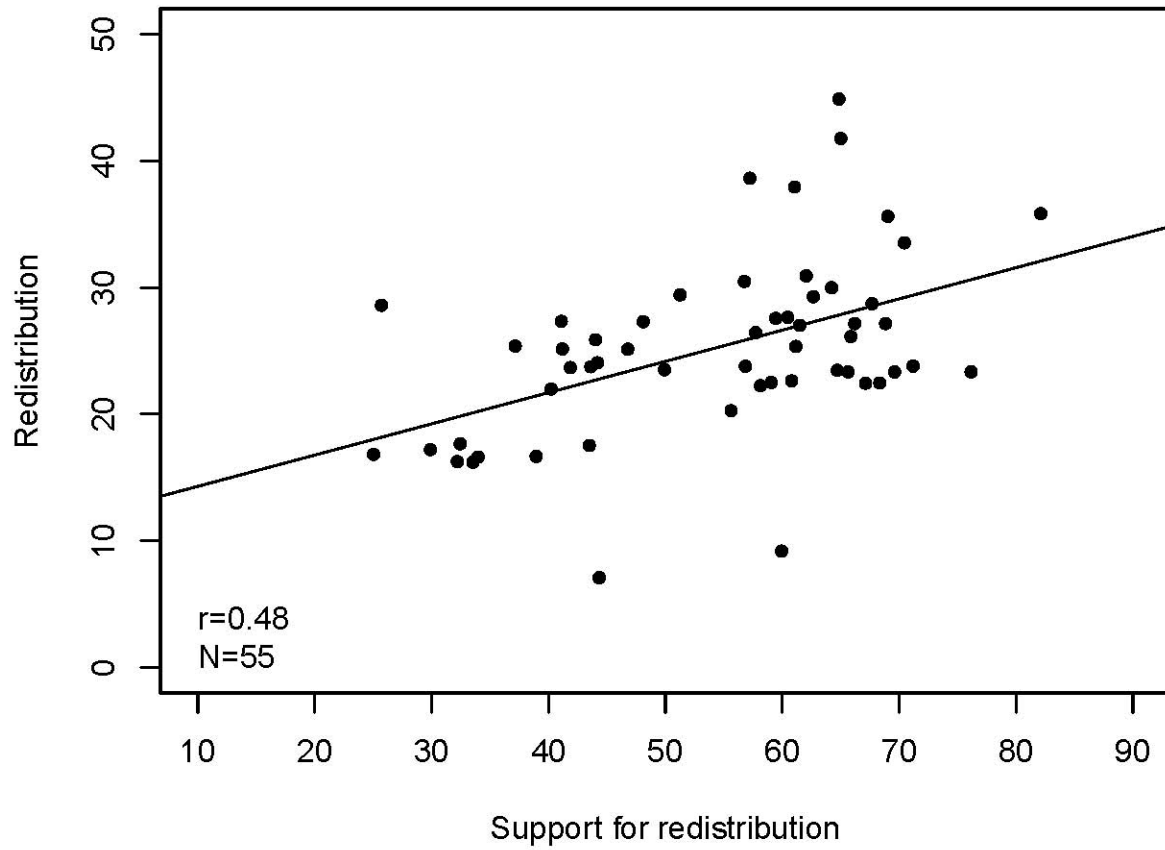
Note: See Appendix A for variable definitions and sources.

**Figure 2.** *Earnings mobility and inequality, 1986-91*



*Note:* Mobility is measured as the percentage of low-paid workers, defined as workers earnings less than 65% of the median wage, in 1986 whose earnings exceed 95% of the median wage in 1991 (OECD 1996: 89-90). Inequality is the 90-10 ratio in 1991 (see Appendix A).

**Figure 3.** *Public opinion and redistribution*





**Appendix A. Definitions and sources of variables**

| <b>Variable</b>            | <b>Definition</b>  | <b>Source</b>   |
|----------------------------|--|---|
| 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  |
| 90-10 ratio                | The earnings of a worker in the 90th percentile of the earnings distribution as a share of the earnings of a worker in the 10th percentile of the earnings distribution  | OECD Database on Relative Earnings  |
| 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  |
| Female labor               | The proportion of working-age women in the labor force   | OECD Labor Force Statistics database  |
| Immigration                | Australia, Canada, US: proportion of the population that is foreign-born; Other countries: non-citizen proportion of the population  | Dancygier dataset   |
| 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)  |
| Pro-redistribution skew    | Ratio of the 90-50 ratio to the 50-10 ratio  | OECD Database on Relative Earnings  |
| 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)   | Kenworthy dataset and Mahler and Jesuit (2006)  |
| Support for redistribution | Proportion of middle-income respondents to ISSP surveys (identified as those falling in the middle third of the distribution of respondent household incomes) who said they “strongly agree” or “agree” 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” | ISSP survey modules (Environment 1993, 2000; Role of Government 1985, 1990, 1996; Social Inequality 1987, 1992, 1996) |
| Unemployment               | Annual rate of unemployment  | Armingeon et al. (2006)   |
| Unionization               | Annual union density measure   | Golden, Wallerstein, and Lange (2006)   |
| Vocational training        | Enrollments in vocational training programs in percent of secondary-school enrollments   | Iversen (2005) and UNESCO database  |
| Voter turnout              | Turnout (as a percentage of eligible voters) in the most recent national election for each year  | Armingeon et al. (2006)   |

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

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| <b>Country</b> | <b>Years</b>   | <b>Minimum</b> | <b>Maximum</b> |
|----------------|--|----------------|----------------|
| Australia      | 1981, 1985, 1989, 1995, 2000, 2003                     | 22.9 (89)      | 27.3 (95)      |
| Belgium        | 1992, 1997   | 36.8 (97)      | 38.9 (92)      |
| Canada         | (1971), 1975, 1981, 1987, 1991, 1994, 1997, 1998, 2000 | 25.6 (94)      | 17.8 (81)      |
| Denmark        | 1987, 1992, 1995, 2000, 2004                           | 38.8 (04)      | 28.2 (87)      |
| Finland        | 1987, 1991, 1995, 2000, 2004                           | 42.5 (95)      | 33.1 (87)      |
| France         | 1979, 1984, 1989, 1994                                 | 27.9 (89)      | 23.9 (79)      |
| Germany        | (1973), 1978, 1981, 1983, 1984, 1989, 1994, 2000       | 27.6 (00)      | 15.4 (78)      |
| Netherlands    | 1983, 1987, 1991, 1994, 1999                           | 33.5 (87)      | 25.8 (91)      |
| Norway         | 1979, 1986, 1991, 1995, 2000                           | 31.5 (95)      | 20.2 (86)      |
| Sweden         | (1967), 1975, 1981, 1987, 1992, 1995, 2000             | 41.7 (92)      | 35.6 (00)      |
| Switzerland    | 1982, 1992, 2000, 2002                                 | 12.7 (02)      | 6.1 (82)       |
| UK             | (1969), 1974, 1979, 1986, 1991, 1994, 1995, 1999       | 27.3 (86)      | 14.2 (74)      |
| USA            | 1974, 1979, 1986, 1991, 1994, 1997, 2000, 2004         | 17.8 (94)      | 15.4 (74)      |

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*Note:* Country-years in parentheses are not included in our analyses.

**Appendix C. Augmented Dickey-Fuller tests for non-stationarity**

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|                         | <b>Before de-trending</b> |                | <b>After de-trending</b> |                |
|-------------------------|---------------------------|----------------|--------------------------|----------------|
|                         | <b>Chi-square</b>         | <b>P-value</b> | <b>Chi-square</b>        | <b>P-value</b> |
| Female labor            | 23.19                     | 0.723          | 47.50                    | 0.012          |
| Immigration             | 8.18                      | 1.000          | 225.37                   | 0.000          |
| Partisanship            | 65.78                     | 0.000          |                          |                |
| Proportionality         | 67.58                     | 0.000          |                          |                |
| Pro-redistribution skew | 56.44                     | 0.001          |                          |                |
| Unemployment            | 49.15                     | 0.008          |                          |                |
| Unionization            | 7.72                      | 1.000          | 57.97                    | 0.001          |
| Vocational training     | 14.70                     | 0.981          | 73.05                    | 0.000          |
| Voter turnout           | 124.36                    | 0.000          |                          |                |

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*Note:* Augmented Dickey-Fuller tests are conducted following Maddala and Shaowen (1999) and with a single lag. The null hypothesis is a non-stationary trending process.