

# INSTITUTIONS, PARTISANSHIP, AND INEQUALITY IN THE LONG RUN

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## INTRODUCTION

**P**OLITICAL scientists are keenly interested in examining cross-country differences in income inequality. Recent quantitative studies have added to this tradition by examining whether institutions like centralized wage bargaining and political factors like government partisanship are correlated with differences in income inequality between advanced industrial countries. Wage bargaining centralization, it is suggested, can directly affect the pretax distribution of income by compressing wage differentials.<sup>1</sup> Partisanship may affect the income distribution to the extent that governments of the left are more likely to engage in redistributive policies. This could involve redistributive policies like taxes and transfers that reduce levels of posttax inequality relative to pretax inequality. While this has been emphasized less in recent political science work, the current partisan orientation of a government may also logically have an effect on future pretax income inequality. This would be true to the extent that policies like progressive

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<sup>1</sup>Centralized wage bargaining will arguably affect earnings inequality and thus household income inequality to the extent that earnings represent a sizable portion of a household's income. While recognizing this distinction, throughout this article we refer more simply to the effect of centralized wage bargaining on "income inequality."

taxation influence capital accumulation or that policies like subsidizing public education allow poorer groups to acquire more human capital. To date, cross-country quantitative studies of income inequality and its political correlates have focused on the period since the beginning of the 1970s. Two reasons in particular account for this choice. First, the fact that income inequality has risen in some OECD countries but not in others is of obvious substantive importance. Contrasting the trends in a number of continental European countries with those in the United States, scholars have been interested in whether the high levels of inequality observed in the U.S. in recent years can be explained by the absence of centralized wage-setting institutions and by the weakness of the political left. A second reason for the focus of existing work is that even if scholars did want to examine the political determinants of inequality over a longer time span, they have lacked the data to do so. The earliest data in the widely used OECD database of earnings dispersion date from 1973. The earliest data from the Luxembourg Income Study date from 1979.<sup>2</sup>

In this article we make use of new data on top incomes, while also drawing on several underutilized long-run series on wage inequality. When combined with existing data on political institutions, as well as with political data that we ourselves have coded, we can investigate whether institutions like centralized wage bargaining are correlated with levels of income inequality for periods prior to the 1970s. We also consider the effect of government partisanship over the long run. One reason for considering a longer time span is that data on top income shares and wage inequality over the course of the twentieth century suggest that there has been significantly more variation within countries over time than there has been between countries. If we are going to suggest that certain partisan and institutional factors explain the current difference in inequality between two countries such as the United States and Sweden, to provide one example, then these same factors ought to be able to account for variation both within and between these two countries. We ought to be able to explain why the United States and Sweden were actually much less different from each other in terms of income inequality in the 1950s than they are today. In addition, by considering a longer time span we are able to examine whether within-country changes in institutions like wage bargaining centralization have

<sup>2</sup> For examples that use data from one of these two sources to consider the political correlates of income inequality, see Beramendi and Cusack 2008; Kenworthy and Pontusson 2005; Rueda and Pontusson 2000; Wallerstein 1999; Golden and Wallerstein 2006a; Pontusson, Rueda, and Way 2002; OECD 1997; Golden and Londregan 2006; and Iversen and Soskice 2006.

been associated with changes in inequality. Ideally, if a country adopted an institution like centralized wage bargaining in the 1930s or 1940s, we would like to know not only whether the presence of this institution was associated with low inequality in the 1980s or 1990s but also whether the initial introduction of the institution appears to have had a significant impact. If this is not the case, then the existing institutional explanation may need to be revisited.

Our empirical results suggest that centralized wage bargaining and left government appear to have little causal effect on inequality over the long run. This raises several possibilities that will require further elaboration. First, it may be that long-run changes in inequality are driven primarily by broad economic trends involving the “race between technology and education” and that political factors are ultimately of secondary importance.<sup>3</sup> A second possibility is that political factors do matter but not in the way that is often suggested. Rather than focusing on the formal centralization or decentralization of bargaining arrangements in two countries like the U.S. and Sweden, it may be that a closer investigation would show that despite their formal institutional differences, similar norms or beliefs prevailed regarding acceptable levels of pay inequality within the labor markets of those two countries during the 1950s. This conclusion is reinforced by at least one recent intriguing analysis of the U.S. postwar experience with income inequality.<sup>4</sup> A third, complementary possibility is that politics matters for inequality, but it may be that the most prominent political effects on inequality are generated by cataclysmic events involving economic crises or wars.<sup>5</sup>

The article proceeds as follows. We first introduce the top income shares data and highlight the potential questions they raise for comparative political economy. We present only a brief overview to aid in interpreting our statistical tests; those looking for a full presentation should consult the volume edited by Anthony Atkinson and Thomas Piketty.<sup>6</sup> We then consider two hypotheses regarding how partisanship and institutions may influence inequality. The first hypothesis is that governments of the left should be associated with lower income inequality. The second hypothesis is that centralized wage bargaining ar-

<sup>3</sup>The original use of the phrase is from Tinbergen 1975.

<sup>4</sup>Levy and Temin 2007.

<sup>5</sup>The potential effects of war on inequality have been emphasized both by those who point to economic effects involving destruction of fortunes, such as Piketty 2003, 2001 and Piketty and Saez 2006, 2003, as well as by those who emphasize the political repercussions of wars, including Gourevitch 1986; Rogowski and MacRae 2004; and Scheve and Stasavage 2008.

<sup>6</sup>Atkinson and Piketty 2007.

rangements should also be expected to reduce inequality by compressing pay differentials.

In the following section we report the results of our empirical tests using the top incomes data. These consider the correlation of centralized wage bargaining and government partisanship with three separate top incomes measures for the period 1916–2000. In our regressions we see no evidence of a statistically significant effect of left government on income inequality except with regard to very top incomes (the top 1 percent). Moreover, the substantive magnitude of this latter effect is very small, implying that it could account for only a small fraction of the large changes in income inequality that have occurred over the course of the twentieth century. For centralized wage bargaining we distinguish between the effect of three different types of wage bargaining: (1) decentralized wage bargaining, meaning wages are set at the firm level or in the absence of collective bargaining, (2) sectoral-level wage bargaining, and (3) centralized wage bargaining, meaning that wages are set at the peak or national level. Existing scholarship has placed the greatest emphasis on the effect of this third type of wage bargaining on inequality, with one reason being that peak-level wage setting can be associated with the adoption of “solidaristic pay” policies. We find some evidence that decentralized wage bargaining has been associated with higher levels of inequality than has sectoral wage bargaining, but this effect is driven by events of the last three decades. During the first three-quarters of the twentieth century there is little evidence of a difference between decentralized and sectoral wage bargaining. In light of previous scholarship, however, the most surprising part of our results is the lack of any evidence of an association of centralized, national-level wage bargaining with income inequality. Finally, in contrast to our finding on centralized wage bargaining, we do find evidence of a negative correlation between union membership and inequality, though further work is needed to establish whether this correlation reflects a causal relationship or instead one in which both the decision to join a union and levels of inequality are simultaneously determined by underlying economic forces.

Our results regarding the absence of an effect of fully centralized (peak-level) wage bargaining on inequality are surprising, given the emphasis that has been placed on this factor in previous work. One possible explanation is that centralized wage bargaining actually matters in the way commonly suggested but that it primarily affects segments of the income distribution not captured by the top incomes measures. However, we cite evidence from specific countries suggesting

that white-collar employees, some of whom could have household incomes in the top decile of the distribution, were often directly affected by centralized wage bargaining agreements. In addition, we demonstrate that a broad top incomes measure focusing on the income of those between the 90th and 99th percentiles of the income distribution is actually very highly correlated with (and therefore an excellent proxy for) the measure of wage inequality most commonly used in existing work: the ratio between the earnings of the 90th and 10th percentiles of the earnings distribution. For an individual country like Sweden, for which we have long series on both top incomes and wages, we can further demonstrate that our broad top incomes measure is remarkably highly correlated with a measure of wage inequality. We can, however, be less confident that our broad top incomes measure will serve as a good proxy for movements at the lower end of the earnings distribution, such as changes in the 50/10 ratio.

In order to further investigate our result regarding centralized wage bargaining, we extend our analysis in the next section by using individual country time series on wage inequality as well as top income shares to take a closer look at whether the adoption of centralized wage bargaining in several countries during the middle years of the twentieth century was associated with a downward structural break in income inequality. This is what one would expect if the idea of a causal effect of centralization on inequality is to be believed. The results are striking for the four countries we consider that adopted centralized wage bargaining during this period. In each of the four cases (Sweden, Denmark, the Netherlands, and Ireland) both wage and income inequality trended downward after the move to centralized bargaining, but it had already been trending downward well before this institutional change; and based on individual country regressions, we see little evidence that this trend was accentuated following the change.

As already noted, our empirical results call for a reconsideration of existing conclusions regarding the effect of government partisanship and centralized wage bargaining on income inequality. In a final section we develop this point by considering potential explanations for our results based on underlying economic or political processes.

#### DATA ON TOP INCOME SHARES

The inequality data used in this article come from two sources: selected time series of variation in wages across different types of workers and measures of top income shares. This section first describes the top in-

comes data that have been collected as part of a project that uses information from income tax returns to calculate the percentage of total pretax income earned by those at the top end of the income distribution in each country. We then consider whether these data are appropriate for testing hypotheses about the effect of government partisanship and centralized wage bargaining on inequality.

### TOP INCOME SHARES

In addition to the edited volume by Atkinson and Piketty, a number of papers on individual countries have already been published using statistics from national tax authorities to measure top income shares.<sup>7</sup> Our analysis makes use of the standardized dataset compiled by Andrew Leigh that uses a number of adjustments to correct (to the extent possible) for heterogeneity in measurement that might otherwise influence inferences about the evolution of top income shares.<sup>8</sup> This new data on income inequality have several advantages over existing measures of income inequality based upon household surveys. First, it results in inequality measures that are more (though certainly not perfectly) homogeneous across countries. Second, the top income shares data provide us with a much longer run view of the evolution of income inequality in different countries than what we find with existing databases. One constraint imposed by using tax data to measure inequality is that prior to World War II in most countries only a small fraction of households was subject to income taxation. This means that it is possible to use tax data to estimate the share of total income earned by the top 10 percent of households and by groups within the top 10 percent (top 1 percent, top 0.1 percent, and so on), but it does not allow one to produce a direct measure of developments at the bottom of the income distribution.

Currently, data on top incomes are available for thirteen advanced industrial countries. Figure 1 presents data on the top 1 percent income share in a single graph that may be useful for identifying trends over time. Several things are immediately apparent. First, as has been noted using numerous other data sources, the last thirty years have seen a significant increase in income inequality in the United States when compared with a number of continental European countries. The top

<sup>7</sup>The full list of original papers is as follows. Australia: Atkinson and Leigh 2007; France: Piketty 2001, 2003; Germany: Dell 2005; Ireland: Nolan 2005; Japan: Moriguchi and Saez 2006; Netherlands: Atkinson and Salverda 2003; New Zealand: Atkinson and Leigh 2005; Spain: Alvaredo and Saez 2006; Sweden: Roine and Waldenström 2008; Switzerland: Dell 2005 and Dell, Piketty, and Saez 2007; United Kingdom: Atkinson 2005; United States: Piketty and Saez 2003.

<sup>8</sup>See Leigh 2007. This includes in particular adjustments for the income unit, personal income total, income definition, dates for the tax year, and the age cutoff.

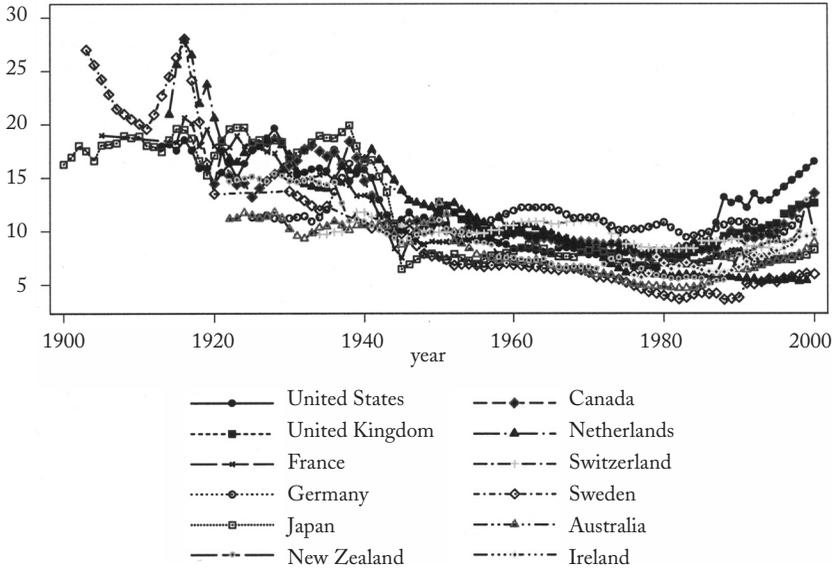


FIGURE 1  
SHARE OF INCOME EARNED BY TOP 1%<sup>a</sup>

<sup>a</sup>The figure reports the adjusted series constructed by Leigh 2007. See text for the original sources. Data for the Top 1% share in Spain are available from Alvaredo and Saez 2006 but only for the period 1981–2000; they are not included in the figure.

income shares data are distinctive, however, in providing us with a view of inequality over a longer time horizon and also allowing us to see the very considerable variation in levels of inequality that has occurred within countries over time. For the immediate postwar period, levels of inequality (as measured by the share earned by the top 1 percent) are remarkably similar across the different countries. When we consider broader measures, such as the top 10 percent income share, changes over the course of the twentieth century are less dramatic than those for the top 1 percent, but they are still very significant. The reality of very significant changes in income inequality within countries over time is potentially troubling for comparative political economy explanations which suggest that certain relatively static features like electoral rules or wage bargaining arrangements account for differences in inequality.

Next, in considering the pre-1945 period, we observe in Figure 1 that levels of inequality were strikingly higher in many countries, whereas in several countries the Second World War appears to have been associated with a drastic reduction in levels of income inequality.

For most countries the magnitude of pre-1945 changes in top income shares makes later changes seem small in comparison. This observation poses a challenge for comparative political economy. It suggests that if variation in factors like partisanship and wage bargaining centralization can account for variation observed between countries since 1970 but not for the larger changes that occurred before 1945, then one may want to reconsider whether, when, and why these factors mattered.

#### ARE TOP INCOME SHARES RELEVANT FOR TESTING POLITICAL HYPOTHESES?

Given that the top incomes data are limited to measuring the share of total income earned by the top 10 percent or smaller gradations therein, before proceeding with our analysis we need to consider whether we would expect this to be a relevant definition of inequality for assessing the effects of government partisanship and of centralized wage bargaining. There are two reasons why this could be the case. The first possibility is that partisanship and centralized wage bargaining directly affect top income shares. The second possibility is that the top income shares can serve as a useful proxy for trends involving other segments of the income distribution.

For partisanship, it is possible to argue that top income shares are a relevant dependent variable both because of the possible direct effect of partisanship and because of the usefulness of top income shares as a proxy. If left governments tend to adopt redistributive policies like progressive taxation and subsidizing education and if these policies influence the accumulation of both human and financial capital, then it seems logical to think that this would affect future pretax top incomes shares for the reasons we have described in the introduction and that we will elaborate in the next section. Further, as we show below, there is substantial evidence that top incomes shares provide a useful proxy for broader inequality measures. Perhaps most important, Andrew Leigh demonstrates relatively strong correlations between the top income measures and existing cross-country inequality measures based on the gini coefficient for the period after 1960.<sup>9</sup>

For centralized wage bargaining, it is a subject of greater debate whether top incomes shares can constitute an appropriate dependent variable. One response might be to suggest that top income shares capture developments only for upper-income groups that are unlikely to be affected by the presence of centralized wage bargaining arrangements,

<sup>9</sup> See Leigh 2007.

even when such arrangements do exist. In addition, while wage bargaining arrangements will have a most direct effect on the distribution of earnings, the top income shares data are based on income of an individual or tax unit (which can be a household) and do not distinguish between income derived from capital and income derived from labor.

Ideally, we would have a full time series on earnings inequality in our sample countries for the entire twentieth century. In the absence of such data we will argue that the top incomes data can nonetheless prove very useful for examining the effect of centralized wage bargaining for two key reasons. First, we know that individuals within the top 10 percent but below the top 1 percent of the distribution earn most of their income from labor, and studies of centralized wage bargaining arrangements in individual countries like the Netherlands and Sweden have suggested that white-collar employees are directly affected by such arrangements, as described in the next section. Second, we can also show that a top income share measure that excludes the top 1 percent of the distribution is actually highly correlated with the principal dependent variable used in the empirical literature on centralized wage bargaining, which is the ratio between the 90th and 10th percentile of the earnings distribution. As such, we think it is informative to use the top incomes data to proxy for earnings inequality.

We can compare data from the OECD on earnings inequality with a top incomes measure that we constructed as follows. Take total income  $y_{90-99}$  earned by individuals (or tax units) between the 90th and 99th percentiles of the income distribution. Then divide this total by the total income earned by all individuals (or tax units) between the bottom and the 99th percentiles of the distribution to get the ratio  $\frac{y_{90-99}}{y_{0-99}}$  or what we call for the rest of the article the *Top10-1* share. This measure excludes the effect of trends in very top incomes (those within the top 1 percent), and we also know that over time, capital income has been heavily concentrated in the top 1 percent. Figure 2 then plots the 90-10 earnings ratio from the OECD database for four time periods (1980-84, 1985-89, 1990-94, 1995-99) against the *Top10-1* income measure for our sample countries. As one would suspect from visual inspection of the graph, the pairwise correlation between the two variables reported in Table 1 is extremely high.<sup>10</sup> In order to control for the possibility raised by the OECD that cross-country differences in the 90-10 ratio may be attributable to differences in definition and coverage rather than to actual differences in earnings dispersion, we should also

<sup>10</sup>The correlation remains very high (0.74) if we exclude all observations from the United States and Sweden as potential outliers.

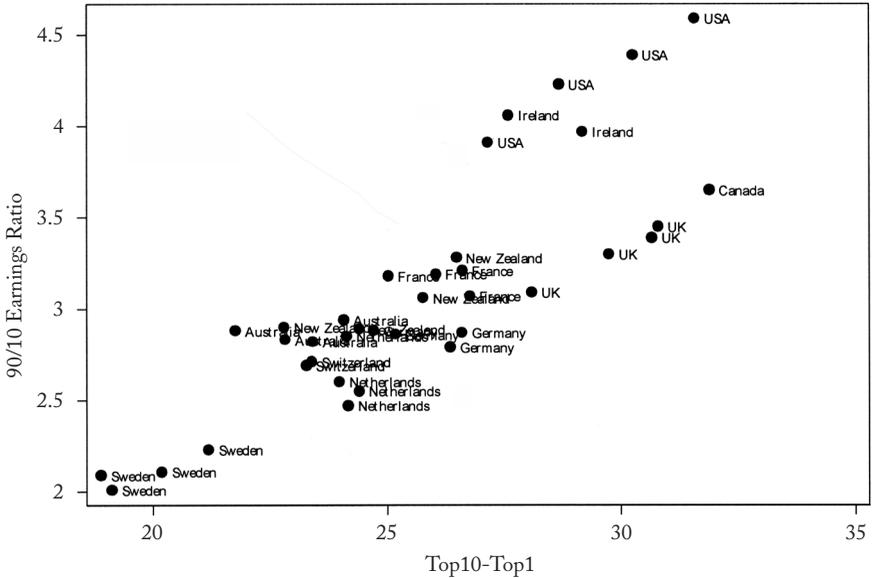


FIGURE 2  
COMPARING TOP INCOMES WITH THE 90/10 EARNINGS RATIO<sup>a</sup>  
(1980–99)

<sup>a</sup>Each country is shown for four separate subperiods: 1980–84, 1985–89, 1990–94, and 1995–99. Earnings data reported in OECD 2004. Top 10–1 income share is from the sources listed in Figure 1.

examine the correlation between these two variables when considering only within-country variation. It turns out that this correlation remains strong even if we subtract country mean values from each variable and then compare within-country variation in the 90–10 earnings and *Top10–1* income measure. In this case the pairwise correlation is 0.63. This further strengthens the case for using the *Top10–1* income share as a proxy measure. As seen in Table 1, we also find that analogous correlations between *Top10–1* income share and the 90–50 earnings ratio are also very strong (0.72 and 0.79, respectively). When we restrict our attention to the bottom half of the earnings distribution by considering the 50–10 earnings ratio, we again observe a strong correlation, but in this case one that weakens significantly when we subtract out country mean values.

While we lack a source like the OECD data for earnings data prior to the 1970s, we do dispose of some data that can help inform us whether the *Top10–1* income share measure is correlated with earnings inequality for earlier periods. First, for Sweden a long-run series on wage inequal-

TABLE 1  
COMPARING TOP INCOMES WITH EARNINGS RATIOS<sup>a</sup>

<i>OECD Earnings Data (1979–95)</i>	<i>Pairwise Correlation Coefficient with Top10-1</i>
90/10 ratio	0.82
90/10 ratio (within-country correlation)	0.63
90/50 ratio	0.72
90/50 ratio (within-country correlation)	0.79
50/10 ratio	0.77
50/10 ratio (within-country correlation)	0.32
Number of observations	121
<hr/>	
<i>Lydall Earnings Data (1930–60)</i>	
90/15 ratio	0.47
90/15 ratio (within-country correlation)	0.46
90/50 ratio	0.65
90/50 ratio (within-country correlation)	0.68
50/15 ratio	0.29
50/15 ratio (within-country correlation)	0.21
Number of observations	39

<sup>a</sup> Table reports pairwise correlation coefficients between the *Top 10-1* measure and earnings ratios for two periods. For 1979–95 the earnings ratios are based on annual data reported in OECD 1996. For 1930–60 the earnings ratios are from selected years reported by Lydall 1968. For both periods the correlations are reported with and without subtracting out the country means. The demeaned correlations measure within-country correlation.

ity that we describe at greater length below is very highly correlated with the *Top10-1* measure as can be seen in Figure 3 (the pairwise correlation between the two series is 0.87). Second, there is an important historical compendium of earnings distribution data produced by Harold Lydall that covers selected years between 1930 and 1960.<sup>11</sup> We used these data to construct the 90/15 earnings ratio (Lydall does not report earnings for the 10th percentile), and as can be seen in Table 1, there are fairly strong correlations with the *Top10-1* income share. The analogous correlations for these country-years with the 90-50 ratio are even stronger (0.65 and 0.68, respectively). There is less evidence in these data of a strong correlation between the *Top10-1* and the 50/15 ratio, but this is not surprising, because of the likely large error with which Lydall was able to measure the bottom of the earnings distribution.<sup>12</sup>

<sup>11</sup> Lydall 1968.

<sup>12</sup> By his own admission, Lydall's data sources frequently truncated the bottom of the distribution if they did not cover individuals who had positive earnings but were not subject to income tax assessments.

Overall, the evidence we have at our disposal suggests that given the absence of long-run time series on the 90-10 earnings ratio for periods before the 1970s, we can nonetheless use the *Top10-1* income measure as a very good proxy, in particular, for movements within the top half of the earnings distribution. This is certainly not to argue that the *Top10-1* measure is a perfect proxy for earnings inequality, and in particular for movements lower down the income distribution. Finally, we should again emphasize that our analysis will present not only evidence using the top incomes data but also evidence using long-run wage inequality measures where available.

### POTENTIAL POLITICAL DETERMINANTS OF INEQUALITY

Our goal is to investigate whether there is empirical support over a long time horizon for the proposition that centralized wage bargaining and left government are correlated with low levels of income inequality. In what follows, we discuss the logic behind each of these two hypotheses.

#### GOVERNMENT PARTISANSHIP

It is widely suggested that when controlling government, political parties that situate themselves on the left of the political spectrum will adopt redistributive policies including greater progressivity in income and estate taxation, more significant transfers, and greater public subsidies for goods like education when compared with their counterparts on the political right. Policies involving taxes and transfers may have an immediate effect on posttax income inequality by creating a wedge between pretax and posttax income. To the extent that the existing literature on partisanship and redistribution considers income inequality, it generally focuses on this mechanism, though the recent work of Larry Bartels has also emphasized the implications of partisan control for changes in the pretax income distribution.<sup>13</sup> But because the existing literature further suggests that left governments will adopt policies that influence the accumulation of human and physical capital, then we should also logically expect that if partisanship matters it should

<sup>13</sup> See Bartels 2008. For recent examples focusing on this redistributive effect, see Iversen and Soskice 2006; and Kenworthy and Pontusson 2005. And for a comprehensive treatment of the relationship between left partisanship and redistributive transfers, see Huber and Stephens 2001. Beramendi and Cusack 2008 consider the effect of partisan and institutional features on both pretax and posttax inequality.

also have an impact on future pretax income inequality.<sup>14</sup> Thomas Piketty has emphasized how high top marginal rates of income tax (a policy often associated with the left) or high top rates of inheritance tax can have a dramatic effect on capital accumulation and thus the pretax income distribution in future periods.<sup>15</sup> To the extent that left governments engage in policies like subsidizing education that facilitate accumulation of human capital for low-income groups, then this provides a second channel through which partisanship should influence the pretax income distribution.<sup>16</sup> The conclusion of this discussion is that if we expect that governments of the left will engage in more redistributive policies and in particular redistributive policies that influence the accumulation of human and physical capital, then we should also logically expect countries in which parties of the political left dominate to have lower levels of pretax income inequality. We should add that in our econometric tests we will take account of the dynamic nature of this suggested relationship.

While the above arguments follow a convincing logic, it is also entirely possible that partisanship might have little effect on the pretax income distribution. The strong common trend in the top 1 percent income share across countries over the course of the twentieth century may be explained by the fact that parties historically on the left or right shift their policy positions to fit current circumstances, constraints, or societal opinions.

#### CENTRALIZED WAGE BARGAINING

While government partisanship may influence the pretax income distribution via redistributive policies, certain labor-market institutions may have a direct effect on pretax inequality via the distribution of earnings. A number of scholars have presented theoretical models and empirical evidence to suggest that in countries where wage negotiations tend to be centralized there will be lower levels of wage dispersion.<sup>17</sup> Centralized bargaining arrangements, it is argued, can reduce the dis-

<sup>14</sup> For the most comprehensive recent empirical treatment of the link between left government and high top marginal rates of income taxation, see Ganghof 2006. For earlier contributions, see Garrett 1998; and Hallerberg and Basinger 1998.

<sup>15</sup> On this point, see in particular Piketty 2003.

<sup>16</sup> See Ansell 2007 for a recent cross-country study providing evidence that governments of the left spend more on public education. Work by Boix 1998 and Huber and Stephens 2001 provide two earlier examples of empirical studies demonstrating a positive correlation between left partisanship and public spending on education.

<sup>17</sup> One of the most powerful statements of this argument is presented by Moene and Wallerstein 2002.

persion of pay between different firms (when bargaining occurs at the industry level), between different industries (when bargaining occurs at the national level), and between different categories of wage earners. In countries such as Sweden centralized bargaining was accompanied by explicit policies emphasizing “solidaristic pay,” paying equal wages for equal work, in addition to reducing differentials between employees at different levels of an organization. The literature has also shown how in countries like Sweden and the Netherlands centralized wage bargaining arrangements have affected both white-collar and blue-collar workers. In some instances this occurred because an agreement set explicit limits on pay for the former category of workers. In other instances this occurred when industrial unions achieved an agreement with a central employers federation and this agreement in turn provided the basis for a subsequent agreement between the employers federation and a white-collar union.<sup>18</sup>

Using data from recent decades, several authors have found empirical evidence of a negative correlation between centralization of wage bargaining and pay inequality in OECD countries.<sup>19</sup> Many of these existing papers also use the ratio between the 90th and 10th percentiles of the earnings distribution as their measures of pay inequality.<sup>20</sup> The notion that wage bargaining centralization reduces earnings inequality is not specific to political science; we also find reference to this empirical result in the work of both economists and international organizations like the OECD.<sup>21</sup> In a broad review of collective bargaining and economic performance, Robert Flanagan suggests that “the search for correlations between economic outcomes and industrial relations institutions has produced one durable relationship. Wage dispersion is negatively correlated with the centralization of collective bargaining, although the relationship somewhat weakened by the late 1990s.”<sup>22</sup>

While the 90/10 ratio is the most commonly used dependent variable in the cross-country empirical literature on wage bargaining centralization, it is important to note that several studies also consider the correlation between centralization and dispersion within more specific parts of the earnings distribution. The study by Jonas Pontusson, David

<sup>18</sup> For the Swedish case, see Olsson 1991; and Swenson 1989. For the Netherlands, see Levenbach 1953; and Windmuller 1957.

<sup>19</sup> See in particular Wallerstein 1999; and Rueda and Pontusson 2000.

<sup>20</sup> A correction and reanalysis of the Wallerstein 1999 results by Golden and Londregan 2006 concludes that wage bargaining centralization has had a statistically significant but small effect on pay dispersion.

<sup>21</sup> OECD 1997.

<sup>22</sup> Flanagan 1999, 1163.

Rueda, and Christopher Way finds a statistically significant negative correlation between bargaining centralization and earnings dispersion both in the upper part of the distribution (90/50 ratio) and in the bottom part of the distribution (50/10 ratio).<sup>23</sup> However, the substantive magnitude of the latter effect was about three times as large. As we have already shown, we can be particularly confident that the *Top10-1* income measure is a good proxy for the 90/50 ratio, but it is less highly correlated with the 50/10 ratio.

Though quantitative studies of the impact of centralized wage bargaining have generally focused on the relationship between bargaining centralization and growing pay inequality since the 1970s, in a number of countries centralized wage bargaining arrangements first emerged during the crises of the 1930s and 1940s. If this is the case, and wage bargaining centralization has had a significant effect on reducing inequality, then we should expect to observe a negative correlation between wage bargaining centralization and inequality for earlier periods and we should also expect to observe that the introduction of centralized bargaining led to a structural break in inequality.

Although arguments about the causal effect of centralized wage bargaining are very plausible, it is also possible to imagine reasons why these institutions, as with partisanship, might not have this effect. A first possibility would be if centralized wage bargaining is correlated with low wage inequality, but the relationship is not causal. As emphasized by Robert Flanagan, it may be that underlying economic forces or underlying social norms about equality determine both which labor market institutions are adopted and what level of pay inequality prevails.<sup>24</sup> A second possibility is that centralized wage bargaining occurs in some countries for the reasons described by Flanagan, but in other cases at similar times broader economic and social forces push inevitably toward lower earnings inequality even without the adoption of formally centralized bargaining institutions.

## INCOME INEQUALITY, 1916–2000

### EMPIRICAL SPECIFICATION

This section empirically evaluates the hypotheses that labor-market institutions and government partisanship are important determinants of income inequality. We show that there is no evidence of a statisti-

<sup>23</sup> Pontusson, Rueda, and Way 2002. See also Blau and Kahn 1996; and Iversen 1999.

<sup>24</sup> Flanagan 1999.

cally significant effect of partisanship on inequality, except regarding the very top incomes (the top 1 percent), and this latter result is small in substantive terms. With regard to wage bargaining, we document, contrary to much of the existing literature, that there is little evidence in our full time series of a robust correlation between centralized wage bargaining and income inequality. Specifically, there is no evidence that national peak-level centralized wage bargaining is associated with lower levels of inequality than is sectoral-level bargaining. Our results throughout this section are based on data from thirteen advanced industrial democracies.

We use three dependent variables for our analysis: *Top10-1*, *Top 10*, and *Top 1*. As explained above, the first measure is designed to take the broadest inequality measure possible using the top incomes data while simultaneously subtracting out the very top incomes that may respond to specific factors. The second and third measures are simply the percentage of national income accruing to the top 10 percent and the top 1 percent of income earners. For testing the hypothesis about centralized wage bargaining, the *Top10-1* is the most relevant measure for the reasons we have already explained. All three dependent variables are relevant for considering the potential effect of partisanship, since redistributive policies like progressive taxation can have an effect on top incomes. We average the data over seventeen five-year periods from 1916 to 2000.<sup>25</sup> Averaging across five-year time periods allows us to examine variation over time without specifying precisely how long it takes for changes in labor-market institutions or government partisanship to affect income inequality. Even with five-year averages, there is evidence of serial autocorrelation in the inequality time series and we model this simply by adding a lagged dependent variable to each of our ordinary least squares specifications. To account for possible panel heteroskedasticity and panel correlations, we report panel-corrected standard errors for all coefficient estimates. We also conducted a set of tests for the existence of a unit root in the different top incomes measures. As discussed in detail in our online appendix, for each top incomes measure, unit root tests exploiting the panel structure of our data rejected the null of all individual country series having a unit root, and we have chosen to perform our estimates in levels.<sup>26</sup> Finally, upon

<sup>25</sup>The starting year of our analysis is determined by data availability. Although we have some information on top income shares for the first decade of the twentieth century, it is rather limited. Consequently, we use data only from 1911 forward in our regression analyses and, as will be elaborated below, due to five-year averaging and the use of a lagged dependent variable, we lose the five years 1911–15 in our actual analysis. Including the data that we have from 1900–1910 produces similar results.

<sup>26</sup>The authors' online appendix for this article is available at [http://pantheon.yale.edu/~ks298/index\\_files/pubs.htm](http://pantheon.yale.edu/~ks298/index_files/pubs.htm).

constructing our data sets, there were nontrivial numbers of missing observations for various variables. We dealt with this issue by multiply imputing missing data using a procedure that is discussed in detail in the online appendix.<sup>27</sup>

To test the hypotheses that government partisanship and labor-market institutions are important determinants of income inequality for the data series from 1916 to 2000, we developed new measures of government partisanship and wage bargaining centralization. This was necessitated by the fact that existing indices of partisanship and wage bargaining do not extend to the period before World War II. As we discuss below, the principal conclusions of this article remain very similar when we instead pursue the alternative strategy of estimating correlates of top incomes shares only for the post-1950 period, for which we can use existing measures of partisanship and wage bargaining.

The variable *Wage Bargaining Centralization* is an index constructed by the authors that takes a value equal to one if wages are primarily determined in a decentralized setting, either without collective bargaining at all or with bargaining at the plant level, equal to two if wages are primarily determined at the industry level, and equal to three if there is national centralized wage setting.<sup>28</sup> As discussed above, the literature suggests that more centralized wage bargaining decreases income inequality through a number of possible mechanisms, leading to an expectation of a negative partial correlation between *Wage Bargaining Centralization* and income inequality. Although the use of indices similar to *Wage Bargaining Centralization* are common in the literature, it is not clear whether each of the unit intervals should be expected to have the same impact on inequality. Consequently, we constructed two dichotomous indicator variables, *Centralized Wage Bargaining* and *Decentralized Wage Bargaining*, based on the values of the index, and we included these measures to evaluate the hypothesis that greater central-

<sup>27</sup> One further potential problem with the analyses reported below is that some specifications include both a lagged dependent variable and country fixed effects in analyses with a relatively small number of time periods. Because the bias from fixed-effects specifications with lagged dependent variables is decreasing in the number of time periods, this source of bias could be substantively important. To address this possibility, we also used the Arellano-Bond generalized method-of-moments estimator, which differences the data to deal with country-specific effects and then uses the panel structure of the data to instrument for the lagged dependent variable to address the correlation between the error term and lagged dependent variable generated by differencing. The results of this analysis were qualitatively similar to those reported in the article. The one variation is that there is much less evidence of a difference in income inequality under decentralized wage bargaining institutions compared with that in sectoral-level bargaining.

<sup>28</sup> We consulted a number of sources to code each country, including Campbell 1992; Ebbinghaus and Visser 2000; Blum 1981; Wallerstein, Golden, and Lange 1997; Iversen 1999; OECD 2004; Swenson 1989, 2002; and Golden and Wallerstein 2006b. For further description of this variable, see Scheve and Stasavage, online appendix (fn. 26).

ization is associated with lower inequality. *Centralized Wage Bargaining* is equal to one if there is national, peak-level centralized wage setting and is equal to zero otherwise. *Decentralized Wage Bargaining* is equal to one if wages are primarily determined in a decentralized setting, either without collective bargaining at all or bargaining at the plant level and is equal to zero otherwise. Since the omitted category is sectoral-level bargaining, the expectation is that *Centralized Wage Bargaining* should be negatively correlated with each of the measures of inequality and that *Decentralized Wage Bargaining* should be positively correlated with these measures.

We also include an alternative measure of the extent of labor-market organization by adding the variable *Union Density* equal to the percentage of the total dependent labor force that are members of unions (less the self-employed).<sup>29</sup>

To measure government partisanship, we constructed a dummy variable, *Left Executive*, equal to one if the country had a prime minister and/or president from a left party in a given year and zero otherwise.<sup>30</sup> Given the expectation that left governments set policies favorable to lower-income voters, this variable should also be negatively correlated with the measures of income inequality.

The economic literature on income inequality suggests a number of control variables that should be included to estimate the partial correlations between *Left Executive*, *Centralized Wage Bargaining*, *Decentralized Wage Bargaining*, *Union Density*, and our measures of income inequality. For our analysis of the complete data series, we include *GDP per Capita*, *Trade Openness*, *Secondary Education Share*, and *Female Participation*, indicating the proportion of women in the economically active population, in all regressions.<sup>31</sup> We also include a control, *Nondemocracy*, equal to one if the country is experiencing a nondemocratic year and zero otherwise, and a control, *Universal Suffrage*, equal to one for all years after which the country had universal suffrage (male and female) and zero otherwise. In all specifications, we include dichotomous indicator variables for the time period of the observation. The time periods allow us to control for common shocks to income inequality in all countries.

<sup>29</sup> The primary source for the period after 1950 is Golden and Wallerstein 2006b. The primary source for the period before 1950 is Kjellberg 1983. When both sources were missing, we used Visser 1989. Additional sources used were OECD 2004; May, Walsh, Harbridge, and Thickett 2003; and Wallace 2003.

<sup>30</sup> Coded based on information in Caramani 2000; and McDonald 2002. For further description, see Scheve and Stasavage, online appendix (fn. 26).

<sup>31</sup> The complete definitions and sources for these variables are described in Scheve and Stasavage, online appendix (fn. 26).

## RESULTS

Table 2 reports the coefficient estimates for the ordinary least squares regression of each of our measures of income inequality on its one-period lag, the key variables of interest, *Left Executive*, *Centralized Wage Bargaining*, *Decentralized Wage Bargaining*, and *Union Density*, and the controls. For each dependent variable two specifications are reported, one without and one with country fixed effects.

The coefficient estimates in Table 2 provide no evidence for an important role for *Centralized Wage Bargaining* in determining national levels of income inequality. In the specifications with or without country fixed effects, the coefficients are small and not statistically significant. Given the presence of the variable *Decentralized Wage Bargaining* in the model, it is important to emphasize that the correct interpretation of this result is that there is no evidence in this data that national peak-level centralized wage bargaining is associated with less income inequality than is sectoral-level bargaining.<sup>32</sup> One potential concern about this finding is that the OLS estimator may be biased if income inequality influences wage bargaining centralization. Specifically, it is plausible that countries experiencing greater equality find it easier to adopt centralized institutions. If this were the case, however, the OLS estimator would be biased but in a negative direction. If anything, this means that we have overestimated the negative effect of centralization on income inequality. Since this null result contradicts an extensive empirical literature based on analyses for the last two to three decades of the twentieth century, in the next section we will revisit this finding by looking at individual cases in greater detail.

The results in Table 2 do indicate a positive and statistically significant correlation between decentralized wage bargaining and each of the three measures of income inequality. These estimates indicate that decentralized wage setting is associated with more income inequality than is sectoral-level bargaining. The result is consistent with the hypothesis that more organized labor markets have less income inequality even if, as our results indicate, peak-level bargaining does not have an effect over and above sectoral wage setting. The long-run impact of decentralized wage setting implied by these estimates is about a 2 to 3 percentage point increase in each of our inequality measures. This is economically important, but it does not suggest that wage bargain-

<sup>32</sup>The absence of a significant correlation between *Centralized Wage Bargaining* and the income inequality measures does not, however, depend on the inclusion of the *Decentralized Wage Bargaining* variable. There is no partial correlation for *Centralized Wage Bargaining* and income inequality with or without the indicator variable for decentralized wage setting.

TABLE 2  
LABOR-MARKET INSTITUTIONS, GOVERNMENT PARTISANSHIP, AND  
INCOME INEQUALITY<sup>a</sup>  
(1916–2000)

	<i>Top10-1</i>		<i>Top10</i>		<i>Top1</i>	
TopX <sub>t-1</sub>	0.537 (0.072)	0.389 (0.071)	0.556 (0.071)	0.468 (0.072)	0.578 (0.069)	0.539 (0.076)
GDP per Capita	0.019 (0.056)	0.005 (0.081)	0.089 (0.069)	0.098 (0.106)	0.103 (0.045)	0.139 (0.076)
Trade Openness	0.019 (0.009)	0.026 (0.014)	0.022 (0.011)	0.024 (0.018)	0.008 (0.007)	0.003 (0.012)
Secondary Education Share	0.442 (1.231)	-1.462 (1.409)	-0.306 (1.408)	-2.095 (1.648)	-0.899 (0.902)	-1.090 (1.074)
Female Participation	-14.822 (5.886)	-11.481 (5.118)	-21.087 (7.539)	-17.734 (7.423)	-11.266 (4.896)	-10.550 (5.622)
Centralized Wage Bargaining	0.100 (0.493)	0.300 (0.499)	0.021 (0.680)	0.079 (0.707)	-0.108 (0.464)	-0.217 (0.486)
Decentralized Wage Bargaining	1.135 (0.457)	1.314 (0.438)	1.554 (0.567)	1.622 (0.559)	0.778 (0.358)	0.732 (0.411)
Union Density	-0.058 (0.015)	-0.071 (0.019)	-0.064 (0.021)	-0.078 (0.027)	-0.017 (0.012)	-0.021 (0.017)
Left Executive	0.252 (0.394)	-0.346 (0.363)	-0.098 (0.521)	-0.673 (0.490)	-0.445 (0.312)	-0.532 (0.318)
Nondemocracy	1.291 (0.818)	1.582 (0.867)	2.271 (0.937)	2.542 (0.991)	1.615 (0.594)	1.638 (0.657)
Universal Suffrage	1.257 (0.392)	1.298 (0.461)	0.953 (0.475)	0.590 (0.574)	-0.200 (0.324)	-0.630 (0.418)
Period fixed effects	yes	yes	yes	yes	yes	yes
Country fixed effects	no	yes	no	yes	no	yes
Number of countries	13	13	13	13	13	13
Number of 5-year periods	17	17	17	17	17	17
Total observations	219	219	219	219	219	219

<sup>a</sup>Table reports the results of OLS regressions for the three measures of income inequality, *Top10-1*, *Top10*, and *Top1*, on *Centralized Wage Bargaining*, *Decentralized Wage Bargaining*, *Union Density*, *Left Executive*, and various control variables for the seventeen five-year periods between 1916 and 2000. Ireland was not an independent country until 1922 and so is not included in the analysis until that five-year period. The table reports multiple imputation estimates of the OLS coefficients for each variable and their PCSEs in parentheses. A constant term is included in each regression but not reported in the table.

ing institutions can account for the large changes over time in income inequality during the twentieth century. As we show below, this partial correlation is not statistically significant for the 1916 to 1975 period and is largely driven by outcomes after the mid-1970s. There is, therefore, little evidence in this data that wage bargaining institutions played

a significant role in the substantial changes in income inequality during the first three-quarters of the twentieth century, though a role at the end of the century is certainly a possibility.

The results for the trade unionism measure indicate a negative correlation between union density and our three top income measures. In the specifications with and without country fixed effects, the coefficient estimates for *Union Density* are negative and statistically significant at the 0.01 level for the *Top10-1* and *Top10* measures of income inequality. The inclusion of both country and period fixed effects means that the fixed-effects estimates are identifying off of within-country variation over time, controlling for common shocks to income inequality experienced by all the countries included in the sample. This is strong quantitative evidence of a correlation between union density and income inequality. The big question, and one that we cannot begin to answer in this article, is whether this result also reflects a causal relationship whereby union density determines inequality. The primary alternative is that union density is an outcome that, like income inequality, is driven primarily by underlying economic or political trends.<sup>33</sup>

The coefficient estimates in Table 2 do not suggest an important role for government partisanship in determining national levels of income inequality. In the specifications with or without country fixed effects, the coefficients are small and not statistically significant. This summary should be qualified somewhat for the *Top1* measure of income inequality for which there is some modest evidence that left partisan control may be associated with less income inequality but that any such effect seems to operate on the incomes at the very top of the income distribution. Moreover, the substantive magnitude of the estimated effect of partisanship is small when compared with the massive changes in the top 1 percent income share that occurred over the course of the twentieth century.<sup>34</sup>

To evaluate the robustness of our empirical results, we conducted a number of sensitivity analyses. First, we reestimated each specification for the 1976–2000 and 1916–75 periods separately. Previous research on the impact of labor-market institutions on income inequality has

<sup>33</sup> See the discussion on this point in Acemoglu, Aghion, and Violante 2001.

<sup>34</sup> Because of space constraints, we do not discuss the results for the control variables in any detail. There is some evidence that richer and more open countries have greater income inequality and that greater female labor-force participation is associated with lower inequality. These estimates, however, are based on specifications including the labor-market organization and partisanship variables; and since these measures may be consequences of the economic variables, they may not be proper controls for evaluating the impact of the economic measures. The results in Table 2 also indicate that nondemocracy is associated with greater inequality.

TABLE 3  
 LABOR-MARKET INSTITUTIONS, GOVERNMENT PARTISANSHIP, AND  
 INCOME INEQUALITY<sup>a</sup>  
 (1976–2000)

	<i>Top10-1</i>		<i>Top10</i>		<i>Top1</i>	
Centralized Wage	0.434	0.348	0.264	0.406	-0.060	-0.218
Bargaining	(0.782)	(0.754)	(1.052)	(1.030)	(0.573)	(0.668)
Decentralized Wage	1.538	1.253	2.145	2.179	1.405	1.088
Bargaining	(0.458)	(0.621)	(0.648)	(0.773)	(0.455)	(0.550)
Union Density	-0.033	-0.049	-0.018	-0.060	0.012	-0.010
	(0.019)	(0.028)	(0.024)	(0.045)	(0.010)	(0.031)
Left Executive	0.460	-0.083	0.341	-0.005	-0.062	-0.137
	(0.495)	(0.435)	(0.632)	(0.610)	(0.333)	(0.318)
Baseline control variables	yes	yes	yes	yes	yes	yes
Period fixed effects	yes	yes	yes	yes	yes	yes
Country fixed effects	no	yes	no	yes	no	yes
Number of countries	13	13	13	13	13	13
Number of 5-year periods	5	5	5	5	5	5
Total observations	65	65	65	65	65	65

<sup>a</sup> Table reports the results of OLS regressions for the three measures of income inequality, *Top10-1*, *Top10*, and *Top1*, on *Centralized Wage Bargaining*, *Decentralized Wage Bargaining*, *Union Density*, *Left Executive*, and various control variables for the five five-year periods between 1976 and 2000. The table reports multiple imputation estimates of the OLS coefficients for each variable and their PCSEs in parentheses. A constant term is included in each regression but not reported in the table.

focused almost exclusively on the prior period due to data availability. It is instructive to determine whether there is heterogeneity in the relationship between these variables over time and thus how sensitive the estimates reported in Table 2 are to the sample period. The results for these regressions, reported in Tables 3 and 4, indicate that there is little evidence of a correlation between *Centralized Wage Bargaining* and *Left Executive* and our measures of income inequality in either sample. The results for *Decentralized Wage Bargaining* suggest that the evidence for the importance of organized labor markets in determining income inequality is largely driven by the empirical record for the 1976 to 2000 period, which has informed the current literature on this question. The Table 3 results for 1976–2000 indicate a positive and statistically significant correlation between *Decentralized Wage Bargaining* and our three measures of inequality, but the Table 4 estimates for 1916–75 are smaller in magnitude and less precisely estimated. In sum, there is little evidence that the centralization of wage bargaining played a significant

TABLE 4  
LABOR-MARKET INSTITUTIONS, GOVERNMENT PARTISANSHIP, AND  
INCOME INEQUALITY<sup>a</sup>  
(1916-75)

	<i>Top10-1</i>		<i>Top10</i>		<i>Top1</i>	
Centralized Wage	0.309	0.759	0.307	0.494	-0.005	-0.204
Bargaining	(0.697)	(0.671)	(0.894)	(0.972)	(0.664)	(0.738)
Decentralized Wage	0.868	0.907	0.987	0.722	0.314	-0.003
Bargaining	(0.560)	(0.587)	(0.670)	(0.734)	(0.422)	(0.615)
Union Density	-0.087	-0.098	-0.111	-0.119	-0.041	-0.044
	(0.023)	(0.027)	(0.028)	(0.035)	(0.018)	(0.026)
Left Executive	0.380	-0.310	0.146	-0.494	-0.347	-0.369
	(0.500)	(0.470)	(0.585)	(0.614)	(0.338)	(0.414)
Baseline control variables	yes	yes	yes	yes	yes	yes
Period fixed effects	yes	yes	yes	yes	yes	yes
Country fixed effects	no	yes	no	yes	no	yes
Number of countries	13	13	13	13	13	13
Number of 5-year periods	12	12	12	12	12	12
Total observations	154	154	154	154	154	154

<sup>a</sup> Table reports the results of OLS regressions for the three measures of income inequality, *Top10-1*, *Top10*, and *Top1*, on *Centralized Wage Bargaining*, *Decentralized Wage Bargaining*, *Union Density*, *Left Executive*, and various control variables for the twelve five-year periods between 1916 and 1975. Ireland was not an independent country until 1922 and so is not included in the analysis until that five-year period. The table reports multiple imputation estimates of the OLS coefficients for each variable and their PCSEs in parentheses. A constant term is included in each regression but not reported in the table.

role in the large changes in income inequality over the first seventy-five years of the twentieth century.<sup>35</sup>

As a second robustness check, we considered the possibility that the cumulative experience of left partisanship and centralized wage bargaining is what matters most for accounting for variation in income inequality. We constructed variables that measured the proportion of years in the last twenty that each country had values of the variables *Left Executive*, *Centralized Wage Bargaining*, and *Decentralized Wage Bargaining* equal to one. The results for the wage bargaining measures are essentially the same. There is no evidence of a correlation between the cumulative version of *Centralized Wage Bargaining* and the income inequality measures while a positive correlation is observed for the cu-

<sup>35</sup> Along these lines, we also reestimated these specifications for the 1951-75 period and found no evidence of a significant positive correlation between decentralization and the income inequality measures. Similarly, there was no evidence of a negative correlation for peak-level national wage setting.

cumulative *Decentralized Wage Bargaining* variable.<sup>36</sup> For partisanship, the hypothesized negative partial correlation is not observed for the *Top10-1* and *Top10* measures.<sup>37</sup> There is further evidence that left partisanship may reduce the share of income that goes to the very highest earners as the cumulative partisan measure is negatively and statistically significantly correlated with the *Top1* variable in the specification with fixed effects (p-value equal to 0.048). As before, however, the substantive magnitude of this effect remains very small as compared with the very large changes that have occurred in the top 1 percent income share over the course of the twentieth century.

As a third robustness check, we estimated specifications that dropped the union density measure to evaluate whether its inclusion might attenuate our estimates for centralized wage bargaining and partisanship. We found that the estimates were substantially similar for our key variables of interest.

Fourth, we considered the possibility that our estimates might be biased due to omitted variables. It is possible that there could be some influence of proportional representation (PR) on income inequality. The inclusion of PR had little impact on our estimates and the coefficient on PR was statistically significant only in specifications without fixed effects for the *Top10-1* and *Top10* measures of income inequality.<sup>38</sup>

Finally, we considered the possibility that our results may be biased due to poorly measured variables. To address this, we investigated each hypothesis further with alternative data, arguably better measured, that are available from standard sources for the period 1950 to 2000. In this analysis, we measure government partisanship as equal to the government's left-right position as determined by the weighted (by seats in parliament) left-right positions of the parties in government; and we use Golden and Wallerstein's coding of wage bargaining centralization to form our indicator variables for labor-market organization. Perhaps our most striking finding in this analysis is the evidence of a partial correlation between this measure of partisanship and income inequality for the more recent 1976–2000 period but very little evidence for this correlation in the 1951–75 period. This result resonates both with

<sup>36</sup> The partial correlation is somewhat weaker for the cumulative version of the *Decentralized Wage Bargaining* variable, but it is still statistically significant at the 0.05 level in the specifications with fixed effects.

<sup>37</sup> For these measures without fixed effects, the cumulative partisan coefficient is actually also statistically significant at the 0.05 level but the sign is positive and therefore inconsistent with the usual partisan hypothesis.

<sup>38</sup> We also reestimated each specification dropping one of our thirteen countries at a time. This also did not substantially change our estimates for any of the main variables.

the existing literature that has focused on the more recent period and with our main results, which do not find evidence of a robust correlation between partisanship and income inequality. The results are also consistent with our main findings for wage bargaining centralization. For both periods, there are no significant partial correlations between peak-level centralized bargaining and the income inequality measures. Further, the positive correlations between decentralized wage setting and income inequality are observed most strongly in the 1976–2000 period and are not present at all in any of the fixed-effects specifications.<sup>39</sup>

### INSTITUTIONAL TURNING POINTS

Our regression results based on a pooled analysis of thirteen countries suggest that when one considers a long time horizon (one extending back to the pre-1970 period), there is relatively weak evidence that the presence of a left government has been associated with lower income inequality. Nor is there clear evidence of an association between centralized wage bargaining and low levels of income inequality. Given the conclusions of previous scholarly work, it makes sense to take a closer look at the relationship between inequality and the establishment of peak-level centralized wage bargaining at the country level. Doing so may suggest whether our findings are influenced by the fact that we are pooling across a heterogeneous set of countries. It can also allow us to investigate whether the initial introduction of centralized wage bargaining in several countries was associated with a drop in inequality. Further, although we have provided considerable evidence that the top income share measures, particularly the *Top10-1* variable, is a good proxy for the most common inequality measures used in the literature, our focus in this section on the individual countries allows us to use wage inequality measures as well as top income shares to investigate the effect of peak-level centralized wage bargaining.

In this section we consider evidence from individual country cases in order to provide a closer look at the effect of centralized wage bargaining, and in particular shifts toward greater centralization, on inequality. Seminal work in the field of political economy has emphasized the im-

<sup>39</sup> For a more complete description of this analysis, see online appendix (fn. 26). To further examine possible measurement problems, we also reestimated our baseline specifications for alternative coding rules for wage bargaining centralization. We examined, for example, the consequences of coding Germany as decentralized rather than centralized for the 1933–45 period. These changes in coding rules did not substantially effect our estimates.

portance of the crises of the 1930s and 1940s in leading to the adoption of corporatist bargaining arrangements in several states. Peter Katzenstein has argued that the economic crisis of the 1930s led to the development of corporatist bargaining in the smaller European states but not in their larger neighbors, because of different political conditions, one of which was the prior adoption of proportional representation.<sup>40</sup> For Peter Gourevitch the crisis of the 1930s and the subsequent war-time experience helped lead to formal centralized bargaining arrangements in some states; he emphasizes as well how these twin crises drove all states by 1945 toward a politics of accommodation between business and labor irrespective of the formal arrangements for bargaining.<sup>41</sup> His arguments regarding the postwar period in the United States closely parallel the more recent analysis of Frank Levy and Peter Temin.<sup>42</sup>

Sweden in the early 1950s is the most frequently cited example of a major shift toward formal centralized bargaining, but the literature also identifies similar historical turning points in countries like Denmark and the Netherlands. In this section we investigate whether these suggested historical turning points are associated with structural breaks in income inequality. Previous quantitative studies of centralized wage bargaining and its effects have not considered whether the steps taken toward centralized wage bargaining during the 1930s, 1940s, and 1950s had a significant impact on outcomes. We argue that there is in fact little evidence that the initial move to centralize wage bargaining led to a downward break in inequality. For each country we instead observe a pattern where inequality was already trending downward before bargaining was centralized, and inequality continued to trend downward at the same rate after this point. This raises questions about the extent to which this institutional change had an independent causal effect on income inequality or whether centralized wage bargaining and lower income inequality were instead outcomes driven by some underlying process at the country level. In what follows we first present visual evidence for Sweden on the extent to which the move to centralized bargaining constituted a structural break. We then consider this issue more formally using data from Sweden, Denmark, the Netherlands, and Ireland—four countries that established centralized bargaining at the national level in the middle of the twentieth century.

<sup>40</sup> Katzenstein 1985.

<sup>41</sup> Gourevitch 1986, chap. 4.

<sup>42</sup> Levy and Temin 2007.

### WAS THERE A STRUCTURAL BREAK?

Among our thirteen sample countries, three established national-level wage bargaining in a democratic context during the 1930s, 1940s, or 1950s. In addition, we also have a separate income inequality series for Denmark that will allow us to examine whether the shift to centralized bargaining in that country was associated with a downward break in income inequality.<sup>43</sup> For Sweden, Denmark, and Ireland we also have long-run series on wage differentials between more highly skilled and less skilled workers. For Sweden we have a long-run series covering the entire twentieth century that represents the ratio of the wage for technicians (at age forty-two to forty-five) to the average wage in manufacturing.<sup>44</sup> For Denmark, we have a series measuring the pay ratio between skilled and unskilled manual workers for the years between 1870 and 1965.<sup>45</sup> Finally, for Ireland we have data showing the ratio between wages for skilled and unskilled workers in three sectors of the Irish economy between 1926 and 1984.<sup>46</sup> Each of these series can be used to examine whether centralization of bargaining produced a break in inequality. An examination of these wage series is also useful because it demonstrates that they are highly correlated with the data on top-income shares.

The Swedish national employers federation (SAF) and the union confederation (LO) took a significant step toward centralized wage bargaining with the 1938 Saltsjöbaden accord, which was an agreement to manage potential labor disputes at the confederal level. From 1953 the SAF and LO bargained in a centralized fashion with pay set according to a “solidaristic wage” policy stating that those performing similar work should receive similar pay irrespective of the firm or sector in which they worked and that wage disparities within firms should be reduced. In Denmark a national collective bargaining agreement was reached in 1934 between the LO union federation and the DA employers federation. This introduced a degree of centralization into the process, though it was not until the late 1950s that the LO and the DA acquired authority to sign legally binding agreements on wages and other issues.<sup>47</sup> In the Netherlands a government decree of 1945 created the Foundation of Labor, a bipartite organization including top employ-

<sup>43</sup> Sorensen 1993 used tax statistics to investigate income inequality over the long run in Denmark, and we will use his measure of income inequality in the structural break tests below.

<sup>44</sup> This has been compiled by Ljungberg 2006, building on the earlier work of Jungenfelt 1966.

<sup>45</sup> This is drawn from Johansen 1985.

<sup>46</sup> This has been compiled by O'Rourke 1994.

<sup>47</sup> See the discussion in Iversen 1999, chap.5; as well as Ebbinghaus and Visser 2000.

ers associations and trade union federations.<sup>48</sup> The foundation had a prominent role in wage setting. Finally, though not having the same history of “democratic corporatism” as the other three countries, Ireland, in 1946, also established a centralized system of wage bargaining with a national wage round.<sup>49</sup>

Figure 3 plots the *Top10-1* income share for Sweden, indicating two potential break points, the Saltsjöbaden accord of 1938 and the establishment of a solidaristic wage policy in 1953. The figure also makes use of the available wage inequality series for Sweden. Two things are apparent. First, the *Top10-1* share and the wage inequality series are very highly correlated and follow a common downward trend until the beginning of the 1980s. Second, there is no apparent evidence for either of these series that the Saltsjöbaden accord of 1938 produced an acceleration of this downward trend. One would draw a very similar conclusion if the break date for the centralization of bargaining was set at a different point, such as the adoption of a solidaristic pay policy in the early 1950s. Interestingly, there is also evidence that the gender wage gap declined in Sweden before equal pay became an official policy of Sweden’s central labor confederation.<sup>50</sup> In other words, we see a similar pattern to overall wages; institutional (or policy) change took place only after the changes in inequality that institutional change is often presumed to have caused were well under way.

As an addendum to the above discussion, it should also be emphasized that while inequality was trending strongly downward well before the institutional changes in Sweden, Denmark, the Netherlands, and Ireland, this was also true of most countries that did not adopt centralized wage bargaining at this time, as can be seen from the strong overall time trend in Figure 1.<sup>51</sup>

#### TESTS FOR A STRUCTURAL BREAK

The visual evidence for Sweden suggests that both the *Top10-1* income share and a measure of wage inequality declined following institutional changes, but inequality was already declining before this date and there is little evidence of an acceleration of this process. We now investigate this question more formally using available data on both top incomes

<sup>48</sup> Windmuller 1957.

<sup>49</sup> Blum 198, 291.

<sup>50</sup> See Svensson 2004.

<sup>51</sup> More concretely, the difference-in-differences during this period in income inequality between the often contrasted cases of the United States, which did not have centralized wage bargaining, and Sweden, which did, is essentially zero.

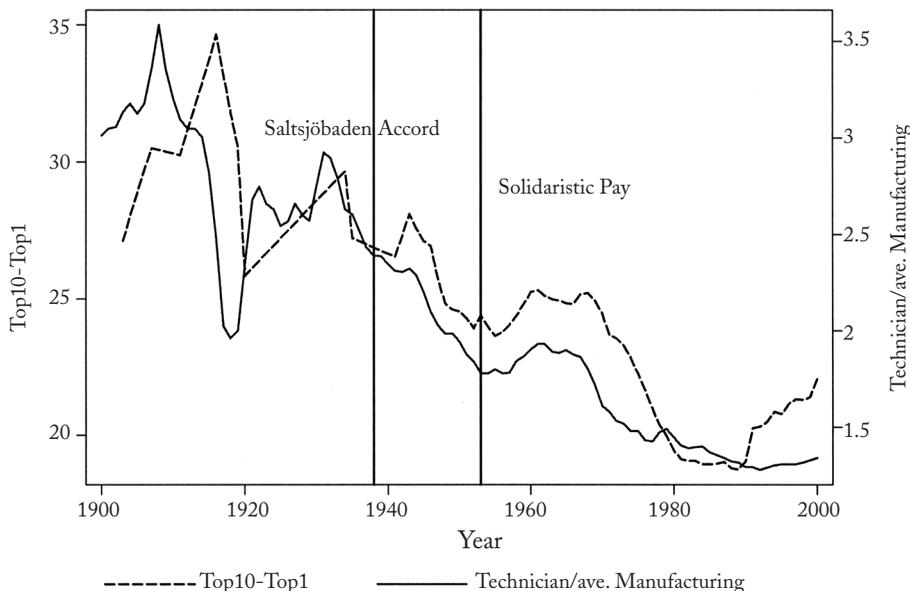


FIGURE 3  
CENTRALIZED BARGAINING AS A STRUCTURAL BREAK IN SWEDEN<sup>a</sup>

<sup>a</sup> Vertical lines indicate dates at which wage bargaining becomes more centralized. Top incomes data from Roine and Waldenström 2006 with the Top 10-1 share on the left axis. The ratio for technicians to the average wage in manufacturing is from Ljungberg 2006 and is on the right axis.

and wages for Sweden, Denmark, the Netherlands, and Ireland.<sup>52</sup> Each series is modeled as a function of a linear trend, allowing for a break at the time that wage bargaining was centralized. We report the results of this regression, as well as regressions setting the break point at five-year intervals up to fifteen years before or after the date that wage bargaining was centralized. Formally, in the equation below the variables  $T_1$  and  $T_2$  refer to separate time trends.  $T_1$  is set equal to  $t$  for all years previous to the chosen break point and it is set at  $T_1 = 0$  for all  $t$  subsequent to the break.  $T_2$  takes a value of zero for all  $t$  prior to the break and a value of  $t$  for all  $t$  after the break. If the adoption of centralized wage bargaining resulted in a structural break in inequality we would expect to find that the coefficient  $\gamma$  is more negative than  $\beta$ .

<sup>52</sup> For Sweden we use data on the *Top10-1* income share from 1903 to 2000 and wage data for the 1900 to 2000 period. For Denmark the MEC inequality series covers the period 1871 to 1965, and the wage series we use is for the period 1870-1965. For the Netherlands we use data on the *Top10-1* income share for 1914-99. For Ireland we use wage data covering the period 1926-1984 but unfortunately lack a top income series other than the Top 0.1 percent share prior to the 1970s.

$$\text{Inequality}_t = \alpha + \beta T_1 + \gamma T_2 + \varepsilon_t$$

For Sweden we set the break at  $T = 1953$  with the initiation of a solidaristic wage policy, though we obtained very similar results when instead setting  $T = 1938$ , the date of the Saltsjöbaden accord. For Ireland we set  $T = 1946$  and for the Netherlands  $T = 1945$ . Finally, for Denmark we set  $T = 1934$  representing the step toward centralization with the accord between the LO and the DA federations.<sup>53</sup>

The estimates of the equation are presented in Table 5.<sup>54</sup> There is strong evidence of a general downward trend in inequality in both the top incomes and the wage series, but there is very little evidence here of an acceleration of this downward trend following the adoption of centralized wage bargaining. This is particularly clear for Sweden, and we also observe that neither of the two Danish series provides evidence of a significant trend shift at any point up to fifteen years before or after the institutional change of 1934. For the Irish wage series there is some evidence of a downward trend shift around the time of the first national wage round in 1946, but the difference between  $\beta$  and  $\gamma$  is small and it is not statistically significant.<sup>55</sup> The Netherlands top income series is the single case where the difference between  $\beta$  and  $\gamma$  is negative, is statistically significant, and occurs around the time of an important institutional change in wage bargaining arrangements. This parallels the observation by John Windmuller that the establishment of the Foundation of Labor in the Netherlands led to a reduction in the skill premium.<sup>56</sup> One might respond that the absence of evidence for a structural break in most of these series results from the fact that we have not correctly identified the date on which wage bargaining became fully centralized. But the results presented in Table 5 show that in almost all cases, if we move the break date either forward or backward by up to fifteen years, there continues to be little evidence of an acceleration in the rate at which inequality was declining.

We can also adopt a more formal procedure for each of the six series, testing whether there is a break in both the trend and the mean of the

<sup>53</sup>The choice of this date for Denmark was in part dictated by the fact that our Danish series end in 1965, making it difficult to assess whether an alternative centralization date of the late 1950s was associated with a structural break in inequality. We should add to this that a visual inspection shows that neither of the two Danish series trends downward from the late 1950s.

<sup>54</sup>We have chosen to account for serial correlation here by using Newey–West standard errors. We obtained similar conclusions regarding the absence of a break in trend (but also less precise estimates of the  $\beta$  and  $\gamma$  coefficients) when including an AR(1) term, or terms for higher-order autocorrelation in the regression.

<sup>55</sup>We failed to reject the null of the two coefficients being identical  $p = 0.39$ .

<sup>56</sup>Windmuller 1957.

TABLE 5  
TESTING FOR A TREND SHIFT AT DIFFERENT BREAK DATES<sup>a</sup>

	<i>Break Set Relative to Centralization Date</i>						
	<i>-15</i>	<i>-10</i>	<i>-5</i>	<i>0</i>	<i>5</i>	<i>10</i>	<i>15</i>
Sweden (Top10-1) (1903–2000)							
$\beta$	-1.101 (.036)	-1.105 (.032)	-.093 (.028)	-.116 (.029)	-.129 (.027)	-.120 (.024)	-.110 (.022)
$\gamma$	-.120 (.016)	-.121 (.016)	-.118 (.016)	-.123 (.016)	-.126 (.016)	-.124 (.015)	-.123 (.015)
Sweden (wage ratio in percentage) (1900–2000)							
$\beta$	-1.634 (0.461)	-1.700 (.403)	-1.844 (0.363)	-2.134 (0.329)	-2.352 (0.293)	-2.290 (0.271)	-2.177 (0.263)
$\gamma$	-1.993 (0.174)	-1.998 (0.184)	-2.023 (0.183)	-2.079 (0.174)	-2.119 (0.166)	-2.102 (0.162)	-2.079 (0.160)
Netherlands (Top10-1) (1914–99)							
$\beta$	-.168 (.050)	-.157 (.052)	-.131 (.046)	-.045 (.030)	-.046 (.032)	-.083 (.034)	-.117 (.028)
$\gamma$	-.145 (.017)	-.144 (.019)	-.137 (.019)	-.111 (.014)	-.110 (.015)	-.121 (.015)	-.132 (.012)
Denmark (MEC) (1871–1965)							
$\beta$	-.183 (.092)	-.171 (.068)	-.173 (.049)	-.164 (.038)	-.159 (.032)	-.165 (.027)	-.181 (.025)
$\gamma$	-.208 (.032)	-.212 (.026)	-.212 (.022)	-.217 (.020)	-.223 (.018)	-.225 (.017)	-.222 (.016)
Denmark (wage ratio in percentage) (1870–1965)							
$\beta$	-.437 (.060)	-.447 (.046)	-.418 (.041)	-.401 (.037)	-.397 (.031)	-.397 (.027)	-.391 (.024)
$\gamma$	-.259 (.043)	-.255 (.039)	-.261 (.040)	-.261 (.040)	-.254 (.039)	-.239 (.036)	-.217 (.032)
Ireland (wage ratio in percentage) (1926–84)							
$\beta$	-.491 (.100)	-.474 (.118)	-.361 (.156)	-.262 (.145)	-.076 (.091)	-.123 (.085)	-.352 (.096)
$\gamma$	-.397 (.051)	-.405 (.063)	-.370 (.079)	-.326 (.080)	-.236 (.054)	-.256 (.052)	-.365 (.058)

<sup>a</sup> Table reports results of OLS estimates of the equation  $Inequality_t = \alpha + \beta T_1 + \gamma T_2 + \varepsilon_t$ .  $T_1$  is set equal to  $t$  for all years previous to the chosen break point and it is set at  $T_1 = 0$  for all  $t$  subsequent to the break.  $T_2$  takes a value of zero for all  $t$  prior to the break and a value of  $t$  for all  $t$  after the break. Newey West standard errors are reported in parentheses.

series against the null hypothesis that the series is characterized by a single trend. The most familiar way to test this hypothesis is a Chow test. Richard Quandt suggested that in the absence of prior knowledge where a break occurs, one could perform a Chow test on each point in the series and choose the point with the largest Chow statistic as the most likely break point.<sup>57</sup> Donald Andrews provided critical values that can be combined with Quandt's method to determine whether there is a statistical break in the series at a specific point.<sup>58</sup> We performed the Quandt-Andrews test for structural breaks on each of the series producing results very similar to those we produced based on the Table 5 regressions. This test shows no evidence of a structural break at any point in any of the six series.

### IMPLICATIONS

Our empirical tests suggest that certain hypotheses about the political correlates of inequality that are supported in analyses for recent decades find much less support from an examination of the evolution of inequality over the rest of the twentieth century. This result is not simply attributable to the fact that if we take a sufficiently long-run view, political factors will always wash out. If we look at earlier periods over the short run, we continue to find little support for the existing arguments. In this section we review our findings regarding government partisanship and wage bargaining centralization, and we then briefly consider three possible explanations for our results.

To begin, our results point first to a need to reconsider the relationship between government partisanship and inequality; a logical next step would be to examine whether our conclusion regarding partisanship also holds when looking at redistributive policies like progressive income and estate taxation for periods before the 1970s. One might find here that in a period like the 1950s and 1960s all governments, irrespective of partisan orientation, implemented highly progressive tax policies. The recent work by Thomas Piketty and Emmanuel Saez shows that a detailed comparison of the progressivity of the tax system in the United States, the United Kingdom, and France over time reveals surprising results as governments dominated by parties of the left have not had a tendency to pursue more progressive taxation.<sup>59</sup> One

<sup>57</sup> Quandt 1960.

<sup>58</sup> See Andrews 1993. The standard critical values for a Chi<sup>2</sup> test are not applicable unless the break date is known *ex ante*.

<sup>59</sup> Piketty and Saez 2003. The work of Atkinson and Leigh 2007 also raises interesting questions in this regard.

may question whether our results regarding partisanship are simply attributable to measurement error. Perhaps existing indicators of partisanship, based on either a simple left-right dichotomy or manifesto coding, are simply insufficient in tracking, for example, the extent to which the Eisenhower administration in the U.S. had a stance on redistributive policy markedly different from that of subsequent Republican administrations. The alternative conclusion one might draw based on our empirical results is that if partisanship has been highly correlated with the extent of redistributive policy (and thus income inequality) in recent decades, the principal story over the long run may be that with regard to redistribution parties of both the left and the right tend to be pushed in the same direction by underlying forces or common processes. The question would, of course, then become what these underlying forces or processes are.

In addition to our findings regarding partisanship, we also observe little evidence of a robust correlation between top income shares and centralized wage bargaining. This conclusion is reinforced when we look at individual country time series. Here we observe that income inequality did decline after the introduction of centralized wage bargaining, but it had already been declining beforehand and it continued to decline at a similar rate afterward. Moreover, inequality was also declining in countries that never adopted centralized wage bargaining. This raises questions about the extent to which it is more accurate to think of centralized wage bargaining as an institution that poses constraints on actors and that produces lower inequality, or alternatively whether centralized wage bargaining was simply a policy outcome produced by an underlying process that influenced both inequality and the formal setup of labor-market institutions. Our pooled estimation results also raise important questions about how countries without centralized wage bargaining arrangements may nonetheless have experienced dramatic declines in income inequality during the immediate postwar period.

A first possible explanation for our results is that political factors have had little influence on the evolution of income inequality over the long run, as income differentials have instead been driven by exogenous economic forces involving the race between technology and education.<sup>60</sup> When we consider the case of wage bargaining, it may be that exogenous changes in the relative demand for different skills in an economy

<sup>60</sup> By stating this we would, of course, be distinguishing this argument from one in which political factors involving partisanship or institutions would themselves influence both technology and education.

influence the willingness of highly skilled individuals to participate in centralized bargaining arrangements that pool skilled workers with unskilled workers; it may also be that at the same time the presence or absence of centralized bargaining in turn determines the degree of pre-tax income inequality. In this case wage bargaining centralization could lead to lower inequality, but it would be only a proximate cause.<sup>61</sup> Lars Svensson has recently proposed exactly this type of argument for understanding the emergence and subsequent dismantling of centralized wage bargaining in Sweden.<sup>62</sup> Depending on the precise assumptions made, this type of model could be consistent with the empirical observation of income inequality trending downward prior to a shift toward more centralized bargaining. However, this type of model would not be able to explain why, from relatively similar prior levels of inequality in the 1930s, some states chose to adopt centralized wage bargaining while others did not, yet almost all states experienced a significant reduction in inequality in the immediate postwar period.

A second possible explanation for our results is that income inequality was endogenous to an underlying *political* process that we have not captured with our data. The experiences of economic crisis (for a country like Sweden) or economic crisis combined with a major war during the 1930s and 1940s (for most other countries in our sample) may have been particularly important in this regard. Thomas Piketty and Emmanuel Saez have emphasized the effect of wartime taxation and destruction on large fortunes, suggesting a direct effect of war in narrowing inequality.<sup>63</sup> For the period after the war the establishment of high top marginal tax rates may have prevented prewar fortunes from being reconstituted quickly. This leaves open the question of why the postwar consensus on progressive income taxation was politically sustained. It may be the case that the postwar decrease in inequality would not have occurred without a change in the political climate that was itself the product of wartime experience and mobilization.<sup>64</sup> This political change may have made progressive income taxation and other new redistributive policies like the U.S. GI Bill politically feasible.

A further potential explanation for our results regarding centralized wage bargaining involves the potential effect of norms regarding pay. Centralized wage bargaining may at certain times (for example, the

<sup>61</sup> Acemoglu, Aghion, and Violante 2001 have used this type of theoretical framework to suggest why participation in trade unions and income inequality may both depend on technological changes that alter relative demand for highly skilled labor.

<sup>62</sup> Svensson 2004.

<sup>63</sup> Piketty 2003, 2001; and Piketty and Saez 2006, 2003.

<sup>64</sup> See Scheve and Stasavage 2008.

1980s) appear to have mattered for inequality because differences in the formal degree of centralization between countries like the United States and Sweden also corresponded with underlying differences in norms or beliefs about pay inequality. But at other junctures, such as the 1950s, underlying norms about pay inequality may have been much more similar across countries irrespective of differences in the degree of centralization in their bargaining arrangements. The important recent work by Frank Levy and Peter Temin on postwar U.S. inequality is particularly interesting in this regard, as it suggests how wage-setting norms in the U.S. in the immediate postwar period may have been quite similar to those operating in a number of continental European countries despite sharp differences in the degree of formal wage bargaining centralization.<sup>65</sup> If one wanted to pursue this line of reasoning, the difficult question is how would it be possible to identify the existence of beliefs about pay inequality independent of their effects.

A final alternative explanation, which also focuses on differences over time, is that centralized wage bargaining and partisanship have a causal effect on inequality but only under certain conditions. Identifying these conditions may also be a productive line of inquiry for future research.

### CONCLUSION

In this article we have suggested that while explaining post-1970 differences in income inequality between OECD countries is an important task, it is also the case that convincing comparative political economy hypotheses should be able to account as well for inequality trends in earlier time periods. We have found little evidence that government partisanship and wage bargaining centralization can account for variation in inequality over the long run. When we take a closer look at the evolution of inequality and of wage bargaining in individual countries—in particular in those countries that established a centralized system of wage bargaining in the middle of the twentieth century—we continue to see little evidence of an effect on inequality. In those countries that adopted centralized wage bargaining, income inequality was indeed lower after this institutional change, but inequality had been trending downward well before the change, and the institutional change was not associated with either a onetime shift downward or with a change in this trend. This raises questions about the extent to which centralized

<sup>65</sup> Levy and Temin 2007.

wage bargaining is an institution that has a causal effect on inequality or alternatively whether centralized bargaining is simply an outcome that has, along with income equality, evolved over time in response to an underlying political or economic process. More generally, our results also suggest that while scholars in comparative political economy have been most active in developing hypotheses that explain cross-country differences in inequality, they have focused less on developing effective arguments to account for changes in inequality over time. In other words, while it is important to seek explanations for the current difference in inequality between the United States and Sweden, political scientists also ought to attempt to explain why in the 1950s and 1960s these countries appeared much less different in terms of inequality than they do today.

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