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Comparative Political Economy of Wage Distribution: The Role of Partisanship and Labour Market Institutions

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Through a pooled cross-section time-series analysis of the determinants of wage inequality in sixteen OECD countries from 1973 to 1995, we explore how political-institutional variables affect the upper and lower halves of the wage distribution. Our regression results indicate that unionization, centralization of wage bargaining and public-sector employment primarily affect the distribution of wages by boosting the relative position of unskilled workers, while the egalitarian effects of Left government operate at the upper end of the wage hierarchy, holding back the wage growth of well-paid workers. Further analysis shows that the differential effects of government partisanship are contingent on wage-bargaining centralization: in decentralized bargaining systems, Left government is associated with compression of both halves of the wage distribution.

It is well known that wage inequality has increased dramatically in the United States over the last three decades. From 1973 to 1998, the hourly earnings of a full-time worker in the ninetieth percentile of the American earnings distribution (someone whose earnings exceeded those of 90 per cent of all workers) relative to a worker in the tenth percentile grew by 25 per cent, and the corresponding figure for men only was nearly 40 per cent. As we document in this article, wage inequality has increased in most Organization for Economic Co-operation and Development (OECD) countries, but the extent of this phenomenon varies a great deal, and cross-national differences in levels of wage inequality remain as great as they were in the 1970s. In the United States, the worker in the ninetieth percentile earned 4.63 times as much as the worker in the tenth percentile in 1996. At the other end of the cross-national spectrum, the 90–10 ratio for Sweden was only 2.27.

While political commentators in Europe and the United States alike frequently invoke inegalitarian labour market trends to explain various manifestations of working-class political disaffection (not only support for right-wing populist parties, but also falling turnout among working-class

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voters), recent work by labour economists demonstrates that supply and demand factors alone cannot account for cross-national variation in wage inequality.[1] Wage inequality appears to have political determinants as well as political consequences. On both counts, it deserves to be a central concern of comparative political economy as conceived and practised by political scientists. Within political science, however, the paucity of research on wage inequality stands in sharp contrast to the large number of quantitative studies that take various measures of macroeconomic performance or government spending as their dependent variable. Drawing on a new dataset published by the OECD,[2] which enables us to engage in a pooled cross-section time-series analysis of the determinants of wage inequality in sixteen OECD countries for the period 1973–95, we seek to make up for some of this neglect.[3]

From the perspective of comparative political economy, the manner in which labour economists deal with cross-national differences in government policy and the organization of wage bargaining leaves something to be desired. Most commonly, rather vaguely specified institutional factors are invoked to explain whatever variance remains when the effects of supply and demand have been taken into account. And when economists incorporate cross-national differences into their models, they typically reduce these differences to a single, one-dimensional variable, such as wage-bargaining centralization.

Our analysis demonstrates that it is possible to distinguish discrete effects of several political-institutional variables. Controlling for certain supply and demand conditions as well as country-specific fixed effects, we find that bargaining centralization, public sector employment, union density and left government are all negatively associated with overall wage inequality, measured by 90–10 ratios. In other words, higher values on these independent variables are associated with lower 90–10 ratios. While bargaining centralization and the size of the public sector can be characterized as 'institutional' or 'structural' features of the political economy, union density and government partisanship pertain to the distribution of power between labour and employers, and also to the distribution of power among different categories of wage-earners. Institutions matter for the distribution of wages, and so do politics.

To get a better handle on the causal mechanisms at work, we engage in separate regression analyses of the determinants of the ratio of earning in the

[1] For example, Richard Freeman and Lawrence Katz, eds, *Differences and Changes in Wage Structures* (Chicago: The University of Chicago Press, 1995), pp. 1–24; Francine Blau and Lawrence Kahn, 'International Differences in Male Wage Inequality', *Journal of Political Economy*, 104 (1996), 791–836; and Peter Gottschalk and Timothy Smeeding, 'Cross-national Comparisons of Earnings and Income Inequality', *Journal of Economic Literature*, 35 (1997), 633–87.

[2] OECD, 'Earnings Inequality', *Employment Outlook* (July 1993), 157–84; OECD, 'Earnings Inequality, Low-paid Employment and Earnings Mobility', *Employment Outlook* (July 1996), 59–108.

[3] For similar efforts by other political scientists, see Torben Iversen and Anne Wren, 'Equality, Employment and Budgetary Restraint', *World Politics*, 46 (1998), 527–55; and Michael Wallerstein, 'Wage-Setting Institutions and Pay Inequality in Advanced Industrial Societies', *American Journal of Political Science*, 43 (1999), 649–80.

ninetieth percentile to median earnings (the '90–50 ratio') and the ratio of median earnings to earnings in the tenth percentile (the '50–10 ratio'). Basically, we ask if these political-institutional variables promote a more egalitarian wage structure by holding back wage growth for highly-paid wage-earners at the upper end of the distribution, or by raising the relative wages of people at the bottom of the wage hierarchy. Our results indicate that bargaining centralization and public sector employment have discernible egalitarian effects in both halves of the wage distribution, although both primarily operate in the bottom half of the wage hierarchy. In contrast, the egalitarian effects of union density appear to be entirely confined to the lower half of the wage distribution, and the opposite holds for left government, which operates almost entirely in the upper half of the wage hierarchy.

In the light of conventional wisdom in the comparative political economy literature, the absence of egalitarian effects of left government at the lower end of the wage distribution is puzzling. Much of this literature leads us to expect that labour-affiliated left parties strive to raise the floor for competition among unskilled, low-paid workers by providing generous unemployment compensation and by boosting mandated minimum wages. As we document below, the cross-national association between left government and levels of unemployment compensation is less strong than conventional wisdom implies, and the association between generosity of unemployment compensation and compression of the lower half of the wage distribution is also quite weak. More importantly, the final iteration of our regression analysis tests the hypothesis that the wage-distributive effects of government partisanship are contingent on the degree of wage bargaining centralization. In comprehensive and centralized wage bargaining systems of the Northern European variety, we should not expect minimum wage legislation to have much impact on the distribution of wages. The results of an interaction model support this hypothesis: under decentralized wage bargaining, left government turns out to have egalitarian effects at the lower end as well as the upper end of the wage distribution, but the egalitarian effects of left government at the lower end diminish as bargaining centralization increases.[4]

Our presentation is organized as follows. We begin with a quick look at the cross-national patterns of wage inequality that our analysis seeks to explain. We then review the existing literature, present our hypotheses and independent variables, briefly explicate the methodology of pooled cross-section time-series analysis, and discuss the results of linear regressions with 90–10, 90–50 and 50–10 wage ratios as the dependent variable. Along the lines indicated above, we end by further exploring the wage-distributive effects of government

[4] The interaction argument builds on David Rueda and Jonas Pontusson, 'Wage Inequality and Varieties of Capitalism', *World Politics*, 52 (2000), 350–83, which argues that the determinants of wage inequality differ across political economy types. Exploring how the determinants of wage inequality differ across the wage hierarchy, this article was conceived as a companion to the *World Politics* piece. In future work, we plan to integrate the two approaches.

partisanship, and conclude by speculating about future trends and the political consequences of rising wage inequality.

PATTERNS OF WAGE INEQUALITY

Table 1 summarizes the wage inequality observations which serve as the dependent variable of our analysis. For each country, the table provides the mean value for each specification of the dependent variable over the entire period 1973–95 and also the percentage change from the earliest to the most recent observation. It should be noted at the outset that these inequality measures refer to gross income from employment for individuals: they ignore other sources of income (government transfers, self-employment, income from capital, etc.) and also leave out the distributive effects of taxation and income pooling within households. What follows must not be confused with an analysis of the distribution of disposable household income.

We also note that the OECD dataset on which we rely is restricted to full-time employees, except in the case of Austria. Since part-time employees invariably earn less, on an hourly basis, than full-time employees, the figures in Table 1 understate the extent of wage inequality in the other countries. And since the incidence of part-time employment has increased in most OECD countries since the early 1980s, they also understate the upward trend in wage inequality.[5] Keeping these qualifications in mind, income from employment still accounts for the lion's share of total income in all OECD countries, and wage inequality among full-time employees, measured by 90–10 ratios, still correlates quite closely with broader cross-national measures of income distribution.[6]

Table 1 reveals dramatic cross-national variation in wage inequality. In these sixteen countries, the average both-gender 90–10 ratio for the 1973–95 period was 2.89. In other words, a person in the ninetieth percentile of the wage distribution earned, on average, nearly three times as much as a person in the tenth percentile. Sweden, with an average 90–10 ratio of 2.07, stands out as the OECD country with the most compressed overall wage distribution. With the notable exceptions of France and Austria, the continental European countries included in this dataset (Belgium, Germany, Italy, the Netherlands and

[5] The exclusion of part-time employees does not pose a serious problem for cross-sectional comparison, for median hourly earnings of part-time workers as a percentage of median hourly earnings of full-time workers correlates very closely with 90–10 ratios among full-time workers cross-nationally. But this data restriction does pose a problem from the point of view of comparing changes in wage inequality across countries, since the growth of part-time employment varies considerably across countries. From 1983 to 1998, the incidence of part-time employment actually declined in the United States and the Scandinavian countries (see OECD, *Employment Outlook*, July 1999, p. 24, for data on part-time wages as a percentage of full-time wages and p. 240 for data on the incidence of part-time employment).

[6] Cf. OECD, *Income Distribution in OECD Countries: Evidence from the Luxembourg Income Study* (1995); Gottschalk and Smeeding, 'Cross-National Comparisons of Earnings and Income Inequality'; and Wallerstein, 'Wage-Setting Institutions and Pay Inequality in Advanced Industrial Societies'.

TABLE 1 *Means and Percentage Changes of the Dependent Variables, 1973–95*

Country and years covered	90–10 ratios		90–50 ratios		50–10 ratios	
	Mean	% change	Mean	% change	Mean	% change
Australia (1976–95)	2.81	10.6	1.69	6.6	1.66	3.1
Austria (1980–94)	3.53	6.1	1.79	1.7	1.96	0.0
Belgium (1986–93)	2.34	− 6.7	1.62	− 4.9	1.45	− 1.4
Canada (1973–94)	4.24	12.1	1.82	2.8	2.30	9.1
Denmark (1980–94)	2.18	0.0	1.55	3.3	1.40	− 2.8
Finland (1977–95)	2.45	− 11.7	1.68	− 1.2	1.46	− 10.2
France (1973–95)	3.27	− 10.8	1.96	2.9	1.66	− 5.7
Germany (1984–95)	2.79	− 9.7	1.71	2.4	1.63	− 11.9
Italy (1986–95)	2.32	5.9	1.63	9.9	1.42	− 3.4
Japan (1975–95)	3.07	− 1.3	1.81	5.1	1.70	− 6.3
Netherlands (1977–95)	2.54	9.7	1.63	6.2	1.56	5.8
Norway (1980–94)	2.08	− 3.8	1.49	2.7	1.39	− 6.4
Sweden (1975–95)	2.07	− 1.8	1.56	− 1.2	1.33	0.0
Switzerland (1990–95)	2.72	2.2	1.69	2.4	1.61	0.0
United Kingdom (1973–95)	3.17	13.5	1.77	11.1	1.78	1.5
United States (1973–95)	4.60	22.3	2.04	10.7	2.00	11.0
Average	2.89	2.29	1.72	3.78	1.64	− 1.1
Standard deviation	0.74	9.79	0.15	4.40	0.26	6.38

Notes and sources: The percentage changes measure the variation from earliest to latest available observation in the country series. See OECD, 'Earnings Inequality, Low-paid Employment and Earnings Mobility', pp. 61–2 for all countries except the United States; for the United States, OECD, 'Earnings Inequality', p. 161, and OECD, 'Earnings Inequality, Low-paid Employment and Earnings Mobility, p. 103.

Switzerland) fall within a rather narrow band (2.34–2.79), well below the OECD average. At the opposite end of the spectrum, the United States occupies an even more distinctive position than Sweden, with an average 90–10 ratio of 4.6. Canada, France, Japan and the United Kingdom also turn out to be considerably more inegalitarian than the OECD average.

Turning to change over time, the cross-national variation in the data is equally impressive. From the earliest to the most recent observation available for each country, we observe large increases of 90–10 ratios in the United States, the United Kingdom, Canada, Australia, the Netherlands, Austria and Italy. However, overall wage inequality fell quite significantly in Finland, France, (West) Germany and Belgium over this period, while the remaining countries – Denmark, Japan, Norway, Sweden and Switzerland – are best characterized as cases of stability. For all sixteen countries, the average 90–10 ratio increased by 2.29 per cent. It is tempting to conclude from the 90–10 data in Table 1 that there is no common trend for wage inequality to increase in the OECD countries, but this conclusion may be a bit hasty. In many of these countries, wage inequality declined in the early part of the period covered by these summary measures and subsequently rose. If we measure change since the trough of wage inequality in each country, we observe increases of wage inequality in eleven out of sixteen, and the magnitude of these increases are typically greater than those shown in Table 1. Also, it is noteworthy that the tendency for rising wage inequality becomes broader and more pronounced when one looks at data for men and women separately.[7] In most of these countries, compression of between-gender differentials offset the effects of growing within-gender differentials in the 1980s. The growth of wage inequality among men has been especially strong, and this relates to another feature of Table 1: over the period 1973–95, the tendency for wage inequality to rise was stronger in the upper half than the lower half of the wage distribution. (This is related to the point about male inequality because men are over-represented in the upper half of the distribution.)

CAUSAL HYPOTHESES

Wage inequality measures such as the 90–10 ratio are summary measures of a complex, multidimensional wage structure, which might be decomposed into a number of different kinds of wage differentiation: most obviously, differentials based on education, work experience, gender, race and corporate profitability. Comparable in degree, wage dispersion may take different forms in different countries. In view of this complexity, it would surely be quixotic to look for a very simple explanation of the cross-national patterns of wage inequality described above. A number of variables are bound to matter and

[7] See Jonas Pontusson, David Rueda and Christopher Way, 'The Role of Political-Institutional Variables in the Making of Gendered Patterns of Wage Inequality' (Working paper, Institute for European Studies, Cornell University, 1999).

TABLE 2 *Theoretical Expectations*

Explanatory variables	Direction of overall effect on 90–10 ratios	Relative magnitude of effect across the wage hierarchy: Upper half (90–50 ratio)	Lower half (50–10 ratio)
Political Institutional Variables			
Union density	Negative	Weak	Strong
Bargaining centralization	Negative	Weak	Strong
Public employment	Negative	Weak	Strong
Left government:			
Wage floor variant	Negative	None	Strong
Marginal taxation variant	Negative	Strong	None
Market Forces Variables			
Unemployment	Positive	Weak	Strong
LDC trade	Positive	Weak	Strong
Female labour-force participation	Positive	Weak	Strong
Private service employment:			
Demand for 'food and fun' variant	Uncertain	None	Uncertain
Innovation incentives variant	Positive	Strong	Weak

it seems most appropriate to inquire about their marginal effects. In other words, this is the kind of research question that calls for multiple regression analysis.

Table 2 identifies eight independent variables that are included in our regression models, and also summarizes our expectations of their effects on the distribution of wages. The variables are grouped into two clusters. The first cluster includes union density, wage-bargaining, public sector employment and government partisanship. These political and/or institutional variables are the variables of primary theoretical interest. The second cluster includes the rate of unemployment, trade with low-wage countries, female labour force participation and private service employment. We conceive these as control variables, designed to capture, in an admittedly crude fashion, supply and demand conditions emphasized by labour economists.[8] Leaving definitions and data sources for the Appendix,[9] let us elaborate on our theoretical expectations for each of the eight independent variables.

[8] A direct test of the supply and demand arguments advanced by labour economists would require the use of individual-level data.

[9] The Appendix is available with the website version of this article.

Union Density

Following Freeman, we can distinguish two dimensions of the relationship between unionization and wage distribution.[10] The first dimension concerns the distribution of wages among union members, and how it compares to the distribution of wages among unorganized wage-earners. The second concerns wage differentials between union members and non-members; in other words, the wage premium associated with union membership for workers with equivalent qualifications, experience and other relevant characteristics. Several arguments lead us to expect that the wage distribution of unionized firms or sectors will be more compressed than that of the non-unionized firms or sectors. First, most unions approximate the logic of democratic decision making (one person, one vote) more closely than markets do, and whenever the mean wage exceeds the median wage, we would expect a majority of union members to favour redistributive wage demands. Secondly, as organizations dependent on membership support in conflicts with management, unions have a strong interest in curtailing wage setting based on the subjective decisions of foremen and personnel managers, in order to curtail their ability to discourage activism by rewarding pliant behaviour. Unions and employers operating in the same product markets can both be expected to favour standardization of wage rates across firms – to take wages out of competition – but unions alone have an unambiguous interest in standardizing wage rates within firms.

The issue of union wage premia renders the relationship between unionization and wage distribution more complicated, for the net wage-distributive effects of unionization also depend on the distribution of union membership across the wage hierarchy. Assuming that there is a wage premium for unionized workers, unionism would be a source of wage inequality if high-wage workers were better organized than low-wage workers, whereas the opposite would hold if low-wage workers were better organized. Helping to sort out this question, recent Eurobarometer surveys enable us to estimate union density by income quartile for the member states of the European Union. In 1994, average union density for the EU as a whole was 37.5 per cent in the first (lowest) income quartile, 37.8 per cent in the second quartile, 34.2 per cent in the third quartile and 23.7 per cent in the fourth quartile.[11] While the distribution of union membership across the wage hierarchy differs somewhat from one country to another, this general picture is consistent with conventional wisdom and leads us to expect union density to be negatively associated with wage inequality in our pooled regressions.

Given that we expect union density to be associated with wage compression and that the lower half of the wage distribution tends to be more unionized than

[10] Richard Freeman, 'Unionism and the Dispersion of Wages', *Industrial and Labor Relations Review*, 34 (October 1980), 3–23; and Richard Freeman, 'Union Wage Practices and Wage Dispersion within Establishments', *Industrial and Labor Relations Review*, 36 (October 1982), 3–20. Cf. also Blau and Kahn, 'International Differences in Male Wage Inequality'.

[11] Eurobarometer, June–July 1994.

the upper half, it follows that, everything else being equal, the lower half of the wage distribution should be more compressed than the upper half. However, it does not follow that the egalitarian effect of unionization (the effect of, say, a 5 per cent increase of union density) on the lower half of the distribution should be greater than the effect on the upper half. The question of differential effects across the wage hierarchy requires us to introduce additional hypotheses. In the case of union density, we hypothesize that unions which primarily organize workers in the lower half of the wage distribution tend to be more solidaristic in their wage demands than unions which primarily organize workers in the upper half of the distribution. The logic behind this hypothesis hinges on the proposition that the highly educated workers who occupy the upper end of the wage distribution enjoy a great deal of individual bargaining power in the labour market and stand to gain relatively little by joining a union. Unions which aspire to organize or retain the loyalty of these workers must moderate their pursuit of wage compression. For unions whose best-paid members occupy the middle of the wage hierarchy this problem is much less pronounced, if not absent. Hence we expect union density to have greater egalitarian effects on 50–10 ratios than on 90–50 ratios.

Bargaining Centralization

The standard argument linking centralization to wage compression asserts that centralization facilitates the reduction of inter-firm and inter-sectoral wage differentials since it means that more firms and sectors are included in a single wage settlement. As Rowthorn suggests, this logic presupposes that at least one of the parties to centralized bargaining wants to achieve a reduction of inter-firm or inter-sectoral differentials.[12] Arguably, then, centralization is a facilitating factor or perhaps a necessary but not sufficient condition for wage compression.

In a somewhat different vein, one might argue that centralization produces wage compression by altering the distribution of power among actors. In the Swedish case, low-wage unions insisted on solidaristic measures as a condition for their participation in peak-level bargaining sought by employers in the 1950s.[13] But why should centralization systematically strengthen the relative bargaining power of low-wage unions? The logic of the median voter model might apply here as well: if low-wage and high-wage unions bargain jointly, organizational politics will influence the demands that they pursue – exerting pressure for compression – and market forces will be less influential in determining the distribution of wage increases.[14] Moreover, we hypothesize that centralized bargaining in the extreme – a single settlement for all wage

[12] Bob Rowthorn, 'Corporatism and Labour Market Performance', in Jukka Pekkarinen, Matti Pohjola and Bob Rowthorn, eds, *Social Corporatism* (Oxford: Clarendon Press, 1992).

[13] Peter Swenson, *Fair Shares* (Ithaca, NY: Cornell University Press, 1989), pp. 56–8.

[14] Cf. Wallerstein, 'Wage-Setting Institutions and Pay Inequality in Advanced Industrial Societies'.

earners – renders wage differentials more transparent, and thus politicizes wage-distributive outcomes. In other words, centralization not only empowers low-wage unions, but also makes them more likely to demand redistributive wage settlements.

Centralization is a characteristic of the process of collective bargaining, and its effects depend on the extent to which wages across the employment spectrum are determined through collective bargaining. Because wages at the upper end of the wage distribution are less likely to be regulated by collective bargaining than wages at the lower end, we expect the egalitarian effects of centralization to be stronger for 50–10 ratios than for 90–50 ratios.[15]

Government Employment

While controlling for the strong association between public-sector employment and union density, we nonetheless expect the size of the public sector (government employees as a percentage of the total labour force) to be negatively associated with wage inequality for several reasons.[16] In general, public-sector unions appear to be more inclined to favour wage solidarity than private-sector unions. This is certainly true if one compares unions organizing similar services on different sides of the public–private divide in the 1970s and 1980s. At the same time, public-sector employers have been more inclined than private-sector employers to accommodate union demands for compression or even to initiate compression. While sheltered from competition in product markets, public-sector employers are more directly exposed to political pressures favouring equality and robust wage growth.[17] The egalitarian logic of public-sector wage setting is most pronounced with regard to equal pay provisions for women and minorities. Very well-documented in the case of the United States,[18] this point would seem to hold more broadly across the OECD countries.

These arguments imply that the wage structure of the public sector should be more compressed than that of the private sector, but a straight comparison of public and private wage distributions is problematic because the public sector encompasses a distinct set of economic activities, often characterized by a very wide spread of educational qualifications. Our expectation that the size of the

[15] To model the inertia associated with institutional change, the bargaining centralization value for a given year used in our analysis is the average for that year and the previous four years. Developed by Torben Iversen, the measure of bargaining centralization used here captures not only the level at which collective bargaining occurs ('centralization' in a narrow sense), but also the degree of concentration of union membership at different bargaining levels (see Appendix to the website version of this article for further details).

[16] OECD, 'Trends in Trade-Union Membership', *Employment Outlook* (July 1991), p. 113.

[17] Geoffrey Garrett and Christopher Way, 'Corporatism, Public Sector Employment, and Macroeconomic Performance', *Comparative Political Studies*, 32 (1999), 411–34.

[18] For example, Ronald G. Ehrenberg and Joshua L. Schwarz, 'Public Sector Labor Markets', in O. Ashenfeleter and R. Layard, eds, *Handbook of Labor Economics* (Amsterdam: North Holland, 1986), pp. 1219–68.

public sector has an egalitarian effect rests on the proposition that the wage structure of, for example, government-run health care is more compressed than the wage structure of privately run health care. Also, we expect that wage compression in the public sector spills over into the private sector as private-sector employers compete with public-sector employers for labour. Observing that the differences between public-sector and private-sector wage distributions are smaller in Sweden than in most countries,[19] we should not conclude that public-sector egalitarianism has been particularly weak in Sweden: more likely, the spillover effects of public-sector egalitarianism have been particularly strong in Sweden.

The spillover effects of public-sector egalitarianism are likely to be most pronounced at the lower end of the wage distribution, for the simple reason that private employers compete most directly with public-sector employers for unskilled labour. As we move up the wage hierarchy, the educational backgrounds and career patterns of public-sector and private-sector employees become more distinct. Similarly, the importance of equal pay provisions in the public sector should affect 50–10 ratios more than 90–50 ratios, since women tend to be over-represented in the lower half of the wage distribution. By contrast, unionization and collective bargaining encompass more of the upper end of the wage hierarchy in the public sector, constraining the ability of well-paid civil servants to capitalize on their marketplace power. On balance, we still expect the egalitarian effects of public employment to be strongest at the lower end of the wage distribution.

Left Government

Governments might influence wage distribution through minimum-wage and equal-pay legislation, other forms of incomes policy, arbitration in bargaining disputes, and a variety of measures that strengthen the competitive position of women and other disadvantaged groups (such as immigrants) in the labour market. Less obviously perhaps, tax policy might influence the distribution of primary as well as disposable income. Ideally, we would like to explore the wage-distributive effects of discrete government policies, but we do not have the data necessary to do so within the framework of pooled cross-section time-series analysis (requiring yearly observations for each variable). Instead, we propose to explore the role of government by including Tom Cusack's measure of government partisanship in our regressions.[20]

Two distinct arguments lead us to expect left government to be negatively associated with wage inequality. The first argument – the wage floor variant in

[19] Janet Gornick and Jerry Jacobs, *Gender, the Welfare State and Public Employment* (Luxembourg Income Study Working Paper, no. 168, 1997).

[20] Tom Cusack, 'Partisan Politics and Public Finance', *Public Choice*, 91 (1997), 375–95. The construction of this measure is described in the Appendix on the website version of this article. Note that we have inverted Cusack's index so that higher values signify more leftist government.

Table 2 – hinges on the proposition that the policy preferences of left parties raise the floor for competition in the labour market. If there is a legislated minimum wage, left governments are likely to set the minimum wage closer to the median wage than right governments. They are also likely to favour higher levels of unemployment compensation and to promote other social wage programmes, curtailing the inegalitarian effects of unemployment and, more generally, boosting the relative bargaining power of unskilled workers. The second argument – the marginal taxation variant – hinges on the proposition that left parties favour progressive income taxation. As Hibbs argues, high marginal tax rates reduce the value of an added increment of income to highly paid people, and might discourage wage earners in the upper reaches of the wage hierarchy from taking full advantage of their market power or perhaps using their market power to gain non-wage benefits from their employers instead.[21]

By analysing the determinants of 90–50 and 50–10 ratios, we test these two variants of the left government argument independently. Whereas the first argument holds that left government compresses the wage distribution by raising the relative wages of poorly paid workers, the second argument holds that left government compresses the wage distribution by holding back highly paid workers (of course, the two arguments are not contradictory: both effects may operate simultaneously).

Unemployment

In seeking to assess the wage-distributive effects of the political-institutional variables discussed above, we ought to control for the effects of market conditions, which also vary over time and across countries. The rate of unemployment is perhaps the most obvious control variable. The basic insight of the literature on labour market segmentation is that unskilled, low-paid workers are more readily substitutable than more skilled, high-paid workers, and consequently that their bargaining position is more immediately and more adversely affected by unemployment.[22]

However, there is another side to the relationship between unemployment and wage inequality, also related to labour market segmentation. As many studies have shown, employers are more likely to lay off unskilled than skilled workers during economic downturns. To the extent that it entails a disproportionate loss of low-paid jobs, an increase of unemployment produces wage compression by altering the composition of the labour force. To minimize this statistical effect, the unemployment variable in our regressions is a five-year moving average (the

[21] Douglas Hibbs, 'Fiscal Influences on Trends in Wage Dispersion' (paper presented at the annual meeting of the European Public Choice Society, Reggio Calabria, 1987); and Douglas Hibbs and Håkan Locking, 'Wage Compression, Wage Drift, and Wage Inflation in Sweden', *Labour Economics*, 77 (1996), 1–32.

[22] Cf. James Galbraith, *Created Unequal* (New York: Free Press, 1998); and Katherine Bradbury, 'Rising Tide in the Labor Market: To What Degree Do Expansions Benefit the Disadvantaged?' *New England Economic Review*, 32 (May–June 2000), 3–33.

observation for each year is the average for that year and the preceding four years). With this specification, we expect sustained higher rates of unemployment to be associated with higher levels of wage inequality, and the inegalitarian effects of unemployment to be concentrated in the lower half of the wage distribution.[23]

Trade with Less Developed Countries

Wood argues that much of the trend towards increased wage inequality in the OECD countries in the 1980s can be attributed to increased manufacturing trade with Less Developed Countries (LDCs).[24] The basic logic of Wood's analysis is that relative supplies of skilled and unskilled labour are a function not only of domestic conditions, but also of the factor content of trade. By importing less skill-intensive goods from low-wage countries, OECD countries are essentially importing low skill labour, so that the effective supply of unskilled labour relative to skilled labour has grown, putting downward pressure on the relative wages of the unskilled. To provide at least a partial test of this thesis, which resonates much of the popular literature on globalization, our regressions include trade with LDC countries that are not producers and exporters of oil (non-OPEC), expressed as a percentage of gross domestic product (GDP). Wood's argument implies that the inegalitarian effects of this variable should primarily manifest themselves in the lower half of the wage distribution.

Female Labour-Force Participation

To the extent that women are on average less educated and/or have less work experience than men, an increase of the proportion of the total labour force made up of women represents an increase of the relative supply of unskilled or less skilled labour.[25] Everything else being equal, we expect female labour-force participation to be associated with more wage inequality, especially in the lower half of the wage distribution. However, there are clearly countervailing forces at work here. Women acquire skills through labour-force participation, and higher levels of female labour-force participation should eventually reduce the skill gap between men and women. As women increasingly come to occupy full-time jobs, moreover, the union density gap between men and women should erode. As a consequence, women become a more important constituency for

[23] Using yearly observations rather than a moving average, Rueda and Pontusson do not find any wage-distributive effects of unemployment in 'Wage Inequality and Varieties of Capitalism'.

[24] Adrian Wood, *North–South Trade, Employment and Inequality* (Oxford: Clarendon Press, 1994).

[25] Cf. Robert Topel, 'Wage Inequality and Regional Labor Market Performance in the United States', in Toshiaki Tachibanaki, ed., *Labour Market and Economic Performance* (New York: St Martin's Press, 1994), pp. 101–32; Lennart Svensson, *Closing the Gender Gap* (Lund: Ekonomisk-historiska föreningen, 1995).

unions, encouraging them to pursue a reduction of wage differentials based on gender. The data at our disposal do not allow us to explore these countervailing effects with much precision. Still, given the dramatic increase in female labour-force participation in recent decades, it makes sense for us to control for women's share of total employment.[26]

Private Service Employment

Relative to manufacturing and public services, private services became an increasingly important source of employment in all the OECD countries over the period covered by this analysis (1973–95). Based on the American experience, it is commonplace to argue that wage inequality and private service employment are associated. Treating wage inequality as a precondition for employment growth, the standard version of this argument asserts that the scope for productivity growth in personal services is inherently limited, that pricing closely reflects labour costs, and that demand for these services is highly price sensitive.[27] In countries with a high wage floor, the expansion of 'food and fun' sectors, employing largely unskilled labour, will be sluggish at best. Though the causal arrows are reversed, the implication of this argument is that the coefficient for private service employment in our regressions should be positive and largest in the regression that uses 50–10 ratios as the dependent variable.

If we relax the assumption that the production of personal services with a high content of unskilled labour is tightly constrained by labour costs, we might expect the opposite association between wage inequality and private service employment. After all, the expansion of such services entails a relative increase in demand for unskilled labour, boosting the bargaining position of workers at the bottom of the wage hierarchy. In a somewhat different vein, it seems plausible to argue that productivity growth in high-end services involves radical innovation and depends on wage incentives to a much greater extent than productivity growth in manufacturing, which tends to be piecemeal and closely linked to capital investment. By this reasoning, the inegalitarian effects of private service employment should manifest themselves primarily in the upper half of the wage hierarchy.

As with female labour-force participation, our goal here is not to sort out the complex relationship between wage inequality and private service employment. Again, we are first and foremost interested in the wage-distributive effects of our four political-institutional variables; the other four variables are included as controls.

[26] On similar grounds, we would like to be able to control for immigration, but the data that we would need to do that is simply not available. Also, we have been unable to locate comparable cross-national data on the distribution of educational qualifications (data on average years of education are available, but not theoretically relevant to our research question).

[27] Cf. Iversen and Wren, 'Equality, Employment and Budgetary Restraint'.

METHODOLOGY

Pooled cross-section time-series analysis has become common practice in quantitative comparative political economy in recent years.[28] In this type of analysis, country-years are the units of observation of dependent and independent variables; in other words, regressions are run on multiple observations for each country, allowing us to take advantage of variation between, say, Sweden 1990 and Sweden 1991 as well as the variation between Sweden 1990 and Germany 1990. Pooling is particularly attractive when time series are short and/or the number of cross sections small, as is often the situation for comparative political economists because our theories and data pertain to a small number of countries (typically, ten to twenty OECD countries). By incorporating over-time variations, pooling dramatically increases the total number of observations, and this in turn enables us to determine the statistical significance of results with greater precision and to test more complex causal models by including more variables in regression.

However, pooled data analysis must not be seen simply as a technical solution to the small-*N* problem of comparative political economy. This methodology is inextricably linked to the idea that cross-national variations and changes over time have common determinants. As Shalev points out, it is common for cross-sectional and time-series regressions with the same variables to produce different results.[29] By engaging in pooled data analysis, we ask: What explains the observed variance across both space and time?

Ordinary least squares (OLS) regression rests on a number of assumptions about the data-generating process producing the observed values on the independent variables and the error term. The pooling of data is likely to violate some of these assumptions for reasons articulated by, among many others, Beck and Katz.[30] Beck and Katz show that one common solution to some of these problems, the Parks–Kmenta method, consistently underestimates parameter variability. Following their recommendations, now accepted by many comparativists, we report OLS estimates of the coefficients and panel-corrected estimates of their standard errors.

In our regression models, we deal with dynamics and auto-correlation by including lagged values of the dependent variable on the right-hand side of the equation.[31] On the assumption that the effects of a one-unit change in a particular

[28] See Alexander Hicks, 'Introduction to Pooling', in Thomas Janoski and Alexander Hicks, eds, *The Comparative Political Economy of the Welfare State* (New York: Cambridge University Press, 1994), pp. 169–88; and Nathaniel Beck and Jonathan Katz, 'What to Do (and Not to Do) with Time-series Cross-section Data', *American Political Science Review*, 89 (1995), 634–47, for citations and general discussion.

[29] Michael Shalev, 'Limits of and Alternatives to Multiple Regression in Macro-comparative Research' (paper presented at conference on 'The Welfare State at the Crossroads', Stockholm, 1998).

[30] Beck and Katz, 'What to Do (and Not to Do) with Time-series Cross-section Data'.

[31] See Nathaniel Beck and Jonathan Katz, 'Nuisance vs. Substance: Specifying and Estimating Time-series Cross-section Models', in John Freeman, ed., *Political Analysis* (Ann Arbor: University

variable persist indefinitely, the total effect of a change – that is, its effects over an infinite period of time – can be computed by dividing the value of the coefficient for the variable of interest by one minus the coefficient for the lagged dependent variable.[32] In what follows, we report long-run as well as short-run effects for each variable.

Our models also include dummy variables for each of the countries in our dataset, though we do not report the coefficient estimates for these variables.[33] Put somewhat crudely, the country dummies control for the values that all observations for a given country share by representing the variance unique to that country. Controlling for omitted variable bias, the inclusion of country dummies in the regression facilitates the estimation and interpretation of the coefficients by clearing out the influences of country-specific factors.

Our regression models, then, take the following general form:

$$Y_{it} = \beta_0 Y_{i(t-1)} + \beta_1 X_{1it} + \beta_2 X_{2it} + \ldots + \beta_n X_{nit} + \alpha_1 C_{1t} + \alpha_2 C_{2t} + \ldots \alpha_{16} \mathrm{C}_{16t} + \varepsilon_{it}$$

where $i = 1, 2 \ldots, i$ refers to the cross-sectional units, $t = 1, 2, \ldots, t$ refers to the time units; Y_{it} is the dependent variable; $Y_{i(t\text{-}1)}$ is the lagged dependent variable; $X_1, \ldots, X_\mathrm{n}$ are the other explanatory variables; $\beta_0, \beta_1, \ldots, \beta_n$ are the slopes of the explanatory variables; $C_1, \ldots, C_{16}$ are the country dummy variables; $\alpha_1, \ldots, \alpha_{16}$ are the intercepts for each country; and ε_{it} is an independent random error term normally distributed around a mean of 0 and with a variance of σ^2.

It should be noted, finally, that we applied logarithmic transformations to all our variables (except for the country dummies) before running the regressions reported below.[34] The slope estimates yielded by these regressions should therefore be understood as percentage changes or elasticity measures representing the relationship between the variables. In other words, the regression coefficients should be interpreted as the percentage change in the dependent variable associated with a 1 per cent change in the independent variable in question.

(F'note continued)

of Michigan Press, 1996), vol. 6, pp. 1–36. As with all time-series data, the possibility of non-stationarity must be considered. Dickey–Fuller tests for the pooled data revealed no evidence suggesting that the inequality series was non-stationary: the null hypothesis of non-stationarity was rejected at better than the 95 per cent level in tests with and without a time trend. The results of Breusch–Godfrey tests indicate that there is no significant auto-correlation in the reported regressions after the inclusion of the lagged dependent variable. See William Greene, *Econometric Analysis* (Englewood Cliffs, NJ: Prentice Hall, 1990), pp. 426–8.

[32] George Box and Gwilyn Jenkins, *Time Series Analysis* (Oakland, Calif.: Holden-Day, 1976), chap. 1.

[33] Technically, this means that we estimate Least Squares Dummy Variable (LSDV) models. We are able to avoid perfect colinearity with a full battery of country dummies because we estimate our models without a general regression intercept. In this set-up, the coefficients for the country dummies represent the (unique) intercept of the country in question. The results of *F*-tests and Wald tests confirm that country dummies belong in the specification of our regression models. See Greene, *Econometric Analysis*.

[34] In addition, we note that the analysis includes linearly interpolated data for a handful of missing observations in the wage inequality series. We did not interpolate across gaps of more than three years, and interpolated observations account for only nineteen out of 211 used in the analysis.

LINEAR REGRESSION RESULTS

Table 3 presents the results of the regression with 90–10 ratios as the dependent variable, and Table 4 presents the results of separate regressions with 90–50 and 50–10 ratios as the dependent variable. Looking at the latter table, we are interested not only in the sign, size and significance of individual coefficients, but also in the differences between coefficients for the same variable in the two regressions. Accordingly, Table 5 reports the results of tests for the equality of coefficients across the wage hierarchy. Joint significance tests of the differences in coefficients – for the entire set of variables and the political/institutional subset – strongly reject the hypothesis that the variables have similar effects across the wage hierarchy. Moreover, the differences in magnitude for all four political/institutional variables not only accord with our expectations, but are significant at better than the 5 per cent level for bargaining centralization, public sector employment and government partisanship. While the difference for union density is not significant by standard cut-points, it is in the right direction (with a greater effect in the lower half of the wage hierarchy), and a joint test of the two labour market institution variables is significant at better than the 5 per cent level.

Let us begin by considering the effects of the four control variables. The rate of unemployment (averaged over five years) turns out to be the only one of these variables that has any statistically significant effects on the overall distribution of wages. As expected, the effects of persistent unemployment are indeed strongly inegalitarian. Contrary to our expectations, however, there is no significant difference in the effects of unemployment across the wage hierarchy. The results reported in Table 4 suggest that a given increase of unemployment weakens the bargaining power of median-income workers relative to the bargaining power of workers in the ninetieth percentile just about as much as it weakens the bargaining power of workers in the tenth percentile relative to median-income workers.

In the 90–10 regression, the signs of the coefficients for low-wage trade and private service employment are contrary to our expectations. However, the coefficients for these variables, and for female labour-force participation as well, do not come close to statistical significance. Even when analysing the upper and lower halves of the wage distribution separately, we still do not observe any wage-distributive effects of trade with low-wage countries, nor can we discern differential effects at different ends of the wage spectrum. A more disaggregated analysis of the evolution of wages within and across economic sectors might well show that low-wage trade has distributive effects, but the results obtained here suggest that Wood exaggerates the significance of this factor in the rise of wage inequality over the last two decades.[35]

[35] Similar findings are reported by Vincent Mahler, David Jesuit and Douglas Roscoe, 'Exploring the Impact of Trade and Investment on Income Inequality', *Comparative Political Studies*, 32 (1999), 363–95. Leamer challenges the 'factor-content-of-trade' approach adopted by Wood and others (see

TABLE 3 *The Determinants of Wage Inequality (90–10 Ratio), 1973–95*

Variable	Coefficients and standard errors	*p*-values	Long-run effects
Lagged dependent variable	0.684 (0.048)	*< 0.001*	–
Unemployment	0.023 (0.007)	*0.001*	0.073
LDC trade	− 0.008 (0.007)	*0.217*	− 0.025
Female labour-force participation	0.025 (0.037)	*0.683*	0.079
Private sector services	− 0.026 (0.034)	*0.447*	− 0.082
Union density	− 0.028 (0.012)	*0.027*	− 0.089
Bargaining centralization	− 0.028 (0.008)	*< 0.001*	− 0.089
Public sector employment	− 0.094 (0.024)	*< 0.001*	− 0.297
Left government	− 0.019 (0.006)	*0.002*	− 0.060

Notes and sources: All entries are least squares dummy variable estimates with panel-corrected standard errors in parentheses. Approximate *p*-values are two-sided. See Appendix in website version of this article for data sources and description. $N = 211$.

In contrast, unpacking the wage hierarchy does clarify the effects of female labour-force participation and private sector services. The coefficients for both are closer to statistical significance in the regression that uses 50–10 ratios as the dependent variable. Contrary to our expectations for female labour-force participation but consistent with the 'demand for food and fun' variant of the private services argument, the signs of both coefficients are negative. In other words, higher levels of female labour-force participation and private sector employment are associated with less rather more wage inequality in the bottom half of the distribution. In the 90–50 regression, the coefficients for these variables have the opposite sign, but only female labour-force participation approaches significance. Moreover and most importantly, the differences in the coefficients across the wage hierarchy are significant at around the 5 per cent

(F'note continued)
Edward Leamer, 'Wage Inequality from International Competition and Technological Change', *American Economic Review*, 86 (1996), 309–14). Appealing to the Stolper–Samuelson theorem, Leamer argues that the liberalization of North–South trade alters relative output prices, and hence relative wages, regardless of what happens to the volume of trade. The results reported here do not tell us anything about the wage-distributive impact of trade as conceived by Leamer.

TABLE 4 *The Determinants of Wage Inequality in the Upper and Lower Halves of the Wage Distribution (90–50 and 50–10 Ratios), 1973–95*

Variable	90–50 ratio			50–10 ratio		
	Coefficients (s.e.)	*p*-values	Long-run effects	Coefficients (s.e.)	*p*-values	Long-run effects
Lagged dependent variable	0.629 (0.064)	*< 0.001*	–	0.539 (0.059)	*< 0.001*	–
Unemployment	0.007 (0.004)	*0.091*	0.019	0.009 (0.005)	*0.047*	0.020
LDC trade	0.002 (0.005)	*0.676*	0.005	− 0.004 (0.005)	*0.456*	− 0.009
Female labour-force participation	0.039 (0.025)	*0.114*	0.105	− 0.032 (0.026)	*0.224*	− 0.069
Private sector services	0.022 (0.022)	*0.991*	0.059	− 0.037 (0.020)	*0.073*	− 0.080
Union density	− 0.010 (0.008)	*0.273*	− 0.027	− 0.026 (0.009)	*0.005*	− 0.056
Bargaining centralization	− 0.010 (0.005)	*0.048*	− 0.027	− 0.027 (0.006)	*< 0.001*	− 0.059
Public sector employment	− 0.031 (0.015)	*0.033*	− 0.084	− 0.086 (0.015)	*< 0.001*	− 0.187
Left government	− 0.017 (0.005)	*0.001*	− 0.046	− 0.003 (0.004)	*0.518*	− 0.007

Notes and sources: All entries are least squares dummy variable estimates with panel-corrected standard errors in parentheses. Approximate *p*-values are two-sided. See Appendix in the website version of this article for data sources and description. $N = 211$.

TABLE 5 *Tests for Equality of Coefficients Across the Upper and Lower Halves of the Wage Distribution (90–50 and 50–10 Ratios)*

Variable	*p*-value
Joint test for equality of effects of all explanatory variables	< 0.001
Joint test for equality of effects of political/institutional variables	0.008
Unemployment	0.756
LDC trade	0.406
Female labour-force participation	0.049
Private sector services	0.053
Joint test for labour market institution variables (union density and centralization)	0.043
Union density	0.203
Bargaining centralization	0.024
Joint test for government variables (partisanship and public employment)	0.006
Public sector employment	0.008
Government partisanship	0.042

Note: p-values are for Wald tests for the equality of coefficients for 90–50 and 50–10 ratios.

level for both variables, indicating notable differential effects across the wage hierarchy.

The results reported in Tables 3 and 4 contradict the notion that the expansion of private services entails downward adjustment of the wages of unskilled workers, and suggest that the production of personal services with low skill content is not as tightly constrained by labour costs as commonly supposed. With respect to female labour-force participation, our findings call into question the proposition that men are, on average, better trained and/or more experienced than women and might be invoked to support the argument that the expansion of personal services, both public and private, entails a shift in the kinds of skills that employers value. In any case, it seems questionable to construe female labour-force participation itself as a *causal factor* compressing the lower half of the wage distribution. More likely, female labour-force participation is correlated with compression of gender differentials and, because women are over-represented in the lower half of the wage distribution, this is where compression of gender differentials manifests itself most strongly. The determinants of female labour-force participation, its relationship to public and private service employment and its impact on the politics of wage distribution constitute a topic for further research.

Turning to the variables of primary theoretical interest for our present purposes, our results indicate that unionization has a very significant egalitarian impact on the overall distribution of wages, and that this impact is primarily concentrated in the lower half of the distribution. The coefficient for union density in the 90–50 regression also has a negative sign, but it is much smaller (less than half) the size of the coefficient in the 50–10 regression, and does not satisfy conventional criteria of statistical significance. Consistent with our expectations, unions compress the wage distribution primarily by boosting the relative wages of poorly paid workers. The long-run egalitarian effect of a given increase in union density is more than twice as large in the lower half of the wage hierarchy.

The results for bargaining centralization and public employment are similar to the results for union density, and also support the hypotheses developed above. As Table 3 indicates, both variables have an egalitarian effect on the overall distribution of wages, and are statistically significant at better than the 1 per cent level. When we analyse the upper and lower halves of the wage distribution, the egalitarian effects of bargaining centralization and public employment appear in both regressions, but they are considerably larger in the regression with 50–10 ratios as the dependent variable. For both variables, the egalitarian effects are roughly twice as large at the bottom of the wage hierarchy, and the difference in the 90–50 and 50–10 coefficients is significant at better than the 5 per cent level.

Perhaps the most interesting results reported in Tables 3–4 concern the wage-distributive effects of government partisanship. As expected, the effect of left government on the overall distribution of wages is egalitarian. In contrast to the other three political-institutional variables, however, the egalitarian effects of left government are most pronounced in the *upper* half of the wage distribution. In the 50–10 regression, the coefficient for left government is still negative, but very small (about one-sixth of the coefficient for left government in the 90–50 regression) and far from statistical significance. These results are consistent with the hypothesis that left parties in government discourage wage growth in the upper half of the wage distribution by raising marginal income tax rates, but they offer little support for the hypothesis that left parties promote relative wage gains for poorly paid workers by setting a floor for competition in the labour market.

It is particularly intriguing to note that the differential effects of left government across the wage hierarchy are the opposite of those we observe for union density. In all the countries covered by our analysis, left parties have historically been allied with trade unions and rely on the votes of union members to a greater extent than any other parties. To the extent that unions dictate the policies of left parties in this arena, it would appear that unions pursue one distributive objective through collective bargaining and another distributive objective through politics.

LEFT GOVERNMENT AND WAGE DISTRIBUTION

The preceding discussion raises two questions. First, does our hypothesis about marginal tax rates correctly specify the mechanism whereby left government is associated with compression of 90–50 ratios? Secondly, why do we not

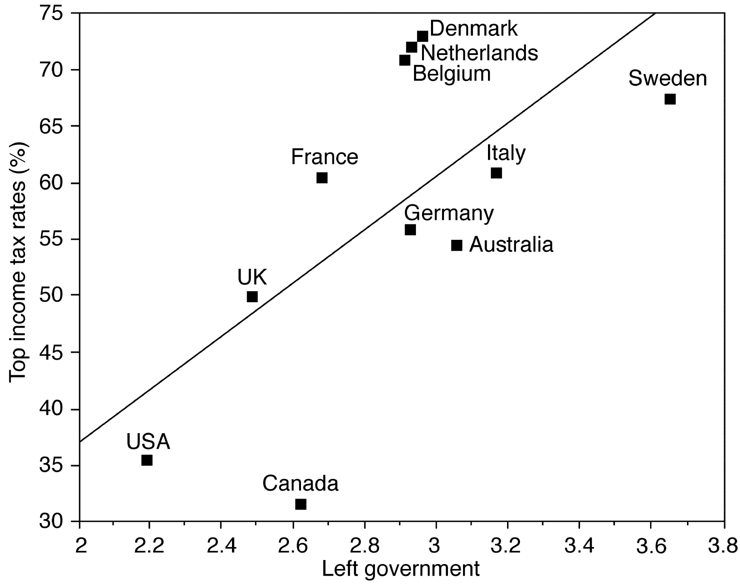

Fig. 1. The relationship between left government and top marginal income tax rates

observe the negative association between left government and 50–10 ratios predicted by the wage floor hypothesis?

Starting with the first question, we have data on top marginal income tax rates for early 1980s and late 1980s for eleven of the sixteen countries included in our analysis.[36] Figure 1 plots the average of these two observations against each country's average score on Cusack's (inverted) index of government partisanship for 1970–90, and Figure 2 plots each country's 90–50 ratio in 1991 against its average top income tax rate in the 1980s. Figure 1 reveals that left government is indeed associated with higher tax rates at the top of the income distribution, and Figure 2 indicates that high marginal tax rates are in turn associated with lower 90–50 ratios. Taken together, the two figures provide rather thought-provoking, though impressionistic, evidence in favour of the argument that left government compresses the upper half of the wage distribution via high marginal taxes.

In Figures 3 and 4, we repeat this exercise with data relevant to the wage floor hypothesis. First, we plot average income replacement rates provided by public schemes during the first year of unemployment in 1985–91 against each

[36] Provided by Geoffrey Garrett (Department of Political Science, Yale University), these data are based on Arnold Heidenheimer, Hugh Heclo and Carolyn Teich Adams, *Comparative Public Policy*, 3rd edn (New York: St Martin's Press, 1990), Table 6.7 and the OECD data series on the tax/benefit positions of production workers.

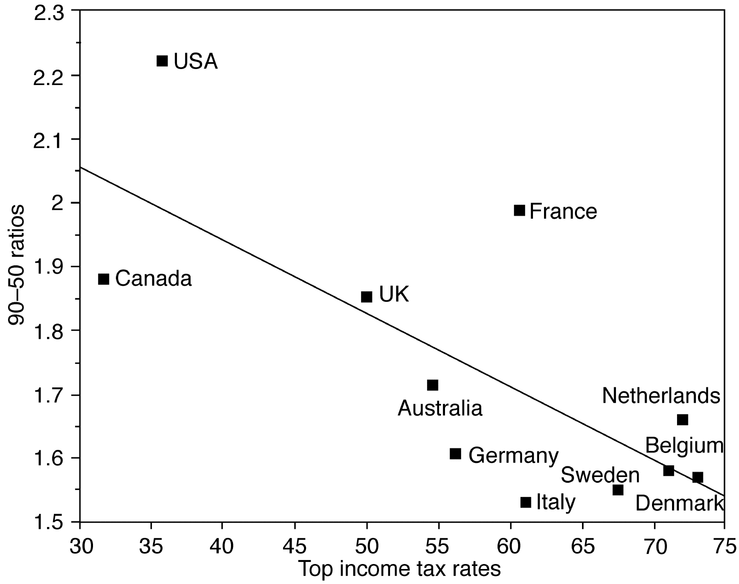

Fig. 2. The relationship between top income tax rates and 90–50 ratios

country's average partisanship score for 1970–90.[37] We then plot 50–10 ratios in 1991 against average unemployment replacement rates. Figure 3 suggests an association between left government and the generosity of unemployment compensation, but this association is considerably weaker than the association between left government and high marginal tax rates in Figure 1. The relationship between unemployment compensation and 50–10 compression is weaker still, partly but not only due to the role of the United States and Canada as outliers. Excluding the United States and Canada produces a better fit in Figure 4, but the regression line becomes nearly flat: at best, there is a discernible but very weak relationship between replacement ratios and 50–10 ratios. In other words, both steps in the argument linking left government to egalitarianism via a wage-floor effect seem to falter.

Comparing Figures 1 and 3, one might conclude that Social Democratic parties are distinguished from their competitors on the centre-right of the political spectrum, particularly Christian Democratic parties, by a greater commitment to income redistribution, and not so much by a greater commitment to high levels of basic income security. However, it must be noted that our data

[37] Based on OECD, 'Unemployment Benefit Entitlements and Replacement Rates' (electronic database), these unemployment replacement rates refer to someone who earned the equivalent of the average production worker at the time that he or she became unemployed and includes various income maintenance programmes (in addition to unemployment insurance in the narrow sense). See OECD, 'Unemployment and Related Benefits', *The OECD Jobs Study: Evidence and Explanations* (1994), part II, chap. 8, for more details.

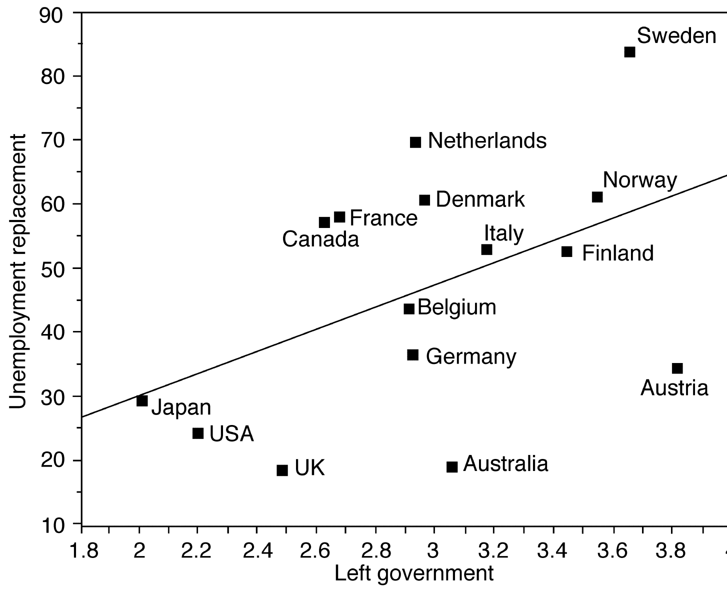

Fig. 3. The relationship between left government and unemployment benefit income replacement rates

on unemployment compensation do not take into account the percentage of unemployed workers who are covered by income replacement programmes in question, and that income replacement for the unemployed constitutes but one component of the 'reservation wage'. Conceivably, the wage-floor hypothesis would fare better with a broader measure of basic income support.

A somewhat different but complementary solution to the puzzle posed by the absence of any strong association between left government and 50–10 compression holds that the wage-distributive effects of government partisanship are contingent on labour market institutions. This line of argument is developed by Rueda and Pontusson, who show that left government is associated with overall wage compression (measured by 90–10 ratios) in liberal market economies, such as the United States and the United Kingdom, but not in the social market economies of Northern Europe.[38] In the latter countries, collectively negotiated wage agreements typically apply to all workers in a company or sector, whether or not they are union members, and wage developments in different companies and sectors are closely linked to each other. Arguably, these arrangements constrain the ability of governments to influence the distribution of wages. Put more positively, they enable unions to negotiate effective wage floors, and therefore reduce their reliance on minimum wage legislation and other forms of government intervention for this purpose.

[38] Rueda and Pontusson, 'Wage Inequality and Varieties of Capitalism'.

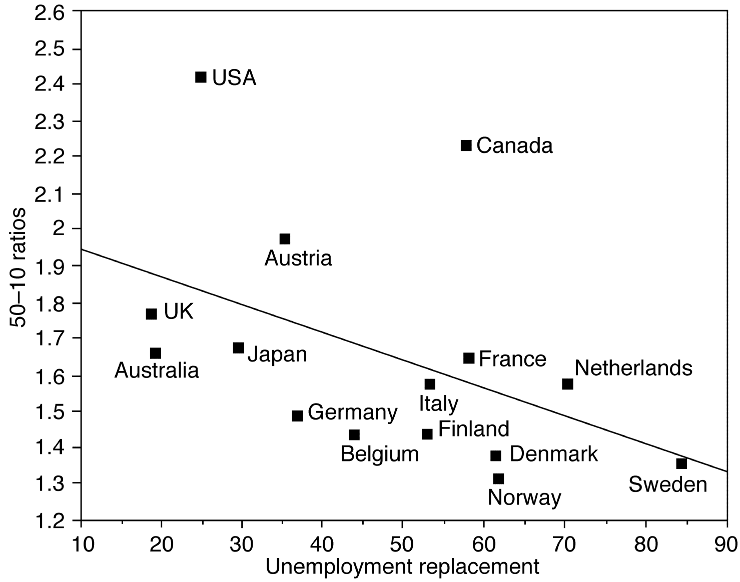

Fig. 4. The relationship between unemployment benefit income replacement rates and 50–10 ratios

Within the framework of the preceding analysis, we test this argument by rerunning the 90–50 and 50–10 regressions reported above with an interaction term for left government and bargaining centralization. While Rueda and Pontusson argue that social market conditions are conceptually distinct from bargaining centralization, it is nevertheless true that social market economies invariably score higher on Iversen's centralization index than other countries. To simplify the test, therefore, we use Iversen's index as a proxy for social market conditions. This argument implies that left governments would only produce egalitarian effects at the lower end of the wage hierarchy where bargaining is decentralized, whereas centralization would attenuate compression at the upper end of the hierarchy to a much lesser degree. In accord with these expectations, the interaction term of left government and bargaining centralization has a positive coefficient in both regressions – indicating the two are to some extent substitutes – but the coefficient is much larger (0.013 compared to 0.003) and more significant ($p = 0.06$ compared to $p = 0.14$) in the regression that uses the 50–10 ratio as the dependent variable.[39] However, as always with interaction models, the conditional coefficients and standard errors are of primary interest, since these allow us to assess the effects of left

[39] Note that we did not obtain significant interaction effects for left government with any of the other political-institutional variables in our analysis (union density, bargaining centralization and public-sector employment).

TABLE 6 *Egalitarian Effects of Left Government Conditional on Bargaining Centralization*

Level of bargaining centralization index	90–50 ratio	50–10 ratio
Decentralized (United States, France)	− 0.019 (0.011) [0.035]	− 0.022 (0.009) [0.011]
Moderately decentralized (United Kingdom, Italy)	− 0.017 (0.007) [0.005]	− 0.013 (0.005) [0.008]
Moderately centralized (Belgium, Germany)	− 0.016 (0.006) [0.001]	− 0.005 (0.005) [0.182]
Centralized (Sweden, Austria)	− 0.015 (0.008) [0.035]	0.004 (0.008) [0.318]

Notes: Panel-corrected standard errors in parentheses. Approximate *p*-value from one-sided *t*-tests in square brackets.

government conditional on bargaining centralisation. Accordingly, Table 6 summarizes the conditional coefficients of left government and their standard errors across the range of scores on the centralization index.

Table 6 conveys a very clear picture revealing statistically significant conditional relationships of the type hypothesized. The effects of government partisanship at the lower end of the wage distribution are strongly contingent on the degree of bargaining centralization and only statistically significant at moderate to low levels of centralization, but this is only marginally true for the effects of government partisanship at the upper end of the wage distribution, which decline very gradually with centralization but remain significant across the entire range of centralization scores. In countries where wage formation is highly decentralized, such as the United States and France, the egalitarian effects of left government are actually greater in the lower than the upper half of the distribution. As bargaining centralization increases, the wage-floor effects of left government disappear while the marginal-tax effects of left government remain.[40]

These results point to the salience of minimum wage legislation and other forms of government intervention for wage-distributive outcomes under decentralized labour-market conditions, a result that was obscured in both the

[40] In countries with highly centralized wage bargaining, such as Austria and Sweden, the 50–10 regression yields a positive conditional effect for left government, although the coefficient is statistically indistinguishable from zero.

overall 90–10 and separate 90–50 and 50–10 regressions. Furthermore, they suggest that wage growth at the upper of the distribution remains beyond the reach of solidaristic unions even under centralized bargaining. While attempting to influence their political allies' pursuit of redistributive tax policies, solidaristic unions operating under conditions of centralized bargaining are able to boost the relative wages of low-paid workers on their own and also recognize that the floor for competition in the labour market cannot be set too high.

CONCLUSION

There can be little doubt that market forces have tended to produce more wage inequality in advanced capitalist societies over the last two or three decades. In our regression models, persistent mass unemployment emerges as the most important factor generating inequality, but there are probably other forces at work here as well, not captured by our crude controls for supply and demand conditions.

While market forces have tended to generate more inequality, there is nonetheless no uniform or universal trend towards more overall wage inequality among full-time employees across the OECD. Strong unions, centralized wage bargaining, a large public sector, and left government have muted and sometimes overcome inegalitarian tendencies. In this constellation of countervailing forces, unions primarily promote the relative wages of poorly paid workers. Under conditions of decentralized wage bargaining, the policies associated with left government also boost relative wages at the bottom of the wage hierarchy, but the universal effect of left government is to hold back wage growth among highly paid workers. Bargaining centralization and public sector employment have egalitarian effects across the wage hierarchy, though they are strongest in the lower half.

Can the countries that have bucked the inegalitarian trend continue to do so in the future? Our analysis offers little comfort on this score, for political-institutional developments since the mid-1980s have not been favourable to wage equality. In most countries, union membership has declined under the pressures of globalization and/or as a result of the growing share of the labour force employed in private services.[41] In many countries, centralized wage bargaining has also fallen into disfavour with powerful employer groups and even some unions, leading to more decentralized forms of bargaining or, at least, more 'flexible' wage settlements at the national level, leaving more room for differentiation of wage increases at lower levels of bargaining. Finally, the near-universal expansion of public employment in the 1970s and 1980s came

[41] See Torben Iversen and Jonas Pontusson, 'Comparative Political Economy: A Northern European Perspective', in Torben Iversen, Jonas Pontusson and David Soskice, eds, *Unions, Employers and Central Banks* (New York: Cambridge University Press, 2000), pp. 1–37.

to a halt or was reversed in the 1990s.[42] While some public services have been privatized, governments have more generally sought to introduce market-oriented management principles into the public sector. More market-oriented pay-setting practices figure prominently in such efforts to reform public services. Not only has the public sector's share of total employment declined in many countries; quite likely, the egalitarian effects of public employment have also diminished.

From the perspective of the preceding analysis, the most obvious countervailing trend is the recent string of electoral victories for left parties in Western Europe. It is too early to tell if this development will last and if the new generation of left politicians retain the policy commitments necessary to foster equality in the labour market. Arguably, the minimum wage introduced by the Blair government will offset some of the inegalitarian effects of the decline of British unions over the last two decades. However, progressive income taxation does not appear to be high on the agenda of the Blair government or its continental counterparts.

Our analysis suggests that if recent political-institutional developments persist, the OECD-wide increase in inequality widely but erroneously perceived to have characterized the 1980s and early 1990s may become reality. In turn, this rising inequality can be expected to have important political consequences. In particular, there may be a symbiotic relationship between moderate levels of wage inequality and support for the universalistic welfare state. With widening income inequality, support for means-testing may well increase as the perception that many flat-rate benefits go increasingly to those who do not really need them grows. Thus, under wider income inequality, governments may be tempted to embark upon a programme of greater redistribution through more generous but means-tested benefits for the neediest matched by a continued and steady erosion of insurance-based benefits. In this light, the experience of the United Kingdom is instructive, as increasing inequality over the past twenty years has fuelled a steady drift away from universalism and insurance-based programmes and towards greater means-testing, not only under successive Conservative governments, but indeed under the current Labour government as well.[43] The stakes in rising wage inequality are potentially high for anyone committed to broad, inclusive welfare state programmes. Whether or not rising inequality undermines the foundations of welfare-state universalism and whether or not it promotes right-wing populism, the political economy of wage distribution is a topic that deserves greater attention from political scientists than it has received thus far.

[42] See Richard Clayton and Jonas Pontusson, 'Welfare-state Retrenchment Revisited: Entitlement Cuts, Public Sector Restructuring and Inegalitarian Trends in OECD Countries', *World Politics*, 51 (1999), 67–98.

[43] Cf., e.g., Nicholas Timmins, 'Death to Universalism', *Financial Times*, 25 January 1999.

APPENDIX: DEFINITIONS OF INDEPENDENT VARIABLES AND DATA SOURCES

Union density: Employed union members as a percentage of employed labour force ('net density') for all countries but Canada. The Canadian figures include unemployed and retired people who retain their membership in the numerator and the unemployed in the denominator ('gross density'). *Source*: Jelle Visser, 'Unionization Trends Revisited' (Centre for Research of European Societies and Industrial Relations, Amsterdam, 1996).

Centralisation: Wage-bargaining centralization as measured by Torben Iversen (Department of Government, Harvard University). Higher figures signify more centralization. Iversen classifies country-years according to the relative weight of three levels of bargaining (local, industry and national), and multiplies these weights by a measure of the concentration of union membership at each level. Thus there are two distinct sources of variation in Iversen's index: (1) index scores increase as the relative significance of higher levels of bargaining increases; and (2) scores increase as union membership becomes more concentrated at any of these bargaining levels (especially at those that are more significant). See Torben Iversen, *Contested Economic Institutions* (New York: Cambridge University Press, 1999), for a complete specification. These figures have been lagged, so that the value for a given year is the average of the actual value for that year and the previous four years.

Public sector employment: Government employees (not including employees of state-owned enterprises) as percentage of total employed labour force. *Source*: OECD, 'Historical Statistics' (electronic database).

Left government: Partisan cabinet composition as measured by Tom Cusack (Wissenschafts-zentrum, Berlin). Cusack groups parties into five families, multiplies each family's share of cabinet portfolios by its weight, and sums the products. In his weighting scheme, 1 = radical Left, 2 = moderate Left, 3 = Centre, 4 = moderate Right, and 5 = radical Right. See Cusack, 'Partisan Politics and Public Finance' for further details. Year-by-year figures up to 1990 were provided directly by Cusack; post-1990 figures were calculated based on cabinet data in *Europa Yearbook*, and Cusack's party classifications. The index has been inverted so that higher scores signify more leftist government.

Unemployment: Average rate of unemployment (unemployed as percentage of total labour force) for the year in question and the preceding four years. *Source:* OECD, 'Historical Statistics'.

LDC trade: Trade with less developed countries (LDCs) as percentage of GDP, not including trade with OPEC countries. For all countries but Belgium, figures up to 1990 were provided by Geoffrey Garrett (Department of Political Science, Yale University). Belgian and post-1990 figures were calculated on the basis of OECD, *Monthly Trade Statistics*.

Female labour-force participation: Female labour force as percentage of total labour force. *Source:* OECD, 'Historical Statistics'.

Private service employment: Service employment as percentage of total employment minus government employment as a percentage of total employment. *Source:* OECD, 'Historical Statistics'.