WAGE INEQUALITY AND VARIETIES OF CAPITALISM

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THE immediate goal of this article is to explore the determinants of wage inequality in advanced capitalist economies. Following a period in which the distribution of wages tended to become more compressed, most OECD countries have experienced some increase in wage inequality since 1980. However, the magnitude of change varies significantly across these countries: the rise of wage inequality began earlier and/or lasted longer in some countries than in others. As we shall see, the United States stands out as the OECD country that has experienced the most sustained rise of wage inequality, lasting at least a quarter of a century. With countries entering the 1980s at very different levels of wage inequality, the persistence of cross-national diversity remains a conspicuous feature of the data that we present below. In 1995 someone occupying the 90th percentile of the U.S. earnings distribution (the bottom of the top 10 percent) had an income that was 4.6 times larger than the income of someone in the 10th percentile (the top of the bottom 10 percent). At the opposite end of the spectrum, the 90–10 earnings ratio in Sweden was only 2.2 in 1995.

To date, most of the literature on the comparative political economy of labor markets has taken macroeconomic performance as the dependent variable and focused on the issue of wage restraint, or the trade-off between inflation and unemployment. In the corporatist tradition inspired by rational choice thinking, wage restraint is viewed as a public good, subject to familiar collective action problems, and divergent outcomes are typically explained in terms of institutional arrangements, which determine the ability of unions and/or employers to coordinate their wage-bargaining behavior. As we turn to explore wage-distributive outcomes, this line of thinking seems less compelling, for any number of (particular) wage distributions satisfy the conditions of Pareto

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optimality. If the politics of wage restraint is essentially about coordination, trust, and perhaps sanctions to avoid a suboptimal outcome, the politics of wage distribution is more accurately described in terms of a continuous process of negotiating temporary settlements among competing interests.

In the real world the politics of wage restraint and the politics of wage distribution are, of course, opposite sides of the same coin, but as our analytical focus shifts from the former to the latter, we might expect that variables which capture the power resources of organized interests, such as union density and government partisanship, take on greater explanatory significance relative to more formal institutional variables, such as the degree of centralization of wage bargaining. Equally important, this shift of focus invites us to think differently about institutional arrangements. If centralization of bargaining matters to wage-distributive outcomes, this is not because it provides for coordination but rather because it affects the distribution of power among actors in the bargaining process.[1]

Quantitatively inclined comparativists are turning increasingly to pooled cross-section time-series regression analysis because it allows them to increase the total number of observations and to test relatively complex causal models with aggregate data from a small number of countries. Nevertheless, for all the sophistication of this recent quantitative work, more qualitatively inclined scholars still cling to the essential historical-institutionalist objection to regression analysis—that it presupposes but does not prove that the relationship between independent variable X and dependent variable Y is the same across all units of observation. The radical version of this objection holds that each country should be conceived as a unique context, determining the relationship between X and Y. More commonly in comparative political economy, we are told that countries cluster into a few broad, historically constituted institutional configurations. The arguments advanced by Katzenstein, Hall, Esping-Andersen, and Soskice imply that changing the value of X will have certain effects in one set of countries and quite different effects in another.[2]

[1] Cf. Michael Wallerstein, "Wage-Setting Institutions and Pay Inequality in Advanced Industrial Societies," *American Journal of Political Science* 43 (July 1999).

[2] Peter Katzenstein, *Small States in World Markets* (Ithaca, N.Y.: Cornell University Press, 1985); Peter Hall, *Governing the Economy* (New York: Oxford University Press, 1986); Gøsta Esping-Andersen, *The Three Worlds of Welfare Capitalism* (Princeton: Princeton University Press, 1990); David Soskice, "Wage Determination," *Oxford Review of Economic Policy* 6 (Winter 1990); and idem, "Divergent Production Regimes," in Herbert Kitschelt et al., eds., *Continuity and Change in Contemporary Capitalism* (New York: Cambridge University Press, 1999).

Behind our immediate goal of exploring the determinants of wage inequality, the ulterior purpose of this article is to show that regression analysis can and should incorporate the insights of the approach associated with these scholars.[3] Exploring the common determinants of wage inequality across nations and over time we seek to determine whether and how these determinants vary across varieties of capitalism. Drawing on a new data set collected by the OECD, we engage in two rounds of pooled regression analysis. In the first round we regress levels of wage inequality on unemployment rates, trade with low-wage countries, female labor-force participation, union density, centralization of wage bargaining, the public sector's share of total employment, and government partisanship. Although the first three variables are meant to capture supply-and-demand conditions, none of them turns out to be a consistent predictor of the observed variance in wage inequality; the other four variables, however, all have statistically and substantively significant coefficients. Our second regression setup adds further institutional complexity by distinguishing between social market economies (SMEs) and liberal market economies (LMEs).

To anticipate, our empirical results indicate that varieties of capitalism matter. We find some support for the proposition that SME conditions mute the impact of market forces on the distribution of wages. More importantly, our results show that SME conditions significantly affect either the direction or the magnitude of most of our political and institutional variables. Of particular interest to political scientists is the finding that the wage-distributive effects of government partisanship are contingent on institutional context. While leftist governments are associated with less wage inequality in liberal market economies, this is not the case in social market economies. Union density emerges as the single most important factor influencing wage inequality across institutional contexts; its effects are consistently egalitarian and they are greater than those of any other independent variable within the country clusters. Our results thus support the contention that the politics of wage distribution involves conflicts between unions and employers as well as distributive conflicts between different firms and different categories of wage earners.

[3] In seeking to incorporate institutional insights into regression analysis, we follow the lead of Bruce Western, *Between Class and Market* (Princeton: Princeton University Press, 1997); Geoffrey Garrett, *Partisan Politics in the Global Economy* (New York: Cambridge University Press, 1998); and Torben Iversen, *Contested Economic Institutions* (New York: Cambridge University Press, 1999). However, none of these works explicitly tests the idea that causal dynamics vary by political economy type. Previous work in this particular vein includes Philip O'Connel, "National Variations in the Fortunes of Labor," in Thomas Janoski and Alexander Hicks, eds., *The Comparative Political Economy of the Welfare State* (New York: Cambridge University Press, 1994); and Duane Swank, "Political Institutions and Welfare State Restructuring," in Paul Pierson, ed., *The New Politics of the Welfare State* (New York: Oxford University Press, 2000).

We begin by describing the wage-distributive outcomes that we seek to explain. We then review the literature on the determinants of wage inequality and generate causal hypotheses pertaining to the discrete independent variables identified above. Third, we elaborate on the distinction between social and liberal market economies and specify how we expect this distinction to affect our causal hypotheses. The fourth section briefly addresses methodological issues, and the fifth section presents our empirical results. By way of conclusion, we address the general implications of our analysis.

I. Patterns of Wage Inequality

The dependent variable in our analysis is a summary measure of the distribution of gross income from employment. The particular measure we use, the ratio of earning at the 90th percentile to earnings at the 10th percentile, is dictated by the OECD data set on which we rely. The 90–10 ratio, a measure of the distance between two points, certainly does not tell us everything about the overall shape of the distribution, but it is a commonly used measure and easy to interpret.

The reader should keep in mind that our inequality measure ignores important sources of income, such as self-employment, income from capital, and government transfers. It also ignores the distributive effects of taxation and income pooling within households. Moreover, the OECD data set on which we rely is restricted to full-time employees (except in the case of Austria). What follows, then, must not be confused with an analysis of the overall distribution of income in OECD countries. This said, income from employment accounts for the lion's share of income in all OECD countries, and the distribution of income from employment, as measured by 90–10 ratios, correlates quite closely with broader cross-national measures of income distribution.[4]

Figures 1 and 2 provide a graphic summary of the wage-distributive outcomes that our analysis seeks to explain. The figures reveal both persistent variations in levels of wage inequality among the sixteen OECD countries and considerable change over time. Starting with the persistence of cross-national variations, the U.S. and Canada clearly constitute a group unto themselves, distinguished by very high levels of wage inequality throughout the time period covered by our data. At the other

[4] Cf. Wallerstein (fn. 1), 649; OECD, *Income Distribution in OECD Countries: Evidence from the Luxembourg Income Study* (OECD, 1995); and Peter Gottschalk and Timothy Smeeding, "Cross-National Comparisons of Earnings and Income Inequality," *Journal of Economic Literature* 35 (June 1997). For further specifications of the wage data used here, see OECD, "Earnings Inequality, Low-Paid Employment and Earnings Mobility," *Employment Outlook* (July 1996).

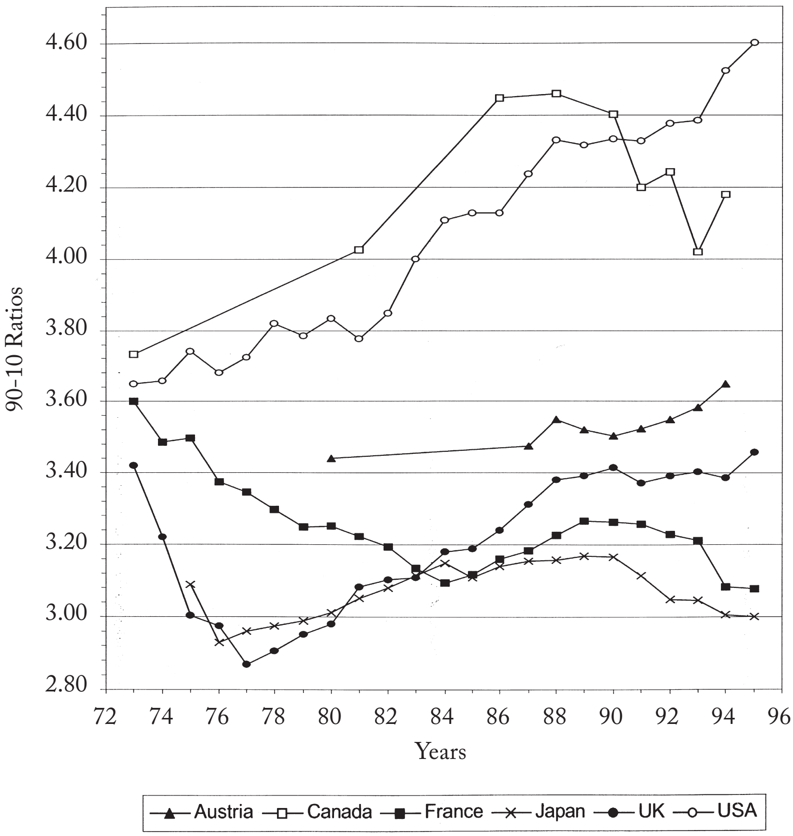

FIGURE 1

SOURCE: See appendix.

end of the spectrum the Scandinavian countries (Sweden, Norway, and Denmark) stand out as the OECD countries with the most egalitarian distribution of wages. The Scandinavian countries might be viewed as part of a broader low-inequality band that would also include Belgium, Finland, Italy, Switzerland, the Netherlands, Germany, and possibly Australia. Alternatively, the latter group of countries might be viewed as a separate, low-to-average band. In any case, Japan, France, the United Kingdom, and Austria make up another band, characterized by comparatively high levels of wage inequality, though not nearly as high as those of Canada and the U.S.[5] While relative rankings within these

[5] The high level of wage inequality in Austria is partly attributable to the fact that the underlying wage data include part-time employees, but other data sources also indicate that the Austrian distribution of wages is quite inegalitarian by continental European standards. See Bob Rowthorn, "Corporatism and Labour Market Performance," in Jukka Pekkarinen, Matti Pohjola, and Bob Rowthorn,

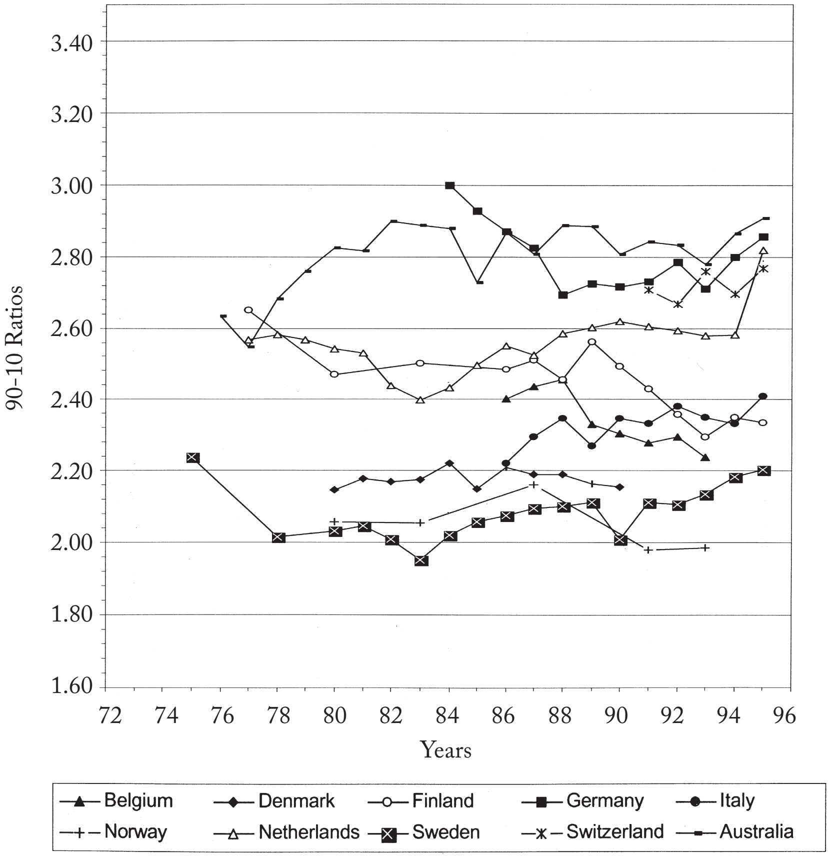

FIGURE 2

SOURCE: See appendix.

bands changed, the country composition of the bands remained remarkably stable over the 1973–95 period.

Looking at change over time, seven countries have experienced a continuous rise in wage inequality in recent years—most dramatically, the U.S. and the U.K, but also Austria, Germany, the Netherlands, Italy, and Sweden. In another four countries (Canada, France, Japan, and Australia), we observe notable increases of wage inequality in the 1970s and the 1980s, followed by subsequent decreases or, in the case of

eds., *Social Corporatism* (Oxford: Clarendon Press, 1992); and Jonas Pontusson, "Wage Distribution and Labor Market Institutions," in Torben Iversen, Jonas Pontusson, and David Soskice, eds., *Unions, Employers and Central Banks* (New York: Cambridge University Press, 2000). The use of country dummies and the small number of Austrian observations (7) renders our analysis essentially unaffected by the exceptional nature of the Austrian wage data.

TABLE 1
PERCENTAGE INCREASE IN 90–10 RATIOS FROM ALL-TIME TROUGH TO MOST RECENT OBSERVATION

Country	*% Change*	*Trough*	*Most Recent*
United States	26.0	1973	1995
United Kingdom	20.6	1977	1995
Netherlands	17.5	1983	1995
Australia	14.1	1977	1995
Sweden	12.8	1983	1995
Canada	12.1	1973	1994
Italy	8.6	1986	1995
Germany	6.3	1988	1995
Austria	6.1	1980	1994
Switzerland	3.7	1992	1995
Japan	2.4	1976	1995
Finland	1.7	1993	1995
Belgium	0.0	1993	1993
Denmark	0.0	1990	1990
France	0.0	1995	1995
Norway	0.0	1993	1993

SOURCE: See appendix.

Australia, a period of stability. Norway, Denmark, and Switzerland seem best characterized as cases of long-term stability while Finland and Belgium stand out as the two countries in which wage inequality declined continuously in the 1980s and 1990s. Measuring the percentage change in 90–10 ratios from their all-time trough to the most recent observation available, Table 1 brings out the widespread trend toward increased wage inequality, as well as important cross-national variations in its duration and magnitude.[6]

II. LITERATURE REVIEW AND HYPOTHESES

The labor economics approach to the problem of explaining wage-distributive outcomes focuses on relative demand for and supply of dif-

[6] The trend for wage inequality becomes more impressive if we look at the wage distribution among men and women separately. See OECD (fn. 4, 1996); and 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). In most countries increases of within-gender inequality have been offset by continued compression of between-gender wage differentials.

ferent types of labor. Looking at the demand side, labor economists typically argue that the growth of wage inequality in the U.S. and elsewhere reflects technological changes, which have rendered more-educated workers more valuable to employers than less-educated workers. With regard to the supply side, labor economists frequently point out that the compression of wage differentials prior to the 1970s coincided with rapid growth of university enrollments, that is, with an increase in the relative supply of educated labor. As the growth of university enrollments decelerated, the supply of better-educated labor subsequently failed to keep up with demand, giving rise to sharply increasing returns to education. At the same time, immigration and, more broadly relevant to all OECD countries, the massive increase in women's participation in the labor force since the 1970s represent an increase in the relative supply of unskilled labor to the extent that immigrants and women are on average less educated and less experienced than "native" men.[7]

In a similar vein, Wood argues that much of the trend toward increased wage inequality in the OECD countries in the 1980s can be attributed to increased manufacturing trade with less developed countries.[8] Though his empirical analysis is controversial, the logic of his argumentation is consistent with the supply-and-demand framework of the labor economics approach: as imports of less skill-intensive goods from low-wage countries increase, the effective supply of unskilled labor relative to skilled labor increases, putting downward pressure on the relative wages of unskilled workers.

Labor economists who engage in cross-national comparison typically find that supply-and-demand factors alone cannot explain observed variations in wage inequality across the OECD countries; they conclude that institutions matter.[9] When economists speak of "institutions," they have in mind not only codified rules or formalized organizational arrangements but also government policy and the distribution of power among organized interests. Commonly, the sectoral distribution of employment and other dimensions of industrial structure are also referred to as institutional variables. From the perspective of comparative political economy, the interesting question is not whether in-

[7] 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); and Lars Svensson, *Closing the Gender Gap* (Lund: Ekonomisk-Historiska Föreningen, 1995).

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

[9] Gottschalk and Smeeding (fn. 4); Rowthorn (fn. 5); Francine Blau and Lawrence Kahn, "International Difference in Male Wage Inequality," *Journal of Political Economy* 104 (August 1996); and various contributions to Richard Freeman and Lawrence Katz, eds., *Differences and Changes in Wage Structures* (Chicago: University of Chicago Press, 1995).

stitutions in this broad sense matter but rather which institutions matter and how (why) they matter. In this spirit, we seek to isolate the causal effects of different political-institutional variables and to explore interactions among those variables.

As a first cut, we derive four relevant political-institutional variables from the literature: unionization, centralization of wage bargaining, the size of the public sector, and the partisan composition of government. Ideally, we would want to control for all the relative supply-and-demand factors identified by labor economists as we estimate the effects of these variables. Measuring relative demand shifts associated with new technology is notoriously difficult, however, and we have not been able to find any such measure that would suit our purposes. Moreover, the existing OECD data on immigration flows are not comparable across countries and OECD data on educational qualifications exist only for the second half of the time period covered by our wage-inequality data.

Thus we are left with two control variables meant to capture cross-national variations and variations over time in the relative supplies of skilled and unskilled labor: non-OPEC LDC trade as a percentage of GDP and women's share of total employment. For the reasons set out above, we hypothesize that both variables are positively associated with wage inequality. With respect to female labor-force participation, some preliminary qualifications are in order, for we know the 1970s and early 1980s were characterized by a significant increase of female labor-force participation, as well as a significant decline of wage inequality in many OECD countries. Our baseline hypothesis implies that other factors, such as increased female participation in higher education and the expansion of public sector jobs, offset the inegalitarian effects of female labor-force participation in this period, but it is also plausible that the wage-distributive effects of female labor-force participation itself are contingent or contradictory. As women acquire skills through labor-force participation, higher levels of female labor-force participation should be associated with a smaller skill gap between men and women.

The regressions reported below also include the rate of unemployment to capture the wage-distributive effects of aggregate demand fluctuations (as distinct from relative demand shifts). It is a commonplace that unskilled, low-paid workers are more readily substitutable than skilled, high-paid workers and that their bargaining position is more immediately and more adversely affected by unemployment. By this logic, we expect the rate of unemployment to be positively associated with wage inequality.[10]

[10] Cf. James Galbraith, *Created Unequal* (New York: Free Press, 1998).

However, there is another side to the relationship between unemployment and wage inequality, for employers are more likely to lay off unskilled workers than to lay off skilled workers during economic downturns. To the extent that an increase in unemployment entails a disproportionate loss of low-paid jobs, it should be associated with less rather than more wage inequality (though it would still be associated with more overall inequality of income). For both female labor-force participation and unemployment, then, our theoretical expectations are somewhat ambiguous. Still, it is desirable to include these variables as controls. Let us now turn to the variables of primary theoretical interest.

Union Density

Following Freeman, we should distinguish two dimensions of the relationship between unionization and wage distribution: one concerns the distribution of wages among union members and how it compares with the distribution of earnings among unorganized wage earners; the other concerns wage differentials between union members and nonmembers, that is, the wage premium associated with union membership for wage earners with equivalent qualifications, experience, and other relevant characteristics.[11] With respect to the first dimension, there are several reasons to expect the wage distribution of the union sector of an economy to be more compressed than that of the nonunion sector. To begin with, 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 favor redistributive wage demands. Moreover, because unions are dependent on membership support in conflicts with management, they have a strong interest in curtailing wage setting based on the subjective decisions of foremen or personnel managers.

While employers competing in the same product markets can be expected to favor standardization of wage rates across firms, thereby taking wages out of competition, they also have an interest in maintaining unilateral control of intrafirm differentials.[12] Moreover, union interest in interfirm standardization is probably deeper and more general than that of employers. Even when market conditions differ among firms, so that price discrimination by the union is possible, the need to sustain mobilizational capacity inclines unions to restrict the scope of interfirm differentials.

[11] Richard Freeman, "Unionism and the Dispersion of Wages," *Industrial and Labor Relations Review* 34 (October 1980); idem, "Union Wage Practices and Wage Dispersion within Establishments," *Industrial and Labor Relations Review* 36 (October 1982); cf. also Blau and Kahn (fn. 9).

[12] Cf. Freeman (fn. 11, 1980).

To the extent that unionized workers earn more than equivalent nonunionized workers, the relationship between unionization and wage distribution becomes more complicated, for the wage-distributive effects of unionization now come to depend in part on the distribution of union membership across the wage hierarchy. Unionism would be a source of wage inequality if highly paid wage earners were better organized than low-paid workers, and the opposite would hold if low-paid wage earners were better organized. Though we lack detailed information on the distribution of union membership by income, the available evidence suggests that in most countries any wage premiums that accrue to union members are likely to compress the wage distribution.[13] Overall, then, we expect union density to be negatively associated with wage inequality.

Wage-Bargaining Centralization

A number of recent studies establish that countries with more centralized wage-bargaining systems consistently tend to have a more compressed distribution of wages than countries with less centralized wage-bargaining systems.[14] Centralization facilitates the reduction of interfirm and intersectoral wage differentials, since it means that more firms and sectors are included in a single wage settlement, but this argument presupposes that at least one of the parties of centralized bargaining wants to achieve a reduction of interfirm or intersectoral differentials. Wallerstein articulates two mechanisms whereby centralization produces egalitarian outcomes: there is a political mechanism through which centralization alters "the influence of different groups in the wage-setting process," and then there is an ideological mechanism whereby centralization affects norms of fairness.[15]

In the paradigmatic Swedish case, low-wage affiliates of the powerful confederation of blue-collar unions (Landsorganisationen) insisted on solidaristic measures as a condition for their participation in peak-level bargaining sought by employers in the 1950s.[16] But why should centralization systematically strengthen the relative bargaining power

[13] The Eurobarometer of June–July 1994 yields the following figures for average union density by income quartile in the EU member states: lowest quartile, 37.5 percent; second-to-lowest quartile, 37.8 percent; second-to-highest quartile, 34.2 percent; and highest quartile, 23.7 percent. For all countries but Canada the union-density data used in our regressions refer to net rather than gross union density ("employed union members as a percentage of the employed labor force" rather than "union members as a percentage of the total labor force").

[14] Wallerstein (fn. 1); Iversen (fn. 3), 2–5; Rowthorn (fn. 5); Blau and Kahn (fn. 9); and OECD, "Economic Performance and the Structure of Collective Bargaining," *Employment Outlook* (July 1997).

[15] Wallerstein (fn. 1), 674.

[16] Peter Swenson, *Fair Shares* (Ithaca, N.Y.: Cornell University Press, 1989), 56–58.

of low-wage unions? Wallerstein's political explanation essentially reproduces Freeman's argument about a single union that formulates wage demands on the basis of some form of majoritarian decision making: if low-wage and high-wage unions bargain jointly, organizational politics will influence the demands that they pursue, and market forces will be less influential in determining the distribution of wage increases. Consistent with Wallerstein's ideological explanation, we also hypothesize that centralized bargaining—in the extreme, a single settlement for all wage earners—renders wage differentials more transparent and thus politicizes wage-distributive outcomes. By this logic, centralization not only empowers low-wage unions but also makes them more likely to demand redistributive measures.

To test these hypotheses, which imply a negative association between centralization and wage inequality, we rely on an index of wage-bargaining centralization developed by Iversen.[17] This index takes into account the degree of union concentration at different levels of wage bargaining, as well as the relative importance of local, sectoral, and national bargaining.[18]

Government Employment

Kahn observes a negative association between the size of the public sector (relative to total employment) and wage inequality and explains this association by arguing that countries expand government employment in response to the employment-dampening effects of wage compression.[19] By contrast, we want to suggest that the size of the public sector may be construed as a cause of wage-distributive outcomes. Conceived in this fashion, the relationship between government employment and the overall distribution of wages is twofold, just like the relationship between unionization and wage distribution: first, there is the question of how the distribution of wages in the public sector compares with the distribution of wages in the private sector, and second there is the question of wage differentials between these sectors and how they affect the overall distribution of wages.

There are several reasons to expect wages in the public sector to be more compressed than wages in the private sector, aside from the fact

[17] For a detailed specification, see Iversen (fn. 3), 83–86.

[18] To capture the inertia associated with institutional change, these yearly figures were lagged so that the value for a given year used in our regressions is the average for that year and the previous four years. It should also be noted that we have extrapolated centralization values for the last two years of our time series.

[19] Lawrence Kahn, "Wage Inequality, Collective Bargaining and Relative Employment 1985–94" (Working paper, Institute for Labor Market Policies, Cornell University, 1999).

that the public sector is more heavily unionized than the private sector in most OECD countries.[20] On average, public sector unions appear to be more inclined than their private sector counterparts to favor wage solidarity. This generalization certainly holds for the 1970s and 1980s, when public sector employers also appear to be 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 in an egalitarian direction, being directly accountable to elected officials. By the same token, public sector unions have had less reason to worry about any potential trade-off between wage compression and employment growth than their private sector counterparts have had.[21]

As with union density, then, we expect the size of the public sector to be negatively associated with wage inequality because wage differentials are more compressed within the public sector. However, public sector wage premiums are much less likely to be egalitarian than are union wage premiums.[22] After all, government employees include a great many well-educated and highly paid civil servants and other professionals. Hence our expectations for the overall impact of the size of the public sector on wage inequality are somewhat uncertain. It seems likely that the effects of this variable are contingent on other institutional variables.

Government Partisanship

There are good reasons to expect leftist parties to pursue redistributive policies when they hold government power.[23] The effects of government partisanship will manifest themselves primarily in terms of redistribution via government taxation and spending, but government policies also affect the distribution of market incomes in general and of wages in particular. Indirectly, government policies affect the distribution of wages via their effects on unemployment, the size of the public sector, and union density. More directly, governments influence the distribution

[20] OECD, "Trends in Trade-Union Membership," *Employment Outlook* (July 1991), 113.

[21] Focusing on wage restraint rather than wage distribution, Geoffrey Garrett and Christopher Way present similar arguments about the distinctive dynamics of public sector bargaining; see Garrett and Way, "Public Sector Unions, Corporatism and Wage Determination," in Iversen, Pontusson, and Soskice (fn. 5).

[22] Drawing on microdata for the early 1990s from the Luxembourg Income Survey, Gornick and Jacobs report very sizable public sector wage premiums for men and women alike in Belgium, Canada, Germany, the Netherlands, the U.K., and the U.S., but not in Sweden. See Janet Gornick and Jerry Jacobs, *Gender, the Welfare State and Public Employment*, Luxembourg Income Study Working Paper no. 168 (1997).

[23] Cf., e.g., Garrett (fn. 3).

of wages through minimum-wage and equal-pay legislation, other forms of income policy, and a variety of measures that strengthen the competitive position of women and other disadvantaged groups (for example, immigrants) in the labor market. Partly for lack of good quantitative measures of such policies, we test the partisanship hypothesis by including Cusack's index of the cabinet center of gravity in our regressions.[24] As this index ranges between a score of 1 for a pure government of the radical left to a score of 5 for a pure government of the radical right, we expect the index values to be positively associated with wage inequality.

III. The Twist: Varieties of Capitalism

As indicated at the outset, this article explores whether the idea that the advanced capitalist countries form clusters with distinctive causal dynamics is applicable to the problem of explaining wage inequality. The typology of advanced capitalist countries used here hinges on the distinction between social market economies (SMEs) and liberal market economies (LMEs). The Northern European countries referred to as "social market economies" are essentially the same as the countries Katzenstein refers to as "corporatist," but the substantive connotations of the SME concept are different from those of the corporatism concept.[25] Katzenstein's well-known definition of corporatism emphasizes the formation of government policy through bargaining among centralized interest groups and the ideology of social partnership. Following Soskice, our conceptualization focuses more directly on the regulation of markets (government outputs rather than inputs).[26] However, our distinction between social and liberal market economies is narrower than Soskice's distinction between coordinated and liberal market economies and therefore yields a somewhat different grouping of countries.

Three basic features distinguish social market economies from liberal market economies. First, social market economies are characterized by comprehensive, publicly funded social welfare systems. Though the degree of redistribution varies, public welfare programs provide a relatively high "reservation wage" for the jobless and reduce workers' dependence on particular employers by providing for portability of employment benefits, retraining opportunities, sick pay insurance, and parental leave insurance. Using Esping-Andersen's terminology, we

[24] Tom Cusack, "Partisan Politics and Public Finance," *Public Choice* 91, no. 3–4 (1997).

[25] Katzenstein (fn. 2); and idem, *Policy and Politics in West Germany* (Philadelphia: Temple University Press, 1987).

[26] E.g., Soskice (fn. 2, 1990 and 1999).

might say that all these countries have reached a critical threshold of labor decommodification via the public provision of social welfare.[27]

Second, social market economies are characterized by government regulation to standardize employment conditions and to provide for a high degree of employment security. The details of such labor-market regulation vary from one country to another, but the effects are broadly similar: increased costs for employers to shed labor and greater standardization of employment conditions across sectors and categories of labor. While pay scales may be more or less compressed, overt pay discrimination based on gender, race, or legal status (in the case of immigrants) is less common in social market economies than in liberal market economies. The employment conditions of full-time and part-time employees are also more alike.

Third, social market economies are distinguished by a high degree of institutionalization of collective bargaining and coordination of wage formation. By coordination of wage formation, we mean that wage developments in different sectors of the economy are more tightly coupled than they are in liberal market economies. As suggested by the Japanese case, employers may be able to coordinate the wage-formation process by themselves under certain circumstances: what is distinctive about the social market economies is that coordination occurs through collective bargaining, which gives unions a central role in the process.

While the public provision of welfare and government regulation of employment conditions clearly introduce new considerations into our analysis, the institutionalization of collective bargaining may appear to be nothing but the sum of union density and wage-bargaining centralization. These variables are undoubtedly correlated on a cross-national basis, but they are conceptually distinct. Union density pertains to the balance of power between unions and employers, centralization pertains to the formal organization of the wage-bargaining process, and institutionalization describes the influence of collectively bargained wages on actual wages across the whole economy. Estimated by recent OECD studies, the percentage of the labor force covered by collective bargaining agreements provides a rough and ready measure of institutionalization. Controlling for union density, coverage rates tell us about the extent to which nonunion workers are affected by the employment terms achieved by unions through collective bargaining.[28]

[27] Esping-Andersen (fn. 2).

[28] Collective bargaining coverage rates exceed union density either because governments decide to extend negotiated agreements to firms or sectors that were not party to the agreements (the French case) or because employers are better organized than unions and collective agreements encompass all employees of the firms that are party to the agreement (the German case).

TABLE 2
FEATURES DISTINGUISHING SOCIAL AND LIBERAL MARKET ECONOMIES
(INDEX SCORES OR PERCENTAGES WITH OVERALL RANKS IN PARENTHESES)

	Welfare State		*Collective-Bargaining*	*Employment Protection*
Country	*Decom 1980*	*Spend/GDP 1990*	*Coverage 1990*	*1989*
SMEs				
Austria	31.1 (6)	23.9 (8)	98 (1)	9.0 (7)
Belgium	32.4 (5)	26.6 (5)	90 (4)	10.5 (3)
Denmark	38.1 (3)	28.3 (3)	69 (11)	3.3 (11)
Finland	29.2 (8)	25.3 (7)	95 (2)	10.5 (3)
Germany	27.7 (9)	23.8 (9)	90 (4)	12.0 (2)
Netherlands	32.4 (4)	28.8 (2)	71 (10)	7.3 (9)
Norway	38.3 (2)	26.9 (4)	75 (9)	9.8 (5)
Sweden	39.1 (1)	32.6 (1)	83 (6)	8.5 (8)
SME average	33.5	27.0	83.9	8.9
LMEs				
Australia	13.0 (16)	13.1 (15)	80 (8)	3.3 (11)
Canada	22.0 (14)	18.0 (12)	38 (14)	1.7 (15)
Japan	27.1 (11)	12.4 (16)	23 (15)	3.7 (10)
Switzerland	29.8 (7)	17.4 (13)	50 (12)	1.8 (14)
UK	23.4 (13)	19.8 (11)	47 (13)	2.3 (13)
USA	13.8 (15)	14.2 (14)	18 (16)	0.4 (16)
LME average	21.5	15.8	42.7	2.2
MIXED				
France	27.5 (11)	26.0 (6)	92 (3)	9.5 (6)
Italy	24.1 (12)	23.0 (10)	82 (7)	14.3 (1)

SOURCES: Decommodification index: Esping-Andersen (fn. 2), 52; total social spending as a percentage of GDP: OECD, *New Directions in Social Policy in OECD Countries* (Paris: OECD, 1994); and idem, *Social Expenditure Statistics in OECD Member Countries*, Labour Market and Social Policy Occasional Papers, no. 17 (1996); collective bargaining coverage rates: OECD (fn. 14), 71; employment protection index: OECD, *The OECD Jobs Study: Evidence and Explanations*, pt. 2 (Paris: OECD, 1994), 74.

Table 2 provides quantitative indicators of the three features that distinguish SMEs from LMEs for our sixteen countries. These figures are meant to be illustrative: in our conceptualization the distinction between social and liberal market economies is a categorical one and cannot be derived simply from the sum of these measures. To be categorized as a social market economy, moreover, a country must partake of all three features of "SME-ness."

Drawing on the typologies of other scholars, as well as on the data presented in Table 2, the majority of our countries are easily categorized as either SMEs or LMEs. France and Italy are the exceptions. Thus, although Italy's welfare state is smaller and less decommodifying than the countries we categorize as SMEs, it has the strictest employment protection legislation of all the countries included in our analysis, and its level of collective-bargaining coverage exceeds that of several SMEs. Ranking sixth in terms of social spending as a percentage of GDP, France would seem to have an even better claim to be coded as an SME. In view of the disorganized and fragmented character of French industrial relations, however, we do not believe that collective-bargaining coverage represents a meaningful measure of institutionalization in the French case.[29] While Italy lacks the welfare-state characteristics associated with SME-ness, France falls short when it comes to the institutionalization of collective bargaining. Nevertheless, it would surely be misleading to categorize these countries as liberal market economies. Following Hall, therefore, we resolve this problem by creating a separate category—"mixed economies"—for France and Italy.[30]

The typology presented in Table 2 diverges most clearly from conventional wisdom in its assignment of Japan and Switzerland to the category of liberal market economies. With respect to Japan, the reader must keep in mind that the criteria underlying our typology pertain specifically to government regulation of labor markets and/or joint regulation by employers and unions. Unquestionably, the Japanese case fails to conform to the model of a liberal market economy in matters of finance, corporate governance, and industrial policy. As for Switzerland, its score on Esping-Andersen's decommodification index falls well within the range of the social market economies, but on our other three indicators it clearly falls within the range of the liberal market economies.[31]

Table 3 summarizes how we expect the distinction between social and liberal market economies to affect the causal relationships hypothesized in the previous section. In general, we expect SME conditions to mute or neutralize the inegalitarian impact of market forces, measured here by unemployment, LDC trade, and female labor-force participation. The argument about unemployment is perhaps the most obvious one: if laid-off workers receive generous unemployment compensation and other welfare benefits, an increase in unemployment is less likely to

[29] The coverage rates estimated by the OECD do not necessarily pertain to wage agreements. A government decision to extend a collective agreement on summer holidays could account for France's exceptionally high coverage rate.

[30] Peter Hall, "Organized Market Economies and Unemployment in Europe" (Manuscript, Center for European Studies, Harvard University, 1998).

[31] As we explain in the conclusion, the question of how to code Japan and Switzerland turns out to be of no consequence for our main findings.

TABLE 3
SUMMARY OF EXPECTED ASSOCIATIONS WITH WAGE INEQUALITY, TAKING SME/LME INTERACTIONS INTO ACCOUNT

Variable	*SMEs*	*LMEs*
Unemployment	– / 0	0 / +
LDC trade	0 / +	+
Female labor-force participation	0	+
Union density	–	–
Bargaining centralization	–	– / 0
Government employment	–	0
Government partisanship		
hypothesis 1	0	+
hypothesis 2	–	0

exert downward pressure on wages at the lower end of the wage hierarchy. With respect to LDC trade, we proceed from the proposition that LMEs are more segmented than SMEs. In a segmented economy, an increase of LDC imports will put downward pressure on the wages of the workers who produce the goods that are exposed to new competition from low-wage producers. In a more closely integrated or coordinated economy, we expect low-wage competition to translate into cuts in real wages for all workers, rather than into increased wage inequality. Finally, SME conditions are likely to alter the wage-distributive consequences of female labor-force participation, because women tend to be less skilled than men and often end up in precarious forms of employment for other reasons as well (most obviously child-rearing responsibilities). Outside the public sector, feminization of the workforce typically correlates with low levels of unionization. As noted above, the institutional arrangements that distinguish SMEs from non-SMEs serve to standardize employment conditions and provide for the extension of union-negotiated wage contracts to nonunion members. This should mitigate the inegalitarian effects of an increase of female labor-force participation.

Turning to the political-institutional variables that constitute our primary interest, the institutional arrangements characteristic of SMEs enhance the significance of collective bargaining relative to market forces and, as a result, the characteristics of collective bargaining should matter more to wage-distributive outcomes. In SMEs, where reservation wages are high and bargaining coverage is extensive, the equalizing effect of bargaining centralization should be especially large.

The implications of SME conditions for the wage-distributive impact of unionization are less straightforward. On the one hand, one might suppose that those aspects of social market economies that enhance employment security would serve to desensitize unions and their members to the potential employment costs of wage compression and therefore render them more prone to pursue (support) egalitarian wage demands. By this logic, a one-unit increase in unionization should be associated with a larger reduction of wage inequality in SMEs than in LMEs. On the other hand, the extension of union-negotiated wage contracts to nonunion workers would seem to imply that levels of wage inequality in SMEs are less sensitive to changes in union density than they are in LMEs. As indicated in Table 3, our prior assumption is that these countervailing effects cancel each other out and that the wage-distributive effects of union density are not affected by the SME-LME distinction.

With respect to government employment, we expect that its egalitarian impact will be greatest under SME conditions. In both types of economies, private sector employers who compete with public sector employers for labor at the lower end of the labor market must follow suit if public sector employers raise the relative wages of the lowest paid employees. By linking wage developments in different sectors more closely, the institutional arrangements characteristic of SMEs should expand the spillover of public sector wage compression into the private sectors.

Finally, there are two competing ways to think about the interaction between social market conditions and government partisanship. On the one hand, SME arrangements might be conceived as constraints that limit the effects of partisanship (or ideology) on government policy. On the other hand, governments would seem to be significantly more involved in wage formation in SMEs than in non-SMEs, and therefore the preferences of governing parties may have more important consequences for wage formation. According to the latter perspective, the association between leftist government and wage compression should be stronger in SMEs than in non-SMEs; according to the former perspective, it should be weaker.

IV. Methodology

In pooled cross-section time-series analysis, "country-years" are the units of observation of dependent and independent variables. By incorporating over-time variations, pooling dramatically increases the total number of observations and enables us to test more complex causal models against data from a relatively small number of countries. At the same time this methodology is inextricably linked to the idea that cross-national variations and changes over time have common determinants; more precisely, it is linked to the goal of ascertaining the common determinants of cross-national variations and changes over time.

We address the autocorrelation problem associated with time-series data by including the lagged dependent variable on the right-hand side of the equation. Following Beck and Katz, we include a lagged dependent variable because it "makes it easier . . . to examine dynamics and allows for natural generalizations in a manner that the serially correlated errors approach does not."[32] As the distribution of wages changes only marginally from one year to the next, the coefficient for levels of wage inequality in the previous year turns out to be highly significant. On the assumption that the effects of a one-unit change in a particular variable persist, the long-term effects of such a change can be computed by dividing the value of the coefficient for the variable of interest by 1 minus the coefficient for the lagged dependent variable.[33] In what follows, we estimate both long-term and immediate effects for each variable.

Our regression models also include dummy variables for each of the countries in our data set. We do this to eliminate the bias resulting from the effects of country-specific omitted variables.[34] Put somewhat crudely, the country dummies control for the values that all observations for a given country share. While we are not interested in why particular countries fit our regression more or less closely and therefore do not report the coefficient estimates for the country dummies below, we believe that it is essential to control for country-specific effects in this manner. Scholars engaged in cross-national comparison sometimes eschew the use of country dummies on the grounds that they simply tell us that countries are different, when the interesting question is how or why they are different. Yet there is every reason to suspect that outcomes such as ours are influenced by country-specific historical or cultural factors, which cannot be measured on a cross-national basis (for instance, the influence of milltown culture on the priority assigned to wage solidarity by the Swedish labor movement). The results of F-tests

[32] Nathaniel Beck and Jonathan Katz, "Nuisance vs. Substance," in John Freeman, ed., *Political Analysis* (Ann Arbor: University of Michigan Press, 1996), 6:1. The results of Breusch-Godfrey tests indicate that there is no significant autocorrelation in our regressions. In tests with a variety of lags, we could not reject the null hypothesis (nonexistence of autocorrelation) at a level even close to the 90 percent traditional significance threshold. See William Greene, *Econometric Analysis* (Englewood Cliffs, N.J.: Prentice Hall, 1997), 595.

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

[34] See Cheng Hsiao, *Analysis of Panel Data* (New York: Cambridge University Press, 1986).

strongly confirm that country dummies belong in the specification of our regression models.[35]

Nickell demonstrates that, with short panel data, OLS estimation of models with lagged dependent variables and fixed effects produces biased coefficients.[36] We address this problem by computing the instrumental variable (IV) estimator suggested by Anderson and Hsiao and using a two-stage IV procedure.[37]

In what follows, we present first the results of a linear regression model and then the results of a model in which our independent variables are interacted with dummy variables for social market economies (SME), liberal market economies (LME), and mixed economies (MIX). We use the independent variables, a one-year lag of the independent variables, and the country dummies as the instruments and treat the lagged dependent variable as endogenous.[38] For the linear model we estimate the following equation:

$$y_{it} = \chi \hat{y}_{i,t-1} + \sum_{k} \beta_k x_{kit} + \alpha_i + \varepsilon_{it} \qquad (1)$$

Where y_{it} is the dependent variable, the x's are the independent variables, i refers to the cross-sectional units, t to the time units, k to the number of independent variables, α_i refers to the separate intercepts for each country (there is no common intercept), β_k refers to the slopes of the explanatory variables, χ to the slope of the lagged dependent variable predicted in the first stage of the IV procedure, and ε_{it} is a random error term normally distributed around a mean of 0 with a variance of σ^2.

The setup of our interaction model is the same except that we introduce three additional dummy variables (SME, LME, and MIX), which we interact with the x variables. Again we use the independent variables, a one-year lag of the independent variables and the country dummies as the instruments and treat the lagged dependent variable as endogenous. We estimate the following equation:

[35] For the regressions reported in Tables 2 and 3 the results of F-tests show that the country dummies are significant at better than the 99 percent level.

[36] Stephen Nickell, "Biases in Dynamic Models with Fixed Effects," *Econometrica* 49 (November 1981).

[37] See T. W. Anderson and Cheng Hsiao, "Estimation of Dynamic Models with Error Components," *Journal of the American Statistical Association* 76 (September 1981); and idem, "Formulation and Estimation of Dynamic Models Using Panel Data," *Journal of Econometrics* 18 (January 1982). This is similar to the method suggested by Hibbs and employed by Álvarez, Garrett, and Lange. See Douglas Hibbs, "Problems of Statistical Estimation and Causal Inference in Dynamic Time Series Models," in Herbert Costner, ed., *Sociological Methodology 1973/1974* (San Francisco: Jossey-Bass, 1974); and Michael Alvarez, Geoffrey Garrett, and Peter Lange, "Government Partisanship, Labor Organization and Macroeconomic Performance," *American Political Science Review* 85 (June 1991).

[38] In both the linear and the interaction models, the instruments we use turn out to be excellent predictors of the lagged dependent variable. The R^2 obtained in the first-stage IV regressions was higher than .95.

$$y_{it} = \gamma\hat{y}_{i,t-1} + \sum_k \lambda_k * x_{kit} * SME + \sum_k \omega_k * x_{kit} * LME + \sum_k \delta_k * x_{kit} * MIX + \tau_i + \eta_{it} \tag{2}$$

where τ_i refers to the separate intercepts for each country (there is no common intercept), λ_k, ω_k, and δ_k refer to the slopes of the explanatory variables, γ to the slope of the lagged dependent variable predicted in the first stage of the IV procedure, and η_{it} is a random error term normally distributed around a mean of 0 with a variance of σ^2.[39]

For each x variable, this equation yields separate coefficient estimates for the countries we have coded as social market economies, liberal market economies, and mixed cases. The same coefficient estimates would be obtained if the linear model were run separately for each set of countries, but the results of our interaction model are more directly comparable to the results of our linear model, since the total number of observations is the same in the two models.

Two other features of our analysis should be noted before we turn to the empirical results. First, the analysis includes interpolated observations of wage inequality for some of the years with missing observations in Figures 1 and 2. Observations for missing years were interpolated on the assumption that any change between two years was evenly distributed across the intervening years, but we decided not to interpolate data across more than three missing years. The analysis therefore uses only 20 interpolated observations out of a total of 217.[40] Second, we engaged in logarithmic transformations of all variables other than the dummy variables before running the regressions. When variables on both sides of the regression equation are logged, the regression coefficients can be interpreted as measuring the percentage change in the dependent variable associated with a percentage change in the independent variable. The slope values yielded by our regressions thus represent elasticity measures of the relationship between the variables.

[39] Nathaniel Beck and Jonathan Katz propose a method for deriving consistent standard error estimates in the presence of panel-heteroscedastic errors that has been widely adopted by students of comparative political economy, e.g., Garrett (fn. 3); and Iversen (fn. 3); see Beck and Katz, "What to Do (and Not to Do) with Time-Series Cross-Section Data," *American Political Science Review* 89 (September 1995). When we ran our regressions with the Beck-Katz procedure (estimating panel-corrected standard errors without instrumental variables), we obtained results that were essentially the same as those reported below (available upon request). None of our substantive findings are affected by the choice of one or the other of these setups.

[40] In terms of overall numbers, the interpolations make up for the loss of observations entailed by the use of a lagged dependent variable. More importantly, interpolation enables us to include Norway in our analysis.

V. Empirical Results

Table 4 reports the results of the linear regression model specified above and Table 5 reports the results of our interaction model. Table 6 pertains to the results of the interaction model, specifically, whether or not the observed differences between SMEs and LMEs are statistically significant.

Let us consider first the results of the linear model. While all four of the political-institutional variables in our analysis turn out to have a statistically significant association with levels of wage inequality, this is true for only one of the three variables intended to capture supply-and-demand conditions. The countervailing effects of unemployment, putting downward pressure on the wages of the unskilled workers but also removing a disproportionate number of unskilled workers from the employed labor force, apparently cancel each other out. As for LDC trade, the empirical results simply do not bear out our expectation that this variable would be positively associated with wage inequality. By contrast, female labor-force participation does have a sizable, positive coefficient, which readily clears conventional thresholds of statistical significance. This finding supports the hypothesis that increased female labor-force participation typically entails an increase in the relative supply of unskilled labor. As expected, union density, bargaining centralization, and government employment are associated with less wage inequality in the model, whereas rightist government is associated with more wage inequality. The size of the long-term coefficients for all these variables is considerable.

Setting aside the question of substantive significance for the time being, the results of our interaction model indicate that the determinants of wage inequality do indeed differ across clusters of advanced capitalist economies. Some of the generalizations supported by Table 4 appear quite misleading in view of the results reported in Table 5. To begin with, none of the variables in our analysis has statistically significant effects on the distribution of wages in the two countries we code as mixed cases (France and Italy). Thus we can concentrate on the theoretically interesting question of the differences between SMEs and LMEs.

Comparing SMEs with LMEs, we find that unemployment and LDC trade are equally irrelevant to wage-distributive outcomes while union density has a strong egalitarian impact in both sets of countries. More clearly than in the linear model, union density here emerges as the single most important variable explaining the observed variance of wage inequality in our data set. For the remaining four variables, we observe statistically significant differences between SMEs and LMEs (compare Table 6).

TABLE 4
THE DETERMINANTS OF WAGE INEQUALITY IN 16 COUNTRIES
(1973–95)

Variable	*Coefficients and Standard Errors*	*P-values*	*Long-Run Effects*
Lagged dependent variable	.499 (.153)	<.001	—
Unemployment	–.002 (.005)	.386	–.004
LDC trade	.0004 (.009)	.484	.0008
Female labor-force participation	.094 (.041)	.012	.188
Union density	–.037 (.018)	.020	–.074
Bargaining centralization	–.048 (.019)	.006	–.096
Government employment	–.099 (.042)	.009	–.198
Government partisanship	.022 (.007)	.001	.044

All entries are two-stage least squares estimates. P-values are one-sided. N = 217.

The coefficient for bargaining centralization is negative in both LMEs and SMEs, but the egalitarian impact of centralization is more than three times as great in SMEs according to these results. Indeed, the centralization coefficient for LMEs fails to clear the 90 percent level of statistical significance. On this count, our results diverge sharply from Wallerstein's. Whereas he writes that "it is difficult to find other variables that matter once the institutional variation in wage-setting is controlled for,"[41] we find not only that other political-institutional variables matter but also that the egalitarian effects of centralization are largely, if not entirely, contingent on SME conditions.[42]

The differences between SMEs and LMEs are more striking still with respect to the wage-distributive effects of female labor-force participation, government employment, and government partisanship, for the signs of the coefficients of these variables actually differ in the two sets of countries. According to our interaction model, the inegalitarian effects of female labor-force participation occur only under LME condi-

[41] Wallerstein (fn. 1), 650.

[42] Our analysis is based on a much larger number of observations and covers a longer time period than Wallerstein's. Different measures of bargaining centralization constitute another potential source of divergent results.

Table 5
The Determinants of Wage Inequality in Social Market, Liberal Market, and Mixed Economies (1973–95)

	Social Market Economies			*Liberal Market Economies*			*Mixed Economies*		
Variables	*Coeffic. and Standard Errors*	*P-values*	*Long-Run Effects*	*Coeffic. and Standard Errors*	*P-values*	*Long-Run Effects*	*Coeffic. and Standard Errors*	*P-values*	*Long-Run Effects*
Lagged dependent variable	.495 (.144)	*<.001*							
Unemployment	−.0004 (.006)	*.475*	−.0008	−.002 (.014)	*.433*	−.004	−.006 (.084)	*.472*	−.012
LDC trade	−.003 (.010)	*.368*	−.006	.013 (.015)	*.192*	.026	−.005 (.033)	*.442*	−.010
Female labor-force participation	−.059 (.071)	*.202*	−.117	.229 (.106)	*.016*	.453	−.374 (.641)	*.280*	−.741
Union density	−.114 (.048)	*.009*	−.226	−.119 (.046)	*.006*	−.236	−.060 (.105)	*.284*	−.119
Bargaining centralization	−.085 (.023)	*<.001*	−.168	−.024 (.021)	*.126*	−.048	−.024 (.051)	*.318*	−.048
Government employment	−.082 (.046)	*.040*	−.162	.104 (.052)	*.024*	.206	.006 (.508)	*.495*	.012
Government partisanship	−.005 (.011)	*.321*	−.010	.033 (.011)	*.002*	.065	.012 (.023)	*.305*	.024

All entries are two-stage least squares estimates. P-values are one-sided. N = 217.

TABLE 6
TESTS FOR EQUALITY OF COEFFICIENTS IN SMEs AND LMEs

Variable	*P-Value*
Unemployment	.898
LDC trade	.360
Female labor-force participation	.042
Union density	.934
Bargaining centralization	.005
Government employment	.018
Government partisanship	.015

P-values are for F-tests for the equality of the coefficients in SMEs and LMEs.

tions. The argument that social market conditions cushion the inegalitarian impact of the entry of new categories of unskilled labor into the workforce by extending union-negotiated wage contracts to nonunion workers provides a reasonable explanation of the absence of inegalitarian effects in the SMEs. However, this argument does not explain why female labor-force participation appears to be associated with less wage inequality in the social market economies of Western Europe, though the association is not quite significant by conventional criteria.

At least in part, the negative coefficient for female labor-force participation under social market conditions reflects the fact that the Scandinavian SMEs are characterized by higher rates of female participation as well as more compressed wages than are their continental counterparts. Arguably, the negative association between these variables is spurious, with female labor-force participation serving as a proxy for a complex of government policies that strengthens the position of women in the labor market and thereby reduces wage differentials between men and women. In other words, we might be picking up the effects of Scandinavian welfare states being more women friendly than the welfare states of continental Europe. But the negative association between female labor-force participation and wage inequality in SMEs might also be related to differences in the timing of increases in female labor-force participation. As noted earlier, the effects of female labor-force participation are likely to become egalitarian over time, as women acquire experience and skills at their jobs and as their position in the labor market becomes less precarious. As a group, the SMEs are more bifurcated on this score than the LMEs. Thus, the Scandinavian countries experienced a sharp increase of female labor-free participation as early as the 1960s, while female participation in the continental SMEs remained below the OECD average in the 1990s.

The results for government employment are similar to the results for female labor-force participation. In this case, the egalitarian effects observed in the linear regression model occur only among the SMEs. For the LMEs, the coefficient for government employment is positive—and statistically significant. The distinctively egalitarian dynamics of public sector wage bargaining appear to be contingent on SME conditions. Along the lines suggested above, our results are also consistent with the propositions that public sector wage premiums typically have inegalitarian effects and that the coordinated wage bargaining characteristic of social market economies curtail the scope of public sector wage premiums.

Finally, the results of our interaction model strongly support the hypothesis that SME conditions constrain the effects of government partisanship. In liberal market economies rightist government is associated with more wage inequality at a high level of statistical significance. In social market economies, by contrast, the association between these variables is far from significant.

The substantive significance of these findings is perhaps best illustrated by simulating the effects of a change in one or several of the statistically significant variables for a specific country. If we change the value of a particular variable to that of a country with more wage inequality, the simulated effect of this change can be compared with the actual difference in wage inequality between the two countries. Table 7 reports the results of simulations that assign German values to Sweden, using data for 1990 and the regression coefficients yielded by our interaction model. These two countries were chosen because Sweden, with a 90–10 ratio of 2.01 in 1990, stands out as the social market economy with the most compressed wage distribution, while Germany, with a 90–10 ratio of 2.72 in 1990, falls at the other end of the SME spectrum on our dependent variable.[43]

The first row of Table 7 contains our estimate of how the Swedish 90–10 ratio would have changed had Swedish union density suddenly dropped to the German level in 1990, with the values of all other variables remaining constant. Our regression results imply that such a change would have translated into a 24 percent increase of Swedish wage inequality over a ten-year period. In the second and third rows we report the results of repeating this exercise for the other two variables that proved to have a statistically significant association with wage inequality in SMEs, bargaining centralization, and government employ-

[43] In Figures 1 and 2 Austria has an even more dispersed wage distribution than Germany, but the fact that the Austrian wage data include part-time employees makes the Sweden-Germany comparison more appropriate.

TABLE 7
SIMULATED EFFECTS OF ASSIGNING GERMAN VALUES TO SWEDEN
(ACTUAL 90-10 RATIO IN 1990: 2.01)

			Simulation	
Variable[a]	*Swedish Level, 1990*	*German Level, 1990*	*90-10 Ratio after 10 Years*	*% Change in Wage Inequality*
Union density	84.0	32.9	2.48	+24
Bargaining centralization	.456	.318	2.13	+6
Government employment	31.6	15.1	2.26	+13
All 3 variables			2.97	+48

[a] When all three variables are changed at once, the total effect is realized in nine years, with the bulk of the effect realized in five years (90–10 ratio of 2.94 after five years).

ment. The fourth row, finally, contains our estimate of how Swedish wage inequality would have changed had all three variables taken on German values in 1990 (still holding all other variables in our analysis constant). When we change all three variables at once, Sweden's 90–10 ratio increases by 48 percent.

Similarly, Table 8 reports the results of simulations based on assigning U.S. values to Australia. As with Sweden and Germany, these countries were chosen because they represent opposite ends on the LME spectrum on our dependent variable. In 1990 the 90–10 ratio for Australia was 2.84, as compared with 4.33 for the U.S. Among LMEs there are four variables that have a statistically significant association with wage inequality: female labor-force participation, union density, government employment, and partisanship. Interpreting the results in Table 8 is complicated by the fact that as a pair Australia and the U.S. do not conform to our finding that there exists a positive relationship between the size of the public sector and wage inequality in LMEs. Assigning U.S. values to Australia on this variable actually reduces wage inequality in Australia (and increases the wage inequality gap between Australia and the U.S.). Nonetheless, changing the values of all four variables at once yields a 28 percent increase in wage inequality in Australia.

In a slightly different vein the British experience of the 1980s provides a convenient real-world illustration of the substantive significance of our estimates of the wage-distributive effects of government partisanship in liberal market economies. When the Conservatives took over the reins of government from the Labor Party in 1979, govern-

TABLE 8
SIMULATED EFFECTS OF ASSIGNING U.S. VALUES TO AUSTRALIA
(ACTUAL 90–10 RATIO IN 1990: 2.84)

			Simulation	
Variable[a]	*Australian Level, 1990*	*U.S. Level, 1990*	*90–10 Ratio after 10 Years*	*% Change in Wage Inequality*
Female labor-force participation	41.3	44.9	2.95	+4
Union density	46.6	15.6	3.68	+29
Government employment	23.0	14.6	2.59	−9
Government partisanship	2	4	2.97	+5
All 4 variables			3.64	+28

[a] When all four variables are changed at once, the total effect is realized in ten years, with the bulk of the effect realized in five years (90–10 ratio of 3.61 after five years).

ment partisanship, as measured by Cusack's index, increased by 100 percent (from 2 to 4). The results of our interaction model imply that the results of this enduring change raised the British 90–10 ratio by about 5 percent (from 2.95 to 3.09) over the subsequent decade, but in fact, the British 90–10 ratio increased by 15.6 percent from 1979 to 1990. In a loose sense, our regression results thus suggest that almost one-third of the increase of British wage inequality might be attributed to Conservative government policies.

VI. CONCLUSION

While our analysis provides no support whatsoever for the contention that unemployment is a source of wage inequality or for the contention that trade with low-wage countries is a source of wage inequality, it lends some support to the hypothesis that female labor-force participation promotes wage inequality by increasing the relative supply of unskilled labor. However, this hypothesis holds only for countries characterized as liberal market economies. Should we conclude, then, that supply-and-demand conditions are largely irrelevant to the evolution of wage inequality in advanced capitalist economies? Our results certainly do not warrant such a far-reaching conclusion, which flies in the face of a great deal of empirical work (as well as theorizing) by labor economists. Variables such as the aggregate unemployment rate and

LDC trade as a percentage of GDP are probably too crude to capture the impact of market forces. At most, our analysis suggests that the wage-distributive effects of such variables are less straightforward than commonly supposed and that we should be wary of exaggerating their significance. In particular, the wage-distributive effects of trade with low-wage countries are likely smaller than suggested by some scholars (notably Wood) and by many media commentaries.[44]

Our findings concerning the role of political-institutional variables are less tentative and also more interesting. We have identified several discrete variables that fall under this general heading and we have shown that it is possible to obtain statistically significant results for all these variables in a pooled regression framework. We have shown furthermore that the distinction between social and liberal market economies has important implications for our understanding of the determinants of wage inequality. As with female-labor force participation, the wage-distributive effects of bargaining centralization, government employment, and partisanship differ across varieties of capitalism. Only one of the variables in our analysis, union density, has a significant association with wage inequality that is unaffected by the distinction between SMEs and LMEs.

The finding that the effects of partisanship are contingent on broad institutional constellations should be of particular interest to political scientists. Students of comparative politics have long argued over whether or not—or to what extent—the partisan composition of government matters to real economic and social outcomes, with the skeptics emphasizing that governments are constrained by some combination of political-institutional arrangements, structural economic conditions (specifically, the critical importance of private investment), and the preferences of the median voter. In recent literature on the evolution of government spending and social policy in advanced capitalist countries, this debate has been recast as a debate over whether or not partisanship still matters. While Pierson and Stephens, Huber, and Ray argue that partisan effects on spending have diminished since the 1970s, Garrett finds that these effects have actually become more pronounced in the era of globalization.[45]

[44] Wood (fn. 8). The more fine-grained analysis by Vincent Mahler, David Jesuit, and Douglas Roscoe also fails to establish any clear and consistent pattern of association between wage inequality and various dimensions of "globalization." See Mahler, Jesuit, and Roscoe, "Exploring the Impact of Trade and Investment on Income Inequality," *Comparative Political Studies* 32 (May 1999).

[45] Garrett (fn. 3); Paul Pierson, "The New Politics of the Welfare State," *World Politics* 48 (January 1996); John Stephens, Evelyne Huber, and Leonard Ray, "The Welfare State in Hard Times," in Kitschelt et al. (fn. 2); and Garrett (fn. 3).

Our analysis suggests that it may be useful to introduce the varieties-of-capitalism idea into the debate about partisanship: the extent to which partisanship matters and the extent to which partisan effects have changed over time may depend on how the political economy (as a whole) is organized. This said, we hasten to point out that the interaction effect we have established here pertains specifically to the determinants of wage inequality. The fact that partisanship matters to the distribution of wages in LMEs but not in SMEs reflects the broad scope of institutionalized wage bargaining and the concomitant absence (or insignificance) of minimum wage legislation and other forms of direct government regulation of wages in SMEs. It does not follow from this finding that the effects of partisanship in the realms of social policy or macroeconomic management are smaller in SMEs than in LMEs. Further research and analysis are clearly necessary to determine the broader implications of the varieties-of-capitalism approach for quantitative comparative political economy.

Needless to say, the results of our interaction model depend on our prior coding of countries as SMEs, LMEs, and MIXs. Most obviously, the decision to code Japan and Switzerland as liberal market economies might be questioned. By the criteria set out above, Japan and Switzerland certainly cannot be categorized as social market economies, but perhaps they should be categorized as mixed cases rather than liberal market economies. Focusing on public or collective regulation of labor markets, the picture presented in Table 2 entirely misses the informal practice of lifetime employment by large Japanese companies, for instance. Rerunning our interaction model with Japan and Switzerland coded as mixed cases (along with France and Italy), the regression results obtained for the mixed cases were quite different, but the results for the SMEs and the LMEs were essentially the same as those reported above. For the SMEs and the LMEs, no variable became either significant or insignificant at the 90 percent confidence level and differences in coefficient estimates between the SMEs and the LMEs that were initially significant at the 90 percent level remained so when we recoded Japan and Switzerland.[46] Practically speaking, then, our conclusions do not depend on whether Japan and Switzerland are considered liberal market or mixed economies.

[46] Among the statistically significant variables, the single largest coefficient change we obtained was for government employment in LMEs: this coefficient fell from .104 to .086 when we recoded Japan and Switzerland. For the mixed cases, the coding change produced statistically significant coefficients for wage bargaining centralization (negative, as in SMEs), government employment (negative, as in SMEs) and government partisanship (positive, as in LMEs). Results available from the authors upon request.

The time-invariant quality of the dummies we use to capture varieties of capitalism represents another potential pitfall. Perhaps recent changes in advanced capitalist political economies have rendered the SME-LME distinction less meaningful. Substantive knowledge of the country cases is the only real check on this problem. SMEs and LMEs alike underwent institutional changes during the time covered by our analysis (1973–95)—to some extent, captured by our union density and wage-bargaining centralization variables—but there is precious little evidence of generalized convergence between SMEs and LMEs and we simply cannot think of any country that can be said to have moved from the SME camp to the LME camp or vice versa.

More fundamentally, two alternative understandings of these clusters of advanced capitalist political economies should be noted here. Consider the matter in terms of our dummy variable for SMEs. On the one hand, this variable might be viewed as a proxy for a set of discrete variables—welfare-state decommodification, institutionalization of collective bargaining, and employment protection—for which we do not have year-to-year observations. Some of these variables may be difficult to measure, but in principle they are all measurable, and if we had observations for each of these variables, the SME dummy would be superfluous. On the other hand, we might think of the SME dummy as capturing the way that discrete variables are configured within a coherent whole, that is, as referring to an institutional context with emergent properties that cannot be reduced to discrete variables. Setting aside the specificities of the SME concept, the tension between these two perspectives animates current theoretical debates in comparative political economy. While we are inclined to take the idea of emergent properties seriously, we want to emphasize the empirical nature of the questions at stake in this debate. The most obvious way to determine whether varieties of capitalism are reducible to discrete variables is to try to achieve such a reduction.

Finally, the importance of union density as a determinant of wage inequality deserves to be underscored one more time. Across the SME-LME divide, the effects of union density are consistently egalitarian, and for each cluster of countries they are greater than those of any other independent variable in our analysis. This finding suggests that conflicts of interest between unions and employers constitute an important dimension of the politics of wage distribution. If wage-distributive outcomes were primarily an expression of employer preferences, as Swenson's revisionist account of the Swedish story of wage solidarity would have it, it would be difficult to make sense of the association between union density and wage compression.[47] On this score, the general implication of our analysis is that we should be wary of embracing a wholly employer-centered approach to comparative political economy.

In no way does our analysis deny that distributive conflict among wage earners also constitutes an important dimension of the politics of wage distribution. The empirical data at our disposal simply do not allow us to capture this dimension. To do so, we would need to be able to measure gender differences in union density or the distribution of union membership across the hierarchy of wages and skills. (By shifting the focus to intersectoral wage differentials, the sectoral distribution of union membership would also be of interest.) The labor force surveys from which such measures might be derived are becoming available, but only for recent years. They thus do not allow very sophisticated analyses using countries as the unit of observation. Tackling this methodological challenge is an important part of the research agenda that flows from the preceding analysis. In future research we also want to explore the interface between distributive labor-market outcomes and the redistributive effects of the welfare state.

Appendix: Data Sources

90–10 ratio. data provided by the OECD, Directorate for Education, Employment, Labour and Social Affairs.

Unemployment. OECD, *Historical Statistics* (electronic database).

LDC trade. For all countries but Belgium, data on LDC trade were provided by Geoffrey Garrett (Yale University); Belgian and post-1990 figures were calculated on the basis of OECD, *Monthly Statistics of Foreign Trade.*

Female labor-force participation. OECD, *Historical Statistics* (electronic database).

Government employment. OECD, *Historical Statistics* (electronic database).

Union density. The pre-1990 figures were taken from Jelle Visser, "Unionization Trends Revisited," Centre for Research of European Societies and Industrial Relations (Amsterdam, 1996); post-1990 figures were provided by Bernhard Ebbinghaus (Max-Planck Institute).

[47] Peter Swenson, *Labor Markets and Welfare States* (forthcoming).

Centralization. The wage-bargaining centralization index was provided by Torben Iversen (Harvard University). For a complete specification, see Iversen (fn. 3).

Government partisanship. Figures were provided by Tom Cusack (Wissenschaftszentrum, Berlin). For details, see Cusack (fn. 24).