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# *Wage-Setting Institutions and Pay Inequality in Advanced Industrial Societies*

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The distribution of pay differs significantly across countries and over time among advanced industrial societies. In this paper, institutional and political determinants of pay inequality are studied in sixteen countries from 1980 to 1992. The most important factor in explaining pay dispersion is the level of wage-setting, i.e., whether wages are set at the level of the individual, the plant, the industry, or the entire private sector. The impact of centralization is the same whether centralization occurs via collective bargaining or via government involvement in private-sector wage-setting. The concentration of unions and the share of the labor force covered by collective bargaining agreements also matter. After controlling for wage-setting institutions, other variables such as the governing coalition, the size of government, international openness, and the supply of highly educated workers have little impact. Economic, political, and norm-based explanations for the association of centralization with egalitarian outcomes are discussed.

## 1. INTRODUCTION

There are large differences in the distribution of wages and salaries across advanced industrial societies and, in some countries, significant change over time in the recent past. In the United States, a worker who somehow managed to rise from the 10th decile of the wage distribution to the 90th decile would have received a pretax wage gain of 440 percent in 1990. To accomplish the same feat in 1980 would have taken a wage gain of only 380 percent. Both figures are in sharp contrast to the 98 percent increase that a Norwegian worker would obtain in going from the 10th to the 90th decile in the wage distribution in 1990. While countries may differ even more in the distribution of income from capital or transfer payments, the preponderance of labor earnings in total income is such that differences in the distribution of wages and salaries account for most of the cross-national variation in measures of the distribution of income among the nonelderly.<sup>1</sup>

This paper draws upon data that were collected as part of a project funded by the National Science Foundation entitled "Union Centralization Among Advanced Industrial Societies" directed by Peter Lange, Miriam Golden, and Michael Wallerstein. Financial support came from the National Science Foundation, SES-9309391 and SES-9108485 to UCLA and SES-9110228 and SBR-9309414 to Duke University. I thank Jelle Visser and Duane Swank for allowing me to use unpublished data they collected and Lyle Scruggs for his help in putting the data set together. I thank Miriam Golden, Peter Lange, Karl Ove Moene, Matthew Rabin, John Stephens, Margaret Levi, David Olson, Bruce Western, and Douglas Hibbs for useful discussions and comments on earlier drafts.

<sup>1</sup>In the United States, the correlation between labor earnings and total income, defined to be revenue from all sources before taxes but after transfers, is .938 (Díaz-Giménez, Quadrini, and Ríos-Rull 1997, 6).

In the United States, the growth of wage inequality since 1980 blunted the usual impact of economic growth on poverty alleviation. The prolonged economic expansion that began in 1982 had little effect on the proportion of the US population with incomes below the poverty line until the mid 1990s, in sharp contrast to the significant declines in poverty that occurred during earlier economic expansions in the postwar period (Blank 1997). While employment increased strongly, the gain in hours at work was more than offset by declining real wages for low-wage workers. With less income and more hours at work, the welfare of the poor unambiguously declined.<sup>2</sup>

Before 1980, when the distribution of income was relatively stable, wage inequality attracted relatively little scholarly attention. In response to the growth of wage inequality since 1980 in the United States, however, the study of the determinants of wage inequality has acquired greater urgency with a large and growing literature that is increasingly comparative in scope.<sup>3</sup> My purpose here is to document and discuss the importance of wage-setting institutions for the distribution of earnings and, hence, for the distribution of income. In particular, the data strongly indicate that the more wages are determined in a centralized fashion, whether through centralized collective bargaining or parliamentary action, the more equal the distribution of earnings. Conversely, the more wages are set in decentralized bargaining between unions and firms at the plant level or between individual workers and their employers, the more unequal the wage distribution. In fact, it is difficult to find other variables that matter once the institutional variation in wage-setting is controlled for.

The existence of an empirical relationship between wage-setting via collective bargaining and the compression of pay differentials has been noted in the literature on wage inequality. Freeman (1980) and Freeman and Medoff (1984) observed that unions in the United States reduce inequality both within unionized establishments and between unionized establishments. More recent studies by Card (1996), Freeman (1996), Dinardo, Fortin, and Lemieux (1996), and Fortin and Lemieux (1997) converge in estimating that the decline in union density in the United States can account for about 20 per-

<sup>2</sup>The political consequences of growing income inequality are less clear than the social consequences. Lipset (1960), among many others, argued that high levels of income inequality are associated with high demand for redistributive policies and political instability. Yet, in the recent past, growing income inequality in the US has been accompanied by reductions in welfare spending (Gramlich, Kasten, and Sammartino 1993; Moffitt, Ribar, and Wilhem 1998). See Moffitt, Ribar, and Wilhem (1998) and Moene and Wallerstein (1998) for studies of the impact of inequality on political support for welfare programs.

<sup>3</sup>Levy and Murnane (1992) is an early survey of the literature on the growth of earnings inequality in the US. The more recent survey by Gottschalk and Smeeding (1997) compares changes in earnings and household inequality in the US with other OECD countries.

cent of the rise in wage inequality during the 1980s. In a comparison of wage dispersion in the United States with wage dispersion in nine other countries using micro-data sets, Blau and Kahn (1996) conclude that the greater wage inequality in the United States exists in spite of supply and demand conditions that would have produced the opposite result if wage-setting were equally decentralized everywhere.

In comparing wage inequality across countries, the share of the work force covered by collective bargaining is less important than cross-national differences in bargaining institutions, in particular, cross-national differences in the centralization of wage-setting and the concentration of unions. Hibbs and Locking (1995) document the dramatic impact of the egalitarian wage policy pursued by the Swedish unions through centralized bargaining on the distribution of wages in Sweden, as well as the rise in inequality after bargaining was decentralized in the early 1980s. Erickson and Ichino (1994) demonstrate a large compression of wages in Italy due to the combination of the cost of living index, the *scala mobile*, and significant levels of inflation between 1975 and 1983. Freeman (1988) considers the relationship between centralized wage-setting and wage equality to be so close that he uses measures of wage dispersion as a proxy for the centralization of bargaining.

Here, the determinants of the inequality of pay are studied using a new data set that includes much better information regarding cross-national and temporal differences in wage-setting institutions for sixteen advanced industrial societies over the period 1950–1992. The central advantage of the new data set is the availability of time-series data on a variety of different aspects of wage-setting institutions, including the involvement of union and employer peak associations in wage-setting, government involvement in (private-sector) wage-setting, the level at which wages are set, and the concentration of membership both within and between union confederations, on a yearly or five-yearly basis for most of the postwar period.

With richer institutional data, the relationship between wage-setting institutions and wage inequality can be established with greater accuracy. With separate series on different aspects of wage-setting institutions for sixteen countries, we can investigate which institutional differences matter for wage equality and which do not. In addition, the impact of institutional differences in wage-setting can be compared with other possible determinants of wage inequality such as international openness, the partisan composition of government, the size of government, or the supply of highly-educated labor. Finally, with time-series data on wage-setting institutions, we can investigate the extent to which institutional change can explain the cumulative changes in the wage distribution from 1980 to 1992 among the major developed countries.

In the literature, it is common to refer to systems of centralized wage-setting as "corporatist." In this case, the conflation of centralization and corporatism is misleading. As is demonstrated below, centralization by means of parliamentary intervention in wage-setting has the same impact on wage equality as centralization by means of collective bargaining between peak associations of unions and employers. While bargaining between peak associations is largely confined to the countries conventionally labeled as corporatist, government intervention in wage-setting is not. Thus, centralization of wage-setting is a more precise description of what matters for wage inequality.

The body of my paper is organized as follows. The measures of wage inequality, institutional variation and other independent variables are described in Section 2. The estimation procedure is outlined in Section 3. Section 4 presents the empirical results on the determinants of pay inequality. Section 5 discusses the extent to which the empirical model can account for the change in wage inequality over time in those countries with the greatest change in wage equality. The implications of the empirical findings for understanding how institutions shape the wage distribution are discussed in Section 6. Section 7 concludes. The data sources are described in the appendix.

## 2. MEASURES OF THE DEPENDENT AND INDEPENDENT VARIABLES

Neither the equality of pay nor the variety of institutional differences in wage-setting are simple concepts to measure. Thus, it is necessary to begin with a discussion of how the dependent and independent variables are defined.

### 2.1 Pay Inequality

The measure of pay inequality that is used as the dependent variable is derived from the ratio of the wage received by the worker at the 90th percentile,  $w_{90}$ , to the wage received by the worker at the 10th percentile,  $w_{10}$ , for both sexes reported in OECD (1996) from 1980 through 1992.<sup>4</sup> No single statistic can encompass all important aspects of the wage distribution. For the purpose of studying the impact of wage-setting institutions on wage compression, however, a statistic like the  $w_{90}/w_{10}$  ratio has two important advantages over alternative measures such as the Gini coefficient or the variance of the log of wages. The first is that measurement error is most serious in the upper and lower tails of the wage distribution.<sup>5</sup> The  $w_{90}/w_{10}$  ratio is,

<sup>4</sup>To be precise, the data refer to gross (pretax) wages and salaries received by full-time workers. Nonwage benefits are not included.

<sup>5</sup>The wage data comes from either labor market surveys or administrative data such as collected by the social security system, depending on the country. In either case, the data is generally "top-coded" where reported earnings above some threshold  $x$  are recorded as being equal to  $x$ . The problem at the bottom of the wage distribution is the absence of reliable measures of earnings that are hidden from the tax authorities.

therefore, a more reliable statistic than measures that are sensitive to reported wages at the top and bottom of the wage scale.

The second advantage of the  $w_{90}/w_{10}$  ratio stems from its insensitivity to wage differentials among observationally equivalent workers. With centralized wage-setting, workers in similar positions with similar credentials and seniority must be treated identically since information about individual employees is not available at levels of centralization higher than the firm. The standardization of wages for workers with identical credentials in identical job categories lowers the variance of the wage distribution, but does not affect the  $w_{90}/w_{10}$  ratio. Thus, if centralization has an effect on the  $w_{90}/w_{10}$  ratio, that effect is due to something other than the reduction of inequality that occurs solely because of informational constraints.

Figure 1 presents the  $w_{90}/w_{10}$  ratios for the sixteen countries included in the study in three different years: 1980, 1986, and 1992.<sup>6</sup> The correlation of the  $w_{90}/w_{10}$  ratio and the  $w_{50}/w_{10}$  ratio, where  $w_{50}$  is the median wage, is very high (around .93). The distribution of  $w_{90}/w_{10}$  ratios is positively skewed, with wage inequality in the US and Canada substantially higher than the other fourteen countries. Moreover, the  $w_{90}/w_{10}$  ratio cannot be less than one by definition, which is incompatible with the standard assumption that the error term has a normal distribution. Therefore, for the purposes of estimation, the  $w_{90}/w_{10}$  ratio was transformed according to the formula

$$y = \ln \left( \frac{w_{90} - w_{10}}{w_{10}} \right) \quad (1)$$

to obtain a measure of pay inequality that can take any value between negative infinity (indicating that the 90th percentile worker and the 10th percentile worker receive equal pay) and positive infinity.<sup>7</sup> A value of  $y = 0$  indicates that  $[(w_{90} - w_{10})/w_{10}] = 1$  or that the wage differential between the 90th and 10th percentile workers is 100 percent, roughly the level of wage dispersion found in Norway and Sweden.

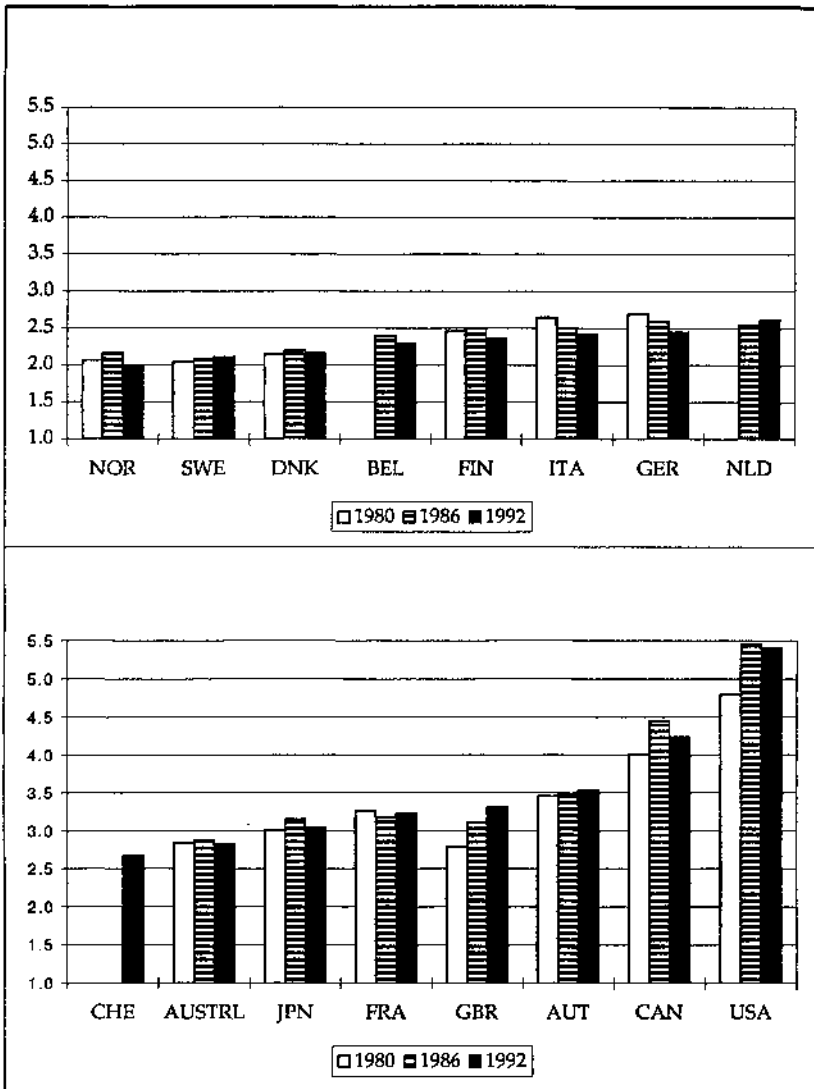
## 2.2 Measures of Institutional Differences in Wage-Setting

The set of independent variables contains measures of confederal and government involvement in wage-setting, the level at which wages are

<sup>6</sup>The countries in Figure 1 are Norway (NOR), Sweden (SWE), Denmark (DNK), Belgium (BEL), Finland (FIN), Italy (ITA), (West) Germany (GER), Netherlands (NLD), Switzerland (CHE), Australia (AUSTRAL), Japan (JPN), France (FRA), Great Britain (GBR), Austria (AUT), Canada (CAN), and the United States (USA). Data for the years 1980, 1986, and 1992 were used whenever possible. In cases with missing data, I used data from the closest available year. (See the appendix for details.) There is no data for Belgium or the Netherlands before 1985. There is no data for Switzerland before 1991. The data source is OECD (1996, Table 3.1, 61–62).

<sup>7</sup>The results presented below are not sensitive to the logarithmic transformation. Similar results are obtained with  $w_{90}/w_{10}$  as the dependent variable.

Figure 1. 90/10 Wage Ratio



predominantly set, the concentration of union membership within and between union confederations, union density, and union coverage. Additional details regarding the data and the list of data sources are contained in the appendix.

*The centralization of wage-setting*

As scholars of comparative industrial relations have frequently noted, there are large cross-national differences in the level at which wage agreements are negotiated and in the role of the peak associations of unions and employers in collective bargaining. First, there is the general distinction between systems of industrial relations in which wage contracts are largely negotiated at the plant level (the US, Canada, Great Britain, and Japan prior to the initiation of industry-wide coordination through the annual spring offensive) or at the industry-level (all of the other countries in the sample). Among countries with industry-wide bargaining, there is a wide range with respect to the role played by the peak associations of unions and employers, ranging from none to peak-level negotiation of a centralized wage agreement.

Moreover, to focus on the collective bargaining system alone is too narrow. Wage-setting can also be centralized via parliamentary action. All governments have an impact on the distribution of pay via their role as employer of a significant share of the work force. Most governments also have some role in private-sector wage-setting. At one end of the spectrum, governments do no more than legislate a minimum wage or extend the terms of collective agreements to nonunion workers. At the other end of the spectrum, governments directly determine private sector wages through arbitration or the imposition of mandatory wage controls.

Table 1a and 1b presents the scales devised by Golden, Lange, and Wallerstein to measure confederal and government involvement in private-sector wage bargaining.<sup>8</sup> Confederal involvement is measured for every bargaining round, usually every two years. Government involvement is measured annually. The rankings of categories reflect both the role of the central confederations (or government) and the degree to which central agreements constrain wage negotiations at lower levels.

Central agreements generally impose a floor on wages. In the absence of an industrial peace clause, industry and local-level negotiators are free to bargain and to strike for additional wage increases above the increase specified in the central agreement. With an industrial peace clause, bargaining at lower levels is permitted but strikes and lockouts are prohibited for the duration of the central agreement. Although other forms of industrial action, such as go-slow or work-to-rule actions, may be allowed and no clause in a contract can prevent wildcat strikes, the existence of an industrial peace clause significantly increases the ability of central wage-setters to control the aggregate wage increase in multi-level bargaining (Moene, Wallerstein, and

<sup>8</sup>For descriptions of the changes in wage-setting institutions during the postwar period revealed by the data, see Wallerstein, Golden, and Lange (1997), Wallerstein and Golden (1997), and Golden, Wallerstein and Lange (1998).



**Table 1a. Index of Confederal Involvement in Wage-Setting**

- 
1. Confederation(s) uninvolved in wage-setting in any of the subsequent ways.
  2. Confederation(s) participates in talks or in formulation of demands for some affiliates.
  3. Confederation(s) participates in talks or in formulation of demands for all affiliates.
  4. Confederation(s) negotiates non-wage benefits.
  5. Confederation(s) negotiates a part of the wage agreement, such as the cost-of-living-adjustment.
  6. Confederation(s) represents affiliates in mediation with centralized ratification.
  7. Confederation(s) represents affiliates in arbitration.
  8. Confederation(s) bargains for affiliates in industry-level negotiations.
  9. Confederation(s) negotiates national wage agreement without peace obligation.
  10. Confederation(s) negotiates national wage agreement with peace obligation.
  11. Confederation(s) negotiates national wage agreement with limits on supplementary bargaining.
- 

**Table 1b. Index of Government Involvement in Wage-Setting**

- 
1. Government uninvolved in wage setting.
  2. Government establishes minimum wage(s).
  3. Government extends collective agreements.
  4. Government provides economic forecasts to bargaining partners.
  5. Government recommends wage guidelines or norms.
  6. Government and unions negotiate wage guidelines.
  7. Government imposes wage controls in selected industries.
  8. Government imposes cost of living adjustment.
  9. Formal tripartite agreement for national wage schedule without sanctions.
  10. Formal tripartite agreement for national wage schedule with sanctions.
  11. Government arbitrator imposes wage schedules without sanctions on unions.
  12. Government arbitrator imposes national wage schedule with sanctions.
  13. Government imposes national wage schedule with sanctions.
  14. Formal tripartite agreement for national wage schedule with supplementary local bargaining prohibited.
  15. Government imposes wage freeze and prohibits supplementary local bargaining.
- 

**Table 1c. Index of the Level of Wage-Setting**

- 
1. Local wage-setting.
  2. Industry-level wage-setting.
  3. Centralized wage-setting without sanctions.
  4. Centralized wage-setting with sanctions.
- 

Hoel 1993, chapter 12). Thus, agreements that contain peace clauses are considered to be more centralized than those that do not.

In the scale for government involvement, presented in Table 1b, wage-setting by Parliament is judged to be more centralized than wage-setting by a government-appointed arbitrator. An arbitrator's mandate is to craft an

agreement that is acceptable to both unions and employers, thereby avoiding industrial conflict. For Parliament, in contrast, the government's macroeconomic goals may be the primary consideration. The scale distinguishes between a government-imposed wage contract, government participation in tripartite talks in which a wage contract is negotiated as part of a broader package, and government attempts to influence the wage agreement without formally participating in the wage negotiations.

In addition to using separate indices for confederal and government involvement, two other summary measures of centralization were constructed. The first combines the confederal and government involvement scores by rescaling both indices to have a common range of [0,1] and then using the maximum of the two. A second summary measure of centralization, summarized in Table 1c, is a four-category scale indicating the level at which wages are predominantly set for each country for each year. A score of three indicates centralized wage-setting with sanctions on lower-level bargaining whether by centralized collective bargaining or government action. A score of two indicates centralized wage-setting without sanctions on subsequent lower-level bargaining. A score of one indicates wage-setting at the level of the industry while a score of zero indicates the predominance of wage-setting at the level of the firm or the individual employee-employer pair. The country means of the index of the level of wage-setting over the time period 1950-1992 are displayed later in Table 5 in the appendix.

### *Union concentration*

Centralization measures explicit coordination of wage-setting among workers in different firms or different industries. However, there can be a substantial degree of implicit coordination that may achieve much the same outcome in the absence of a centralized procedure. A particular union, the German metalworkers for example, may act as the wage leader. If the wage agreement signed in the leading industry is quickly adopted in other industries, and the wage negotiators in the leading industry understand that the terms of their agreement will rapidly spread throughout the economy, the outcome may be a wage schedule that is not very different from what would result from the direct negotiation of a centralized agreement covering the private sector as a whole.

An important determinant of the ability of unions and employers to coordinate wage settlements implicitly in the absence of a formal centralized agreement may be the extent to which the union side is dominated by a small number of actors or the degree of concentration of union membership (Golden 1993). In the data set, concentration is measured along two dimensions. The first is between confederations, or the extent to which union members belong to a single confederation rather than being divided among

multiple confederations. The measure used is the Herfindahl index between confederations:

$$H_B = \sum_{j=1}^N (S_j)^2 \quad (2)$$

In Equation 2,  $S_j$  is the share of total union members who belong to confederation  $j$ , and  $N$  is the total number of confederations. The Herfindahl index of concentration between confederations is the probability that two union members, picked at random, will belong to the same confederation.<sup>9</sup>

Another dimension of concentration is the extent to which the membership of a single union confederation is concentrated within a small number of affiliates. To measure concentration within each confederation, an approximate Herfindahl index was constructed using the membership of the three largest affiliates and the total number of affiliates. The approximate Herfindahl index of concentration within confederations is defined as:

$$H_w = \sum_{j=1}^3 (s_j)^2 + \left[ 1 - \sum_{j=1}^3 s_j \right]^2 \frac{1}{n-3} \quad (3)$$

where  $s_j$  is the share of the confederation's members who belong to the  $j$ th largest union, and  $n$  is the total number of affiliates in the confederation.  $H_w$  represents an approximation of the probability that two members of confederation  $i$ , picked at random, will belong to the same affiliate. The formula for  $H_w$  in Equation 3 implicitly assumes that the fourth through  $n$ th unions are the same size. Thus, Equation 3 is an underestimate of the true Herfindahl index, but the underestimate is not large in practice. The approximate Herfindahl was calculated for the main blue-collar confederation, where there is a dominant blue-collar confederation. In countries with multiple major blue-collar confederations, such as Italy or the Netherlands, the average Herfindahl index for each blue-collar confederation was used, weighted by the confederation's share of total union membership.<sup>10</sup> Unfortunately, membership by affiliate is unavailable in France, so the Herfindahl index of concentration within confederations is only available for fifteen countries. Both

<sup>9</sup>Members of unions unaffiliated with a confederation are not included, due to the difficulty of obtaining reliable membership figures for independent unions in all sixteen countries.

<sup>10</sup>The Belgian figure for within-confederal concentration reflects only the Catholic confederation, as membership figures by affiliate are unavailable for the socialist confederation. Since the two confederations have close to the same (small) number of affiliates, the two are assumed to be equally concentrated.

measures of concentration were calculated every five years from 1950 though 1990. (See Table 5 in the appendix for the country means of  $H_w$  from 1950–1990.)

### *Union density and coverage*

Studies of the influence of unions in the US almost always use union density as the measure of union influence. Here, density is defined as union members who work as employees divided by the total number of wage and salary earners. Thus, the definition of union density excludes workers who are retired, unemployed, or self-employed from both the numerator and the denominator.

A different measure of the extent to which unions influence the aggregate wage distribution is coverage, defined to be the share of the work force covered by a collective agreement. In some countries, such as the US, Canada, Japan, and Great Britain, there is a close correspondence between union density and coverage. In other countries, coverage far exceeds density for a variety of reasons. In France and Belgium, coverage is frequently extended by government decree. In Germany and Austria, a labor agreement signed by the employers' association is binding on all affiliated firms whether or not the firm's employees are union members, and most employees work for employers who belong to the employers' association. In fact, membership of firms in the relevant employers' association is mandated by law in Austria. Not only is union coverage much higher than union density in continental Europe, but coverage has remained stable since 1980 in continental Europe, even in countries where union density has declined.<sup>11</sup> While union density figures are available for the entire postwar period, union coverage figures are available only for 1980 and 1990. The data for union coverage in 1990 is presented in Table 5 in the appendix.

## **2.3 Other Independent Variables**

### *Partisan variables*

Given the extensive government involvement in private-sector wage-setting in many advanced industrial countries, the ideology or the constituency of the government might have an impact on wage equality. In particular, social democratic governments might push for greater wage equality than conservative governments. Left government is measured by the share of cabinet portfolios held by socialist, social democratic, or labor parties as a proportion of all cabinet portfolios. Alternatively, it is sometimes argued that the

<sup>11</sup>In 1990, the unweighted average of union density was 46.8 percent while the unweighted average of union coverage was 76 percent for the eleven countries of the sample in continental Europe.

important political divide is not between the socialist camp and the rest, but between both left and center parties, on the one side, and conservative parties on the other. Therefore, the proportion of cabinet portfolios held by conservative parties was also included, following the classification of Castles and Mair (1984).

### *International openness*

Wood (1994) argues that increasing international trade is responsible for much of the rise in the inequality of pay in the US and, to the extent that the wage-setting system allows wages at the bottom to fall, in other advanced industrial societies as well. Trade dependence, as measured by imports plus exports as a share of GDP, was included to test the direct impact of trade on pay inequality.<sup>12</sup>

### *The size of the public sector*

Wage equality might be affected by the size of the public sector. Katz and Krueger (1991) show that public sector wages are more compressed than private sector wages in the US. If public sector wages are generally more compressed than private sector wages, then public sector employment as a share of total employment might be an important determinant of aggregate wage inequality. The size of the public sector in terms of government spending as a share of GDP may also be important. On the spending side, generous welfare policies may increase the bargaining power of low-wage workers by providing better options outside the labor market. On the revenue side, Hibbs and Locking (1996) suggest that high marginal tax rates reduce the cost of wage compression for high wage workers, since much of the extra income that high wage workers would receive with a less compressed wage scale would be taxed away.

### *Education*

Finally, one might expect to find more compressed wage scales in countries with a relatively large supply of highly educated labor.<sup>13</sup> Two measures of the supply of educated workers on the wage distribution are used. The

<sup>12</sup>In fairness, it should be noted that what matters for the wage distribution, according to standard trade theory, is the impact of trade on relative prices in the domestic economy which may not be accurately proxied by the ratio of imports plus exports to GDP. For a review of the debate over the importance of international trade in explaining the rise of pay inequality in the US, see Freeman (1995), Richardson (1995), and Wood (1995).

<sup>13</sup>At the same time, if additional years of schooling have a relatively small effect on future earnings, the number of persons who seek higher education may be reduced (Edin and Topel 1995). Thus, it is impossible to say a priori whether the correlation between a large supply of educated workers and wage equality should be positive or negative.

first is the mean number of years of higher education among persons between the ages of 15 and 64. The second is the mean number of years of education at all levels for the same age bracket.

### 3. STATISTICAL PROCEDURES

The adjustments to the wage distribution that occurs from one year to the next are always small, whatever the wage-setting system. Workers' relative wages in year  $t$  are strongly dependent on their relative wages in year  $t-1$ . The most compelling empirical specification of the evolution of the distribution of pay is the error-correction model. The underlying assumption is that there is an equilibrium wage distribution associated with a given set of wage-setting institutions and other determinants of relative pay, but that the adjustment to the equilibrium distribution is not instantaneous. The actual wage distribution is modeled as a weighted average of the equilibrium wage distribution, which changes as the wage-setting environment changes, and the wage distribution that existed in the previous period.

Formally, let  $y_i^*(t)$  be the equilibrium wage dispersion for country  $i$  at time  $t$ . The equilibrium wage dispersion is the level of wage dispersion that, once obtained, would not change provided there was no change in the exogenous variables. Different wage-setting institutions and other exogenous factors are assumed to affect the equilibrium wage distribution in the standard linear fashion

$$y_i^*(t) = X_i(t)' \beta + v_i^*(t) \quad (4)$$

where  $X_i(t)$  is a vector of independent variables associated with country  $i$  at time  $t$  and  $v_i^*(t)$  is a random error term associated with the equilibrium wage distribution.

The actual wage dispersion at time  $t$ , denoted  $y_i(t)$ , is assumed to be equal to the wage dispersion at time  $t-1$  plus an adjustment that depends on the difference between the wage dispersion at time  $t-1$  and the equilibrium wage dispersion at time  $t$ , or

$$y_i(t) = y_i(t-1) + (1-\lambda)[y_i^*(t) - y_i(t-1)] + v_i(t) \quad (5)$$

where  $\lambda$  is a parameter indicating the speed at which the distribution of wages adjust and  $v_i(t)$  is a random error term. Combining Equations 4 and 5, we have

$$y_i(t) = \lambda y_i(t-1) + (1-\lambda)X_i(t)' \beta + u_i(t) \quad (6)$$

where  $u_i(t) = [v_i(t) + (1 - \lambda)v_i^*(t)]$ . I assume, initially, that  $E[u_i(t)^2] = \sigma^2$  for all  $i$  and all  $t$  and that  $E[u_i(t)u_j(s)] = 0$  if either  $i \neq j$  or  $t \neq s$ .<sup>14</sup>

Rather than work with annual data, the statistical analysis is performed with the three cross-sectional panels of data displayed in Figure 1. There were several reasons for this choice. First, one of the institutional variables that turns out to be important is only available in five-year intervals. Second, the annual change in the wage distribution is small relative to the measurement error in the data on wage dispersion.<sup>15</sup> Cumulating the change in dependent and independent variables over approximately six years has the effect of increasing the variance of the independent variables relative to the noise in the data.

Let  $t_0$  be the first year for which we have wage dispersion data for country  $i$ . This is usually 1980, but in some countries it is 1981 or even, in the case of Germany, 1983. Let  $t_1$  be 1986, unless 1986 data is missing for country  $i$ , in which case  $t_1 = 1987$ . Finally, let  $t_2$  be the last year for which we have wage dispersion data. This is 1992 for most countries, but 1991 for Norway and Italy and 1990 for Denmark. Then, for country  $i$ , repeated substitution using Equation 6 yields

$$y_i(t_1) = \lambda^{(t_1 - t_0)} y_i(t_0) + \sum_{k=0}^{t_1 - t_0 - 1} \lambda^k \left[ (1 - \lambda) X_i(t_1 - k)' \beta + u_i(t_1 - k) \right] \quad (7)$$

for  $y_i(t_1)$  and a similar expression for  $y_i(t_2)$ ). Since data for the independent variables are available for many years prior to  $t_0$ ,  $y_i(t_0)$  can be estimated without knowing  $y_i(t_{-1})$  by making repeated use of Equation 6 to obtain

$$y_i(t_0) = \sum_{k=0}^{\infty} \lambda^k \left[ (1 - \lambda) X_i(t_0 - k)' \beta + u_i(t_0 - k) \right] \quad (8)$$

In practice, the Golden, Lange, Wallerstein data set only goes back to 1950. Therefore, in estimating Equation 8, it was assumed that  $X_i(1980 - k) = 0$  for  $k > 30$  for all independent variables except the constant. With regard to the variables pertaining to wage-setting institutions, this is equivalent to as-

<sup>14</sup>Both fixed period effects and fixed country effects are introduced later.

<sup>15</sup>Some sense of the extent of measurement error can be obtained by comparing  $w_{90}/w_{10}$  ratios derived from different surveys of the same country in the same year (OECD 1996, annex 3A: 100-103). The correlation of the annual change in wage inequality derived from different surveys of the same country is frequently close to zero. See Rueda and Pontusson (1998) for a study of wage inequality based on annual pooled time series data. The major difference between Rueda and Pontusson's results and the results in this paper are discussed below.

suming that the distribution of wages was determined in decentralized markets in all sixteen countries prior to 1950. Such an assumption is clearly false, but to the extent that the assumption matters, it biases the coefficients toward zero given the high correlation between the centralization of bargaining in the postwar period and the centralization of bargaining before 1950.<sup>16</sup>

In sum, for each country  $i$  the following system of equations was estimated:

$$\tilde{y}_i = \tilde{X}_i \beta (1 - \lambda) + \tilde{u}_i \quad (9)$$

where

$$\tilde{y}_i = \begin{pmatrix} y_i(t_0) \\ y_i(t_1) - \lambda^{(t_1 - t_0)} y_i(t_0) \\ y_i(t_2) - \lambda^{(t_2 - t_1)} y_i(t_1) \end{pmatrix}$$

$$\tilde{X}_i = \begin{pmatrix} \sum_{k=0}^{t_0 - 1950} \lambda^k X_i(t_0 - k) \\ \sum_{k=0}^{t_1 - t_0 - 1} \lambda^k X_i(t_1 - k) \\ \sum_{k=0}^{t_2 - t_1 - 1} \lambda^k X_i(t_2 - k) \end{pmatrix}$$

and

$$\tilde{u}_i = \begin{pmatrix} \sum_{k=0}^{\infty} \lambda^k u_i(t_0 - k) \\ \sum_{k=0}^{t_1 - t_0 - 1} \lambda^k u_i(t_1 - k) \\ \sum_{k=0}^{t_2 - t_1 - 1} \lambda^k u_i(t_2 - k) \end{pmatrix}$$

<sup>16</sup>An additional problem is that some of the data is not available on an annual basis. The Herfindahl indices between and within confederations are available every five years starting in 1950. Before cumulating, these data series were completed by linear interpolation between data points. The change in Herfindahl indices is small, so the use of interpolated data has little effect on the results. Coverage is available only in 1980 and 1990. Again, linear interpolation was used to fill in the data after 1980. To fill in the rest of the series, coverage was assumed to be constant in the three decades prior to 1980. While this is a strong assumption, the data do show that coverage was nearly constant in the decade following 1980. Nevertheless, coverage is measured with a great deal more error than the other independent variables used in this study.



Stacking the set of equations for each country, the variance structure of the error term in Equation 7 can be written succinctly as<sup>17</sup>

$$E(\tilde{u}\tilde{u}') = I \otimes \Omega$$

where  $I$  is a  $16 \times 16$  identity matrix and

$$\Omega = \frac{\sigma^2}{1-\lambda^2} \begin{pmatrix} 1 & 0 & 0 \\ 0 & 1-\lambda^{2(t_1-t_0)} & 0 \\ 0 & 0 & 1-\lambda^{2(t_2-t_1)} \end{pmatrix}$$

Once  $\lambda$  is determined, the system of equations in 9 can be estimated by GLS. The parameter  $\lambda$  was estimated by maximum likelihood.<sup>18</sup>

#### 4. EMPIRICAL RESULTS

The impact of wage-setting institutions on the inequality of wages and salaries is presented in Table 2. The value of  $\lambda$  used in the regressions,  $\lambda = .95$ , is the value that maximizes the likelihood function with a weighted constant (as described in the notes at the bottom of the table), the level of wage-setting and the Herfindahl index of concentration within confederations as independent variables. The maximum likelihood value of  $\lambda$  changes only slightly with the inclusion of different sets of independent variables.<sup>19</sup>

Consider, first, the first three equations of Table 2. When the index of confederal involvement and the index of government involvement are added to the regression equation separately, both coefficients are negative, indicating that greater centralization reduces wage inequality, but only the coefficient on confederal involvement is bounded away from zero with standard confidence intervals. Combining the two indices by taking the value of whichever is larger substantially increases the fit of the regression line. But the simple fourfold index of the level of wage-setting produces an even better fit, explaining more than 70 percent of the variance with a single variable. The level of wage-setting consistently outperforms the other measures of centralization whatever other independent variables are added, including the addition of dummy variables for both time periods and countries. When

<sup>17</sup>The symbol  $\otimes$  indicates the Kronecker product.

<sup>18</sup>Assuming that the error term is normally distributed, the GLS estimate maximizes the likelihood function for a fixed  $\lambda$ . Thus, the procedure outlined here would be identical to finding the maximum likelihood estimators for Equation 7 if the estimate of  $\lambda$  was recalculated for every specification of the exogenous variables. See Greene (1993, chapter 13) for a discussion of the properties of the GLS estimator.

<sup>19</sup>The robustness of the results with respect to different values of  $\lambda$  is investigated below.

**Table 2. Centralization, Concentration, Coverage, and Pay Inequality**

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Constant	.918 (9.96)	1.06 (11.8)	1.17 (17.5)	1.33 (10.7)	1.42 (14.2)	1.57 (10.4)	1.48 (4.12)
Confederal Inv.	-.0469 (3.80)						
Government Inv.	-.0165 (0.98)						
Maximum		-.0650 (6.24)					
Level of Wage- Setting			-.0234 (10.3)	-.0165 (5.34)	-.0191 (6.93)	-.0182 (7.03)	-.0159 (3.91)
Concentration Between Conf.				.0171 (1.79)			
Concentration Within Conf.				-.116 (3.01)	-.134 (3.55)	-.164 (4.00)	-.040 (0.34)
Density				-.0122 (0.75)			
Coverage				-.0215 (1.72)	-.0148 (1.45)	-.0196 (2.00)	-.0635 (1.68)
Fixed effects							
Period	no	no	no	no	no	yes	yes
Country	no	no	no	no	no	no	yes
<i>N</i>	44	44	44	41	41	41	40
Buse $R^2$	.431	.481	.716	.818	.796	.844	.959

## Notes:

The dependent variable is  $\ln [(w_{90} - w_{10})/w_{10}]$ . GLS estimation with  $\lambda = .95$ . The absolute value of the *t*-statistics are reported in the parentheses. "Constant" refers to the quasi-differenced constant given by setting  $x_i(t) = 1$  in Equation 9. There are three period dummies and thirteen country dummies, with the country dummies quasi-differenced in the same way as the constant term. Equations 4–7 exclude France for lack of data on concentration within confederations. Equation 7 also excludes Switzerland for lack of time series data on the Swiss wage distribution. The *F*-statistics for period fixed effects in (6) and (7) are 3.54 (with [3,34] degrees of freedom) and 7.46 (with [3,20] degrees of freedom) respectively. The *F*-statistic for country fixed effects in (7) is 4.35 (with [13,20] degrees of freedom). The Buse  $R^2$  is equal to  $1 - (SS_{res}/SS_{const})$  where  $SS_{res}$  is the sum of squares of the residuals of the regression and  $SS_{const}$  is the sum of squares of the residual when "constant" is the only independent variable (Buse 1973).

the level of wage-setting and one of the other measures of the centralization of wage-setting are combined in the same equation, only the level of wage-setting receives an estimated coefficient significantly different from zero.

Specification 4 adds the Herfindahl indices for concentration between and within confederations, along with density and coverage. The results indicate that the concentration of confederation members among a small number of affiliates is associated with greater pay equality. In contrast, the coefficient on the Herfindahl index between confederations is invariably estimated to have the wrong sign regardless of the other variables that are added to the regression equation. Equation 4 also indicates that union density has a substantially smaller impact on pay inequality than the coverage of union contracts. Equation 5 presents the regression equation with the institutional variables that work best: the level of wage-setting, concentration within confederations, and coverage.

To check the robustness of the results, Equation 5 was reestimated with fixed period and country effects in Equations 6 and 7. The point estimate of the coefficient on level of wage-setting is somewhat reduced with the addition of fixed country and period effects, but it still remains significant at the 1 percent significance level in all specifications. Concentration within confederations has very little variance over time within countries. Thus, the addition of country fixed effects removes almost all of the explanatory power of the Herfindahl index. With the data we have, it is impossible to distinguish the impact of union concentration from unobserved country-specific effects. Finally, the estimate of the coefficient on coverage is sharply increased by the addition of fixed period effects.

If we accept the point estimates in Equation 6 of Table 2, the long-term impact of a permanent change in the system of wage-setting from a system of plant or individual-level bargaining (as in Britain, Canada, or the US), to a system of industry-level bargaining (as in Switzerland, Austria, or Germany), is to reduce the wage differential  $(w_{90} - w_{10})/w_{10}$  by 30 percent, since  $\exp(-.018/(1 - \lambda)) = \exp(-.36) \approx 0.70$ . A decline of two steps in the four step scale, for example, a lasting move from highly centralized bargaining, as in Sweden before 1983, to a system of industry-level bargaining would raise the predicted wage differential by 50 percent ( $\exp(-.72) \approx 0.5$ ) in the long run.

The most concentrated union confederations in the data set are the German and Swiss confederations. The estimate of the long-term impact of the difference between the Herfindahl index for the German DGB (about .16 in 1990) and the Herfindahl index of the least concentrated confederation in the data set, the Australian ACTU (about .02 in 1990), on the  $(w_{90} - w_{10})/w_{10}$  wage differential is  $\exp(-(.14)(.16)/(1-\lambda)) \approx 0.63$ , which is roughly the same magnitude as the estimated difference between local and industry-level bargaining. Finally, the impact of increasing the coverage of union contracts from the US level (18 percent in 1990) to the average level of coverage in continental Europe (74 percent in 1990), would reduce  $(w_{90} - w_{10})/w_{10}$  in the US by roughly 20 percent in the long run since  $\exp(-(.02)(.56)/(1-\lambda)) \approx 0.80$ .

**Table 3. Pay Inequality and Other Explanatory Variables**

	(1)	(2)	(3)	(4)
Constant	1.51 (10.4)	1.92 (9.75)	1.39 (6.05)	1.43 (6.28)
Level of Wage- Setting	-.0179 (6.03)	-.0187 (7.58)	-.0192 (5.88)	-.0184 (5.86)
Concentration	-.182 (3.75)	-.201 (4.83)	-.111 (1.90)	-.0184 (2.33)
Coverage	-.0295 (2.43)	-.0267 (2.61)	-.0099 (0.72)	-.0124 (0.92)
Left Government	.0021 (0.23)	.0090 (1.11)		
Right Government	-.0093 (1.27)	-.0117 (1.92)		
Trade Dependence	-.0134 (1.32)	-.0261 (2.61)		
Government Employment	-.0971 (1.83)	-.0845 (1.89)		
Government Spending	.0864 (1.93)	.0720 (1.92)		
Year	0 (0)			
Higher Education			.170 (1.33)	.169 (1.34)
Total Education			-.0811 (1.96)	-.0300 (0.59)
Period Dummies	no	yes	no	yes
<i>N</i>	41	41	26	26
Buse $R^2$	.839	.896	.858	.886

## Notes:

The dependent variable is  $\ln [(w_{90} - w_{10})/w_{10}]$ . GLS estimation with  $\lambda = .95$ . The absolute value of the *t*-statistics are reported in the parentheses. Concentration refers to concentration within confederations. The constant term and the Buse  $R^2$  are described in the notes to Table 2. The *F*-statistics for the period fixed effects in (2) and (4) are 5.31 (with [3,29] degrees of freedom) and 2.22 (with [2,18] degrees of freedom) respectively. France is excluded because of missing concentration data. The 1992 panel is excluded from (3) and (4) because of missing data for mean years of education.

The first two columns of Table 3 present estimates of the impact of the partisan composition of government, economic openness, and the size of the public sector on wage inequality controlling for the level of wage-setting concentration within confederations and coverage. The first column of Table

3 includes a time trend. The second column replaces the time trend with fixed period effects.<sup>20</sup> To save space, Equations 1 and 2 in Table 3 only report the results of adding the six additional independent variables at the same time. Adding subsets of the six additional independent variables does not change the general pattern of results, except for the interaction of government employment and government spending noted below.

The data decisively reject the hypothesis that the participation of social democratic, socialist, or labor parties in government reduces wage inequality or that the participation of conservative parties in government increases wage inequality. These results indicate that partisan composition of government does not have a *direct* effect on wage inequality. To the extent that left governments are more likely to encourage centralized bargaining by the unions or to intervene directly in wage-setting, however, the partisan composition of government may have an important indirect effect.

In addition, the cross-national data fail to support the claim that the increase of wage inequality in the US since 1980 is a consequence of increased trade. The hypothesis that growing trade dependence, as measured by imports plus exports as a share of GDP, is associated with greater wage inequality is contradicted by the negative point estimate of the coefficient on trade dependence. Countries with high levels of trade dependence tend to have relatively egalitarian wage distributions, even after controlling for the fact that countries with high levels of trade dependence have relatively centralized systems of wage-setting. Nor is there evidence of any time trend in the inequality of pay in the cross-national data. Note that the absence of a positive time trend does not mean that wage inequality has remained constant over time on average. Rather, the zero time trend implies that the change in average inequality of pay is fully explained by the change in the means of the other independent variables.

The impact of increased public sector employment is to reduce wage inequality as expected but only at a given level of government spending as a share of GDP. If government spending is removed from the set of independent variables, the point estimate for the coefficient on public sector employment is close to zero. The hypothesis that government spending reduces wage inequality, in contrast, is rejected. At given level of public sector employment, greater government spending as a share of GDP is associated with higher, not lower, wage inequality. This pattern appears to correspond to Esping-Andersen's (1990) distinction between the social democratic emphasis on the public provision of services and the Christian Democratic emphasis on cash transfers in welfare policy, with the former associated with greater wage equality than the latter.

<sup>20</sup>There are not enough degrees of freedom to add fixed country effects to the set of independent variables in Table 3.

Evidence on the impact of the supply of educated workers on wage equality is presented in Equations 3 and 4 of Table 3. Missing data force the dropping of the last time period which reduces the sample size from 41 to 26. More years of higher education (college and above) in the labor force is associated with greater wage inequality, controlling for mean years of education at all levels. In contrast, more years of education below the college level is associated with lower wage inequality. This pattern of results is consistent with a simple economic model of the labor market in which an increase in the education of the less-educated workers raises wages at the bottom, while the reduction of relative wages for workers at the top of the pay scale induced by centralized wage-setting lowers students' incentives to pursue higher education. By this argument, a reduction in average years of higher education is a consequence of wage compression and should be removed from the set of independent variables. Equation 4 in Table 3 indicates that the estimate of the impact of average years of education at all levels is sensitive to the inclusion of fixed period effects.

In sum, while independent variables like government employment, government spending, and the number of years of schooling appear as significant determinants of wage inequality in some specifications, only the level of wage-setting is robust in the sense that the estimated coefficient is stable and both statistically and substantively significant no matter what else is included on the right hand side of the regression equation. If we exclude fixed country effects, the estimate of the impact of concentration within confederations is also robust. Finally, if we include fixed period effects, the coverage of collective agreements is almost always significant. Moreover, there is a sense in which the coverage of union contracts must be important if collective bargaining institutions are important. Highly centralized bargaining that covered a vanishingly small fraction of the work force could not have a large impact on the aggregate wage distribution.

Table 4 illustrates how the estimates of the coefficients on the level of wage-setting, concentration within confederations and coverage change with different values of  $\lambda$ , the parameter for the speed of convergence to the equilibrium wage distribution. At  $\lambda = .95$ , a transitory shock to the wage distribution loses half its impact in 13.5 years. Increasing the value of  $\lambda$  to  $\lambda = .975$  doubles the half life of a transitory shock to 27 years while reducing  $\lambda$  to  $\lambda = .9$  cuts the half life of a transitory shock by roughly 50 percent to 6.5 years. Reducing  $\lambda$  further to  $\lambda = .8$  reduces the half life of a temporary shock to only three years. To compare the magnitudes of the estimated coefficients with different values of  $\lambda$ , the estimates are divided by  $(1-\lambda)$ , labeled "long-term coefficients" in Table 4. The long-term coefficient is equal to the eventual impact of a permanent one unit change in the independent variable on  $\ln [(w_{90,-} w_{10})/w_{10}]$ . The lower and upper bounds of the 95 percent confidence interval of the long-term coefficient are also reported. As Table 4 indicates,

**Table 4. Robustness of Results with Respect to  $\lambda$** 

	Coefficient	t-statistic	Long-term Coefficient	(Lower, Upper)
Level of Wage-Setting				
$\lambda = .8$	-.0522	6.52	-.261	(-.341, -.181)
$\lambda = .9$	-.0306	7.62	-.306	(-.386, -.226)
$\lambda = .95$	-.0182	7.03	-.364	(-.468, -.260)
$\lambda = .975$	-.0128	5.91	-.512	(-.688, -.336)
Concentration				
$\lambda = .8$	-.610	5.37	-3.05	(-4.19, -1.91)
$\lambda = .9$	-.313	5.51	-3.13	(-4.27, -1.99)
$\lambda = .95$	-.164	4.00	-3.28	(-4.92, -1.64)
$\lambda = .975$	-.101	2.83	-4.04	(-6.90, -1.18)
Coverage				
$\lambda = .8$	-.0775	2.51	-.388	(-.697, -.079)
$\lambda = .9$	-.0342	2.27	-.342	(-.644, -.040)
$\lambda = .95$	-.0200	2.00	-.400	(-.800, 0)
$\lambda = .975$	-.0124	1.49	-.496	(-1.160, .170)

**Notes:**

The dependent variable is  $\ln \{(w_{90} - w_{10})/w_{10}\}$ . GLS estimation with  $\lambda$  as indicated in the first column. All regressions include the three variables listed in the table, a constant term as described in the notes to Table 2 and fixed period effects. Concentration refers to concentration within confederations. The absolute value of the t-statistics are reported in the third column. The long-term coefficient is equal to the coefficient listed in the second column divided by  $(1 - \lambda)$ . The fifth column presents the 95 percent confidence interval for the long-term coefficient.

the main effect of increasing the assumed value of  $\lambda$  is to raise the estimate of the magnitude of the impact of the level of wage-setting and to reduce the precision of the estimates of the coefficients on concentration and coverage.

Finally, we test the argument of Rueda and Pontusson (1998) that the determinants of wage inequality differ in two sets of countries, the "social market economies" (SMEs) consisting of Austria, Germany, Finland, Denmark, Sweden, Norway, the Netherlands, and Belgium and the rest. Dividing the sample in two and redoing the estimates with a constant level of wage-setting, concentration within confederations, and coverage as independent variables, the F-statistic that tests the null hypothesis that the SME and nonSME samples are drawn from the same population is 1.43 with (4,33) degrees of freedom without fixed period effects and 1.77 with (7,27) degrees of freedom with fixed period effects. Neither statistic is significant at the 5 percent significance level. Thus, the null hypothesis that the institutional determinants of wage inequality are the same in the social market economies and the nonsocial market economies is not rejected by the data.

## 5. CHANGE OVER TIME

There is much greater variation in both wage-setting institutions and wage inequality between countries than over time within a twelve-year period. One implication is that the close relationship between institutions and pay inequality revealed in the statistical analysis is largely driven by the cross-sectional variation. Nevertheless, there is sufficient longitudinal variation for the impact of the level of wage-setting and the coverage of union contracts to be evident, even when cross-national differences are removed from the data using fixed country effects as shown in the last column of Table 2.

How well do the institutional variables highlighted in this study explain the changes that occurred in the wage distribution between 1980 and 1990 in the US and elsewhere? In the US, the 90/10 wage differential ( $(w_{90} - w_{10})/w_{10}$ ) increased by 15.7 percent between 1980 and 1992. The decline in coverage, from 26 percent in 1980 to 18 percent in 1990, can only explain an increase in the wage differential of about 1 percent. The fading effects of past periods of more centralized bargaining explains another 2 percent increase.<sup>21</sup> Together, the changes in wage-setting institutions only account for around 3/16 or 19 percent of the increased inequality in the US, an estimate that is very close to the conclusions of the single-country studies of the US discussed in the introduction.

The increase of wage inequality was even greater in Great Britain than in the US in the 1980s in percentage terms. Between 1980 and 1992, the British 90/10 wage differential rose from 1.79 to 2.31, an increase of 29 percent. Coverage declined sharply in the same period, from 70 percent to 47 percent, which can explain an increase of about 3 percent. More significant for the distribution of wages was the decentralization of wage-setting. From 1965 through 1979, wage-setting was frequently centralized in Britain through a series of incomes policies. According to the estimates in Table 2, the abandonment of income policies by the Conservative governments after 1980 and the decentralized bargaining that followed would be expected to

<sup>21</sup>The counterfactuals of this section were computed by calculating counterfactual values of the relevant independent variables and comparing the difference in the predicted values using the point estimates of Equation 6 in Table 2. To determine the counterfactual scores for coverage in the US and Britain, the cumulative value of coverage was recalculated assuming that no change in coverage occurred after 1980. To determine the importance of the declining impact of past episodes of more centralized bargaining in the US and Great Britain, the cumulative score of bargaining level in 1980 was converted into a constant annual bargaining level score by multiplying by  $(1-\lambda)$ . The constructed annual bargaining level score was then assumed to be the actual annual bargaining level score from 1980 through 1992. Finally, to determine the impact of changes in bargaining level in Italy, Norway and Sweden between 1980 and 1992, the cumulative score for bargaining level was recalculated under the assumption that no change in bargaining level occurred after 1980.



increase the 90/10 wage differential in 1992 by around 10 percent. Together, decentralization and declining coverage account for close to half of the increase in wage inequality that occurred.

Sweden, like Britain, witnessed a significant decline in the centralization of wage-setting, albeit from a much more centralized starting point in 1980. The famous Swedish pattern of highly centralized wage agreements negotiated by union and employers' confederations came to an end in 1983. In subsequent years, the role of the peak associations and government in wage-setting fluctuated from bargaining round to bargaining round, but centralized wage-setting with an industrial peace obligation was never reestablished during the sample period. Wage inequality did increase in Sweden, but less than would be predicted on the basis of the decentralization of wage-setting that occurred. According to the point estimates in Table 2, the decline in bargaining level in Sweden raised the expected wage differential by approximately 17 percent. According to the OECD data, the actual increase in the 90/10 wage differential was around 6 percent.<sup>22</sup>

Not all change in wage-setting institutions was in the direction of greater decentralization between 1980 and 1992. In Italy, government involvement in wage-setting increased in 1984 with the legislative enactment of the cost-of-living index. In Norway, centralized wage negotiations between the union and employer's confederations were resumed in 1988 after a period of less centralized bargaining in the early and mid-1980s. In both countries, the decline in wage inequality that occurred between 1980 and 1992 was close to what would be predicted by the increased centralization of wage-setting.

## 6. DISCUSSION

One can think of wage-setting institutions as varying along a continuum from bilateral negotiations between an individual employer and an individual employee to centralized negotiations covering the entire wage distribution conducted by elected representatives, whether the negotiators are office-holders in the union and employers' confederations or members of Parliament. The data strongly suggest that the more the wage schedule is determined collectively, whether the coordination is achieved by the explicit centralization of wage-setting or through the implicit cooperation of a small number of actors, the more egalitarian the distribution of pay. Collective choice of wages leads to greater wage equality than decentralized wage-setting.

<sup>22</sup>The data on blue-collar wages in Swedish manufacturing industries reported in Hibbs and Locking (1995) show a much larger increase in wage inequality after 1983 than the OECD data. See Iversen (1996), Pontusson and Swenson (1996), and Wallerstein and Golden (1997) for contrasting accounts of the change in bargaining institutions in Sweden and other Nordic countries in the 1980s.

It might appear that a connection between collective pay-setting and greater wage equality must exist, almost by definition, since a wage agreement covering a work force of any size must specify a general rule rather than a list of individual pay levels. However, it is easy to write general rules for pay raises that do not compress relative wages. For example, a collective agreement or legislative act stating that all wages should increase by  $x$  percent per year would preserve existing differentials. The strength of the relationship between collective pay-setting and relatively egalitarian outcomes should be seen as an important fact begging for an explanation.

There may be multiple reasons for the strong association of collective wage-setting with relatively egalitarian wage distribution. In fact, three different types of explanation can be distinguished: "economic" explanations that are based on considerations of economic efficiency, "political" explanations that refer to the way wage-setting institutions affect the relative influence of different groups of workers, and "ideological" explanations that point to the impact of wage-setting institutions on the application of widespread norms.

The economic explanations start from the premise that the wage differentials that emerge from decentralized interactions among employers and employees in the labor market are inefficient in some way. Consider, for example, an economy with decentralized wage-setting institutions in which strong unions exist in some industries but not in others or in some plants but not in others in the same industry. Even equally strong unions will differ in terms of the tradeoff they face between wage increases and employment that stem from differences in the productive process and the elasticity of demand for output. Among employers of nonunion labor, some firms may have substantial monopsony power while other firms have none. In such an economy, wages for equivalent workers in the unionized sector would differ according to workers' share of the monopoly rents, which vary across both industries and individual firms, while wages in the nonunionized sector would differ according to the monopsony power of employers.

Both in markets where workers' wages are higher than the competitive wage, due to monopoly power, and lower than the competitive wage, due to monopsony power, employment is inefficiently low and the relative price of output too high. There is both a misallocation of labor and a misalignment of prices. In this scenario, centralized wage-setting, by imposing a rule like equal pay for equal work, generates a wage distribution that may be closer to the textbook model of a perfectly competitive labor market than does decentralized wage-setting in actual markets. Although some workers and some firms would be worse off, aggregate income may be higher if local rents are reduced through a process of centralized wage-setting (Moene and Wallerstein 1997).

Another example of a model in which the wage differentials associated with decentralized wage-setting are inefficiently large is provided by the winner-take-all reward structure described by Rosen (1981) and Frank and Cook (1995). Winner-take-all markets are markets in which workers' rewards depend, at least in part, on their performance relative to other workers. Thus, the best musicians earn much more money than musicians with only slightly less talent, since, with modern audio technology, we can all listen to the best. The huge rewards that are obtained by the best musicians induce many to enter the competition, although logic dictates that almost all who compete to be best will fail. There is a social gain from recording the best musician rather than the second best, but the gain is only the possibly small difference in quality between the two. The private gain to being best, in contrast, may be huge. Thus, there is too much entry into winner-take-all markets with decentralized wage-setting. If, by centralized agreement, the prize from winning in winner-take-all markets were reduced, there would be fewer entrants in winner-take-all contests which could increase the efficiency of the allocation of labor in the economy as a whole.

There are many other reasons why the reduction of wage differentials might lead to lower rather than higher efficiency. Moreover, even in circumstances where greater equality is more efficient, an explanation in terms of efficiency is insufficient in the sense that a change that increases total income but not everyone's income may be blocked by those who would lose. Nevertheless, the possibility that wage compression within some range yields efficiency gains in some dimensions that offset efficiency losses in others may be an important part of the explanation of why institutions matter so much for the distribution of pay.<sup>23</sup> If there were a large, self-evident, economic cost from imposing changes in relative pay through collective processes of wage-setting, workers in countries with compressed wage scales would be paying a high price to satisfy their desire for greater equality. If pay inequality has little net effect on productivity within a wide range, however, institutions promoting wage compression could persist even in the absence of widespread willingness to accept lower incomes for the sake of greater equality.

The political explanation of the association of centralized wage-setting with egalitarian wage distributions is simply that centralization alters the influence of different groups in the wage-setting process. Freeman and Medoff (1984) argue that the wage structure in a nonunion labor market is shaped by

<sup>23</sup>As suggested by a study of the impact on productivity growth of wage compression in Sweden by Hibbs and Locking (1995), the reduction of wage differentials probably has multiple effects on economic performance, some beneficial and some not, such that the net effect varies over time and place.

the preferences and outside options of mobile workers who employers are trying to attract or retain while the structure of wages under collective bargaining reflects the preferences of the median voter in elections for union leadership or contract ratification. In general, a mean-preserving reduction in wage inequality will raise the pay received by the median wage-earner (and all other workers with wages below the mean) given the positively skewed shape of the wage distribution.

Moreover, as Moene and Wallerstein (1997) show, employers as well as low-wage workers may benefit from a wage policy that raises the wages of low-wage workers and lowers the wages of high-wage workers even when wage compression is inefficient. Moene and Wallerstein examine a model with heterogeneous employers and heterogeneous workers in which wage differentials arise from competition among employers to obtain more skilled employees. It is demonstrated that the wage differentials associated with the decentralized equilibrium are both efficient, in the sense that social surplus is maximized, and unjust, in the sense that differences in pay exceed differences in workers' abilities or efforts. In this model, reducing the wage differential between high and low-skilled workers increases both profits and the wages of low-wage workers as long as the average wage is kept low enough to clear the labor market. The possibility that employers can benefit from wage compression is important in understanding the history of centralized bargaining in Northern Europe. As Swenson's (1989, 1991) research has documented, the centralized wage-setting procedures in Scandinavia were created with the active support of the employers' associations.

The ideological explanation starts from the premise, well documented in the experimental literature, that people care about fairness as well as about their own income (Thaler 1989; Rabin 1998). The fact that people care about fairness does not, by itself, separate decentralized from collective decision making. Even with completely decentralized wage-setting, firms that disregard workers' concern with fairness when designing their pay policies suffer the consequences in terms of the morale and productivity of their work force.

Although concerns with fairness exist whatever the institutional environment, the centralization of wage-setting may have a large impact by altering how the norm of fairness is applied. In decentralized bargaining, the norm of equal sharing results in a wage that depends on the worker's usefulness to the firm and his or her alternative opportunities. In centralized pay-setting, the same norm of equal sharing results in pay that depends on the importance of the work force as a whole and their outside opportunities as a group. The larger the fraction of workers who are considered as a group in the wage-setting process, the more egalitarian the potential impact of applications of equal sharing rules. The association of pay equality and collective or political processes of wage determination may be due to the

way wage-setting institutions shape the application of norms of fairness as much as to the way wage-setting institutions affect the ability of different groups to pursue their self-interest.

## 7. CONCLUSION

The most important institutional factors in explaining the degree of wage inequality in advanced industrial societies are (a) whether wages are set locally, at the industry-level or at the level of the economy as a whole, (b) the extent to which wage-setting is dominated by a few large unions who are able to coordinate informally, and (c) the extent to which union contracts cover the labor force. The more wages and salaries are set in a centralized manner, the more egalitarian the distribution of wages and salaries.

This finding points to three lines of research, at least, that deserve further work. The first is to examine the economic costs and benefits of the reduction of wage differentials associated with centralized wage-setting. In particular, it is common to attribute a significant part of the high levels of unemployment in many European countries to the high wages received by relatively low-wage workers. However, some European countries with much more compressed wage scales than the US, such as the Netherlands and Norway, have unemployment rates as low as in the US. Thus, the nature of the connection between wage compression and macroeconomic performance is far from settled.

The second area for future research is to ascertain the generality of the egalitarian bias in centralized decision-making procedures. One would guess, for example, that public insurance programs are more egalitarian than private insurance or that even the earnings-related component of public pension systems is more egalitarian than private pensions. The third area for future research is to better understand the source of the egalitarian bias of collective decision making that is so strongly evident in the process of wage determination.

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## APPENDIX Data Sources

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All data used in this paper are available from the author in an Excel file upon request. Wage Differentials: Data for all countries except the US is from OECD (1996, 61–62).

The US data is from OECD (1993, 159–161) for 1980–1989 and from OECD (1996, 61–62) for 1990–1992. Data for the years 1980, 1986, and 1992 was used whenever

possible. In cases with missing data, I used data from the closest available year. Thus, the earliest year of data is 1981 for Canada and 1983 for Germany. For Austria and Norway, the 1986 data actually refer to 1987. The last data point is 1991 for Norway and Italy and 1990 for Denmark.

**Wage-setting institutions:** Data on the level of wage-setting and concentration within and between confederations are from the Golden, Wallerstein, and Lange data set on unions, employers' associations, and collective bargaining procedures for sixteen countries from 1950–1992. Concentration data are available only at five-year intervals. A complete series was created by assuming no change between 1990 and 1992 and constant linear change in each five-year period. For a few countries, missing data at the beginning of the series were filled in by extending the first available value back to 1950. The Golden, Wallerstein, and Lange data set is available on the web at [www.shelley.polisci.ucla.edu/data](http://www.shelley.polisci.ucla.edu/data). Selected data on the level of wage-setting, concentration within confederations, and coverage are presented in Table 5.

**Union density:** Data are from Visser (1996) for 1970–1992 and from Visser (1992) for 1950–1969. To make the two series compatible, Visser's (1992) figures were

**Table 5. Selected Data on Centralization, Concentration, and Coverage**

	Average Level 1950–92	Cumulative Level 1992	Average Concentration 1950–90	Cumulative Concentration 1992	Coverage 1990
Australia	2.02	36.2	.017	0.31	.80
Austria	1.09	18.3	.099	1.82	.71
Belgium	1.33	26.5	.100	1.73	.90
Canada	0.21	4.2	.034	0.68	.38
Denmark	2.58	39.9	.130	2.20	.74
Finland	2.02	33.6	.075	1.34	.95
France	1.02	18.0			.92
Germany	1.00	17.8	.144	2.72	.76
Great Britain	0.67	10.7	.049	0.87	.47
Italy	2.16	41.0	.080	1.34	.83
Japan	0.67	15.5	.095	1.52	.21
Netherlands	2.30	34.5	.136	2.65	.60
Norway	2.70	47.9	.069	1.60	.75
Sweden	2.53	42.8	.093	2.02	.83
Switzerland	1.00	17.8	.169	3.12	.43
United States	0.14	2.2	.022	0.40	.18

Notes:

"Level" is the level of wage-setting index. "Concentration" is the Herfindahl index of concentration within confederations ( $H_w$ ). Columns 1 and 3 present the average value over the period for which data was collected. Columns 2 and 4 present the cumulative value in 1992, defined by the formula  $X_i^*(1992) = \sum_{k=0}^{42} (.95)^k X_i(1992 - k)$  where  $X_i(t)$  is the value of variable  $X$  for country  $i$  in year  $t$ . Data sources are described in the appendix. Data for concentration within confederations is missing for France.

adjusted to remove unemployment union members from the numerator and unemployed workers from the denominator.

Coverage: Data are unadjusted coverage figures from Traxler (1994). Data are only available for 1980 and 1990. To create a complete series, the 1990 value was extended forward through 1992, the 1980 value was extended backward to 1950, and the figures between 1980 and 1990 were filled in through linear extrapolation.

Partisan composition of government: Data are from Duane Swank's Eighteen Nation Pooled Time Series Data Set. See Swank (1992).

Trade dependence: Data are from the Summers and Heston data set, described in Summers and Heston (1991).

The size of the public sector: Data on government outlays (at all levels) as a share of GDP are from the OECD, *Economic Outlook*, various years. Government outlay data begins in 1960. Data on public employment as a share of the work force are updated data originally published in Cusack, Notermans, and Rein (1989).

Stock of human capital: Data on the mean number of years of tertiary education and the mean number of years of schooling at all levels is from the Nehru, Swanson, and Dubey data set described in Nehru, Swanson, and Dubey (1995).

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