

**WAGE AND SALARY INEQUALITY  
IN U.S. MANUFACTURING, 1973-1981:  
THE EFFECTS OF ECONOMIC RESTRUCTURING**

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Vassar College Economics Working Paper # 6

January, 1990

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## I. INTRODUCTION

According to a variety of studies, the behavior of the distribution of income and earnings in the U.S. departed from its 20th Century tradition of stability and began to widen sometime in the late 1960s or early 1970s. By one account, the income share of the middle third of the earnings distribution decreased by 8 percentage points between 1960 and 1975 (Stanback and Noyelle, 1982), while new employment in the service sector became increasingly concentrated in the lower earnings tier.<sup>1</sup> Other studies have found changes in the distribution of earnings or of family incomes within the manufacturing as well as the service sector.

This anomalous trend has been attributed to a variety of factors, with the multiplicity of hypotheses to explain increasing inequality due in part to the fact that widening inequality seems to characterize both the distribution of wage and salary earnings and the distribution of total income, including transfer payments, for *both* families and individuals. Among the many hypotheses advanced to explain the problem are the increase in single parent household heads in tandem with an increasing number of two-earner households, rising mean education levels, increasing proportions of the work force made up of inexperienced workers in the 1970s, the shift from manufacturing to service employment, slowed productivity growth, and changes in labor relations and the strength of unions.<sup>2</sup> It has been argued in particular that these latter workplace phenomena have intensified a tendency toward increasing bifurcation in employment opportunities caused primarily by the restructuring of American industry.<sup>3</sup>

The purpose of this paper is to contribute to an understanding of changes in the distribution of

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<sup>1</sup>This result and others like it (see, for instance, Bluestone and Harrison, 1986), contributed to the "good jobs, bad jobs" debates of the 1988 presidential campaigns.

<sup>2</sup>Contributors to this literature are, for example, Henle and Ryscavage, 1980; Dooley and Gottschalk, 1982; Plotnick, 1982; Lawrence, 1985; Bluestone and Harrison, 1986; Bradbury, 1986; McMahon and Tschetter, 1986; Horrigan and Haugen, 1988; Grubb and Wilson, 1989.

<sup>3</sup>See Bluestone and Harrison, 1982 and 1988.

earnings with a particular emphasis on the role played by the restructuring of employment in the manufacturing sector of the economy. Since the spatial reorganization of employment has played a significant part in the discussion of restructuring (Bluestone and Harrison, 1982; Massey and Meegan, 1982), it is somewhat surprising that little attention has been paid to the possibility of a parallel spatial dimension to changes in the distribution of earnings. The research presented here will attempt to correct this omission in the literature by relating the regional dimensions of economic restructuring to factors influencing the distribution of earnings.

## II. ECONOMIC STRUCTURE AND THE DISTRIBUTION OF EARNINGS

It can be expected that economic restructuring may affect structural factors that underlie the distribution of earnings. Structural factors may be broken down into two general categories: 1) industry structure, which would include establishment or firm size, and market power; and 2) labor force structure, including unionization rates, racial composition, and occupational hierarchy.<sup>4</sup> Tests of these structural effects on inequality for the U.S. economy have generally been based on cross-sectional data across urban areas or other regional groupings such as states.<sup>5</sup>

How will the restructuring process affect these two categories? One approach suggests that labor force structure is affected by restructuring because labor's position is weakened by the cost-cutting and reorganizational efforts of corporations, which would be reflected in ineffective

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<sup>4</sup> A third category of structural influence on the distribution of income would be spatial structure, which might include such factors as urban size, agglomeration effects on productivity, or regional differences in industry mix (which frequently reduces in practice to differences in the percentage of manufacturing employment). A partial sampling of such research includes tests of the effects of the space economy by Long, Rasmussen, and Haworth (1977), Garofalo and Fogarty (1979), Hirsch (1982), and Kennedy and Nord (1984).

<sup>5</sup> Many of these models were inspired by the work of Kuznets (1955) and Kravis (1960) which suggested a negative relationship between economic development as measured by income levels and income inequality. Kravis in particular emphasized the effects of factors such as race discrimination and political and social forces on the distribution of income. See, for example, Al-Sammarie and Miller (1967), Aigner and Heins (1967), Jonish and Kau (1973), and Jacobs (1982).

or shrinking unions and a decompression of the wage structure due to increasing fragmentation in occupational hierarchies, for example. In addition, the effect of higher rates of unemployment in some areas resulting from corporate decisions to shift investment or employment to other locations would suggest a relative decline in wages among the less skilled workers whose employment is likely to be least secure. Consequently, restructuring may tend to increase the inequality of earnings.<sup>6</sup>

Industry structure may affect inequality indirectly through its relationship to labor force structure. Following such researchers as Bluestone and Harrison, one could argue that spatial reorganization of production disrupts existing relations between labor and management within firms as well as disrupting community cohesiveness and thus could lead to a worsening of earnings inequality within firms as well as within regions. In so far as industrial characteristics affect firms' incentives and abilities to restructure, they will indirectly affect the structure of the labor force and consequently inequality. Markusen (1985) argues, for instance, that high degrees of market concentration allow corporations to postpone internal reorganization of production which would otherwise be necessary to survive were the industry more competitive. Ultimately, however, oligopolistic industries facing competitive threats from such quarters as foreign producers, will eventually have to respond, and they then "make up for lost time with a vengeance" (Markusen, 1985, p. 47). If this is so, then the degree of disruption both within the firm and within local communities may be deeper and have greater consequences for changes in earnings structures than would otherwise have been the case.

The foregoing approach suggests that for a period such as the 1970s, which was marked by inter-sectoral and interregional (as well as international) shifts in employment and by the reorganization of several key manufacturing industries in the U.S., such as automobiles and

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<sup>6</sup>For an extended development of this argument, see Van Wagner (1989). This argument is related to Gordon, Edwards, and Reich's (1982) discussion of changing social structures of accumulation. The geographic dimensions of restructuring and its effects on earnings and conditions of work have been analyzed by Bluestone and Harrison (1982 and 1988) in the United States, and Massey and Meegan (1982) in the United Kingdom, as well as others.

steel, we would expect to be able to explain some part of increasing inequality in the distribution of wage and salary earnings by the restructuring process. A vector of characteristics could reflect the restructuring process as well as exercise direct effects on the distribution of earnings. For instance, restructuring may be accompanied by the following: alterations in occupational hierarchies; changes in the level and duration of local unemployment rates; and growing instability of employment, reflected in hours of work. The following is the specification and estimation of an empirical model of the relationship between inequality and restructuring which incorporates these variables.

### III. A MODEL OF EARNINGS INEQUALITY

In this section, a regression model of earnings inequality will be developed and estimated. The model is ultimately intended to test the effects of restructuring on the distribution of earnings, which, in practice, involves testing the impact of structural factors on the distribution of wages and salaries. As suggested by the discussion in Section II, restructuring may affect inequality through the impact of production and location decisions on the occupational and earnings structure within firms and industries as well as the impacts of these decisions on surrounding communities and regional labor markets. Therefore, the relevant level of a model of inequality that wishes to capture restructuring effects would be at once intra-industrial and intra-regional. The following discussion addresses both of these dimensions.

Section III.1 presents a very general discussion of the model which identifies the structural variables considered most important in the analysis of restructuring and inequality. Section III.2 describes the data used in the estimation of the model. Section III.3 contains the model specification, and the estimation is presented in Section III.4.

#### III.1 SPECIFYING THE DETERMINANTS OF EARNINGS INEQUALITY

In its most general form, the empirical model regresses levels of inequality within industries and regions on two vectors of explanatory variables, one containing observations on labor force characteristics and the other containing observations on characteristics of industries.<sup>7</sup>

The degree of strength of employee organizations in their dealings with management will clearly have an impact on the distribution of earnings. While union membership is an imperfect measure of employee strength, it is the one most appropriate to statistical analysis.<sup>8</sup> The effect of unionization on inequality in previous studies has not been entirely consistent (Ford, 1977; Hirsch, 1982; Hyclack, 1979). Although the presence of unions may tend both to increase wage levels and to compress their distribution, it is also possible that unionization may result in a greater differential between the earnings of workers in unionized and nonunionized industries; however, available evidence suggests that the "within-sector" effect dominates widened earnings differentials between the unionized and nonunionized sectors so that unionization is, on balance, an equalizing force (Freeman, 1980). Also, the presence of threat effects may result in higher wage levels in non-union sectors than would exist in the absence of unions. We generally expect that the latter effect dominates and that unionization is negatively related to inequality in the cross-section. In the context of restructuring, high relative levels of unionization may accelerate cost-cutting and confrontational efforts on the part of management and therefore be associated with rapid increases in inequality over time.

The set of variables measuring industrial structure would ideally include market concentration and the size of establishments, both of which capture to some degree the ability of firms to weather both conflicts with workers and competitive threats from other enterprises, as well as

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<sup>7</sup>An ideal model relating restructuring and inequality would place emphasis on changes in inequality over time. In this case, unfortunately, data constraints on the length of the time series available (to be discussed below) were such that the some of the dynamic elements of the model had to be sacrificed. The dependent variable is therefore the level of inequality by regional industry groups rather than changes in inequality.

<sup>8</sup>Work stoppages are perhaps a better measure since they indicate militancy, but data on work stoppages were not consistently available over the time period studied.

reflecting the degree to which the local economy is split into primary and secondary sectors. The expected effects of these measures of industry structure on inequality are not unambiguous in the literature. In general, it may be expected that dominant firms in oligopolistic industries will tend to pay higher wages because of high profits and the ability to pass off some portion of higher labor costs to consumers; the result may be a relative compression of wages within such industries (Dalton and Ford, 1977 and 1978; Jacobs, 1982). It may be, however, that oligopolies contract out ("outsource") to the competitive sector for various functions. Thus, on balance, concentration may increase the disparity between mean wages in the oligopolistic and competitive sectors. Overall, however, areas with a relatively large proportion of oligopolistic firms should have a lower proportion of earners in the low-paying, more competitive sector of the economy resulting in a lower intra-regional level of inequality, if earnings within concentrated industries tend to be clustered around the median (Jacobs, 1982). As suggested by the discussion of the profit cycle, earnings inequality in these industries may be prone to more rapid change over the period studied. Unfortunately, market concentration data are inappropriate to the regression model. Since the data vary little over time and are not available by region, market concentration would be almost perfectly collinear with industry.

Jacobs (1982) suggested that establishment size is a measure of the degree of efficiency and stability of enterprises. He argues that smaller size implies lower profitability due to inefficiency in production technologies and limited access to credit markets, resulting in lower wages to employees and consequently a greater differential between the bottom and middle of the earnings distribution, leading to greater overall inequality in areas with relatively high proportions of earners in small establishments. Alternatively, Podgursky's research (1986) supports the hypothesis that union threat effects may be smaller in smaller establishments. For these reasons, one would expect the proportion of workers employed in small establishments to exert a positive impact on inequality within a geographic area.

Macroeconomic activity may influence earnings inequality indirectly through the impetus that the progression of the business cycle gives to the restructuring process. Contractions in industrial output and associated downturns in overall economic activity may afford firms the opportunity to restructure by a reorganization of the labor process made possible by layoffs and slack capacity. If standardized production processes requiring less skilled labor are the most immediately affected by cost-cutting measures, during downturns the distribution of earnings *within* the industry may in fact improve as a consequence of layoffs occurring in the bottom earnings tier. Upswings should be associated with greater earnings inequality as the restructuring process bears fruit. In this case increases in output should be positively associated with intra-industrial inequality. An alternative hypothesis is that the creation of tight labor markets resulting from rapid growth would be expected to decrease inequality among earners on a cyclical basis (Blinder and Esaki, 1978).

Restructuring involves both *in situ* rationalization of production (that is, rationalization at existing sites) and employment shifts which originate both within the firm as production processes are relocated and within the industry. Changes in a region's share of national employment in a particular industry should therefore reflect the process of restructuring as well as exert an effect on inequality. The direction of this effect is difficult to predict *a priori*. If spatial reorganization of production leads to shifts in the location of low-wage oriented production processes from higher-wage, lower-inequality regions, then a declining share of industry employment within those regions should lead in the short run to lower inequality of earnings within the industry/region group since the lower tail of the distribution is effectively being removed. At the same time, unemployment resulting from industrial decline in a region should have a depressing effect on wages within the region across industries. In low-wage, high-inequality regions that are gaining employment share, the expansion of low-wage employment may lead to high levels of inequality within the industry; it is unlikely that gains in employment share would lead to lower levels of inequality in the absence of external forces to compress the wage distribution. On balance, increases in employment share should have a



positive effect on inequality.

Clearly, simultaneity can be expected between the determinants of inequality and shifts in employment if the factors that compress the earnings distribution also impose costs on firms that can be avoided through the restructuring process. If the restructuring process can be interpreted at least in part as an outgrowth of class conflict, (as in Gordon, et al., 1982), and if the level of inequality can be interpreted at least in part as an index of working class strength, then areas with low levels of inequality, high relative wages, and high relative union membership rates would tend to be more prone to restructuring through the reorganization of production and employment losses. As an indicator of restructuring, changes in employment share will be related to unionization rates as well as relative wage levels and pressure emanating from competition among firms, such as increased levels of import penetration. This issue has been addressed elsewhere.<sup>9</sup>

#### *Additional Hypotheses*

Characteristics of the labor force exert a considerable impact on inequality. Increased levels of education and experience in particular are associated with greater equality in the work force (except, under certain circumstances, in human capital theory). The neoclassical approach posits that education and experience are both contributors to marginal productivity, wage levels, decreased poverty rates, and decreased inequality. While rejecting the strictly marginalist approach of the neoclassical model, other research on labor markets also finds a negative relationship between education levels and experience on one hand, and inequality on the other (e.g. Al-Sammarie and Miller, 1967; Jonish and Kau, 1973; Ruthenberg and Stano, 1977).

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<sup>9</sup>See Van Wagner, 1989.

Discrimination within labor markets is also hypothesized to be a factor in inequality. Discrimination on the basis of both race and gender results both in a portion of the population receiving lower wages than white males in jobs requiring similar levels of training and experience, and in female and nonwhite workers being channeled into lower-skilled, less stable work. Reich (1981) found that inequality among whites is negatively related to racial inequality (black-white earnings ratios); greater discrimination therefore leads to greater inequality both between races and among whites. A similar relationship will be assumed to hold for discrimination based on sex.

### III.2 THE DATA

The data for this project were drawn from several sources. The bulk of the variables (earnings and labor force characteristics) were taken from the U.S. Census Bureau's annual March Current Population Survey (CPS). The CPS is designed to serve primarily as a source for cross-sectional analysis, or limited longitudinal analysis (i.e. over the course of the two-year cycles). As a result, surveys are not necessarily designed for consistency over time in either the coding of variables or the questions asked. Obviously, this makes time-series analysis on CPS data inconvenient.<sup>10</sup> To address this problem, CPS data files from the March surveys were transformed into a uniform format by Professor Robert Mare of the University of Wisconsin and it was these files which were the source for this empirical analysis. Industry and occupation categories had to be recoded according to comparability charts issued by the Bureau of the

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<sup>10</sup>There are, of course, other limitations to the CPS data. Questionnaires are administered to only one household member, who supplies information for all members of the household. Thus, distortions and inaccuracies may result. In addition, individuals may tend to overstate such things as occupational status, while deflating such things as income (for fear of the Internal Revenue Service) and understating age. And while the CPS sample is designed to mirror the characteristics of the population, the sample size necessarily makes estimations of small cross-sections inaccurate. The data are weighted according to population characteristics derived from the most recent decennial census, so in decades marked by demographic change (such as the 1970s) the weights may become more inaccurate as distance from the last census increases.

Census in order to attain inter-decadal continuity (U.S. Bureau of the Census, Census Comparability Charts C and D).

The independent variable for this study is Theil's index of inequality (Theil, 1967 and 1972). The Theil index is a statistical measure of entropy based on information theory and meets certain criteria for a measure of inequality that have been deemed crucial; these are scale invariance, the strong principle of transfers, and additive decomposability (Allison, 1978; Nygard and Sandstrom, 1981). The use of a scale invariant inequality measure eliminates the need for the price deflation of data, while preserving the distortionary effects of any non-proportionate changes in the distribution of income due to inflation. Conformity to the strong principle of transfers means that the Theil index will increase as income is transferred from low-income individuals to high-income individuals and will be more sensitive to the transfer of a given money amount at low income than at high incomes, reflecting the diminishing marginal utility of money.

The Theil index is defined as follows:

$$T = \frac{\frac{1}{n} \sum_{i=1}^n y_i \log y_i - \bar{Y} \log \bar{Y}}{\bar{Y}}$$

where  $y_i$  = individual income, and  $\bar{Y}$  = the overall mean. Its upper and lower bounds are 0 and  $\log n$ : 0 when all individuals receive the same income and  $\log n$  when one individual receives all the income.

The CPS is the primary source in the U.S. of unionization figures by industry affiliation and geographic region, but the question requesting union information appeared until 1981 in the May, rather than the March, survey so the union data were unavailable in their raw form for

this study.<sup>11</sup> There are two primary sources of published union data for recent years based on the CPS: Freeman and Medoff (1979), and Kokkelenberg and Sockell (1985). Freeman and Medoff provided only single estimates by selected categories for the period 1973-75. Kokkelenberg and Sockell provided annual or three-year-moving averages from 1973-1981. Neither of these published sources included unionization cross-classified by both region and industry, but Kokkelenberg and Sockell published separate one-way classification tables for these categories, as well as providing unionization estimates for race, sex, and education groups. The Kokkelenberg and Sockell figures were the ones selected for this analysis.

The unionization estimates should be interpreted with caution. The published union data contained estimates for unionization by industry based on three-year moving averages while the state data were presented annually. In order to match the two without further loss of annual observations, values were imputed for industries for 1973 and 1981. Also, the sample sizes in the CPS are such that multi-way disaggregation easily results in missing cell values, or in few observations per cell. Two problems arise from this, the first of which is that standard errors for unionization estimates in virtually any study using CPS union figures will be fairly high, a fact which is of particular concern in interpreting the results of studies which focus on small geographic units or population subgroups. The second is that unionization percentages for particular states or industries are frequently unstable over time when sampling results in only a few unweighted observations in each cell. Aggregation of state unionization rates to the nine Census Divisions ameliorated this problem. Finally, the unionization data presented in Kokkelenberg and Sockell were in percentage terms, although they included the unweighted sample frequencies. In order to aggregate from states to regions, the union percentages were transformed into frequencies using the appropriate weighted frequencies from the March CPS file. In theory, the weighted sample frequencies from month to month should result in approximately the same population estimates, but in practice, applying population estimates

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<sup>11</sup>After 1982, the CPS again collected information on unionization but the method of data collection changed. See Kokkelenberg and Sockell, p. 498, footnote 2.

from one month's sample to estimates from another's undoubtedly introduces measurement error.

The scope of this analysis was therefore limited by the availability of union data in two ways. The length of the time series, originally intended to cover the late 1960s through the mid 1980s (1968-1985) was limited to nine years, 1973-1981. The planned spatial detail of the analysis, which would have included 22 state groupings, was also limited because it was necessary to aggregate the data to Census Divisions.

Data availability also limited the empirical analysis to manufacturing industries. Information on industrial characteristics such as establishment size, market concentration, and international trade for non-manufacturing industries is scarce if not unavailable. The data for variables designed to capture industrial and structural characteristics were gathered from a number of government sources, including the *National Income and Product Accounts* and *County Business Patterns*. Industry data provided by government bureaus outside of the Census Bureau are coded according to the Standard Industrial Classification (SIC) system, and had to be recoded to conform with the Census classifications (CIC). According to government comparability charts, SIC and CIC manufacturing codes largely conform at the two-digit level.<sup>12</sup>

The data base for this study was created to include a variety of variables for each explanatory category (labor force, industrial characteristics, and the business cycle). For the labor force, a number of measures were constructed to reflect education, labor force experience, full- or part-time status, unemployment and unionization (discussed above) in addition to demographic variables such as sex and race. The education measure used was mean years of completed education. Labor force experience was calculated as the minimum of two alternate specifications, age minus years of education minus five and age minus fifteen, in order to avoid

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<sup>12</sup>See U.S. Bureau of the Census, 1980, *Classified Index of Industries and Occupations*.

ascribing years of experience to those who may have dropped out of school before being old enough to work. A variable to reflect hours worked per week or year was not included on the Mare CPS tapes, so full-time and part-time work was reflected by variables which measure the percentage of workers in the subgroup who were employed part-time for either the full year or for part of the year, or who were employed simply less than the full year. Unemployment was measured by the percentage of the group who were looking for or layed off from work for a substantial length of time (over 4 weeks ).

Discrimination on the basis of race and sex was measured by black/white and female/male earnings ratios, following Reich (1981, pp. 148-150) and Ford (1977). The alternative would have been the use of race and sex composition of the population as proxies for discrimination, which is the convention in *ad hoc* studies of inequality. However, these proxies introduced a high degree of collinearity into the estimation, particularly between the percentage of females in the labor force and part-time work. Additionally, while both the population composition and the wage ratio proxies for discrimination are conceptually flawed, it was not clear to this researcher that the presence of greater or lesser proportions of nonwhite or female workers in the labor force would in any way reflect greater or lesser degrees of discrimination. The population composition proxies may instead be indirectly measuring the relatively low earnings of nonwhites and women, which may be the result of a combination of discrimination in the labor market and other "premarket" factors such as the quality of education. The wage ratio proxies capture these forces more directly than the population composition does.

### III.3 ESTIMATING THE DETERMINANTS OF EARNINGS INEQUALITY

The theoretical approach discussed above suggests that the forces influencing earnings inequality operate in both spatial and industrial dimensions. Since the discussion of restructuring has a strong time dimension as well, estimating this model implies a pooled cross-

section and time-series regression equation which captures regional and industrial effects as well as effects over time.

The model to be estimated is a single equation regressing the level of inequality on groups of labor force, industrial, and regional characteristics across time. The model takes the following form:

$$T_{irt} = \alpha_0 + \beta_1 MEANED_{irt} + \beta_2 MEANEXP_{irt} + \beta_3 WRATIOS_{irt} + \beta_4 WRATIOR_{irt} + \beta_5 PCPT_{irt} + \beta_6 PCPY_{irt} + \beta_7 PCUN5_{irt} + \beta_8 PCWHCL_{irt} + \beta_9 PLT20_{irt} + \beta_{10} PCD3GNP_{irt} + \beta_{11} CHSH3_{irt-(t-3)} + \beta_{12} RIUN_{irt} + \epsilon_{irt}$$

Where:

- i, r, t = subscripts denoting industry, region, and time, respectively.
- T = the Theil index of inequality for each of 20 2-digit manufacturing industries in each of 9 Census Divisions.
- MEANED = mean years of education completed by the labor force for each of the 180 industry/region groups.
- MEANEXP = mean years of labor force experience as measured by the minimum of age - 15 or age - (years of completed education - 5).
- WRATIOS = female/male mean earnings ratios, designed to capture sex-based wage discrimination.
- WRATIOR = non-white/white earnings ratios, designed to capture race-based wage discrimination.
- PCPT = the percent of the labor force in each group employed part time.
- PCPY = the percent of the labor force in each group employed part year.
- PCUN5 = the percent of the labor force in each group laid off or looking for work for 5 weeks or more.
- PCWHCL = the percent of the labor force employed in white-collar occupations.
- PLT20 = the percent of the labor force employed in establishments with less than twenty employees.
- PCD3GNP = the percent change in industry output at the national level over the preceding three years.
- CHSH3 = the change in the share of U.S. employment for each

industry/region group over the preceding three years,  
calculated as  $(N_{it}/N_{uit}) - (N_{it-3}/N_{uit-3})$

UNION = the percentage of unionization of the labor force in each industry/region group.

As discussed above, five of these variables are expected to exert a negative effect on inequality: MEANED, MEANEXP, WRATIOS, WRATOR, and UNION. The rest are expected to have positive effects, although the expected effect of changing employment shares (CHSH3) is ambiguous. Both wage ratios are constructed with the mean wage of the group which is expected to suffer from discrimination in the numerator; therefore higher degrees of discrimination should be associated with a smaller wage ratio and higher levels of inequality. It has already been explained that the specification of unionization rates was problematic. Since unionization rates by industry/region groups were unavailable, the specification of unionization in this regression was limited to three choices: unionization by industry, by region, or by the multiplicative combination of the two. The inclusion of any of these requires strong and generally untenable *a priori* assumptions about the relationship between unionization at the industrial and the regional level. Including only unionization rates by region would require the assumption of much more uniformity in worker organization within regions than is likely to exist. Since regional unionization rates are calculated on the basis of the entire labor force, they also understate the degree of unionization in manufacturing. Similarly, including only industry unionization rates assumes uniformity of worker organization within industries across regions which is clearly not the case in the United States. From an econometric point of view both specifications of the union variable would result in much less variation in unionization than is likely to exist, biasing the estimated relationships and possibly introducing an undesirable degree of collinearity. The third specification, the product of industrial and regional unionization rates, is perhaps more appropriate to the actual variation of unionization rates by region and industry. However, this specification requires an assumption of independence between industrial and regional unionization which may not be particularly realistic.



Theory offers little guidance in choosing among suboptimal econometric specifications. Inclusion of the purely regional specification of unionization was rejected because it encompasses unionization in all industries and because it seemed likely to create the most pronounced collinearity problem. Separate regressions were estimated for each of the two remaining unionization specifications, PUNION (unionization by industry) and RIUN (the product of PUNION and regional unionization rates). Since PUNION is calculated on the basis of national industrial rates, the variable will underestimate unionization in industries in the traditionally highly unionized regions and overestimate it in traditionally weakly unionized regions. Regression coefficients for PUNION may tend to be biased as a result. RIUN, on the other hand, will tend to underestimate unionization in manufacturing industries in all regions given that worker organization can be expected to be higher in manufacturing than in non-manufacturing industries even in areas with low rates of union membership. This may also result in a bias in the RIUN coefficient.

An F-test was conducted on the model in order to determine whether it would be appropriate to pool the data into one equation, and the test indicated that this would be acceptable.<sup>13</sup> Consequently, a dummy variable model was used and the mean effect of industrial structure and other unobservable or unmeasurable variables is reflected in intercepts varying over both

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<sup>13</sup>The least restricted specification of this model would be a separate regression for each of the 180 industry/region groupings. However, constructing an F-test on this basis would have resulted in 180 equations with only 9 observations and 12 explanatory variables in each. Therefore, two "next best" specifications were developed for the unrestricted sums of squares: one set of equations for the 20 industries, and another for the 9 regions. Restrictions were then placed on the specifications by pooling the data and 1) including dummy variables for the 9 regions and 20 industries, and 2) excluding the dummy variables entirely. The estimation of the industry equations was complicated by the high number of missing values for the establishment size variable (PLT20), which resulted in matrices of less than full rank for some industries. The F-tests determined that differences among the restricted and unrestricted sets of equations were insignificant at conventional 95% levels of confidence in the absence of the PLT20 variable, with F ratios uniformly less than unity. The F value of the test comparing the pooled equation to the individual industry equations was a mere 0.5948. The F value for the comparison of the pooled dummy equation with the industry equations was an even lower 0.4390. In addition, further F tests were conducted to determine if there were significant differences between the restricted equations when PLT20 was included and when it was excluded. Since these tests also demonstrated insignificant differences between the specifications, and since it is unlikely that the missing variable would systematically change the statistical results, it was concluded that it was acceptable to include the variable in the final specification despite its absence from the test for pooling.

regions and industries.<sup>14</sup>

#### III.4 RESULTS OF THE ESTIMATION

Table 1 summarizes the results of the OLS estimation of the single equation model for both specifications of the unionization variable. The columns of the table list coefficients and t-statistics for the 27 industry and region dummy variables and the 12 explanatory variables.<sup>15</sup>

The overall fit of the model as shown by the F-values at the end of Table 1 is quite good, as these are significant at a .0001 confidence interval. The coefficient of determination of .53 for both equations is acceptable for a model that incorporates many cross-sectional relationships. For both models, the signs of most of the variable coefficients are as hypothesized, and most are significantly different from zero at the 90% confidence level or better. The wage discrimination variables, WRATIOS and WRATIOR are negative and highly significant, as are both PCPT and PCPY. The variables designed to capture restructuring, PCD3GNP and CHSH3, are positive and significant at the 95% level. A few variables do not perform as expected in both equations. MEANED and MEANEXP are insignificant and MEANED has the wrong sign. The coefficients of the union variables are also insignificant, although of the hypothesized sign, and RIUN performs considerably better than PUNION. RIUN is significant at the 80% level, while the coefficient of PUNION is not significantly different from zero; aside from these differences in the significance of the coefficients on the union variables, there is very little difference between the two equations. The establishment size variable, PLT20, is highly insignificant. In addition, unemployment (PCUN5) is significant and of the wrong sign.

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<sup>14</sup>In this inequality model, the unobservable effects cannot be assumed to be uncorrelated with explanatory variables; unobservable characteristics of both regions and industries are likely to be related to the characteristics of the relevant labor force. For this reason, the error components model is considered to be an inappropriate estimation technique (see Judge, et al., 1985).

<sup>15</sup>In this case, the excluded industry dummy variable is textiles, and the excluded regional variable is the South Atlantic.

The coefficients of the dummy variables indicate the marginal contribution of each variable to the overall intercept of the regression. The dummy variables which were necessarily excluded were chosen on the assumption that they would tend to represent a region (the South Atlantic) and an industry (textiles) which could be expected to have positive marginal contributions to the intercept term. The signs on the remaining dummy variables indicate the marginal effects of unobservable factors pertaining to each dummy. In both estimations, the industry dummies tend to follow a general pattern of positive marginal contributions for nondurables and negative contributions for durable goods, although there are significant exceptions in the case of petroleum and chemicals. In Model 2 (with RIUN) seven of the nineteen industry coefficients were significant at a 90% level of confidence or better, while five industry dummies were significant in Model 1. In Model 2, the regional dummies show positive marginal contributions for New England, the Middle Atlantic, the East North Central, and Pacific regions, with negative contributions elsewhere. Regional dummy signs differ from those in Model 2 for New England, the East North Central, and the Pacific Divisions. However, none of the regional dummy coefficients are significantly different from zero except for the East South Central in Model 1.

**TABLE 1**  
**Single equation regression results**  
**on industry/region Theil indices**  
**1973-1981**

Independent Variables	Model 1 (PUNION)	Model 2 (RIUN)
MEANED	0.002 (0.493)	0.002 (0.450)
MEANEXP	-0.001 (-1.312)	-0.001 (1.304)
WRATIOS	-0.103 (-9.693)***	-0.103 (-9.672)***
WRATIOR	-0.029 (-5.416)***	-0.029 (-5.418)***
PCPT	0.322 (7.425)***	0.322 (7.417)***
PCPY	0.269 (9.569)***	0.269 (9.647)***
PCUN5	-0.094 (-2.776)***	-0.096 (-2.859)***
PCWHCL	0.114 (4.937)***	0.112 (4.873)***
PLT20	0.000 (0.043)	0.000 (0.092)
PCD3GNP	0.030 (2.276)**	0.033 (2.477)**
CHSH3	0.148 (2.193)**	0.149 (2.221)**
PUNION	-0.029 (-0.398)	n.a.
RIUN	n.a.	-0.190 (-1.297)

*(continued on following page)*

Table 1 continued: Nondurable goods dummies

Independent Variables	Model 1 (PUNION)	Model 2 (RIUN)
Food	0.025 (0.737)	0.029 (1.099)
Tobacco	-0.004 (-0.173)	-0.001 (-0.074)
Apparel	0.035 (2.034)**	0.037 (2.894)**
Paper	-0.012 (-0.353)	-0.007 (-0.383)
Publishing	0.002 (0.098)	0.004 (0.247)
Chemicals	-0.037 (-1.894)*	-0.0134 (-2.331)**
Petroleum	-0.061 (-2.471)**	-0.057 (-3.338)***
Rubber	0.003 (0.164)	0.006 (0.468)
Leather	0.002 (0.152)	0.004 (0.436)

*(continued on following page)*

Table 1 continued: Durable goods dummies

Independent Variables	Model 1 (PUNION)	Model 2 (RIUN)
Lumber	0.023 (1.495)	0.025 (1.927)*
Furniture	-0.006 (-0.405)	-0.005 (-0.380)
Stone, Clay & Glass	-0.014 (-0.475)	-0.009 (-0.567)
Primary Metals	-0.029 (-0.774)	-0.024 (-1.219)
Fab. Metals	-0.024 (-0.519)	-0.009 (-0.610)
Machinery	-0.032 (-1.628)*	-0.029 (-2.115)**
Electrical Equip.	-0.006 (-0.313)	-0.004 (-0.262)
Trans. Equip.	-0.053 (-1.596)	-0.048 (-2.664)***
Instruments	-0.019 (-1.416)	-0.018 (-1.429)
Misc.	0.062 (4.337)***	0.064 (5.266)***

(continued on following page)

Table 1 continued: Regional Dummies

Independent Variables	Model 1 (PUNION)	Model 2 (RIUN)
New England	-0.001 (-0.098)	0.006 (0.650)
Middle Atlantic	0.002 (0.252)	0.015 (1.200)
E. N. Central	-0.000 (-0.065)	0.012 (1.004)
W. N. Central	-0.009 (-1.225)	-0.003 (-1.197)
W. S. Central	-0.007 (-0.999)	-0.007 (-1.081)
Mountain	-0.008 (-0.880)	-0.004 (-0.480)
Pacific	-0.004 (-0.558)	0.005 (0.512)
Adj. R <sup>2</sup>	0.53	0.53
R <sup>2</sup>	0.54	0.54
F-Value	37.56	37.64

*\*, \*\*, \*\*\* = .10, .05, .01 confidence intervals, respectively*

Interpreting the marginal effects of the dummy variables is difficult since we do not have full information on the unobservable effects. However, it is plausible to assume that industry dummies are capturing the differences in industrial concentration which could not be included directly in the model, as well as other industrial characteristics. If this is the case, then the negative contributions of durable goods and petro-chemicals to the intercept may reflect the offsetting effect of the greater degree of market concentration which tends to prevail in these groups of industries. However, aggregation of industrial sectors to 2-digit groupings may effectively dilute the effects of market concentration and the intercept terms may be reflecting some other characteristics of these industries not captured directly by the explanatory variables.

The results tend to support the hypothesis that restructuring of the economy, as represented by employment shifts, has had a significant effect on inequality within industrial groups over this period. The results also indicate that the inclusion of structural variables such as part-time work and discrimination tend to reduce the importance of education and experience in the determination of the distribution of earnings. Podgursky's (1986) hypothesis of the effects of establishment size on inequality within a region is not supported here. However, the results of the estimation must be interpreted in light of the econometric problems presented by the data. The Durbin-Watson  $d$  statistic of 1.88 suggests an absence of autocorrelation when there are 39 regressors and 1265 observations. However, there remain problems of collinearity among some of the variables. Condition indices and eigenvalues calculated for these equations indicate a high degree of collinearity between the union variables and education, which could be responsible for the poor results for both variables. There may be additional problems in the data due to measurement error, which would result in a downward bias of the  $\beta$  coefficients (Johnston, 1984, pp. 428-430). As mentioned above, the sample size and sampling technique of the CPS may lead to inconsistencies over time and across units when the data are cross-tabulated. Some sectors in some regions occasionally had very few unweighted observations leading to unreliable population estimates. In addition, the establishment size data provided by



*County Business Patterns* contained a large number of missing values due to disclosure rules. This may have affected the estimation of the relationships for some industries over time, a problem which was indicated by the inability to obtain full-rank matrices for regionally-concentrated industries such as tobacco. And while the labor force experience variable used here is a better estimate of experience than the frequently-used mean age, estimation, it is still a very rough approximation. Measurement error in the unionization variable has already been discussed. Another source of bias may arise from aggregation; Census Divisions are quite large and mask a great deal of geographic diversity, while two-digit level aggregation of manufacturing industries may overstate the similarity among industrial groupings. The aggregation problem is, of course, one that consistently haunts studies of both geographical and industrial economic behavior. Finally, specification error may be present because of the absence of industrial concentration from the estimation.

The negative sign for the unemployment variable remains puzzling nonetheless. One would not expect unemployment to negatively affect inequality. Other specifications of this model used in the F-test for pooling show that the unemployment variable is not robust. When regressions were run separately for Census Divisions and industries, the sign and significance level of PCUN5 was inconsistent across estimations. It is possible, however, that including the effects of employment shifts, changes in output, and industry and region effects in the pooled dummy variable model removes inequality-exacerbating effects from the unemployment variable. Or, it may be that the negative effect of unemployment on inequality may be capturing a skill or wage-level bias in industry-specific regional unemployment. As mentioned in the discussion of the possible effects of changes in industrial output and employment share above, the immediate effects of some aspects of restructuring may be to narrow intra-industry earnings inequality rather than widen it if unemployment results in a removal of the bottom tier of the earnings distribution within an industry.

With the caveats to interpretation of this model in mind, however, the hypothesized relationships between restructuring and inequality appear to be supported by the estimation of the single-equation pooled model. Employment shifts and changes in output are both significantly and positively related to inequality. In addition, the hypothesis that the effects of restructuring are felt most strongly at the bottom tier of the earnings distribution is weakly supported given the interpretation of the relationships among output changes, employment shifts, and unemployment.

#### IV. Conclusions

This paper has attempted to establish a relationship between economic restructuring and sources of earnings inequality in the United States during the 1970s. We argued that the analysis of the trend toward a greater dispersion of earnings discovered over the period spanning the 1960s to the 1980s should take into account the structural changes within economic sectors rather than focusing exclusively on demographic factors or intersectoral employment shifts. The reorganization of production and investment over space carries with it implications for such factors as occupational composition, unemployment, unionization, and other forces influencing earnings distributions.

In the empirical model, variations in inequality among regional industrial groupings are partially explained by variables suggested in the theoretical analysis of restructuring, specifically, changes in industry output and in regional shares of industrial employment. In addition, labor force and employment conditions such as part-time or part-year work, wage differentials between whites and blacks and between men and women, and the occupational status of the labor force are significant in explaining variations in inequality in the pooled model, while education levels are not.

Unfortunately, it was not feasible to conduct a full test of the restructuring hypothesis due to data limitations. Factors such as industrial concentration, the long-term effects of economic structural change, and the accelerating influence of cyclical fluctuations could not be adequately addressed by this model. More seriously, the length of the time series precluded estimating the impacts of the independent variables on *changes* in inequality, which are of more immediate concern than the level of inequality. This is a subject for further research. Further research may also address some potential problems caused by possible simultaneity within the single equation model.<sup>16</sup>

The important question which cannot be answered here is whether the inequality patterns that have been evident during the period studied are temporal deviations from a long-term trend toward stable, if not narrowing, earnings differentials within the population, or whether the underlying distribution has suffered a long-term shift toward greater inequality. The former argument would hold if observed widening of the distribution is due primarily to demographic factors or even due to the process of shifting employment from manufacturing to services. The latter argument, that the distribution of earnings has become structurally more unequal, is one that deserves further study and that requires a more extensive time series. On the basis of the results presented here however, we are hopeful that further study of this issue will recognize the importance of shifts in the location of employment in the determination of differentials in earnings inequality.

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<sup>16</sup>This issue has been addressed in part in Van Wagner, 1989.

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