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The Impact of Regional Difference in Unionism on Employment

by Edward Montgomery

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Introduction

Almost 20 percent of the people in the work force are union members. Just in terms of numbers, trade unions are an important influence in the labor market and in the U.S. economy. Further, unions are widely believed to play a major role in determining workers' standard of living and how work is done and in affecting firms' profitability. Freeman and Medoff (1984) recently presented evidence suggesting that unions affect labor markets in a variety of ways. The beneficial effects of unions include protection for older workers, reduced quit rates, reduced earnings inequality, and increased productivity. Unions might adversely affect profits and stock prices and might increase the number of workers laid off in cyclical downturns, as well.

Although the impact of unions on these measures of economic performance has been studied, the majority of research on unions concerns how they affect compensation. Freeman and Medoff (1984) show that unions increase fringe benefits, and there is a large body of empirical evidence that suggests unions raise the relative wages of their members.[1] In addition, unions have been found to affect the wages of nonunion members, although the direction and magnitude of this effect is ambiguous. Despite the attention focused on how unions affect wages, little attention has been paid to how this change in the relative cost of unionized labor affects employment—clearly an important part of assessing the welfare costs and benefits of unionism.[2] (By "welfare costs," we mean social or aggregate costs and not simply private costs and benefits to union members.) If unions succeed in raising wages only at the cost of massive employment reductions, as some analysts believe is the case, the welfare implications are radically different than if wage increases could be achieved with little or no impact on aggregate employment.

This study examines whether changes in unionism affect the aggregate level of employment in the economy, and in particular, whether an individual who lives in a standard metropolitan statistical area (SMSA) where unions are rare or weak is more likely to be employed than an individual who lives in an area where unions are strong.

Whether or not unions have a harmful effect on employment is also important to analysts of regional unemployment differences. Murphy (1985), found that differences in sensitivity to demand conditions in the product market and in wage differentials are vital in determining regional differences in unemployment rates. Since unions have been found to affect both of these variables, differences in the extent or impact of unionism could be important in understanding regional unemployment rate differentials.

1 See Parsley (1980) for a review of this voluminous literature.

2 There have been studies of the relative wage effect of unions across industries, occupations, and race and gender groups.

In fact, Freeman and Medoff's study (1984) suggests that unemployment rates are 1.0 percent higher in areas with a high degree of unionism relative to low unionism areas. However, since they also fail to find any correlation between the degree of unionism and the employment rate, a further, more explicit analysis of this question seems to be necessary to determine what effect, if any, unions have on aggregate and regional employment rates.

I. Previous Literature

Most studies of the employment effects of unions have been on the industry level.[3] Industry or firm studies, however, may overestimate the disemployment effect of unions, because they ignore the fact that some or all of the displaced workers may become re-employed in other industries or firms. Consequently, these studies cannot provide estimates of the net or aggregate employment effect of unions.

Lewis (1963 and 1964) provided the first analysis of the relative wage and employment effects of unions on an aggregate basis. Lewis divides the economy into a union and a nonunion sector. Industries with a relatively high degree of unionism, like manufacturing and mining, are part of the unionized sector, while those with a low degree of unionism are part of the nonunion sector.[4] Using time series data, Lewis estimates whether changes in relative employment levels across these two sectors can be attributed to differences in the average union/nonunion wage premium and to the average percent unionized. His results suggest that unions have a significant negative effect on relative employment levels and man-hours worked.

Pencavel and Hartsog (1984) recently updated and extended this seminal work. They failed, however, to find any consistent negative impact of unionism on man-hours. In fact, they conclude that the hypothesis that unionism depresses man-hours can be accepted only for the late 1920s and early 1930s. This basic result is not sensitive to whether the employment and wage effects of unions are estimated with Lewis' reduced form model or with a structural model that they developed.[5]

These results might be ambiguous because aggregate data are not suited to testing the employment effects of unionism. Aggregating industries into two sectors ignores the effects of unions *within* these sectors and, thus, may not yield good estimates of the overall effect of unions on employment and wages. Further, the absence of controls for changes in labor quality across sectors means that these studies might overestimate the impact of unions on wages and underestimate the effects on employment. In other words, if firms respond to the union wage demands by hiring for higher-quality labor, then "quality-adjusted" wages will not rise as much as measured wages.[6] Since firms may substitute skilled for unskilled workers, the effect on total demand for labor could differ from the effect on a particular type of labor.[7]

Kahn (1978), Kahn and Morimune (1979), and Holzer (1982) provide cross-section estimates of the effects of variations in the extent of union membership across SMSAs on employment, hours worked, and unemployment stability. In these cross-section studies, the fraction of employed workers in an SMSA who are union members is used as a measure of union strength, because it is believed that unionism affects all workers in the same labor market, not just those in the same industry. Workers who may be displaced because of union wage demands are likely to seek employment not just in that industry, but throughout the local labor market. Studies with detailed cross-section data, either from the Current Population Survey (CPS) or the Survey of Economic Opportunity (SEO), offer better control for individual characteristics and for labor market variables that affect employment. These cross-section studies avoid some of the aggregation problems that crop up in aggregate time series studies, and thus, are preferable.

Nevertheless, results of these cross-section studies are somewhat inconclusive. Kahn (1978) finds that annual hours worked are significantly reduced for nonunion females, but not for nonunion males; these effects did not differ by race. Holzer (1982),

3 See Lewis (1963) for a review of some of these industry studies.

4 The union sector was made up of mining, construction, manufacturing, transportation, communication, and public utilities; the nonunion sector was made up of all others, except military and government relief.

5 The structural model of the labor market that is used by Pencavel and Hartsog (1984) was developed to test for the wage and employment effects of unions without assuming that employment is unilaterally set by employers or that the union wage premium is exogenous. It should also be noted that their model also differs from that estimated by Lewis (1964) in that they use only the percent organized variable to capture the effect of unionism and not the estimated union wage premium.

6 The potential importance of these biases can be seen by the fact that the estimates of the quality-adjusted union relative wage effect differ substantially from those derived in cross-section studies.

7 See Pencavel and Hartsog (1984, p. 216) for a further discussion of these limitations.

reduction in supply in the nonunion sector that results from the drop in wages.

It can be shown that in a two-sector model with constant factor intensities, the changes in nonunion wages will be a function of the elasticity of labor supply, *e*, the elasticities of labor demand in the union, η_u, and nonunion sectors, η_n, the percent unionized, *k*, and the change in union wages, w_u.[12] Thus:

$$(4) \quad w_n = \frac{-k\ (n_u - \epsilon)\ \dot{w}_u}{(n_u - \epsilon)\ [(1 - k) + \epsilon w_u] + k\epsilon\ (n_u - n_n)\ \dot{w}_u}$$

From equation (4) we see that unless the elasticity of labor supply is zero ($\epsilon = 0$), nonunion wages will not fall enough to prevent average wages from rising and total employment from falling. Falling wages in the nonunion sector cause workers with high reservation wages to withdraw from the labor force, thus causing total employment to decline.[13] Since previous research has found that unions tend to organize industries where the elasticity of labor demand is low, it is interesting to note that the greater the elasticity of labor demand in the nonunion sector relative to the union sector, the smaller the drop in nonunion wages, and the smaller the aggregate employment loss.[14] Using equations (1), (2), and (4), we can express the change in total employment as a function of the union wage change:

$$(5) \quad \frac{\partial E_T}{E_T} = \frac{1}{\mathrm{A}} \left\{ [\eta_u k\ \dot{w}_u]\ A\ -\ [(1\ -\ k)\ \eta_n k\ \dot{w}_u\ (\eta_u\ -\ \epsilon)] \right\}$$

where

$$A = (\eta_u\ -\ \epsilon)\ [(1\ -\ k)\ +\ \epsilon\ \dot{w}_u] + k\epsilon\ (\eta_u\ -\ n_n)\dot{w}_u.$$

The higher the elasticity of supply, ϵ, or elasticity of demand in the union sector, η_u, or the greater the percent organized, *k*, the greater the disemployment effect associated with an increase in union wages. As the percent organized rises, more workers are in the union sector, and hence, are affected by the increase in union wages. However, if labor supply is inelastic, total employment will remain fixed.

In a general equilibrium model with variable factor intensities, the effect of unions on wages in the nonunion sector, and hence total employment, is ambiguous. If the unionized sector is the intensive sector then, as shown in Johnson and Mieszkowski (1979), both the substitution and the scale effect will result in a reduced capital/labor ratio in the nonunion sector, and hence, a reduction in the marginal product of labor and wages.

However, with a capital-intensive unionized sector, nonunion workers will get higher wages if the scale effect is greater than the substitution effect and lower wages if the converse is true. In either case, increases in union wages or in the percent of the labor force that is unionized tends to be associated with an increase in average wages and a drop in total employment, as long as labor supply is not completely inelastic.

The theoretical models discussed in this section imply that increase in either the percent unionized or in the union/nonunion wage differential can lead to a reduction in aggregate employment. The size of the disemployment effect will depend, in part, upon the elasticity of labor supply, where the more elastic the supply, the greater the reduction in employment. As seen in equation (5), the employment effect of unionism depends upon the extent of union strength, which is a function of both the union wage premium and the percent of the work force receiving it. Based on this theory, we would expect an inverse relationship between union strength and employment. We would also expect this effect to be small, if the elasticity of labor supply is near zero.

III. Empirical Results

To test for the employment and unemployment effects of unions, we used data from the 1983 Current Population Survey (CPS) Earnings File and Census data on SMSA characteristics. This data set was chosen, in part, because it contains detailed personal characteristics for each respondent, which allow us to control for differences in worker quality. In addition, it contains earnings and union membership data across individuals in each SMSA. To ensure a

12 See Welch (1974, p. 304, equation [6]), for derivation of a similar result under the assumption that demand elasticities do not vary across sectors.

13 It is possible that the existence of a union wage premium may actually draw more workers into the labor force than exit because of the depressed nonunion wage rate. This will occur, however, only if the turnover rate exceeds the elasticity of demand for labor. As noted earlier, this condition is unlikely to hold in the union sector.

14 See Freeman and Medoff (1984).

sufficient sample size in each of the 44 SMSAs in our sample, we combined the survey responses for each month over the year, yielding a sample of 104,409 observations.[15]

To examine the disemployment effect of unions, we initially looked at the effect of unionism on the probability of an individual in the population being employed. Because displaced workers from the unionized sector may either become unemployed or withdraw from the labor force, the employment and unemployment effects of unionism need not be the same. Since the distinction between unemployed and not-in-the-labor-force may not be pronounced, and since some of those displaced by unions may withdraw from the labor force, the probability of being employed might be a better measure of the "true" disemployment effect of unionism than the probability of being counted as unemployed. An additional benefit from focusing on employment status is that we can examine whether unionism has a different effect on the likelihood of getting part-time work than on getting full-time work. These effects may differ substantially if unionism affects the length of the workweek for those who remain employed.

As shown in section II, the effect of unionism on employment is a function of both the percent organized and the union wage premium. Consequently, the measure of the effect of unionism that we used is the product of the percent of employment in an SMSA that is unionized and the union/nonunion wage differential.[16] This index is similar to the Kaitz index, which is widely used to examine potential disemployment effects of a legislated minimum wage increase. It appears that unions impact aggregate employment via their effect on the average cost of labor. The distortion in labor costs due to unionism is the change in wages—that is, the union wage premium times the number of workers who receive that wage.[17]

Previous cross-section work by Holzer (1982), Kahn and Morimune (1979), and Kahn (1978) has implicitly limited the effect of unions on employment to differences in the percent organized from SMSA to SMSA . This is like constraining the union relative wage effect to be the same across SMSAs, which may be inappropriate for theoretical and econometric reasons.

Recent theoretical work by Lazear (1983) suggests that the percent unionized in an industry or region is not a good measure of union power. He shows that to the degree the cost of running a union differs across industries, different wage/employment packages are negotiated by unions facing the same opportunity locus or having the same strength. That is, unions in industries where costs are high tend to prefer higher wage/lower employment share packages than unions in relatively low-cost markets. Consequently, the percent of employment that is unionized or the union wage premium varies across industries or regions, even though union power is the same.

Greater union strength is indicated by a better wage/employment share package, not just a higher percent unionized. Consequently, it is necessary to control for both the wage premium and the percent unionized to get a measure of union strength across markets. To the degree the union relative wage effect differs across SMSAs, failure to control for differences in the wage premium will yield inefficient and potentially biased estimates. Since the union wage premium may be determined by many of the same exogenous variables that determine employment, this term is likely to be correlated with the independent variables in the model. The result may indicate that the estimated coefficients in previous studies are biased.

To construct our measure of union strength, it was first necessary to derive an estimate of the union/nonunion wage differential in each SMSA. To do this, we estimated separate wage equations for union and nonunion members in each SMSA:

$$(6) \quad \ln W_{ik} = \beta X_{ik} + e_i$$

where W_{ik} is average hourly earnings of individual, i, in SMSA, k, X_{ik} is a vector of individual characteristics that determine wages, and e_i is an error term. In estimating these wage equations, we included controls for schooling, experience,

15 Beginning in 1981, the CPS reduced the number of surveyed individuals and asked detailed employment questions of only one-quarter of the sample each month. As a result, there were fewer than 30 union members in many of the SMSAs in any given month.

16 We restrict our sample to the nonfarm economy when calculating both the union wage premium and the percent of employed who are union members. The sample was restricted to civilians age 16 to 65, working for wages and salary.

17 Because the multiplicative form places strong restrictions on how the percent organized, k, and the union wage premium, z, affect employment, we also estimated our employment equations using several other constructions of the union strength variable. In particular, we estimated an equation where these terms were entered separately and equations with multiplicative indexes that rise more than proportionately with changes in the percent unionized ($zk/(1 - K)$) or with the union wage premium (z^2k). Because of their qualitative nature, our results were not sensitive to the use of these other indexes.

Thus, the fraction of the population employed in an SMSA is inversely related to the extent of unionism and to the union wage premium. The magnitude of this effect can be captured by calculating the change in the probability of being employed for a base case or average worker when the value of the union strength variable changes by one standard deviation from its mean value.[22] The expected probability of being employed declines from 0.829 to 0.825 with this increase in union strength. On the other hand, the probability of the average worker in the SMSA where union strength is highest (San Bernardino, CA) being employed is only about 2 percent less than it is if that worker lived in the SMSA where union strength is the least (Atlanta, GA).[23] Thus, it would appear that changes in the extent of union strength have only a very limited impact on aggregate employment.

Given this reduction in the probability of gaining employment due to unionism, it is of interest to see if unionism also affects the length of the workweek for those who remain employed. If unionism has no effect on hours worked, then the effect on the probability of working full time should be the same as it is on the likelihood of working part time. Conversely, if employers cut their employees' hours, then the union variable should be positive in a regression where the dependent variable is the probability of working part time regression and negative in a regression where the dependent variable is probability of working full time. In regression (2) the dependent variable equals 1 if an individual is employed full time and zero otherwise; in regression (3) the dependent variable equals 1 if an individual is employed part time and zero otherwise.

We found that the union variable was negative and significant in the full-time employment equation, while it was positive but insignificant in the part-time employment equation. In addition, both the point estimate and the degree of significance of the union strength variable are higher in the full-time equation than in the total employment equation. Using these estimated coefficients, a standard deviation increase in union strength leads to a 0.7 percent reduction in the probability of being employed full time and a 1.5 percent increase in the probability of being employed part time.[24] If our base-case worker lived in Cleveland, he would be approximately 2 percent less likely to be working full time, and 4 percent more likely to be working part time than if he lived in the lowest union strength SMSA. Thus, these results suggest that part of the disemployment effect of unions comes through reducing the number of hours worked on that job.

As a further test of this hypothesis, we re-estimated the employment equation with the probability of working part time if an individual was employed as the dependent variable. Unions may reduce the workweek by increasing the relative frequency of part-time jobs. As seen in regression (4), increases in union strength increase the fraction of employment that is part time. A standard deviation increase in union strength increases the likelihood of working part time for the base-case worker by about 3 percent.[25] Given these estimates, the conditional probability that an average worker has a full-time job (as opposed to a part-time job) is about 8 percent less in the Cleveland SMSA than in the lowest union strength SMSA. Thus, these estimates suggest that increases in union wages (or the percent organized) might have a bigger effect on hours worked per week or on the mix of full-time and part-time jobs than on the level of total employment. This shift toward more part-time jobs may occur because unionized workers are more likely to work full time than nonunion workers, and because unionized workers are more likely to accept layoffs than reduced hours.[26] Thus, an increase in the cost of union labor will primarily cause a reduction in the number of full-time jobs in the union sector, because unionized workers tend not to engage in work-sharing arrangements to reduce hours worked. Some of the displaced workers, however, will find employment in the nonunion sector where there are more part-time jobs. Employment will thus tend to fall by less than the drop in the number of full-time jobs.

In section II, it was shown that the disemployment effect of unions was a function of the elasticity of labor supply. The greater the elasticity of supply, the greater the disemploy-

22 The base-case worker is a single white male with 12.6 years of schooling, 18.5 years of experience who lives in the East-North-Central region of the United States in an SMSA with an unemployment rate of 9.4 percent in March, a population of 3,479,000 where 5.5 percent of the population receives AFDC, and the union strength variable equals 0.031.

23 The union strength variable equals 0.0367 in Cleveland and -0.0016 in Atlanta. In Cleveland, the probability of being employed is 0.827, while it is 0.837 in Atlanta.

24 The probability of being employed full time and part time for our base-case workers is 0.707 and 0.104, respectively.

25 The probability that the job a worker has is a part-time one for the base-case worker is 0.1429.

26 See Freeman and Medoff (1984) for a discussion of this issue.

ment effect. Given this, we might expect that the disemployment effect would be largest for groups with a weak labor force attachment or a high elasticity of labor supply. Teen-agers or young people may be more adversely affected than older workers, and females may suffer more than males. To test for differences in the disemployment effect across groups, we estimated separate employment equations for part-time and full-time workers by gender and age group. These results are presented in *appendix II.*

The basic predictions of our theory seem to hold. Based on the point estimates from these regressions, we see that the disemployment effect of unions is smaller for prime-age males than for teen-agers or 20-to 24-year-old males. In fact, prime-age males do not appear to be adversely affected by changes in union strength at all. This probably reflects their strong labor force attachment or the low elasticity of labor supply. Interestingly, the evidence does not support the hypothesis that teen-agers are more adversely affected than 20-to 24-year-olds. As expected, the disemployment effect of unionism is greater for prime-age females than for prime age males.[27] In general, increases in either the union wage premium or the percent organized affect the workweek, or the likelihood of being employed part time, more for females than for males.

IV. Conclusions and Implications

Results of estimates of the effect of changes in union strength on the likelihood of being employed are presented here. They suggest that in areas where the unionized percent of the labor force is large, or where the union/nonunion wage premium is large, workers are less likely to be employed. Besides affecting the number of workers employed, unions reduce the likelihood of an individual having a full-time job by altering the mix of part-time and full-time jobs in the economy. Thus, unions appear to adversely affect the average workweek for those who remained employed. These disemployment effects are felt mainly by females and young men, with little, if any, negative impact on prime-age males.

This disemployment effect was quite small, however. Unionism has a larger effect on the mix of part-time and full-time employment (and hence the workweek) than on the number of jobs. All of these effects are dwarfed in importance by other factors: the state of the local labor market and the level of the individual's human capital, or skills. Changes in schooling, experience, and local labor market conditions have a much greater impact on the likelihood of being employed than does unionism. For instance, a standard deviation increase in the number of years of schooling increases the likelihood of being employed for the base-case worker about 10.6 percent, while a standard deviation increase in the number of years of potential labor market experience increases it by 36.6 percent.[28] Thus, a standard deviation change in these measures of human capital is approximately 10 to 30 times more important than a similar change in union strength. This result implies that differences in union wage differentials, or the percent organized, are not the primary cause of regional differences in employment rates.

Data Appendix

The data for this study come from the Current Population Survey 1983 and from the Bureau of Census, *County and City Data Book, 1982.*

UN is the product of the percent unionized and the union wage premium in each SMSA.

Unemployment Rate is the local unemployment rate for all workers in the SMSA.

Population is the number of people living in the SMSA.

AFDC is the proportion of the population in the SMSA receiving AFDC payments.

Schooling is the number of years of schooling completed by the individual.

Experience is calculated as Age -Schooling -6.

Race is a dummy that equals 1 if the individual is white.

Sex is a dummy that equals 1 if the individual is a male.

In addition to these variables, each regression contains a dummy term that equals 1 if the individual is married, nine regional dummies where the omitted catagory is the East-North-Central region and 11 monthly dummies to control for the month the individual was surveyed. The complete regression results are available from the author upon request.

27 The adverse effect of unionism increases with age for females. Whether this reflects a greater attachment to the labor force is a question for further research.

28 The standard deviation is 2.9 years for schooling and 14.4 years for experience.

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The Changing Nature of Regional Wage Differentials From 1975 to 1983

by Lorie D. Jackson

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Introduction

Over the past 30 years, a great deal of research has been done on regional wage differentials. The subject has received considerable attention for a variety of reasons, notably because of its implications for understanding the degree to which competitive market forces lead to the equilibration of returns to labor, and also because of the possible effects of labor cost differentials on regional economic growth.

For the most part, the work on regional wage differentials has had three goals: (1) to estimate the size of regional wage differentials at a particular date or over time, (2) to identify their sources, and (3) to provide a theoretical explanation for their existence.

Estimates of regional wage differentials vary considerably as a result of variations in data sources, in measures of regional wage differentials, in measures of payments to workers, in geographic divisions, in time periods considered, and in methodologies used. Despite these inconsistencies across studies, most of the empirical work done confirms the view that, while some intermittent convergence has occurred over time, money wages in the northern United States have tended to be significantly greater than those in the South, at least since the beginning of this century.[1]

Most of the recent work on regional wage differentials defines the regional wage differential as the difference in wages that exists after controlling for differences in worker characteristics. This is because what is of interest to most researchers of regional wage differentials is not why workers with different characteristics are paid differently, but rather why workers with similar characteristics are paid differently across regions. Evidence of regional wage differentials is consistently found in the literature even after adjusting for the compositional mix of the work force. These differences reflect differences in the way particular worker characteristics are remunerated across regions due to variations in culture, tradition, degrees of discrimination, the bargaining strength of local unions, amenities, and public goods, as well as to temporal variations in supply and demand pressures. The differences in the way worker characteristics are remunerated across regions are referred to as differences in wage structures.

Several studies have separated the overall regional wage differential into the portion that can be explained by the compositional mix of the work force and into the portion that cannot. This separation makes it possible to isolate the regionally-specific source of the wage differential, and to determine which work force characteristics account for most of the difference in wage structures across regions.

Studies by Sahling and Smith (1983) and by Kiefer and Smith (1977) discuss the importance of differences in race and sex discrimination, and the effects of unionization in the

1 A different conclusion is reached in the study of real regional wage differentials. Recent studies that have adjusted for regional cost-of-living differences (Sahling and Smith [1983]) have found the real wage differential between the North and the South has not only been converging over time, but has been reversed in recent years.

wage structure component of the regional wage differential. To the author's knowledge, however, no study has been done on the changing importance of differences in the compositional mix of the work force and differences in regional wage structures on the overall size of regional wage differentials over time.

The purpose of this article is to estimate wage differentials between the East North Central region and two Southern regions in 1975 and 1983, and to discuss the changing nature of the differential over this period. The Southern regions considered are the East South Central and the South Atlantic. They were chosen to examine the widely held view that wages in the East North Central region are far out of line with wages in the Southern regions, and that this has been a major reason for the relative decline in manufacturing employment in the East North Central region over the past 20 years.

The East North Central area includes Ohio, Indiana, Illinois, Michigan, and Wisconsin. The South Atlantic region includes Delaware, Florida, Georgia, Virginia, West Virginia, North Carolina, and South Carolina. The East South Central area includes Kentucky, Tennessee, Mississippi, and Alabama.

Weighted Mean of Hourly Wage by Division, 1983 (in dollars)

	1983	1975
New England	8.92	4.80
Mid-Atlantic	9.39	5.63
East North Central	9.11	5.49
West North Central	8.56	4.87
South Atlantic	7.76	4.49
East South Central	7.69	4.47
West South Central	8.64	4.85
Mountain	9.02	5.36
Pacific	9.98	5.80

SOURCE: Data from 1983 and 1975 *Current Population Surveys,* Department of Commerce, Bureau of the Census.

TABLE 1

Two different regions of the South are considered in order to investigate the differences in the nature of the wage differentials between each of the two Southern regions and the East North Central region. In order to analyze their changing size and character over time, the differentials in two time periods are considered. The year 1983 was chosen because it was the most recent year for which the data were available. The year 1975 was chosen because the national economy was then at a point in the business cycle fairly similar to where it was in 1983, a fact that eliminates some of the differences in the magnitude of the differential over time due to cyclic variation in the demand for and supply of labor.

I. The Magnitude of Regional Wage Differentials

In the two periods considered, 1975 and 1983, the East North Central region had the third-highest average wage level of the nine census regions, while the South Atlantic and East South Central areas had the two lowest. The average hourly wage of a nonfarm worker between the ages of 25 and 64 in 1975 was $5.49 in the East North Central, compared to $4.47 in the East South Central, and to $4.49 in the South Atlantic. In 1983 the average hourly wage had risen to $9.11 in the East North Central, to $7.69 in the East South Central, and to $7.76 in the South Atlantic (see *table 1*). While money wages in the Southern regions were well below those in the East North Central region in both 1975 and 1983, the absolute percentage differentials declined by 3 percentage points over this period. The absolute wage differential between the East North Central and the South Atlantic regions went from about 18 percent in 1975 to 15 percent in 1983, while the differential between the East North Central and the East South Central regions went from 19 percent to 16 percent.

II. Theoretical Framework

Two basic theories of wage determination are posited to explain the existence of regional wage differentials: the neoclassical theory and the institutional theory. (Unless otherwise stated, the term "wage" will be used throughout this article to represent total labor compensation—wages plus supplemental benefits.)

The simple neoclassical model predicts that wages will be equalized across regions. This prediction rests on the assumption that labor and capital will move to where they can maximize their respective rates of return. Differences in wage levels across regions are expected to exist only in the short run when regional labor markets are out of equilibrium: both capital and labor take time to adjust to changing market signals. Since it is the purchasing power of the wage that is important to individuals, it is generally understood that it is the real, rather than the nominal, wage that neoclassical theory predicts would be equalized across regions (Sahling and Smith [1983]).

Elaborations have been made upon this simple model to bring into the fold nonwage factors affecting the location decision of labor and capital. Workers attempt to maximize their overall utility rather than simply their real wage. Similarly, firms attempt to maximize profits that are affected by more than just labor costs. Examples of nonwage factors affecting an individual's location decision are family considerations, such as employment opportunities for the spouse in a two-income household, amenity levels, and the quality of publicly provided services. Workers may require higher-than-average wages to locate in areas generally considered to have negative characteristics, such as air pollution, high population density, severe climate, and poor public services. Individuals may find that they can maximize their utility in a relatively low-wage region because of compensating nonwage considerations such as mild climate and good schools.

Similarly, firms take many factors into account when making location decisions. Among these factors are differences in the quality of the labor force, access to raw materials and markets, and proximity to the center of industry innovation. A firm may find that it can maximize profits by locating in a high-wage area because of cost and market advantages.

Since individuals and firms take into account nonwage factors when making location decisions, even if wages were driven by competitive forces, the movement of labor and capital would not necessarily equalize wages across regions. Rather, neoclassical theory would predict an equalization of utility and profits, which are composed of some mixture of wages, cost-of-living, amenities, etc. across regions. Because of the importance of nonwage factors, some difference in wages across regions would be expected to exist even in the long run and even after taking into account differences in worker and industry characteristics across regions.[2]

Many economists and industrial relations specialists believe that a satisfactory explanation for large and persistent regional wage differentials must go beyond the neoclassical model discussed above. Over the past 10 years, there has been a growing body of work on the importance of institutional forces on the wage adjustment process. Institutional factors include unions, racial and sexual discrimination, market concentration, and other noncompetitive forces that have a strong bearing on wages.

One common view within this literature is that wage changes, to a certain extent, are transmitted across regions as workers, and in some cases employers, attempt to maintain the wage standing of one group of workers relative to another across regions. These forces occur, both formally through collective bargaining, and informally through custom and convention.

Some researchers argue that one outcome of the existence of institutional factors is that regional wage differentials are decreased through comparisons and parity-bargaining between different groups of workers across regions (Martin [1981]). In some cases, workers adjust their wage expectations to maintain pay positions relative to other worker groups. This process is facilitated by the fact that unions and other labor groups are often organized on an industry-wide basis, or are represented in several industries or firms. While there is currently disagreement among labor economists about whether institutional factors have a long-term or merely a short-term effect on wages, their importance in the short run is widely recognized.

One often-cited institutional factor affecting wage differentials is unionization. Unionization affects an area's wage level to the extent that union workers, and perhaps some share of nonunion workers, can earn a wage that is different from what it would be without unionization. The actual effect of unionization on a region's wage level is the difference between a region's wage level, given the existence of unionization, and the wage level that would exist if there were no unionization. Thus a complete measure of the effect of unionization on regional wage levels should consider not only the difference between the wages of unionized and nonunionized workers, but also the amount of spillover from union wages on the determination of nonunion wages.[3] Capturing the spillover effect of unionization on nonunion wages, however, is a difficult and slippery process that is avoided in most studies of regional wage differentials.[4] Instead, many studies measure the effects of unionization on regional wage differentials as the proportionate union/nonunion wage advantage multiplied by the proportion of the work force that is unionized (Johnson [1983]; and Kiefer and Smith [1977]).

2 Within a competitive model, in order for industries to be competitive over time in regions where workers require wage premiums, there must be compensating cost factors associated with locating in those regions, such as nearness to raw materials, markets, and suppliers.

3 Most of the literature emphasizes the positive spillover effects of unions on nonunion workers when nonunion firms must compete with unionized firms or workers. Positive spillovers are assumed to be most acute for skilled nonunion workers who are costly to locate, hire, and train. Some researchers have also argued that a high degree of unionization in an area may lower the nonunion wage if workers are willing to accept a lower wage (a reservation wage) in a nonunion job in anticipation of future union employment and higher lifetime earnings (Johnson [1983]). Another possibility is that the existence of unions may have little or no effect on the nonunion wage. This may be the case if there is little competition between union and nonunion workers resulting from a low degree of local unionization, from a slack local labor market, or from workers waiting in the queue for union employment choosing unemployment over nonunion employment.

4 For further discussion of measuring the union-nonunion wage differential, see Moore, Newman, and Cunningham (1985).

III. Methods of Approach

As stated earlier, the regional wage differential can be separated into a portion that can be explained by differences in work force characteristics across regions, and a portion that cannot be so explained. The latter portion may reflect more regionally-specific differences, notably differences in the remuneration of particular characteristics. While both portions of the differential are potentially interesting subjects for investigation, the latter portion of the differential particularly concerns those who expect wages for similar workers in different regions to become equalized over time. The methodology used in this study permits a breakdown in the overall differential. It is the same methodology popularized by Oaxaca's 1973 study of the male/female pay differential and has become a standard decompositional approach.

The percentage wage differential between two regions (call them Region 1 and Region 2) can be decomposed into its compositional and wage structure components.[5] In order to decompose the differential, one must determine each region's wage structure. This is done by estimating separate wage equations using multiple regression analysis with the log of the wage as the dependent variable. Worker characteristics are included as the independent variables. The resulting regression coefficients indicate how particular characteristics are rewarded in that region. In order to determine the portion of the differential due to compositional differences, the average wage of Region 1 workers can be compared with the estimated wage of Region 2 workers in the absence of wage structure differences. To determine what portion of the overall differential can be explained by differences in the wage structure, the estimated wage of Region 2 workers, in the absence of wage structure differences can be compared with the actual average wage of workers in Region 2.

Since the actual earnings structure in the absence of regional differentials is not known, it is necessary to make some assumptions about what wage structure would exist if all regional wage structures were alike. There are two extreme possibilities: one is that the structure would be that estimated for Region 1, and the other is that the structure would be that estimated for Region 2. The fact that there is more than one possible estimate of the regional wage differential results in an index number problem. To deal with this problem, some researchers, such as Sahling and Smith (1983), averaged the estimated differentials resulting from using the bases of the two regions being compared. The exact meaning of the average, however, is difficult to interpret. Since the primary concern of this study is the effect of the East North Central's wage structure on regional wage differentials, the results using the East North Central as the base region are emphasized. This avoids the difficulties of interpreting the averages of the two extreme results. The results using the Southern bases will be discussed briefly to provide the reader with an idea of the range in the measures of the regional wage differentials.[6] The procedure is illustrated below:

If the East South Central (ESC) had the same wage structure as the East North Central (ENC), workers in the East South Central would receive:

$$ln\ \widehat{W}_{ESC} = f_{ENC}\ (\overline{X}_{ESC}),$$

$\widehat{W}_{ESC}$ = the estimated wage for ESC workers given the ENC wage structure,

f_{ENC} = the wage structure coefficients estimated for the ENC,

$\overline{X}_{ESC}$ = vector of the mean values of the independent variables for ESC workers.

The portion of the percentage wage differential attributable to differences in worker characteristics is measured by:

5 Many studies of regional wage differentials estimate a national wage equation that includes regional dummy variables. The coefficients on the locational variables are interpreted as the estimated proportionate difference between the wage rate in the region and its value in the nation for comparable workers. One major presumption behind the use of this approach is that regional wage structures are similar to the national wage structure, in other words, that the earnings of persons with the same attributes do not differ among the regions in any systematic way. This view is based on the premise that the United States is, geographically speaking, a single economy, operating within a single set of institutions, consisting of people of different ages, sexes, races, skills, and attachments to the labor market and engaged in a variety of occupations and industries. Regional divisions are presumed to have no significance in and of themselves, but merely to represent different groupings of human and material resources (Hanna [1951]). Hence, regional differences in the composition of these groupings are presumed to be the primary reason for differences in earnings across regions.

The assumption of similar wage equations across regions was questioned by Denison as far back as 1951. Hanushek (1973) performed Chow tests for the equality of coefficients for regions, and homogeneity within broad regions was consistently rejected at the one percent level of significance. In other words, Hanushek found that worker characteristics were compensated differently across regions. With a nationally estimated equation, differences in the way worker characteristics are remunerated are lost in the intercept term.

For further discussion of the appropriate approach for measuring regional wage differentials, see Kiefer and Smith (1977).

6 Decomposition results using the Southern regions wage structures as the base are available on request from the author.

$$ln\ \overline{W}_{ENC} - ln\ \widehat{W}_{ESC},$$

where: $\overline{W}_{ENC}$ = the average wage of ENC workers, and

$\widehat{W}_{ESC}$ = the estimated wage of ESC workers, given the ENC wage structure,

while that portion attributable to differences in the wage structure is measured by:

$$ln\ \widehat{W}_{ESC} - ln\ \overline{W}_{ESC},$$

where: $\widehat{W}_{ESC}$ = The estimated wage for ESC workers, given the ENC wage structure, and,

$\overline{W}_{ESC}$ = the average wage of ESC workers.

IV. Model

In keeping with most studies on wage differentials, a standard human capital earnings model developed by Becker (1975) and Mincer (1970) is estimated. According to this model, individuals attempt to maximize their income through investment in schooling and on-the-job training. This standard human capital earnings model is specified as follows:

$$ln\ W = B_0 + B_1S + B_2S^2 + B_3E + u$$

where:

W = average hourly wage,
S = years of schooling completed,
E = potential years of work experience, and
u = random error term.

The model is also specified to include a squared term for years of schooling to take into account diminishing returns to additional years of schooling.

Other work force characteristics associated with different wage levels are also included in the wage equation. They include a worker's sex, race, facility with the English language, marital status, union status, public or private employment status, full-time or part-time status, and occupation and industry affiliation.[7] Including these variables in the earnings model provides some adjustment for productivity and skill differences, for the existence of discrimination in the labor market, and for the wage effect of unions.

Some studies have attempted to adjust for compensating nonwage factors in individual location decisions, such as cost of living and amenities. Data limitations, however, make it difficult to construct measures of many of these compensating factors, particularly amenity levels. Studies have been done that estimate the wage differential across regions after adjusting for regional differences in the cost of living. Up until 1981, the Bureau of Labor Statistics published family budget indexes by three income categories for about 20 large metropolitan areas in the United States. Because no such data have been published on a census region basis, the data restrict analysis to a limited group of major SMSAs. Studies that have looked at real regional wage differentials have grouped the metropolitan areas for which data is available into broad regional groups (Sahling and Smith [1983]). These studies have thus considered only the real wage differential between regional groupings of large metropolitan areas. Cost-of-living data are not used in this study because they are not available on the desired geographical basis.

V. Data

The data sources used for this study are the 1975 and 1983 *Current Population Surveys* that contain information on worker characteristics and earnings from wages, salaries, commissions, and tips. Subsamples from each year were created to consist only of civilian, non-agricultural, private sector, and government workers between the ages of 25 and 65 years who worked either full time or part time (10 hours a week or more). The subsamples are limited to so-called prime age workers, in order to avoid addressing the unique characteristics of teen-age and elderly worker employment. Only workers who were recorded as working 10 hours or more per week were included because studies have found a large

7 The dummy variables are defined as follows:

Sex:	Dummy variable = 1 if the individual is male, and 0 if female;
Race:	Dummy variables for white, black, and other, with white individuals as the reference group; Spanish origin: Dummy variable = 1 if the individual is of Hispanic origin, and 0 otherwise. Serves as a proxy for not having English as a first language;
Marital status:	Dummy variable = 1 if the individual is married with spouse present, and 0 otherwise;
Full time:	Dummy variable =1 if the individual is a full-time employee, and 0 otherwise;
Class of worker:	Dummy variables for individuals working in the private sector, the federal government, the state government, and the local government, with private sector workers as the reference group;
Union coverage:	Dummy variable = 1 if the individual is either a union member or covered under a union contract, and 0 otherwise;
Occupation:	Dummy variables for U.S. Census one-digit occupations, with operators as the reference group;
Industry:	Dummy variables for U.S. Census one-digit industries, with durable manufacturing as the reference group.

chance of response errors for those registering fewer hours (Sahling and Smith [1983]). The hourly wage rate is estimated using information on usual weekly earnings and usual hours worked per week. The data series does not include information on years of work experience, so the conventional proxy (age, minus years of schooling, minus six) is used instead. Also, because data are not available on a worker's facility with the English language, Hispanic origin is used as a very rough proxy for English language difficulties. While the type of information contained in the 1975 and 1983 surveys is not identical, some general comparisons of the results for the two years can be made.

VI. Decomposition of Wage Differentials for the 1983 Sample

In 1983, the overall logarithmic wage differential between the East North Central and the South Atlantic was 20 percent, while that between the East North Central and East South Central was 18 percent (see *table 2*). Using the East North Central as the base wage structure, we find that differences in compositional mix made up only 30 percent of the wage differential between the East North Central and the South Atlantic, and only about 20 percent between the East North Central and East South Central.

The decomposition indicated that 70 percent of the wage differential between the East North Central and South Atlantic and close to 80 percent of the differential between the East North Central and East South Central are attributable to differences in wage structures. A Chow test verified that the wage structures of the Southern regions are significantly different from that of the East North Central region.

Decomposition of Regional Wage Differentials (East North Central base)

	1983		1975	
	East North Central/ East South Central (S = ESC)	East North Central/ South Atlantic (S = SA)	East North Central/ East South Central (S = ESC)	East North Central/ South Atlantic (S = SA)
Absolute differential ($\overline{W}_{ENC} - \overline{W}_S$)	$1.36	$1.50	$0.89	$0.98
Logarithmic differential ($ln\ \overline{W}_{ENC} - ln\ \overline{W}_S$)	0.18	0.20	0.20	0.23
Portion explained by different characteristics ($ln\ \overline{W}_{ENC} - ln\ \widehat{W}_S$)	0.04	0.06	0.06	0.09
Percent contribution to total logarithmic differential	23%	29%	29%	39%
Portion explained by different wage structures ($ln\ \widehat{W}_S - ln\ \overline{W}_S$)	0.14	0.14	0.14	0.14
Percent contribution to total logarithmic differential	77%	71%	71%	61%

where in 1983:

$\overline{W}_{ENC}$ = $8.27 $ln\ \overline{W}_{ENC} = 2.11$
$\overline{W}_{ESC}$ = $6.91 $ln\ \overline{W}_{ESC} = 1.93$
$\overline{W}_{SA}$ = $6.77 $ln\ \overline{W}_{SA} = 1.91$

$ln\ \widehat{W}_{ESC} = 2.07$
$ln\ \widehat{W}_{ESA} = 2.05$

where in 1975:

$\overline{W}_{ENC}$ = $4.91 $ln\ \overline{W}_{ENC} = 1.60$
$\overline{W}_{ESC}$ = $4.02 $ln\ \overline{W}_{ESC} = 1.39$
$\overline{W}_{SA}$ = $3.93 $ln\ \overline{W}_{SA} = 1.37$

$ln\ \widehat{W}_{ESC} = 1.53$
$ln\ \widehat{W}_{ESA} = 1.51$

TABLE 2

An important limitation of the wage information reported is that it does not include supplemental benefits. Studies have found that supplemental benefits tend to be positively correlated with wages, so the estimated regional differential using wage data alone probably understates the actual differential in total labor compensation across regions.

After taking into account differences in work force characteristics, the wage differential between the East North Central and both the Southern regions is the same, namely, about 14 percent. If the Southern regions are used as the base, the remaining differential between the East North Central and the two Southern regions after adjusting for compositional mix both fell slightly from 14 percent to 13 percent. Regardless of the base used, differences in regional wage structures appear to account for the lion's share of the wage differential.

While this is an interesting result in and of itself, it would also be useful to know the variables responsible for differences in wage structure. Most of the differences in wage structure, however, appear to be buried in the intercept term. This result may be partly explained by the omission of controls for regional differences in the cost of living, in amenities, and in supplemental benefits.

Wage Rate Equations, 1983
(estimated standard errors in parentheses)

Dependent variable: *ln W*	East North Central	East South Central	South Atlantic
Constant	0.9883	0.8019	0.8513
	(0.0245)	(0.0377)	(0.0255)
Education	0.0397	0.0458	0.0413
	(0.0015)	(0.0022)	(0.0015)
Experience	0.0153	0.0149	0.0128
	(0.0010)	(0.0017)	(0.0011)
Experience squared	-0.0002	-0.0002	-0.0002
	(0.0000)	(0.0000)	(0.0000)
Sex	0.2588	0.2780	0.2443
	(0.0068)	(0.0109)	(0.0073)
Race:			
White	——	——	——
Black	0.0003	-0.0900	-0.0997
	(0.0098)	(0.0125)	(0.0083)
Other	-0.0314	-0.0603	-0.0391
	(0.0256)	(0.0695)	(0.0346)
Spanish origin	-0.0309	-0.0467	-0.0859
	(0.0217)	(0.0802)	(0.0165)
Marital status	0.0413	0.0552	0.0494
	(0.0065)	(0.0109)	(0.0070)
Full time	0.1837	0.1105	0.1372
	(0.0195)	(0.0151)	(0.0099)
Class of worker:			
Private sector	——	——	——
Federal government	0.0311	0.1195	0.0688
	(0.0195)	(0.0239)	(0.0177)
State government	-0.0616	-0.0707	-0.0123
	(0.0110)	(0.0174)	(0.0118)
Union coverage	0.1487	0.1755	0.1691
	(0.0068)	(0.0118)	(0.0088)
R^2	0.4373	0.4551	0.4389
N	18,880	7,009	15,702

SOURCE: Data from 1983 and 1975 Current population Surveys, Department of Commerce, Bureau of the Census.

TABLE 3A

Even though the major sources of the differential appear to be buried in the intercept term, differences in returns to a few variables do stand out as important contributors to the wage differential due to structural differences (see *table 3a*).[8] For example, higher returns for full-time employment in the East North Central account for 30 percent of the structural differential between it and the South Atlantic, and 35 percent of the structural differential between the East North Central and the East South Central. Differences in returns for each additional year of experience account for 40 percent of the structural differential between the East North Central and the South Atlantic, while accounting for only 5 percent of the structural differential between the East North Central and East South Central.

There is no simple explanation for why returns to full-time workers would be higher for East North Central workers than for Southern workers, or why returns to experience would be greater for East North Central workers than for South Atlantic workers. It could be that the industries that are concentrated in the East North Central require more experienced, stable, full-time employees than industries concentrated in the Southern regions.

Differences in the degrees of racial discrimination between the North and South also appear to be a fairly important contributor to the structural differentials. The differences in returns between black and white workers account for 14 percent of the structural differential between the East North Central and South Atlantic, and for 8 percent of the differential between the East North Central and East South Central. While differences in the degrees of racial discrimination between the North and the South have long been recognized, it appears that relative to other variables and to the unknown portion of the differential, the contribution of differences in racial discrimination played a small role in the wage structure component of the differential in 1983.

Another interesting result is that the wage premium of unionized workers is very simi-

Wage Rate Equations, 1975
(estimated standard errors in parentheses)

Dependent variable: *ln W*	East North Central	East South Central	South Atlantic
Constant	0.4564	0.0914	0.1866
	(0.0657)	(0.1163)	(0.0769)
Education	0.0452	0.0507	0.0447
	(0.0037)	(0.0065)	(0.0045)
Experience	0.0137	0.0169	0.0214
	(0.0027)	(0.0050)	(0.0033)
Experience squared	-0.0002	-0.0002	-0.0004
	(0.0001)	(0.0001)	(0.0001)
Sex	0.3319	0.3424	0.2626
	(0.0196)	(0.0381)	(0.0241)
Race	-0.0283	0.0919	0.1197
	(0.0290)	(0.0463)	(0.0279)
Marital status	0.0049	0.0388	-0.0390
	(0.0206)	(0.0400)	(0.0275)
Full time	0.1052	0.0526	0.0901
	(0.0245)	(0.0491)	(0.0305)
Union member	0.1148	0.2205	0.2045
	(0.0173)	(0.0372)	(0.0279)
R^2	0.5206	0.5425	0.5069
N	2,069	594	1,299

SOURCE: Data from 1983 and 1975 Current Population Surveys, Department of Commerce, Bureau of the Census.

TABLE 3B

8 Full regression results are available on request from the author.

lar across the three regions observed. In fact, differences in the returns to unionized workers show that in the East North Central, unionized workers have a slightly smaller wage advantage over nonunionized workers than is true in the two Southern regions. The wage premium of unionized workers is about 15 percent in the East North Central, compared to about 18 percent in the East South Central and 17 percent in the South Atlantic. The slightly smaller union premium in the East North Central may result partly from the spillover effects of unions on nonunion wages. This seems probable, given the high degree of unionization and its associated threat effect in the region. But, as stated before, this spillover effect is difficult to measure. The similarities in wage pre-

Mean Values for Independent Variables, 1983
(standard deviations from the mean in parentheses)

Dependent variable: *ln W*	East North Central	East South Central	South Atlantic
Constant	——	——	——
Education	12.9880 (2.6067)	12.3549 (2.9317)	12.5144 (2.894)
Experience	21.2579 (11.6783)	21.2350 (11.6274)	21.3799 (11.6998)
Experience squared	588.2824 (567.2277)	586.1243 (578.9908)	593.9879 (582.5812)
Sex	0.5570 (0.4967)	0.5476 (0.4977)	0.5346 (0.4988)
Race:			
White	0.8967 (0.3044)	0.8247 (0.3802)	0.8047 (0.3964)
Black	0.0916 (0.2885)	0.1711 (0.3766)	0.1875 (0.3903)
Other	0.0117 (0.1075)	0.0042 (0.0648)	0.0078 (0.0878)
Spanish origin	0.0165 (0.1274)	0.0032 (0.0562)	0.0355 (0.1850)
Marital status	0.7343 (0.4417)	0.7553 (0.4299)	0.7175 (0.4502)
Full time	0.8635 (0.3433)	0.8899 (0.3131)	0.8814 (0.3233)
Class of worker:			
Private sector	0.8231 (0.3816)	0.7871 (0.4131)	0.7916 (0.4061)
Federal government	0.0255 (0.1577)	0.0532 (0.2245)	0.0422 (0.2010)
State government	0.0398 (0.1955)	0.0594 (0.2363)	0.0544 (0.2269)
Local government	0.1116 (0.3145)	0.1057 (0.3074)	0.1118 (0.3151)
Union coverage	0.3426 (0.4746)	0.2217 (0.4154)	0.1700 (0.3756)

SOURCE: Data from 1983 and 1975 Current Population Surveys, Department of Commerce, Bureau of the Census.

TABLE 4A

miums to unionized workers across regions may reflect the relative pay-setting practices of unionized workers within industries across regions.

As stated earlier, a popular, although incomplete, measure of unionization's effect on the regional wage level is the proportionate union/nonunion wage advantage, multiplied by the proportion of the work force that is unionized (see *table 4a*). Based on this procedure, the unionization effect in 1983 was 0.05 in the East North Central, 0.04 in the East South Central, and 0.03 in the South Atlantic. Hence, while the wage premium to unionized workers is slightly less in the East North Central than in the Southern regions, the union effect is greater because of the large concentration of unionized workers in this region.

VII. Changes in the Decomposition Over Time

The overall wage differential between the East North Central and each of the two Southern regions appears to have decreased between 1975 and 1983. The overall wage differential between the East North Central and the East South Central went from 20 percent in 1975 to 18 percent in 1983, and from 23 percent in 1975 to 20 percent in 1983 between the East North Central and the South Atlantic (see *table 2*). In contrast to the decline in the overall differential in both regional wage comparisons, the share of the differential due to wage structural differences was higher in 1983 than in 1975. The portion of the wage differential between the East North Central and the East South Central due to wage structure differences rose from 66 percent in 1975 to almost 80 percent in 1983. Over the same period, the portion of the wage differential between the East North Central and the South Atlantic due to differences in wage structures differences rose from about 60 percent to 70 percent.

When the Southern regions are used as the base, differences in wage structures showed similar increases in their contribution to the overall wage differential. One interesting difference in the results using the Southern bases was that, in 1975, differences in compositional mix accounted for almost 50 percent of the wage differential between the East North Central and the Southern regions. Regardless of the base used, differences in compositional mix have become less important in the overall regional wage differentials over time.

In 1975, as in 1983, the major portion of the structural component of the differen-

Mean Values for Independent Variables, 1975
(standard deviations from the mean in parentheses)

Dependent variable: *ln W*	East North Central	East South Central	South Atlantic
Constant	——	——	——
Education	12.3245 (2.7458)	11.4895 (3.2228)	11.6821 (3.1300)
Experience	22.8545 (11.7477)	24.6004 (12.8308)	23.0627 (12.0892)
Experience squared	660.3341 (583.2343)	769.8099 (672.9978)	678.0376 (621.6552)
Sex	0.6247 (0.4342)	0.5735 (0.4946)	0.5613 (0.4962)
Race	0.9304 (0.2544)	0.8804 (0.3690)	0.8406 (0.3885)
Marital status	0.8392 (0.3673)	0.8374 (0.3690)	0.8147 (0.3885)
Full time	0.88334 (0.3210)	0.8924 (0.3099)	0.8767 (0.3288)
Union member	0.3524 (0.4777)	0.2432 (0.4290)	0.1620 (0.3684)

SOURCE: Data from 1983 and 1975 Current Population Surveys, Department of Commerce, Bureau of the Census.

TABLE 4B

tial is not identified in the wage equation. Again, the intercept terms raise the wage structure in the East North Central above that of the Southern regions. There were also similarities in the variables identified in the wage equation that are important contributors to the structural differential in 1975, as was the case in 1983. Differences in returns to full-time workers explain 35 percent of the structural component between the East North Central and the East South Central in 1975, compared to 30 percent in 1983. Differences in returns to full-time workers explain less than 10 percent of the structural component between the East North Central and South Atlantic in 1975, compared to 35 percent in 1983. This result suggests that, between 1975 and 1983, differences in returns to full-time employment became a more important source of the regional wage differential between the East North Central and South Atlantic.

Differences in degrees of racial discrimination were, as one might expect, even more pronounced in 1975 than in 1983. The decline in the role of racial discrimination in explaining wage structure differences may reflect a decline in discriminatory practices in the Southern regions between the two years considered.

Between 1975 and 1983, differences in the degree of unionization across regions persisted, but returns to unionization became more similar. In 1975, the difference in the wage advantage to unionization across regions was considerably greater than it was in 1983 (see *tables 3a* and *3b*). But, in 1975, as in 1983, unionized workers in the South received a greater wage premium than their East North Central counterparts.

The total union effect in 1975 was smaller in the East North Central (0.04), than it was in 1983. It was larger in the East South Central (0.05), and was little changed in the South Atlantic (0.03). The union effect in the East South Central was greater than in the East North Central in 1975 despite the larger share of unionized workers in the latter region. This is because of much higher wage premiums to unionized workers in the East South Central at the time.

Market pressures probably contributed to the convergence in regional wage differentials over the period observed. Between 1975 and 1983, total non-agricultural employment rose by only 3 percent in the East North Central, compared to 27 percent in the South Atlantic and to 13 percent in the East South Central. While both of these Southern regions experienced stronger employment growth than the East North Central, it appears that labor market conditions were even tighter in the South Atlantic. This is suggested not only by the exceptionally strong employment growth in the region, but also by the region's relatively low unemployment rates over the periods considered. For example, in 1983, the unemployment rate in the South Atlantic was 8.5 percent, compared to 12.3 percent in the East South Central. Because of tighter labor market conditions in the South Atlantic, one might expect the regional wage differential to show greater convergence between the East North Central and the South Atlantic than that which exists between the East North Central and the East South Central. Indeed, this appears to be the case. The percentage wage differential between the East North Central and South Atlantic declined by 13 percent between 1975 and 1983, while the differential between the East North Central and the East South Central fell 10 percent. The portion attributable to wage structure differences, however, rose for both sets of regions, as was discussed above. The major reason for convergence appears to be the growing similarities in work force composition between the East North Central and Southern regions.

VIII. Conclusion

This study finds great similarity in the nature of wage differentials between the East North Central and the East South Central and South Atlantic regions. In both 1975 and 1983, structural differences account for most of the wage differential between the East North Central and the Southern regions. There are also similarities in the way that the differential changed between 1975 and 1983. For both regional comparisons, the importance of wage structure differences in the overall regional wage differentials grew over the time period considered. This wage convergence appears to result more from growing similarities in the composition of the work force than from returns to worker characteristics. The characteristics of the populations in the Southern regions have become more similar to those of the East North Central population, causing the importance of compositional differences in the overall wage differential to decline (see *tables 4a* and *4b*). The rise in the importance of the structural component appears to be solely attributable to the declining importance of compositional differences across regions.

While major sources of the differential remain unknown, it is clear that wage differentials continue to exist between the broad regional groupings observed in this study. Furthermore, adjustments for the standard productivity and skill-related variables, degrees of unionization, and the existence of race and sex discrimination, only eliminate about one-quarter of the overall regional wage differentials.

One encouraging result is that the wage differential between the regions considered declined between 1975 and 1983. Even if the decline continues at a rate similar to that experienced over the period (although there is no reason to expect this), nominal regional wage differentials can be expected to persist for some time. This suggests that considerable attention should be given to improving productivity in the East North Central and in other high-wage regions, in order to compensate for the region's higher, although converging, wages. Greater attention should also be given to the importance of nonwage factors that can be affected by regional policies, such as differences in the provision of public goods and services, in the unexplained portion of regional wage differentials.

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Labor Market Conditions in Ohio Versus the Rest of the United States: 1973-1984

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Introduction

This paper presents evidence that contrasts labor market conditions in Ohio and the rest of the United States during the 1973 to 1984 period. The evidence supports the following four propositions:

1. Whether we focus on the entire private sector or just on private manufacturing, Ohio's percentage change in employment was less than the percentage change in employment in the United States as a whole from 1973 to 1984. While this was particularly true in the last five years of the period, it was nearly as true for the first six.

2. The impact of unions on Ohio's relative wages undoubtedly contributed to the fact that Ohio's employment growth was below the national average, but the existing evidence does not support the belief that the direct union wage effect was a key factor.

3. While increases in the price of the U.S. dollar have deservedly received much attention of late, changes in exchange rates were not a significant factor in the *relative* worsening of Ohio's employment situation. The appreciation in the dollar's price hurt every state in the country, but did not hurt Ohio by an above-average amount.

4. Netting out the direct wage effects of unions, Ohio's manufacturing wage rates for a given quality of labor are substantially above the national average today, as they were in 1973. While we do not know exactly why Ohio's nonunion manufacturers pay a great deal more than comparable employers elsewhere in the country, this phenomenon is likely to be one reason why Ohio's employment growth rate was below the national average during the past 10 years.

The evidence presented is based on May Current Population Survey (CPS) microdata for 1973, 1979, 1983, and 1984. These data come from surveys of about 60,000 households conducted by the Bureau of the Census for the Bureau of Labor Statistics. The CPS surveys collect information on such things as employment status, usual hourly earnings, state of residence, union status, years of education, age, sex, race, occupation, and industry.

I. Findings

Table 1 gives unemployment rates for the United States as a whole, for Ohio, for a group of "high-growth" states, and for five states to which Ohio frequently compares itself—Michigan, Pennsylvania, Indiana, Illinois, and New York. The table reveals that, in 1973, Ohio's unemployment rate was slightly below the rate in the United States as a whole. In 1979, the two rates were identical, and in 1984, the Ohio rate was substantially above the national figure. Thus, the unemployment statistics suggest that Ohio's labor market conditions worsened slightly more than conditions elsewhere in the country during the 1973 to 1979 period, and worsened substantially more in the years between 1979 and 1984.

It is now well known that unemployment rates depend greatly on the extent to which the labor force is affected by the business cycle and by various structural factors. Thus, many

Private Sector Union Percentages in 1973, 1979, and 1984

	All sectors 1973	All sectors 1979	All sectors 1984	Manufacturing 1973	Manufacturing 1979	Manufacturing 1984
United States	24	21	16	39	35	27
Ohio	31	31	22	51	54	42
High-growth states	17	15	13	26	22	16
Michigan	36	34	28	58	54	52
Pennsylvania	31	30	20	50	53	43
Indiana	34	30	28	60	53	59
Illinois	31	27	19	44	38	29
New York	30	26	23	41	38	26

NOTE: High-growth states include California, Florida, Georgia, Massachusetts, North Carolina, and Texas.
SOURCE: May *Current Population Survey* data for all years.

TABLE 3

manufacturing sector taken by itself are given in *table 4B*. (Because the sample used to construct usual hourly earnings was cut substantially between the 1979 and 1983 May CPS surveys, the 1983 and 1984 surveys were merged to produce a sample of roughly the same size as was used in 1979.) What these two tables indicate is that the "union wage effect" has been lower in Ohio than elsewhere in the United States throughout the past decade, and that it has become substantially lower throughout the 1973 to 1984 period. The ability of unions to raise their members' wages above those of comparable nonunion employees is today much less in Ohio than it is in the vast majority of states. Furthermore, the fact that the union/nonunion wage differential is conditioned by the impact of unions on nonunion wages has been recognized since measurement of that differential first began.[1]

Tables 5A, 5B, and *5C* provide estimates of the percentage amount by which private sector hourly earnings were higher in Ohio than in comparison states in 1973, 1979, and 1983-84, respectively; *tables 6A, 6B,* and *6C* provide analogous estimates for the manufacturing sector taken by itself.[2] It is instructive to consider the first column in *table 5A*. The first figure in this column indicates that in 1973, usual hourly earnings were 4.8 percent higher in Ohio than in the rest of the country. The second figure in this column indicates that when the compari-

Percentage Amounts by which Union Hourly Earnings Exceeded Nonunion Hourly Earnings in 1973, 1979, and 1983-84

	A. Private Sector as a Whole						B. Private Sector, Manufacturing Only					
	Same worker			Same worker, same industry			Same worker			Same worker, same industry		
	1973	1979	1983-4	1973	1979	1983-4	1973	1979	1983-4	1973	1979	1983-4
United States	29	26	29	23	21	24	17	18	20	14	14	16
	(0.6)	(0.8)	(0.9)	(0.6)	(0.8)	(0.9)	(0.8)	(1.1)	(1.2)	(0.8)	(1.0)	(1.2)
Ohio	25	23	17	18	19	14	14	8.9	5.3	12	4.8	1.5
	(2.4)	(3.1)	(3.9)	(2.3)	(3.0)	(3.7)	(2.7)	(3.5)	(4.7)	(2.7)	(3.4)	(4.7)
High-growth states	30	26	35	25	22	31	16	19	25	13	14	21
	(1.3)	(1.8)	(2.0)	(1.3)	(1.7)	(1.9)	(1.7)	(2.4)	(2.9)	(1.7)	(2.4)	(2.9)
Michigan	27	19	22	19	15	16	14	16	18	6.4	13	9.6
	(2.7)	(3.5)	(4.2)	(2.7)	(3.4)	(4.0)	(3.7)	(4.3)	(5.2)	(3.6)	(4.0)	(5.1)
Pennsylvania	25	15	18	18	8.6	9.8	12	2.2	8.2	7.6	-2.2	7.1
	(2.6)	(3.2)	(3.4)	(2.5)	(3.1)	(3.3)	(2.8)	(4.0)	(4.7)	(2.8)	(3.8)	(4.9)
Indiana	29	24	31	22	18	20	14	10	5.2	8.4	5.0	-0.4
	(3.2)	(4.5)	(5.3)	(3.2)	(4.5)	(5.0)	(3.7)	(5.3)	(5.1)	(3.7)	(5.8)	(5.1)
Illinois	23	21	27	17	17	21	11	7.4	13	10	9.8	14
	(2.5)	(3.4)	(4.1)	(2.4)	(3.4)	(4.1)	(3.1)	(4.7)	(5.5)	(3.1)	(5.1)	(5.8)
New York	16	7.2	16	12	5.8	13	7.1	7.0	-1.1	7.7	9.3	1.8
	(2.1)	(2.7)	(3.1)	(2.0)	(2.7)	(3.1)	(2.9)	(4.2)	(5.6)	(3.0)	(4.4)	(6.0)

NOTES: Numbers in parentheses below percentages are standard errors. The adjective "same" refers to years of education, age and its square, race, sex and occupation (one of eight broad categories). The expression "same industry" denotes one of seven broad categories (in the case of table 4A) and one of 20 two-digit SIC industries in the case of table 4B. High-growth states include California, Florida, Georgia, Massachusetts, North Carolina and Texas.
SOURCE: May *Current Population Survey* data for all years.

TABLE 4

Percentage Amounts by which Private Sector Hourly Earnings Were Higher in Ohio than in Comparison States

A. 1973

Comparison states	United States	High-growth states	Michigan	Pennsylvania	Indiana	Illinois	New York
Total amount	4.8 (1.3)	6.4 (1.4)	-5.5 (1.8)	4.1 (1.7)	3.5 (2.0)	6.1 (1.7)	-8.8 (1.5)
Same workers	1.9 (1.0)	3.1 (1.1)	-8.1 (1.3)	2.8 (1.3)	1.3 (1.5)	-7.7 (1.2)	-8.4 (1.1)
Same workers, net of union premium	0.0 (0.9)	-0.3 (1.0)	-7.2 (1.2)	2.6 (1.3)	1.3 (1.4)	-7.1 (1.2)	-8.1 (1.1)
Same workers, same industry	1.7 (0.9)	2.7 (1.0)	-7.8 (1.2)	3.3 (1.2)	2.0 (1.4)	-7.5 (1.2)	-8.6 (1.1)
Same workers, same industry, net of union premium	0.3 (0.9)	-0.1 (1.0)	-7.2 (1.2)	3.1 (1.2)	1.9 (1.4)	-7.1 (1.1)	-8.3 (1.1)

B. 1979

Comparison states	United States	High-growth states	Michigan	Pennsylvania	Indiana	Illinois	New York
Total amount	3.5 (1.7)	6.2 (1.8)	-7.4 (2.2)	0.3 (2.2)	5.8 (2.7)	-6.9 (2.2)	-0.6 (2.1)
Same workers	2.0 (1.3)	4.8 (1.4)	-8.8 (1.6)	1.5 (1.7)	5.2 (2.0)	-8.4 (1.6)	-0.2 (1.6)
Same workers, net of union premium	-0.0 (1.3)	1.0 (1.4)	-8.6 (1.6)	1.4 (1.7)	4.5 (2.0)	-9.0 (1.5)	-0.3 (1.6)
Same workers, same industry	2.1 (1.3)	4.7 (1.4)	-8.0 (1.6)	2.5 (1.6)	6.0 (2.0)	-8.0 (1.6)	-0.2 (1.6)
Same workers, same industry, net of union premium	0.1 (1.3)	1.5 (1.3)	-7.9 (1.6)	2.4 (1.6)	5.5 (1.9)	-8.5 (1.5)	-0.3 (1.6)

C. 1983-84

Comparison states	United States	High-growth states	Michigan	Pennsylvania	Indiana	Illinois	New York
Total amount	3.3 (1.8)	1.6 (1.9)	-3.6 (2.5)	0.8 (2.4)	4.9 (3.0)	-5.3 (2.3)	-2.6 (2.2)
Same workers	0.7 (1.3)	-0.2 (1.4)	-4.8 (1.9)	-0.4 (1.8)	3.8 (2.3)	-6.7 (1.8)	-4.2 (1.7)
Same workers, net of union premium	-0.8 (1.3)	-2.5 (1.4)	-3.7 (1.9)	-0.2 (1.7)	4.3 (2.2)	-6.9 (1.7)	-3.6 (1.7)
Same workers, same industry	1.1 (1.3)	0.2 (1.4)	-5.4 (1.8)	0.1 (1.7)	4.6 (2.2)	-6.4 (1.7)	-3.7 (1.7)
Same workers, same industry, net of union premium	-0.2 (1.3)	-1.9 (1.4)	-4.5 (1.8)	0.2 (1.7)	4.9 (2.2)	-6.6 (1.7)	-3.2 (1.7)

NOTES: Numbers in parentheses below percentages are standard errors. The adjective "same" refers to years of education, age and its square, race, sex, and occupation (one of eight broad categories). The expression "same industry" means one of seven broad categories. High-growth states include California, Florida, Georgia, Massachusetts, North Carolina, and Texas.
SOURCE: May *Current Population Survey* data for the given year.

TABLE 5

manufacturers pay substantially more for a given type of worker than do employers elsewhere in the country. While this may reflect a desire to "avoid unionization," the evidence to support this contention has not yet been forthcoming.

Even if employers in Ohio have to pay more to attract and retain their workers than do employers elsewhere in the country, Ohio's employment situation can improve. A weakening of the dollar would not help Ohio more than the average state in the country on the employment front, but it clearly would increase the number of jobs in the state. Productivity improvements, on the other hand, would improve both Ohio's absolute and its relative employment situation. In the political arena, where I believe the trade situation can ultimately be improved, and at the worksite, where many productivity-enhancing innovations can be adopted, labor and management should be working together toward a common end — greater competitiveness. I also believe that this cooperation is much more likely if neither party continuously blames the other for today's problems, especially without solid evidence to support the position. Where one of the parties is clearly at fault, it must be willing to work with the other in the name of more and better jobs. Labor and management must be united, not divided, to improve labor market conditions in Ohio and in the rest of the country.

Economic Commentary