Order Number 9129684

Unions and wage determination: Three essays

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University of Hawaii, 1991

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UNIONS AND WAGE DETERMINATION:

THREE ESSAYS

A DISSERTATION SUBMITTED TO THE GRADUATE DIVISION OF THE UNIVERSITY OF HAWAII IN PARTIAL FULFILLMENT OF THE REQUIREMENTS FOR THE DEGREE OF

DOCTOR OF PHILOSOPHY

IN

ECONOMICS

MAY 1991

By

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ACKNOWLEDGMENTS

I would like to express my greatest appreciation to Prof. Moheb A. Ghali, my dissertation chairman, for his patient guidance and constructive advice throughout this research. I am also grateful to the other dissertation committee members, Prof. Hagen Koo, Prof. Marcellus S. Snow, Prof. John G. Bauer, and Prof. Gerard Russo for their valuable comments and suggestions on earlier drafts of my dissertation. I also thank Prof. Luigi Ermini for his useful advice on the time series analysis of this study.

I would also like to acknowledge the assistance of many people who helped my research. I am particularly indebted to Dr. Minja K. Choe, Ms. Victoria Ho, and Ms. Judith Tom at the Population Institute, East-West Center for their practical guidance in using computer programs.

The financial support of the Population Institute, East-West Center is also acknowledged, together with encouragement I have received from Dr. Lee-Jay Cho and Ms. Valerie Wong. I am also thankful to my place of employment, the Economic Planning Board, for extending my studying period abroad.

Finally, deepest gratitude goes to my mother and my late father for their endless love. I thank my wife and my two sons, Sungmin and Rheehoon. Their encouragement and support made it possible for me to successfully complete my doctoral dissertation.

ABSTRACT

Since the pioneering work by Lewis (1963), U.S. studies of unionism have concerned not only with what unions do for organized labor but also with what they do for unorganized labor. The main purpose of this study is to determine whether unions have a significant impact on nonunion wage determination. This paper is composed of three essays.

In the first essay, we have focused on whether union/nonunion wage spillovers exist and in which direction they run using time-series data. The empirical results suggest that union wage behavior spills over into nonunion wages and vice versa, that is, feedback relationship exists between the union and nonunion sector. Therefore, the hypothesis that nonunion wages are determined solely by market forces is rejected.

The second essay addresses the effects of the extent of union organization on the wages of union and nonunion workers using cross-section, individual-worker data. The empirical results demonstrate that geographic area (SMSA) union density has a positive and significant effect not only on union wages but also on nonunion wages while industry union density has a positive impact only on union wages. These results imply that the threat effect for the nonunion sector is determined by area-wide unionism rather than nation-wide industry unionism, and hence the existence of the union-to-nonunion spillover

effect is also supported, in part, by the cross-section data.

On the other hand, some authors have sought to take account of the reverse causality from wage rates to unionism. The endogeneity of union status in wage determination is investigated in the third essay. The empirical results show that the worker's predicted wage differential has a strong positive impact on the likelihood of union membership. This finding does not refute the hypothesis of the endogeneity of union status in wage determination.

To summarize, the role of unions in wage determination is not negligible, but unionism is not independent of the market mechanism.

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CHAPTER 1

INTRODUCTION

1.1 Purpose of The Study

Unions' effect on earnings - whether unions cause earnings to be higher than is determined by market forces alone - has long been one of the most important issues in labor economics. Since the pioneering work by H. Gregg Lewis (1963), numerous studies have produced strong evidence that unions increase the wages of their members relative to the wages of nonunion workers.

However, there is no agreement on what these relative union/nonunion wage differentials imply for the impact of unions on wage inflation. In the literature there exist contrasting views on the role of unionism in aggregate wage determination, which are based on the orthodox neo-classical theory and on the institutionalist school, respectively (Hirsch and Addison, 1986).

The orthodox position is that wage changes are primarily determined by market forces, although unions may impart some degree of inflationary bias to the economic system in the short run. That is, unions are essentially a passive force, responding to but not themselves generating inflation. Representing this view, Lucas and Rapping (1969) suggest that unions merely attempt to maintain a certain position relative to the nonunion sector and that it is therefore appropriate to analyze aggregate wage determination solely in terms of market forces.

On the other hand, institutionalists argue that unions can and do affect aggregate wage determination independently of market forces. The central idea of this approach is that wage changes are transmitted from one sector of the labor market to another not by a market mechanism, but by a spillover mechanism. Eckstein and Wilson (1962) emphasized a number of institutional aspects of the wage determination process using the concept of "key group" of industries that are heavily unionized. They argue that outside the key group of industries, wages are principally determined by wage spillovers from the key group.

Which of these views is a more accurate characterization of wage determination is an empirical question. If the wages in the nonunion sector are determined solely by market forces, we can ignore the influence of unionism on wage inflation because only a small fraction of workers are unionized in the United States. On the other hand, if union wage behavior spills over into the nonunion sector through the wage imitation process, union wage increases would be expected to lead to inflationary pressure throughout the economy. Therefore, the role of unions in wage inflation will depend on the direction in which, and degree to which nonunion wages are influenced by union wages, and the direction of union/nonunion wage spillovers would be an important policy concern.

In addition, there remains another unsettled issue - whether or not unionism is endogenous in wage determination. Recently, it has been recognized that not only can unions procure wage gains, but also the worker's decision to join a union is likely to be influenced by the wage comparison between his union and nonunion alternatives. At the extreme, it has sometimes been argued that unionization is influenced by wage rates but not vice versa and that high-wage industries already paid relatively high wages before the advent of wide-scale unionization in the U.S. (Krueger and Summers, 1987).

If wage rates have a greater effect on unionization, unions should be viewed as playing only a limited role in the wage determination process, and even the observed wage differentials may not be attributed to the role of unionism. For this reason, the endogeneity of unionism in wage determination must be examined as a fundamental issue.

The main purpose of this paper is, by tackling these issues, to throw light on the specific role of unions in wage determination in the United States. The importance of these issues is obvious as the government's effort to use incomes policies and other tools depends on what is believed to be the inflationary influence of unions. If we get the result that unions have a significant impact on the nonunion wage determination independently of market forces, the dominant policy orientation in the 1960s and 1970s during which incomes policies were employed would be supported by the data.

1.2 Organization of The Study

This study is composed of three essays. The first essay (Chapter 2) is a time-series analysis of union/nonunion wage spillovers, that is, whether or not the wage determination in one sector is influenced by the wage behavior of the other sector. In this part, we focus on whether union/nonunion wage spillovers exist and in which direction they run using the time-series data of union and nonunion wages. This topic is directly concerned with the relevance of the incomes policy in the United States.

The second essay (Chapter 3) examines the effects of the extent of union organization on the wages of union workers and nonunion workers using cross-section, individual-worker data. In this chapter, our main concern is to determine whether or not union density has a positive effect on the wages of nonunion workers because it can be regarded as cross-section evidence of the wage spillovers from the union to nonunion sector. This part complements the time-series evidence of the union/nonunion wage spillovers obtained in the first essay.

Finally, in the third essay (Chapter 4), we will investigate the endogeneity of union status in the wage determination model using cross-section, individual-worker data. The main hypothesis is that the worker's decision to join a union is influenced by the wage comparison between his union and nonunion alternatives. If the worker's likelihood of union membership is positively and significantly associated with his predicted wage differential, the hypothesis of the endogeneity of union status in the wage determination model cannot be refuted by the data.

These three essays can enhance our understanding of the specific role of unions in the wage determination process.

CHAPTER 2

TIME-SERIES ANALYSIS OF UNION/NONUNION WAGE SPILLOVERS

It has long been recognized that unions, in addition to directly affecting their members' wages, can indirectly affect the wages of nonunion workers. However, previous empirical studies have not provided clear evidence on the direction of the union/nonunion wage spillovers in the United States. In this chapter, we will investigate the direction of wage spillovers in the context of Granger-causal ordering. In addition, as a next step, empirical tests based on the theoretical framework of the wage spillover mechanism will be done to confirm the results of the causality tests.

2.1 Theoretical Framework of The Wage Spillover Mechanism

The wage spillover mechanism points to a uniformity of wage increases across the labor markets. Institutionalists argue that money wages in any sector is determined by the reference comparisons that sector participants make with an institutionally given set of wages in other sector (Burton and Addison, 1977). The traditional view of this approach is that union wage behavior influences nonunion wage determination but not vice versa.

In the literature there have been several explanations for the microfoundations of the wage spillover mechanism.

Among them the threat effect and morale effect hypotheses have been widely supported by many economists. The threat effect hypothesis, first suggested by Slichter (1954), contends that the possibility of union organization provides an incentive to nonunion employers to increase wage rates as a means of discouraging unionism. More formally, profit-maximizing nonunion employers will set the optimum nonunion wage such that the expected wage cost can be minimized subject to the probability of organization which is regarded as a function of the union/nonunion wage differential as well as the degree of unionization (Rosen, 1969).

Another approach to a direct link between union and nonunion wages focuses on the morale effect of a large value of the union/nonunion wage differential (Johnson, 1977). If an increasing wage differential is perceived by nonunion workers as inequitable, it will lead to a decline in workers' efforts. The morale effect hypothesis is based on the assumption that labor turnover is costly to the firm due to the existence of hiring costs and firm-financed investment in specific training. Assuming that workers respond by either quitting or reducing their efforts when dissatisfied with their wages, nonunion firms will take account of the union wage when setting their own wage offers.

Although the threat effect hypothesis has been more widely cited as a framework of the union-to-nonunion wage spillovers, which of these hypotheses is a more appropriate explanation of the wage spillover phenomenon is an empirical issue.

If we assume that unions seek a respectable wage advantage over nonunion workers and that nonunion wages are influenced by union wage behavior due to the threat and/or morale effects, the spillover mechanism may be written:

(1) the target wage from the point of view of nonunion employers, WN^*, is an increasing function of the observed union wage, but it is presumably lower than the observed wage[1];

$$WN^* = x^* \, WU \ , \qquad 0 < x^* < 1 \tag{2.1.1}$$

(2) the target wage from the point of view of union workers, WU^*, is taken to be higher than the observed nonunion wage;

$$WU^* = y^* \, WN \ , \qquad y^* > 1 \tag{2.1.2}$$

where x^* and y^* indicate the desired wage differential from the point of view of each sector (Ashenfelter, 1978).

Note that x^* and y^* may change over time. The implication for x^* is different between the threat effect and morale effect hypotheses. According to the threat effect hypothesis, the value of x^* will be increasing when the threat of union

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[1] Both employers and employees may prefer a nonunion work environment with equivalent wages due to union dues, strikes, and less flexible work rules (Hirsch and Addison, 1986).

organization is rising or unionism is expanding, while the morale effect hypothesis suggests that x^* is likely to be stable over time since it is largely determined by notions of fairness or equity. On the other hand, y^* is generally assumed to increase with the degree of union militancy or aggressiveness. The most common proxy for union militancy has been the unionization rate in the literature.

2.2 Review of Previous Studies

In the literature, although they are not mutually exclusive, there have been empirical studies on the following three types of wage spillovers: (1) inter-industry wage spillovers; (2) wage spillovers between individual collective bargaining agreements; and (3) union/nonunion wage spillovers (W. Vroman, 1982).

2.2.1 Evidence on Inter-Industry Wage Spillovers

This view is based on the argument that there is a leading sector which plays an important role in other sectors' wage determination as a reference sector. As mentioned earlier, Eckstein and Wilson (1962) provide empirical evidence, based on specific wage rounds, supporting the inter-industry spillovers within the U.S. manufacturing sector. For a 'key group' of industries, wage changes are determined by profit and unemployment rates based on the aggregate rates for all industries within the key group. Outside the key group of industries, wages are principally determined by wage spillovers from the key group.

McGuire and Rapping (1968), using cross-section data, reveals that while market forces play an important role in wage determination, wage spillovers from 'steel' and 'automobile' are also relevant in U.S. manufacturing.

Mehra (1976) conducted follow-up tests using residuals from two-digit U.S. manufacturing wage equations. After removing any autocorrelation in the error term for each industry, cross-industry serial correlation, which Mehra attributes to the existence of wage spillovers, is detected for some industries.

2.2.2 Evidence on Spillovers between Individual Contracts

This type of spillovers has been examined by several studies. Flanagan (1976) found weak evidence of the union-to-union wage spillovers using a limited sample of major collective bargaining agreements.

Mitchell (1980) also surveyed this issue using the 1973 United Auto Workers (UAW) agreements and the 1976 teamsters agreements. He concluded that evidence on wage spillovers between individual contracts was more obvious in the case of the UAW agreements.

Christofides, Swidinsky, and Wilson (1980) analyzed a large body of Canadian data on individual agreements. Their results is that spillover effects exist in the narrowly defined industry classification and that spillover effects, in addition to market forces, also play a significant role in wage determination.

2.2.3 Evidence on Union/Nonunion Wage Spillovers

The empirical validity of union/nonunion spillovers can be tested if extensive time-series data on average union and nonunion wage rates are available. As explained above, the traditional view is that union wage behavior spills over into nonunion wages but not vice versa.

Empirical tests involving union/nonunion wage spillovers have been conducted by Flanagan (1976), Johnson (1977), and Susan Vroman (1980, 1982) in the United States. Their model is based on the following two assumptions: (1) When the union/nonunion wage differential is too large, nonunion firms will respond to this by raising their wage rates because of the threat of organization and/or the morale effects; and (2) when the wage differential is small, unions will push for large wage increases in order to justify their existence to the membership.

In this context, they include the union/nonunion wage differential as an explanatory variable in both union and nonunion wage equations. If the coefficient of the wage differential is positive and significant in the nonunion wage equation, it can be said that nonunion wages are influenced by union wages, i.e.,there is a spillover from the union to the nonunion sector. On the other hand, if the coefficient of the wage differential is negative and significant in the union wage equation, it can be said that union wages are influenced by nonunion wages, i.e., there is a spillover from the nonunion to the union sector.

Johnson (1977) and Flanagan (1976) report, using annual data for the period 1954-71 and 1960-75, respectively, that nonunion wages are determined independently of union wages, but union wages are influenced by the wage differential. So, their conclusion would be that predominant wage spillovers are from the nonunion to union sector.

However, Susan Vroman (1982) reports an opposite wage spillover effect from the union to nonunion sector for the period 1960-78 using the same data set as Flanagan (1976) has used. She (1980, 1982) also points out that the results of Johnson's study are reversed when it is estimated using a different set of wage data[2] and that Flanagan's results are quite sensitive to the data period and choice of model specification.

Thus, it is not clear from these studies whether or not union wage behavior spills over into nonunion wages, i.e., there is no clear time-series evidence on the direction of union/nonunion wage spillovers in the literature. In the

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[2] S. Vroman re-estimates Johnson's model using the same data set as Flanagan has used. The data Johnson (1977) used were his estimated data on the rate of wage change.

following sections, we focus on whether union/nonunion wage spillovers exist and in which direction they run using both statistical tools and wage spillover models.

2.3 Statistical Methodologies and Empirical Testing for Causality

In this section, to identify the direction of wage spillovers we rely on some statistical tools which examine the causal ordering between union and nonunion wages. Empirical testing will be done using three alternative causality tests.

2.3.1 Granger's Definition of Causality

Granger (1969) has proposed a definition of a testable kind of causal ordering which can be applied in empirical work. His definition of causality is that " X causes Y " if and only if Y is better predicted in the mean by using all available information than if the information apart from X had been used.

In order to make Granger's definition operational, Granger makes some restrictive assumptions: (1) the domain of relevant information is restricted to past values of X and Y; (2) X and Y are stationary time series; (3) the predictors are constrained to be linear in past X and past Y; (4) the criterion for comparison of predictors is mean square error (Granger, 1969).

Under these conditions, Granger causality can be expressed in a testable form: " X does not cause Y " if the (minimum mean square forecast error) linear predictor of Y_t based on past Y and past X is identical to the linear predictor based on past X alone. The underlying basis of the Granger concept of causality is that the future cannot cause the past and that a cause contains unique information about an effect which is not available elsewhere (Granger and Newbold, 1986).

If Y causes X and X does not cause Y, it is said that "uni-directional" causality exists from Y to X. If X causes Y and Y causes X, it is said that "feedback" relationship exists between X and Y. If X does not cause Y and Y does not cause X, then we can say that X and Y are either statistically independent or related contemporaneously but in no other way.

2.3.2 Three Alternative Tests of Causality

In the literature we can find the following three alternative statistical tests of causality all of which are based on the Granger concept of causality (Guilkey and Salemi, 1982).

The Granger Test

In a simple two-variable model the Granger causality test can be illustrated: let X_t and Y_t be two stationary time-series with zero means. The Granger test is based on estimation of

$$X_t = \Sigma\ a_i\ X_{t-i} + \Sigma\ b_i\ Y_{t-i} + u_t$$

$$Y_t = \Sigma\ c_i\ Y_{t-i} + \Sigma\ d_i\ X_{t-i} + v_t$$

$$(i= 1,2,..,m)$$

where u_t and v_t are taken to be white noise, i.e., serially uncorrelated error terms (Granger, 1969).

According to the definition of causality given above, the test of the hypothesis that Y does not cause X is a test that $b_i = 0$ for i= 1,2,..,m. Similarly, X is causing Y if the hypothesis that $d_i = 0$ (i= 1,2,..,m) is rejected. The test statistic (F) is calculated by estimating the equation in both constrained and unconstrained form. Note that the Granger test may not be valid when the two variables are not stationary time series.

The Sims Test

Sims (1972) developed a new version of a test for Granger causality which is motivated by his theorem, that is, Y_t can be expressed as a one-sided distributed lag function of current and past X

$$Y_t = \Sigma\ a_i\ X_{t-i} + u_t \qquad (i= 0,1,..,m)$$

with a disturbance, u_t, which is not correlated with any values of past or future X, if and only if, Y does not cause X in Granger's sense (pp.544-545).[3]

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[3] Sims (1972) proved that his definition is equivalent to Granger's definition under the condition that the stochastic process is covariance-stationary with no linearly deterministic component. Later, Hosoya (1977) showed that these conditions are not necessary (see Chamberlain, 1982).

The Sims test for causality is based on estimation of

$$Y_t = \Sigma\ a_i\ X_{t-i} + \Sigma\ b_j\ X_{t+j} + u_t$$
$$(i= 0,1,..,m;\ j= 1,2,..,n)$$

According to his theorem, Y does not cause X, if and only if, the projection of Y involves only current and past X. So, the test of the hypothesis that Y does not cause X is a test that $b_j = 0$ for j= 1,2,..,n.

In the testing procedure, we regress Y on past and future values of X, taking account by generalized least squares (GLS) or prefiltering of the serial correlation. Sims (1972) argues that the prefilter approximately flattens the spectral density of most economic time-series and that regression residuals are expected to be white noise with this prefiltering.[4]

The Modified Sims Test

More recently, Geweke, Meese, and Dent (1983) devise a variant of the Sims test which avoids the need to use a GLS estimator or an ad hoc prefilter to cope with serial correlation. The modified test is based on estimation of

$$Y_t = \Sigma\ a_i\ X_{t-i} + \Sigma\ b_j\ X_{t+j} + \Sigma\ c_k\ Y_{t-k} + u_t$$
$$(i= 0,1,..,m;\ j= 1,2,..,n;\ k= 1,2,..,p)$$

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[4] In his empirical work Sims (1972) used the second-order filter: each variable X_t is replaced by $X_t - \rho_1\ X_{t-1} - \rho_2\ X_{t-2}$.

Interestingly, this equation is similar to a combination of the Granger and the Sims test. They call this a "lagged dependent variable version of the Sims test". As in the Sims test, the test of the hypothesis that Y does not cause X is the test that $b_j = 0$ for j= 1,2,..,n.

Note that in this test the error term, u_t, is white noise by construction (Granger and Newbold, 1986). In other words, the remedy for serial correlation is to include lagged values of the dependent variable Y rather than to use the GLS procedure of the Sims test.

2.3.3 The Application of Causality Tests and Data

Econometric Specification

The wage spillover phenomena between the union and nonunion sectors will be investigated using the three alternative tests of causality. The two time-series analyzed are the union wage (WU) and the nonunion wage (WN). A linear trend variable (t: time trend) is included in the estimating equations, following previous empirical work (Sims, 1972; Sargent, 1976; Salemi; 1980). When quarterly data are used, seasonal dummies (D_1, D_2, D_3) are also included.

The estimating equations can be expressed as follows: The Granger test is

$$\text{(i)} \quad WU_t = \Sigma\, a_i\, WU_{t-i} + \Sigma\, b_i\, WN_{t-i} + \alpha + \beta\, t + u_t$$

$$H_0: b_i = 0 \quad (i= 1,2,..,m)$$

$$\text{(ii)} \quad WN_t = \Sigma\, a'_i\, WN_{t-i} + \Sigma\, b'_i\, WU_{t-i} + \alpha' + \beta'\, t + v_t$$

$$H_0\text{: } b'_i = 0 \quad (i= 1,2,..,m)$$

If the null hypothesis in equation (i) is rejected, we can conclude that causality runs from nonunion wages to union wages. If the null hypothesis in equation (ii) is rejected, we can conclude that causality runs from union wages to nonunion wages. If both null hypotheses are rejected, the conclusion would be that feedback relationship exists between union and nonunion wages.

The Sims test is

$$\text{(i)} \quad WU_t = \Sigma\, a_i\, WN_{t-i} + \Sigma\, b_j\, WN_{t+j} + \alpha + \beta\, t + u_t$$

$$H_0\text{: } b_j = 0 \quad (j= 1,2,..,n)$$

$$\text{(ii)} \quad WN_t = \Sigma\, a'_i\, WU_{t-i} + \Sigma\, b'_j\, WU_{t+j} + \alpha' + \beta'\, t + v_t$$

$$H_0\text{: } b'_j = 0 \quad (j= 1,2,..,n)$$

If the null hypothesis in equation (i) is rejected, we can conclude that causality runs from union wages to nonunion wages. If the null hypothesis in equation (ii) is rejected, we can conclude that causality runs from nonunion wages to union wages. If both null hypotheses are rejected, the conclusion would be that feedback relationship exists between union and nonunion wages.

The modified Sims test is

$$\text{(i)} \quad WU_t= \Sigma\, a_i\, WN_{t-i} + \Sigma\, b_j\, WN_{t+j} + \Sigma\, c_k\, WU_{t-k} + \alpha + \beta\, t + u_t$$

$$H_0\text{: } b_j = 0 \quad (j= 1,..,n)$$

(ii) $WN_t = \Sigma\, a'_i\, WU_{t-i} + \Sigma\, b'_j\, WU_{t+j} + \Sigma\, c'_k\, WN_{t-k} + \alpha + \beta\, t + v_t$

$$H_0:\ b'_j = 0 \quad (j= 1,..,n)$$

Tests for the direction of causation in the modified Sims test are the same as in the Sims test.

The Data

The causality test requires extensive time-series data because it generally includes long length of lag distributions in the estimating equation. In the United States there are two kinds of time-series data on union and nonunion wages which are published by the Bureau of Labor Statistics (BLS).

One is the data from the WDM survey, Wage Developments in Manufacturing, which provide annual rates of change of union wages and nonunion wages in manufacturing. This is the data set which was used by Flanagan (1976) and S. Vroman (1982). The WDM data were first collected in 1959, but the survey was terminated after 1978. So, we can calculate the annual data series of union and nonunion wage indexes from 1958 (letting 1958=100) to 1978.[5]

The other data source is the Employment Cost Index (ECI), which was introduced by the BLS in 1976. The ECI provides quarterly data on union and nonunion wages (seasonally unadjusted wage index; quarter= March, June, September, December). The ECI union and nonunion wage data for the

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[5] These wage series are not annual averages but year-end figures (see Appendix Table A.1 and Flanagan, 1976).

manufacturing sector are available during the period 1976 (III)-1990 (I).

In this study, our main interest is to examine the direction of wage spillovers between the union and nonunion sectors and also to compare our results with previous studies. Because the coverage of the ECI data is not long enough for our purpose of analysis, we construct an extended annual data series for the period from 1958 to 1989 using the two data sets explained above.

The extended data set is composed of the 1958-76 wage series calculated from the WDM data (year-end figures) and the 1977-89 wage series from the ECI December data in the manufacturing sector. The extended annual data series of union and nonunion wages are reported in Appendix Table A.1. For empirical tests both the extended annual data (1958-89) and the ECI quarterly data (1976 III-1990 I) will be used.

2.3.4 Empirical Results of Causality Tests

In this part the direction of causation between union and nonunion wages will be examined by the Granger, Sims, and modified Sims tests using both the annual data (1958-89) and the quarterly data (1976 III-1990 I).

Using Annual Data

We have conducted the Granger causality test using annual data on union and nonunion wages. Table 2.1 summarizes the

F-statistics which are pertinent for testing the null hypothesis of no causality between union and nonunion wages. The full regression results are reported in Appendix Tables A.2 and A.3. The F-statistic is calculated by estimating the equation in both constrained and unconstrained forms. In this procedure, the choice of the length of lag distributions (m) is arbitrary. Considering the fact that most union workers are covered by three-year contracts and that nonunion wages are generally adjusted on a yearly basis, we choose three cases of m: m=2, m=3, and m=4.

Table 2.1

Summary Results of the Granger Causality Test: 1960-89

(i) $WU_t = \Sigma\, a_i\, WU_{t-i} + \Sigma\, b_i\, WN_{t-i} + \alpha + \beta\, t \quad (i=1,2,..,m)$
$H_0\!: b_i = 0$ for all i

(ii) $WN_t = \Sigma\, a'_i WN_{t-i} + \Sigma\, b'_i WU_{t-i} + \alpha' + \beta' t \quad (i= 1,2,..,m)$
$H_0\!: b'_i = 0$ for all i

	m = 2	m = 3	m = 4
union wage	$F_b = 4.65^*$	$F_b = 2.45$	$F_b = 4.16^*$
equation(i)	$\bar{R}^2 = .9995$ (SSE= 186.0)	$\bar{R}^2 = .9994$ (SSE= 184.4)	$\bar{R}^2 = .9995$ (SSE= 127.1)
nonunion wage	$F_{b'} = 4.58^*$	$F_{b'} = 3.46^*$	$F_{b'} = 2.66$
equation(ii)	$\bar{R}^2 = .9998$ (SSE= 62.5)	$\bar{R}^2 = .9998$ (SSE= 56.8)	$\bar{R}^2 = .9997$ (SSE= 52.7)
direction of causation	WU $\longleftrightarrow$ WN	WU $\longrightarrow$ WN	WU $\longleftarrow$ WN

Note: SSE denotes sum of squared errors.
* significant at the .05 level.

As reported by the F-values in Table 2.1, the null hypothesis that nonunion wages do not cause union wages is rejected in the union wage equation when m=2 and m=4, but cannot be rejected when m=3 (F_b= 2.45). Similarly, the null hypothesis that union wages do not cause nonunion wages is rejected in the nonunion wage equation when m=2 and m=3, but cannot be rejected when m=4 ($F_{b'}$= 2.66).

However, if we compare the values of $\bar{R}^2$ and SSE (sum of squared errors) in the three cases of m (m=2,3,4), we can see that the F-statistics drop to nonsignificance due to the inclusion of additional insignificant lagged variables, and thus we may ignore these two insignificant cases. Then, the conclusion of the Granger test would be that both null hypotheses are rejected, and therefore causality runs both ways between union and nonunion wages. Note that the Granger test may not be valid when the variables are not stationary time-series. This makes us turn our focus to the Sims and modified Sims tests.

For the Sims and modified Sims tests we have to choose not only the length of past lags (m) but also the length of future lags (n). Here, to minimize the loss of degrees of freedom we select two cases: (m=2, n=1) and (m=3, n=1).

Table 2.2 demonstrates the regression results of the Sims and modified Sims tests in the union wage equation. Both tests yield that the coefficient of the future nonunion wage (WN_{t+1}) has a sizable positive value and that it is significant

Table 2.2

Results of the Sims and modified Sims Tests: 1960-88
Union Wage Equation

(i) $WU_t = \Sigma\ c_i\ WN_{t-i} + d\ WN_{t+1}\ (+\ f\ WU_{t-1}) + \alpha + \beta\ t \quad (i= 0,1,..,m)$

H_0: d = 0

variable	Sims		modified Sims	
	m = 2	m = 3	m = 2	m = 3
WU_{t-1}			.879* (11.1)	.480* (7.88)
WN_t	.675* (3.14)	.722* (3.02)	.395 (1.07)	.499 (1.25)
WN_{t-1}	.499* (2.34)	.487* (2.08)	-.094 (0.25)	-.067 (0.17)
WN_{t-2}	-.678* (3.20)	-.727* (3.00)	-.683* (3.58)	-.660 (1.68)
WN_{t-3}		-.013 (0.06)		-.047 (0.20)
(F-statistic)	F_c=7.91*	F_c=5.72*		
WN_{t+1}	.662* (2.99)	.730* (3.10)	.523* (2.80)	.480* (2.36)
t	-1.08 (1.10)	-1.74 (1.55)	-.762* (3.27)	-.934* (3.30)
constant	-14.63* (3.50)	-14.56* (3.80)	-.050 (0.03)	-.478 (0.29)
$\bar{R}^2$	.9997	.9997	.9998	.9998
D.W.	2.10	2.08		
Durbin-h			-.63	-.80
direction of causation	WU ----→ WN		WU ----→ WN	

Note: The results of the Sims test are after correcting for second-order serial correlation.
Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level.

at the 5 percent level when m=2; n=1 as well as when m=3; n=1. So, the null hypothesis that union wages do not cause nonunion wages is rejected. Note that in the Sims test equations second-order serial correlation is corrected for by the Cochrane-Orcutt procedure[6], while in the modified Sims test equations serial correlation is cured by the inclusion of a lagged dependent variable.

In addition, Table 2.2 shows that the F-statistics of the current and lagged nonunion wages ($\Sigma\ WN_{t-i}$) are significant in the equations of the Sims test. This implies, according to Pindyck and Rubinfeld (1981), that causality also runs from nonunion to union wages. Thus, we can expect that the coefficient of the future union wage (WU_{t+1}) will be significant in the nonunion wage equation.

Table 2.3 reports the regression results of the Sims and modified Sims tests in the nonunion wage equation. According to both tests, as suggested by Pindyck and Rubinfeld, the coefficient of the future union wage (WU_{t+1}) is positive and significant at the 5 percent level when m=2; n=1 as well as when m=3; n=1. So, the null hypothesis that nonunion wages do not cause union wages is rejected.

In Table 2.3 the Sims test reveals that the current and lagged union wages ($\Sigma\ WU_{t-i}$) are significant as a group in the nonunion wage equation. This is consistent with the results in Table 2.2 that causality runs from union to nonunion wages.

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[6] See Pindyck and Rubinfeld (1981) p.157.

Table 2.3

Results of the Sims and modified Sims Tests: 1960-88
Nonunion Wage Equation

(ii) $WN_t = \Sigma\, c'_i WU_{t-i} + d'WU_{t+1}$ $(+\, f'WN_{t-1}) + \alpha' + \beta' t$ (i= 0,1,..,m)

$H_0 : d' = 0$

variable	Sims		modified Sims	
	m = 2	m = 3	m = 2	m = 3
WN_{t-1}			.891* (12.7)	.714* (8.47)
WU_t	.303* (3.54)	.002 (0.02)	.123 (0.88)	.091 (0.76)
WU_{t-1}	.348* (3.68)	.263* (2.11)	-.113 (0.78)	.110 (0.74)
WU_{t-2}	-.044 (0.44)	-.428* (3.06)	-.073 (0.84)	-.362* (2.70)
WU_{t-3}		.597* (7.35)		.218* (2.75)
(F-statistic)	$F_{c'}$=18.14*	$F_{c'}$=299.7*		
WU_{t+1}	.283* (3.21)	.394* (5.97)	.156* (2.25)	.181* (2.98)
t	-.363 (0.41)	1.13* (6.96)	.347* (2.77)	.549* (4.11)
constant	13.68* (3.67)	11.43* (19.2)	.604 (0.56)	2.57* (2.25)
$\bar{R}^2$	.9998	.9999	.9999	.9999
D.W.	2.29	2.08		
Durbin-h			-1.36	-1.25
direction of causation	WU ←---- WN		WU ←---- WN	

Note: The results of the Sims test are after correcting for second-order serial correlation.
Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level.

Therefore, the overall conclusion from the Sims and modified Sims tests is that causality runs both ways between union and nonunion wages.

On the other hand, when we compare the coefficient of the future nonunion wage (WN_{t+1}) with that of the future union wage (WU_{t+1}), we can see that the coefficient size of WN_{t+1} is considerably larger than that of WU_{t+1} in both Sims and modified Sims tests. Sims (1972) points out that the absolute size of the coefficient is important regardless of its significance and that small coefficients on future values of the independent variable may sometimes be ignored even when they are statistically significant (pp.545-546). In this context, if we interpret the absolute size of the coefficient as the degree of causation, we can say that the causality from union to nonunion wages is stronger than the causality from nonunion to union wages.

The regression results of the three causality tests, using annual data (1958-89), can be summarized as follows: (1) the Granger, Sims, and modified Sims tests generally support the notion that causality runs both ways or feedback relationship exists between union and nonunion wages; (2) if we measure the degree of causation by the absolute size of the coefficient, the results of the Sims and modified Sims tests imply that the causality from union to nonunion wages is stronger than the causality from nonunion to union wages.

Using Quarterly Data

We have also conducted the Granger, Sims, and modified Sims tests using quarterly data for the period 1976 (III)-1990 (I). The summary results (F-statistics) of the three causality tests are reported in Tables 2.4 and 2.5.

For the Granger test we choose the length of past lags (m) as three cases (m=4, m=8, and m=12) because four quarters represent one year. When m=4, as reported by the F-statistics in the first column of Table 2.4, the null hypothesis that nonunion wages do not cause union wages cannot be rejected in the union wage equation, but the null hypothesis that union wages do not cause nonunion wages is rejected in the nonunion wage equation. When m=8 and m=12, both null hypotheses cannot be rejected because the F-statistic is not significant in either of the equation, which implies that union and nonunion wages are statistically independent.

However, if we give more weights to the case of m=4 due to the higher values of $\bar{R}^2$, we may conclude that the Granger test supports uni-directional causality from union to nonunion wages.

Turning to the Sims and modified Sims tests, as shown in Table 2.5, two cases of the length of past lags (m) and future lags (n) are selected: (1) m=4, n=4; and (2) m=8, n=4. In both cases, the the coefficients on the future nonunion wages ($\Sigma\ WN_{t+j}$) are statistically significant in the union wage equation, but those on the future union wages ($\Sigma\ WU_{t+j}$) are not

Table 2.4

Summary Results of the Granger Test: Quarterly Data

(i) $WU_t = \Sigma a_i WU_{t-i} + \Sigma b_i WN_{t-i} + \alpha + \beta t + \text{dummies}$

(ii) $WN_t = \Sigma a'_i WN_{t-i} + \Sigma b'_i WU_{t-i} + \alpha' + \beta' t + \text{dummies}$

H_0: $b_i = 0$ (i= 1,2,..,m) H_0: $b'_i = 0$ (i= 1,2,..,m)

	m = 4	m = 8	m = 12
union wage equation(i)	F_b= 1.96 $\bar{R}^2$= .9995	F_b= 0.87 $\bar{R}^2$= .9992	F_b= 1.97 $\bar{R}^2$= .9992
nonunion wage equation(ii)	$F_{b'}$= 4.53* $\bar{R}^2$= .9997	$F_{b'}$= 1.87 $\bar{R}^2$= .9996	$F_{b'}$= 2.28 $\bar{R}^2$= .9997
causation	WU ----→ WN	none	none

Note: * significant at the .05 level.

Table 2.5

Summary of Sims and modified Sims tests: Quarterly Data

(i) $WU_t = \Sigma c_i\ WN_{t-i} + \Sigma d_j\ WN_{t+j} + \alpha + \beta t\ (+ f\ WU_{t-1}) + \text{dummies}$

(ii) $WN_t = \Sigma c'_i\ WU_{t-i} + \Sigma d'_j\ WU_{t+j} + \alpha' + \beta' t\ (+ f' WN_{t-1}) + \text{dummies}$

H_0: $d_j = 0$ (j=1,2,..,n) H_0: $d'_j = 0$ (j= 1,2,..,n)

	m = 4, n = 4		m = 8, n = 4	
	Sims	modified Sims	Sims	modified Sims
union wage equation(i)	F_d=9.57* $\bar{R}^2$=.9996	F_d= 3.94* $\bar{R}^2$=.9996	F_d= 6.34* $\bar{R}^2$=.9995	F_d= 2.90* $\bar{R}^2$=.9995
nonunion wage equation(ii)	$F_{d'}$=1.52 $\bar{R}^2$=.9997	$F_{d'}$=1.47 $\bar{R}^2$=.9997	$F_{d'}$=0.74 $\bar{R}^2$=.9996	$F_{d'}$=0.85 $\bar{R}^2$=.9996
causation	WU ----→ WN		WU ----→ WN	

Note: * significant at the .05 level

significant in the nonunion wage equation. These results seem to suggest that causality runs from union to nonunion wages but not vice versa.

Thus, the results shown in Tables 2.4 and 2.5 appear to support uni-directional causality from union to nonunion wages, which are not consistent with the results derived from annual data.

However, if we carefully compare the full regression results of each wage equation, we may question the validity of the conclusion obtained in Tables 2.4 and 2.5. First, let us re-examine the results of the Sims test.

Table 2.6 reports the full regression results of the Sims test when m=n=4. In the union wage equation, as seen in column (1) of Table 2.6, the coefficients of the current and lagged nonunion wages ($\Sigma\ WN_{t-i}$) are significant as a group (F_c= 11.50*), which implies that causality runs from nonunion to union wages. But, in the nonunion wage equation, as shown in column (2) of Table 2.6, the coefficients of the future union wages ($\Sigma\ WU_{t+j}$) are not significant as a group ($F_{d'}$= 1.52), which means that causality does not run from nonunion to union wages. These two results are contradictory to each other.

Moreover, as seen in column (2) of Table 2.6, none of the coefficients on the wage variables are significant in the nonunion wage equation, even though $\bar{R}^2$ is suspiciously high ($\bar{R}^2$=.9997). This is a typical pattern of multicollinearity. The estimation of an equation with a substantial number of

Table 2.6

Results of the Sims Test: Quarterly Data with Correction for Multicollinearity

(i) $WU_t = \Sigma\, c_i WN_{t-i} + \Sigma\, d_j WN_{t+j} + \alpha + \beta t + \text{dummies}$ (ii) $WN_t = \Sigma\, c'_i WU_{t-i} + \Sigma\, d'_j WU_{t+j} + \alpha' + \beta' t + \text{dummies}$

H_0: $d_j = 0$ $(j=1,..,4)$ H_0: $d'_j = 0$ $(j=1,..,4)$

variable	WU equation	variable	WN equation	
	(1) ordinary[a]		(2) ordinary[a]	ridge[b]
WN_t	.452* (2.82)	WU_t	.225 (1.53)	.161* (2.89)
WN_{t-1}	.267 (1.55)	WU_{t-1}	.099 (0.68)	.092 (1.70)
WN_{t-2}	.498* (2.82)	WU_{t-2}	-.053 (0.36)	.033 (0.64)
WN_{t-3}	.026 (0.15)	WU_{t-3}	-.003 (0.02)	.022 (0.46)
WN_{t-4}	.051 (0.32)	WU_{t-4}	.049 (0.41)	.001 (0.02)
WN_{t+1}	.377* (2.19)	WU_{t+1}	.124 (0.86)	.140* (2.44)
WN_{t+2}	.162 (0.93)	WU_{t+2}	.256 (1.83)	.121* (2.05)
WN_{t+3}	.101 (0.59)	WU_{t+3}	-.166 (1.20)	-.037 (0.65)
WN_{t+4}	.405* (2.66)	WU_{t+4}	-.099 (0.85)	-.091 (1.52)
t	-1.99* (12.2)	t	.834* (17.2)	.827* (31.4)
$\bar{R}^2$	.9996	$\bar{R}^2$	.9997	.9996
D.W.	1.74	D.W.	1.85	
(F-statistic)	F_c=11.50* F_d=9.57*	(F-statistic)	$F_{c'}$=1.68 $F_{d'}$=1.52	$F_{d'}$=2.71*
direction of causation	WU ---→ WN	direction of causation	none	WU ←--- WN

Note: [a] The results shown are after correcting for second-order serial correlation.
[b] The ridge estimator $\tilde{\beta} = (X'X+\theta I)^{-1}X'Y$, $\theta = 0.0005$.
Other variables included are three seasonal dummies and a constant.
Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level.

lagged explanatory variables is likely to lead to imprecise results due to the presence of multicollinearity. In this case, multicollinearity arises when variables are highly autocorrelated: the variable observed at time t has a close relationship to the variable at time t-1, t-2, etc (Pindyck and Rubinfeld, 1981). The reason multicollinearity occurs only in the nonunion wage equation can be explained by the fact that quarterly union wages included in the nonunion wage equation are more rigid than quarterly nonunion wages because of the presence of multi-year contracts in the union sector.

In this study, we use ridge regression to cope with multicollinearity in the nonunion wage equation. The ridge estimator can be expressed as $\tilde{\beta}= (X'X+\theta I)^{-1}X'Y$, $0<\theta<1$, while the ordinary least squares (OLS) estimator is $\hat{\beta}= (X'X)^{-1}X'Y$. Ridge regression yields a biased estimator with total bias directly related to the magnitude of θ, but it has a smaller covariance matrix than the OLS estimator.[7]

It may be argued that usual t- and F-tests may be misleading in ridge regression because the choice of θ is arbitrary and high values of θ can make all the coefficients significant. Here, we choose the value of θ as follows.

In column (1) of Table 2.6, the coefficients of the future nonunion wages ($\Sigma\ WN_{t+j}$) are significant as a group in the union wage equation, which suggests that causality runs

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[7] Hoerl and Kennard (1970) also show that there always exists a $\theta > 0$ such that the mean square error is less than that for the OLS estimator.

from union to nonunion wages. Thus we can expect that the current and lagged union wages (Σ WU_{t-i}) will also be significant as a group in the nonunion wage equation. In this context, we choose the value of θ which makes at least one of the current and lagged union wages significant in the nonunion wage equation, and by this criterion the minimum value of θ is chosen as θ=.0005. Of course, the F-value increases with θ, but bias is directly related to the value of θ. So, the regression result obtained at a smaller value of θ can be regarded as more reliable.

The result based on the choice of θ=.0005 is reported in the last column of Table 2.6. The 'pseudo' F-statistic[8] shows that the coefficients on future lags (Σ WU_{t+j}) are significant as a group in the nonunion wage equation, and thus the null hypothesis that causality does not run from nonunion to union wages is rejected. Therefore, our revised conclusion is that causality runs both ways between union and nonunion wages.

The regression results of the modified Sims test when m=4, n=4, and p=1 are reported in Table 2.7. Column (2) demonstrates that the coefficients on future lags are not significant as a group in the nonunion wage equation.

However, we can infer that the non-significance of the future coefficients in the nonunion wage equation is, as in the Sims test, caused by the presence of multicollinearity.

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[8] In ridge regression it is called the pseudo t- or F-test because of its biasedness.

Table 2.7

Results of the modified Sims Test: Quarterly Data with Correction for Multicollinearity

(i) $WU_t = \Sigma c_j WN_{t-j} + \Sigma d_j WN_{t+j} + \alpha + \beta t + f\ WU_{t-1}$ + dummies

H_0: $d_j = 0$ $(j=1,..,4)$

(ii) $WN_t = \Sigma c'_j WU_{t-j} + \Sigma d'_j WU_{t+j} + \alpha' + \beta' t + f' WN_{t-1}$ + dummies

H_0: $d'_j = 0$ $(j=1,..,4)$

variable	WU equation (1) ordinary	variable	WN equation (2) ordinary	WN equation ridge[a]
WU_{t-1}	.699*	WN_{t-1}	.574*	.363*
	(5.80)		(3.87)	(7.80)
WN_t	.182	WU_t	.169	.105*
	(0.82)		(0.85)	(2.20)
WN_{t-1}	-.064	WU_{t-1}	-.036	.034
	(0.28)		(0.18)	(0.75)
WN_{t-2}	.280	WU_{t-2}	-.090	.002
	(1.17)		(0.46)	(0.05)
WN_{t-3}	-.350	WU_{t-3}	.079	.012
	(1.45)		(0.41)	(0.30)
WN_{t-4}	.127	WU_{t-4}	-.053	-.006
	(0.77)		(0.43)	(0.14)
WN_{t+1}	.217	WU_{t+1}	-.027	.099
	(0.96)		(0.14)	(1.99)
WN_{t+2}	.106	WU_{t+2}	.359	.128*
	(0.45)		(1.92)	(2.52)
WN_{t+3}	-.169	WU_{t+3}	-.125	-.021
	(0.72)		(0.67)	(0.43)
WN_{t+4}	.403*	WU_{t+4}	-.095	-.069
	(2.52)		(0.78)	(1.35)
(F-statistic)	F_d=3.94*	(F-statistic)	$F_{d'}$=1.47	$F_{d'}$=2.95*
t	-.644*	t	.362*	.523*
	(2.44)		(2.76)	(12.7)
$\bar{R}^2$	.9996	$\bar{R}^2$	.9997	.9997
Durbin-h	.57	Durbin-h	$\rho=0$[b]	
direction of causation	WU ---→ WN	direction of causation	none	WU ←--- WN

Note: [a] The ridge estimator $\tilde{\beta} = (X'X+\theta I)^{-1}X'Y$, θ= 0.0005.
[b] Durbin-h statistic cannot be computed. So, we use an alternative test (ρ=0) by creating a lagged residual variable.
Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level.

The reason is that the same independent variables are included in the modified Sims test as in the Sims test except for the lagged dependent variable.

After correcting for multicollinearity by ridge regression (θ=.0005), the coefficients on future lags, as reported in the last column of Table 2.7, become significant as a group in the nonunion wage equation. Thus, the revised conclusion from the modified Sims test is that causality runs both ways between union and nonunion wages, and this is consistent with that of the Sims test.

Turning back to the Granger test, it may be argued that multicollinearity is also involved in the Granger test. However, it is not clear whether non-significance of the lagged nonunion wages is caused by multicollinearity in the union wage equation, because the Granger test includes the same independent variables for both union and nonunion wage equations. Thus, the use of ridge regression may not be appropriate for the Granger test. Although we have some reservations on the empirical validity of the Granger test result due to the possibility of multicollinearity, we may interpret the Granger test result as indicating stronger causation from union to nonunion wages.

Therefore, our conclusion from the three causality tests, using quarterly data, would be: (1) causality runs both ways or feedback relationship exists between union and nonunion wages; and (2) the degree of causation from union to nonunion wages seems to be stronger than that from nonunion to union wages.

To summarize, our findings from the causality tests are as follows: the results of the three causality tests lend general support to the notion that causality runs both ways between union and nonunion wages, i.e., union wage behavior spills over into nonunion wages and vice versa. This conclusion seems to be robust with respect to the use of annual data (1958-89) or quarterly data (1976 III-1990 I). Also, the causality tests provide some evidence which implies that the degree of causation from union to nonunion wages is stronger than that from nonunion to union wages.

It may be argued that the power of the causality test relies on the stationarity assumption of the data. So, we have also conducted comparable causality tests, after the union and nonunion wage variables are detrended by first logarithmic differences. The results are qualitatively similar to those obtained above: feedback relationship is supported by the Sims and modified Sims tests (the Granger test produces no significant result). Note that the Sims and modified Sims tests are, as explained earlier, not sensitive to the stationarity assumption by construction.

In the next section we will re-examine these results using wage spillover models which are based on the theoretical framework of the spillover process. This approach can help us understand the underlying mechanism of the wage spillovers.

2.4 Wage Spillover Models and Empirical Testing for The Direction of Wage Spillovers

Although the concept of Granger causality can be usefully applied in empirical work, some objection has been raised against it in that it involves only the concept of predictability and not the concept of cause and effect in a philosophical sense. And, it is also proved that the concept of Granger-noncausality and exogeneity serve different purposes (Engle et al., 1983; Wu, 1983). In this respect, the results of the causality test may be interpreted as arbitrary unless economic theory is capable of providing a unique model form to explain the causal relationship (Judge et al.,1985; Jacobs et al., 1979). Therefore, as a next step, we will examine the direction of wage spillovers in the context of the theoretical framework of the wage spillover mechanism.

2.4.1 Two References of Wage Spillovers

In the literature we can find two different ideas in models of wage determination which are based on whether workers are concerned about wage relativities or the target rate of wage growth (Burgess, 1988). The former involves the wage claim on comparisons with the wage levels of other workers while the latter refers to a rate of wage change workers desire or demand.

If we apply these ideas to the union/nonunion wage spillover mechanism, two references can be chosen as channels of the wage spillover process: (1) the union/nonunion wage differential and (2) the comparison of rates of wage change between the two sectors. If workers care about wage relativities, the wage determination in each sector will be affected by the observed wage differential. If workers are more concerned about the rate of wage growth, the wage determination in each sector is likely to be influenced by the comparison of rates of wage change between the two sectors, or more probably by last year's difference in the rate of wage change. The union/nonunion wage differential can be regarded as a more comprehensive reference because it results from the cumulative comparison of rates of wage change between the two sectors. On the other hand, it is usually more difficult to obtain information on the average wage or total compensation in the other sector than to get information on the rate of wage change in the other sector. Thus, we expect that both references can play their own roles in the wage determination process.

In this context, we can construct two types of wage spillover models in which either the absolute size of the wage differential or last year's difference in the rate of wage change is used as a reference of wage determination. Note that the time span of the second reference is restricted to only last year to minimize potential similarity between the two references.

2.4.2 Derivation of Wage Spillover Models

In this section we develop wage spillover models based on the two references explained above. We call the model with the wage differential Wage Spillover Model (I) and the model with the difference in the rate of wage change Wage Spillover Model (II).

Wage Spillover Model (I)

In this part we construct Wage Spillover Model (I) by assuming that workers are concerned about wage relativities between the union and nonunion sectors. As a first step we begin with a general form of a wage adjustment equation. In a competitive labor market, it is generally hypothesized that the rate of change of money wages (ΔW) depend on labor market conditions as well as on the expected rate of change of money wages (ΔW^e):

$$\Delta W_t = f(U_t) + \Delta W^e{}_t$$

where U is the unemployment rate. In labor market equilibrium $\Delta W = \Delta W^e$, and thus $f(U) = 0$. We call the unemployment rate which satisfies this equilibrium condition the natural rate of unemployment (Phelps, 1970).

The most common proxy for the expected rate of wage change has been the expected rate of price inflation (ΔP^e), which is based on the assumption that workers seek to maintain their real standard of living. Thus, the nonunion wage equation in the absence of wage spillovers may be written:

(2.4.1) $$\Delta WN_t = f(U_t) + \gamma\ \Delta P^e{}_t$$

On the other hand, the wage determination in the union sector can be assumed to depend not only on unemployment and expected price inflation but also on the degree of union militancy (Johnson, 1977; S. Vroman, 1982). The union wage equation in the absence of wage spillovers may be expressed:

(2.4.2) $$\Delta WU_t = g(U_t, M_t) + \delta\ \Delta P^e{}_t$$

where M denotes the degree of union militancy. Note that union militancy changes over time as union environment changes. The most common proxy for it has been the unionization rate in the literature.

Now let us accommodate the effect of wage spillovers on the wage determination in each sector when the union/nonunion wage differential is used as a spillover reference. We can find, in the work of Ashenfelter et al. (1972), some discussions of a wage spillover model in which the union/nonunion wage differential is used as a reference of wage spillovers. Previous empirical studies which analyzed the direction of wage spillovers have relied on this model.

Derivation of this model begins with the wage spillover mechanism between the union and nonunion sectors. As explained earlier, it is assumed that the target wage (WN^*) in the nonunion sector is an increasing function of the observed union wage and that the target wage (WU^*) in the union sector is taken to be higher than the observed nonunion wage. It may be written:

$$(2.4.3) \qquad WN^*_t = x^* + WU_{t-1} \quad \text{and} \quad WU^*_t = y^* + WN_{t-1}$$

where x^*, y^* indicate the desired wage differential from the point of view of each sector, and variables are transformed by taking logarithms.

If partial adjustments to equilibrium are hypothesized in both sectors, the wage change process can be expressed:

$$(2.4.4) \qquad \begin{aligned} WN_t - WN_{t-1} &= a\ (WN^*_t - WN_{t-1}), \quad 0 < a < 1 \\ WU_t - WU_{t-1} &= b\ (WU^*_t - WU_{t-1}), \quad 0 < b < 1 \end{aligned}$$

Substituting equations (2.4.3) into (2.4.4) yields:

$$(2.4.5) \qquad \Delta WN_t = a\ (x^* + D_{t-1}) \quad \text{and} \quad \Delta WU_t = b\ (y^* - D_{t-1})$$

where $\Delta WN_t = WN_t - WN_{t-1}$, $\Delta WU_t = WU_t - WU_{t-1}$, and $D_{t-1} = WU_{t-1} - WN_{t-1}$. Here, D_{t-1} approximates the proportionate union/nonunion wage differential at time t-1 since variables are transformed by taking logarithms. Equations (2.4.5) imply that the wage determination in each sector is influenced by the absolute size of the observed wage differential D_{t-1}.

Suppose that wage changes are determined not only by market forces but also by the spillover effects from the other sector. Then, by combining equations (2.4.5) with (2.4.1) and (2.4.2), the following form of testable equations of Wage Spillover Model (I) can be derived:

$$(2.4.6) \qquad \Delta WN_t = f(U_t) + \gamma\ \Delta P^e_t + a\ x^* + a\ D_{t-1}\ , \qquad 0 < a < 1$$

$$\Delta WU_t = g(U_t,\ M_t) + \delta\ \Delta P^e_t + b\ y^* - b\ D_{t-1}\ , \qquad 0 < b < 1$$

Note that x^*, y^*, and union militancy M may change over time.

In this model, the test of the hypothesis that nonunion wages are not influenced by union wages is a test that the coefficient of the wage differential a=0 in the nonunion wage equation. The test of the hypothesis that union wages are not influenced by nonunion wages is a test that the coefficient b=0 in the union wage equation. If a>0 and b=0, the conclusion would be that there is a spillover from union to nonunion wages but not vice versa. On the other hand, if a=0 and b>0, the conclusion would be that there is a spillover from nonunion to union wages but not vice versa. If a>0 and b>0, then it can be regarded as a feedback relationship between union and nonunion wages.

Wage Spillover Model (II)

In this part we construct Wage Spillover Model (II) by assuming that workers are mainly concerned about the target rate of wage change, instead of the wage differential. The wage equations in the absence of spillover effects are the same as above.

As shown earlier, the wage spillover mechanism between the union and nonunion sectors may be expressed:

(2.4.7) $$WN^*_t = x^* + WU_{t-1} \quad \text{and} \quad WU^*_t = y^* + WN_{t-1}$$

where WN^*, WU^* are the target wages in the nonunion and union sector, respectively, and variables are transformed by taking logarithms.

Suppose for the moment that the wage change in each sector is determined solely by the comparison of the rate of wage change between the two sectors. From equations (2.4.7), the target rate of wage change in each sector may be written:

(2.4.8) $$\Delta WN^*_t = \Delta x^* + \Delta WU_{t-1} \quad \text{and} \quad \Delta WU^*_t = \Delta y^* + \Delta WN_{t-1}$$

where $\Delta WN^*_t = WN^*_t - WN_{t-1}$ and $\Delta WU^*_t = WU^*_t - WU_{t-1}$.

If the target rate of wage change in the nonunion sector (ΔWN^*) is greater than the expected rate of wage change (ΔWN^e) that would be occurred in the absence of wage spillovers, the difference ($\Delta WN^* - \Delta WN^e$) will positively influence the wage determination in the nonunion sector. If the target rate is the same as the expected rate, it will have no effect on wage determination. If the target rate is smaller than the expected rate, the spillover forces might have a negative effect. Thus, the spillover reference of wage determination can be regarded as the difference between the target rate of wage change and the expected rate that would be occurred in the absence of wage spillovers, and the reference in each sector may be written:

(2.4.9) nonunion sector: $\Delta WN^{*}_{t} - \Delta WN^{e}_{t}$

union sector: $\Delta WU^{*}_{t} - \Delta WU^{e}_{t}$

We have already defined ΔWN^{*}_{t} and ΔWU^{*}_{t} in equations (2.4.8). If we assume that $\Delta WN^{e}_{t} = \Delta WN_{t-1}$ and $\Delta WU^{e}_{t} = \Delta WU_{t-1}$, the spillover reference in each sector may be rewritten:

(2.4.10) nonunion sector: $\Delta x^{*} + (\Delta WU_{t-1} - \Delta WN_{t-1})$

union sector: $\Delta y^{*} + (\Delta WN_{t-1} - \Delta WU_{t-1})$

It means that the wage determination in each sector is influenced by last year's difference in the rate of wage change between the two sectors.

If we add these spillover components in the wage equations, the following form of Wage Spillover Model (II) can be derived:

(2.4.11)
$$\Delta WN_{t} = f(U_{t}) + \gamma'\Delta P^{e}_{t} + c\ \Delta x^{*} + c\ (\Delta WU_{t-1} - \Delta WN_{t-1})$$
$$\Delta WU_{t} = g(U_{t}, M_{t}) + \delta'\Delta P^{e}_{t} + d\ \Delta y^{*} - d\ (\Delta WU_{t-1} - \Delta WN_{t-1})$$

where $0 < c, d < 1$. Also, note that Δx^{*}, Δy^{*} may change over time.

In this model, the test of the hypothesis that nonunion wages are not influenced by union wages is a test that the coefficient c=0 in the nonunion wage equation, and the test of the hypothesis that union wages are not influenced by nonunion wages is a test that the coefficient d=0 in the union wage equation. If both null hypotheses are rejected, the conclusion would be that feedback relationship exists between union and nonunion wages.

Wage Spillover Model (III)

As noted earlier, the spillover references in Models (I) and (II) are not mutually exclusive. Thus, a more generalized form of wage spillover model can be derived by including both spillover references in the wage equations. We call it Wage Spillover Model (III), and it may be expressed:

(2.4.12) $$\Delta WN_t = f(U_t) + \gamma\ \Delta P^e{}_t + a_1\ D_{t-1} + a_2\ \Delta W_{t-1} + a_1\ x^* + a_2\ \Delta x^*$$
$$\Delta WU_t = g(U_t,\ M_t) + \delta\ \Delta P^e{}_t - b_1\ D_{t-1} - b_2\ \Delta W_{t-1} + b_1\ y^* + b_2\ \Delta y^*$$

where $D_{t-1} = WU_{t-1} - WN_{t-1}$, $\Delta W_{t-1} = \Delta WU_{t-1} - \Delta WN_{t-1}$, and $0 < a,\ b < 1$.

In this model, the test of the hypothesis that nonunion wages are not influenced by union wages is a test that $a_1 = a_2 = 0$ in the nonunion wage equation. The hypothesis that union wages are not influenced by nonunion wages is a test that $b_1 = b_2 = 0$ in the union wage equation. If both hypotheses are rejected, the conclusion would be that feedback relationship exists between union and nonunion wages.

2.4.3 Econometric Specification and Data

Model Specification

To provide an empirical specification we need some restrictive assumptions on the wage spillover models derived above. First, concerning the functional form of unemployment in wage equations, a simple linear and negative relationship is assumed between wage inflation and unemployment. Phillips (1958) argues that there is a trade-off between the rate of wage change and unemployment and that it has a non-linear nature, which suggests the use of the reciprocal of the unemployment rate when estimating the wage equation. However, the linear specification performs better according to $\bar{R}^2$ and the standard error of the regression. So, we adopt the simple linear specification.

Secondly, it is necessary to specify how expectations for price changes are formed. Following the adaptive expectations hypothesis, we assume that $\Delta P^e{}_t = \Sigma\ a_i\ \Delta P_{t-i}$ where a_i is the weight given to the previous price change and $a_i > 0$. Usually the lag structure of price changes is supposed to be longer in the union sector than in the nonunion sector because of the presence of multi-year contracts in the union sector.

Thirdly, two dummy variables are included to adjust for periods during which incomes policies were employed. The Kennedy-Johnson guidepost dummy, I_1, has a value of one for the years 1962-66 and a value of zero otherwise (Pierson, 1968; S. Vroman, 1982). In mid-August 1971, the Nixon administration initiated another incomes policy. The policy period is represented by a dummy variable, I_2, taking the value of one in 1972 and 1973 and zero in all other years (Flanagan, 1976; S. Vroman, 1982).

Finally, we have to choose some proxy for x^*, y^*, (Δx^*, Δy^*), and union militancy M. As mentioned earlier, these variables may change over time as the union environment changes. Rosen (1969) suggests that x^* is positively associated with the degree of unionization because the threat of organization increases when unionism is expanding. Also, y^* and union militancy are generally assumed to be positively related to the unionization rate. Hines (1964) argues that the rate of change of unionization can be used as proxy for union militancy because union militancy would be highest when unions are gaining new members rapidly. Although some people argue that various measures of strike activity can be used as alternative proxies for union militancy, the most common proxy has been the unionization rate in the literature.

If we choose the extent of unionization, T, as proxy for x^*, y^*, and union militancy M, the empirical specification of the equations of Wage Spillover Model (III) can be expressed as follows:

$$\Delta WN_t = a_0 + a_1 U_t + \Sigma a_{2i} \Delta P_{t-i} + a_3 D_{t-1} + a_4 \Delta W_{t-1} +$$
$$a_5 I_1 + a_6 I_2 + a_7 T_{t-1} + a_8 \Delta T_{t-1} + \varepsilon^n{}_t$$
$$H_0\text{: } a_3 = a_4 = 0, \quad H_1\text{: } a_3, a_4 > 0$$

$$\Delta WU_t = b_0 + b_1 U_t + \Sigma b_{2i} \Delta P_{t-i} + b_3 D_{t-1} + b_4 \Delta W_{t-1} +$$
$$b_5 I_1 + b_6 I_2 + b_7 T_{t-1} + b_8 \Delta T_{t-1} + \varepsilon^u{}_t$$
$$H_0\text{: } b_3 = b_4 = 0, \quad H_1\text{: } b_3, b_4 < 0$$

where T is measured as union membership divided by non-agricultural employment; $\Delta T_{t-1} = T_{t-1} - T_{t-2}$; and ε^n, ε^u are random disturbances. Here, instead of T_t, lagged union density T_{t-1} is chosen as proxy for x^*, y^*, and union militancy because it performs better in the regression. ΔT_{t-1} is assumed as a proxy for Δx^* in the nonunion wage equation and a proxy for Δy^* in the union wage equation.

However, quarterly data on union density are not available in the United States. And, the annual union density data used above are not for the manufacturing sector[9] although wage data used are for manufacturing. For these reasons, we adopt another proxy variable.

One possible alternative is the use of the time trend, t. That is, we assume that x^*, y^*, and union militancy are a function of time trend, h(t). In the United States, as shown in Table 2.8, unionism was stable from the mid-1950s to the

[9] The union density series in manufacturing is discontinuous in the United States. See Troy and Sheflin (1985).

mid-1970s and has been continuously declining since the mid-1970s. If these trends in union density reflect changes in union environment over time, it may be assumed that x^*, y^*, and union militancy remained constant for the years from 1958 to the mid-1970s and have declined since the mid-1970s. Then, the functional form of the time trend can be specified as follows[10]:

$$h(t) = \alpha_0 + \alpha_1 (t*S) + \alpha_2 S$$

where S denotes a dummy variable which has a value of unity for the period from the mid-1970s to 1989 and zero otherwise. In this specification the constant α_0 represents the stable x^*, y^*, and union militancy for the years from 1958 to the mid-

Table 2.8

Trends in Union Density in the United States

	1956-59	1960-63	1964-67	1968-71
union density (%)	30.5	29.4	30.0	29.2
	1972-75	1976-79	1980-83	1984-87[a]
union density (%)	28.6	25.9	22.1	18.1

Note: [a] 1984-87 data are from the Bureau of Labor Statistics. Union density figures shown are annual averages during the period.
Source: Troy and Sheflin (1985), *U.S. Union Sourcebook*.
Hirsch and Addison (1986).
BLS, *Current Wage Developments*, February, 1985-88.

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[10] In this specification Δx^* and Δy^* cannot be identified since $\Delta t = t - (t-1) = 1$. They are assumed to be included in the constant.

1970s, and the expected sign of α_1 is negative, reflecting the decline in union density since the mid-1970s. The last term (α_2 S) is introduced to remove unnecessary restrictions on the constant term.

On the other hand, when quarterly data (1976 III-1990 I) are used, this specification is simplified to

$$h(t) = \alpha_0 + \alpha_1 t$$

Here, the expected sign of α_1 is negative, reflecting the continuing decline in unionism during this period.

Therefore, when the time trend (t) is used as proxy variable, the specification of Wage Spillover Model (I) may be expressed:

When we use annual data (1958-89),

$$\Delta WN_t = a_0 + a_1 U_t + \Sigma a_{2i} \Delta P_{t-i} + a_3 D_{t-1} + a_4 I_1 + a_5 I_2 + a_6 (t*S) + a_7 S + \varepsilon^n_t$$

$$H_0: a_3 = 0 , \quad H_1: a_3 > 0$$

$$\Delta WU_t = b_0 + b_1 U_t + \Sigma b_{2i} \Delta P_{t-i} + b_3 D_{t-1} + b_4 I_1 + b_5 I_2 + b_6 (t*S) + b_7 S + \varepsilon^u_t$$

$$H_0: b_3 = 0 , \quad H_1: b_3 < 0$$

When we use quarterly data (1976 III-1990 I),

$$\Delta WN_t = a_0 + a_1 U_t + \Sigma a_{2i} \Delta P_{t-i} + a_3 D_{t-4} + a_4 t + \varepsilon^n_t$$

$$H_0: a_3 = 0 , \quad H_0: a_3 > 0$$

$$\Delta WU_t = b_0 + b_1 U_t + \Sigma b_{2i} \Delta P_{t-i} + b_3 D_{t-4} + b_4 t + \varepsilon^u_t$$

$$H_0: b_3 = 0 , \quad H_0: b_3 < 0$$

where $\Delta WN_t = WN_t - WN_4$, $\Delta WU_t = WU_t - WU_4$, $D_{t-4} = WU_{t-4} - WN_{t-4}$, $\Delta W_{t-4} = \Delta WU_{t-4} - \Delta WN_{t-4}$, $U_t = (UN_t + UN_{t-1} + UN_{t-2} + UN_{t-3})/4$, and $\Sigma\ \Delta P_{t-i}$ is ΔP_{t-4} , ΔP_{t-8} , etc.

The specifications of these variables are chosen because wage rates are generally adjusted on a yearly basis while the data used are quarterly data.

The Data

For empirical tests, we use both annual (1958-89) and quarterly (1976 III-1990 I) data. The union and nonunion wage data used here are the same as used in the causality tests. Annual and quarterly average unemployment rates are from the Bureau of Labor Statistics, *Employment and Earnings*, October in each year. The quarterly unemployment data are seasonally adjusted rates.

Price inflation is calculated using the Consumer Price Index (CPI). Annual rates of price changes are December-to-December changes of the CPI. Quarterly data on the CPI are March, June, September, and December indices, and those are from the Bureau of Labor Statistics, *CPI Detailed Report*.

The annual union density data for the non-agricultural sector are available from Hirsch and Addison (1986, pp.46-47), which are provided by Leo Troy. But these data cover only up to 1984. Union density data for the period 1985-1989 are from the Bureau of Labor Statistics, *Current Wage Developments*, February, 1986-90. Since the two data series show a small difference in magnitude, some adjustments are done to make them a consistent data series.[11]

2.4.4 Empirical Results of Wage Spillover Models

Using Annual Data

Table 2.9 presents the regression results of Wage Spillover Models (I) and (III), in which the unionization variable is used as proxy for x^*, y^*, and union militancy. First, we estimate the equations of Wage Spillover Model (III), which can be regarded as a more generalized form of spillover model, and the results are reported in the first and second columns of Table 2.9. The coefficient of the wage differential (D_{t-1}) is positive and significant at the 5 percent level in the nonunion wage equation but not significantly different from zero in the union wage equation. In contrast, the coefficient of the difference in the rate of wage change (ΔW_{t-1}) is not significant in either of the union and nonunion wage equations. This implies that the wage differential is a more appropriate reference in the wage spillover process when annual data are used.

The regression results of Model (I), including only the wage differential variable, are reported in the third and fourth columns of Table 2.9. The results show that the wage differential (D_{t-1}) has a positive and significant effect on

[11] See Appendix Table A.1.

the wage determination in the nonunion sector but no significant effect on union wages. Thus, the direction of spillovers is from union to nonunion wages but not vice versa.

Note that D_{t-1} is calculated by assuming that the wage differential is zero in 1958. The value used for 1958 does not affect the coefficient estimates except for the constant.

In Table 2.9, we can find that the union sector is less responsive to current labor market conditions, comparing the coefficients of the unemployment rate between the union and nonunion sectors. This result occurs not only because multi-year contracts are prevailing in the union sector but also because unions are more interested in the employment of their current membership than in the total labor force.

The coefficient of first-year lagged price change is positive and significant in all cases, and the coefficient size is larger in the union sector than in the nonunion sector. The second-year lagged price change variable is not significant in either union or nonunion wage equation. This seems not consistent with our expectations that the lag structure of price changes is supposed to be longer in the union sector. However, similar results were found by Flanagan (1976) when union wages are measured as effective wages which include all wage adjustments receiving during the year.[12] In this study the data used are effective wages.

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[12] Effective wage adjustments include first-year adjustments, deferred increases, and cost-of-living adjustments (COLA).

Table 2.9

Regression Results of Wage Spillover Models: 1960-89 with Unionization Variables (T, ΔT)

variable	Model (III)		Model (I)	
	ΔWN eq.	ΔWU eq.	ΔWN eq.	ΔWU eq.
U_t	-.713*	-.553*	-.707*	-.440*
	(4.39)	(2.45)	(4.91)	(2.13)
ΔP_{t-1}	.170*	.272*	.169*	.244*
	(2.81)	(3.23)	(3.03)	(3.03)
ΔP_{t-2}	.045	.136	.038	.102
	(0.68)	(1.49)	(0.64)	(1.20)
D_{t-1}	.418*	.033	.403*	-.088
	(3.18)	(0.17)	(3.58)	(0.52)
ΔW_{t-1}	-.036	-.297		
	(0.22)	(1.29)		
(F-statistic)	F=5.96*	F=0.95		
I_1	-.777	-.370	-.794	-.377
	(1.48)	(0.51)	(1.60)	(0.53)
I_2	-.852	-.341	-.830	-.473
	(1.54)	(0.44)	(1.63)	(0.65)
T_{t-1}	.277*	.224	.260*	.176
	(2.14)	(1.13)	(2.30)	(0.97)
ΔT_{t-1}	-.042	.066		
	(0.25)	(0.27)		
constant	-1.65	.162	-1.10	1.61
	(0.44)	(0.03)	(0.34)	(0.32)
$\bar{R}^2$	.886	.861	.896	.863
D.W.	2.13	1.96	2.13	2.11
SER	.643	.903	.615	.896
spillovers	WU ------→ WN		WU ------→ WN	

Note: The results are after correcting for serial correlation.
SER denotes the standard error of the regression.
Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level in a two-tailed test.

The effects of incomes policies are captured by the coefficients of the dummy variables for the Kennedy-Johnson guidepost program (I_1) and the Nixon controls program (I_2). They are expected to have downward effects on both union and nonunion wage determination. However, neither of them is significant in the wage equations although they have negative coefficients as expected.

Turning to the unionization variables, the coefficient of union density (T_{t-1}) is positive and significant in the nonunion wage equation. This implies that the desired wage differential in the nonunion sector, x^*, is positively associated with the degree of unionization, and thus the threat effect hypothesis is supported by the data. However, the union density variable (T_{t-1}) has a positive but not a significant coefficient in the union wage equation. This means that the union militancy hypothesis is not supported by the data. Parsley (1980) points out, in his survey of the literature, that the adequacy of unionization to portray union militancy is limited since it excludes such factors as the political structure of the union and the attitude of the union leadership.

The results shown above are based on the assumption that the unionization rate is a good proxy for union environment. Now, we re-examine the direction of wage spillovers when the time trend (t) is used as proxy for union environment.

Table 2.10 demonstrates the regression results for union and nonunion wage equations of Wage Spillover Model (I) with the time trend variable. As explained above, U.S. union density remained stable from the mid-1950s to the mid-1970s and has been declining considerably since the mid-1970s. Because the exact point in time when unionism began to decline is not clear, we estimate three cases of regressions assuming that the turning point was 1975, 1976, and 1977. The regression results when we assume that the turning point was 1976 are reported in the first and second columns of Table 2.10. That is, we set S=1 for the period 1976-89 and S=0 otherwise. After correcting for serial correlation, the coefficient of the wage differential D_{t-1} is positive and significant in the nonunion wage equation, and it is negative and significant in the union wage equation. This implies that union wages influence nonunion wages and vice versa, i.e., feedback relationship exists between them. Note that the coefficient size of D_{t-1} indicates that the degree of spillovers is stronger from union to nonunion wages.

On the other hand, when we set S=1 for the years 1977-89 and S=0 otherwise, as shown in the third and fourth columns of Table 2.10, the coefficient of the wage differential D_{t-1} is positive and significant in the nonunion wage equation, while it is negative but not significant in the union wage equation. This result suggests that the direction of spillovers is from union to nonunion wages but not vice versa. When we set S=1 for the years 1975-89 and S=0 otherwise, the regression

Table 2.10

Regression Results of Wage Spillover Model (I): 1960-89
with the Time Trend (t)

variable	S = 1 for 1976-89		S = 1 for 1977-89	
	ΔWN eq.	ΔWU eq.	ΔWN eq.	ΔWU eq.
U_t	-.727* (4.36)	-.152* (2.86)	-.728* (5.08)	-.274 (1.72)
ΔP_{t-1}	.138 (2.04)	.575* (20.1)	.165* (3.01)	.431* (6.28)
ΔP_{t-2}	.060 (0.83)	-.120* (3.02)	.097 (1.51)	.135 (1.55)
D_{t-1}	.412* (3.05)	-.090* (2.62)	.311* (2.72)	-.111 (0.98)
I_1	-.786 (1.46)	-.622* (3.47)	-.637 (1.30)	-.630 (1.20)
I_2	-.979 (1.72)	2.14* (9.16)	-.796 (1.56)	1.09 (1.71)
t*S	-.223 (1.83)	-.409* (16.4)	-.304* (2.83)	-.385* (5.57)
S	-.144 (0.15)	4.27* (18.1)	1.40 (1.84)	3.43* (4.26)
constant	6.63* (5.36)	3.62* (15.1)	6.41* (6.58)	4.29* (5.57)
$\bar{R}^2$	.886	.962	.900	.901
D.W.	2.05	2.26	2.08	2.08
SER	.645	.474	.602	.764
direction of spillovers	WU ←——→ WN		WU ——→ WN	

Note: S is a dichotomous variable (S=1 or S=0).
The results are after correcting for serial correlation.
SER denotes the standard error of the regression.
Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level in a two-tailed test.

results (not reported here) also indicate that the direction of spillovers is from union to nonunion wages but not vice versa.

The result that the coefficient of the wage differential D_{t-1} is sensitive to the specification of the time trend dummy variable S seems to add support to the argument by S. Vroman (1982) that the regression results of the wage spillover model are sensitive to the data period and choice of model specification.

In Table 2.10, the effect of incomes policy is ambiguous in that the Nixon controls program (I_2) has a positive effect on the wage determination in the union sector. Table 2.10 also shows that the coefficient of the time trend since the mid-1970s is generally negative and significant in both union and nonunion wage equations. It appears to support the threat effect and union militancy hypotheses. However, this may not be clear evidence for these hypotheses because the time trend may capture other unmeasured productivity changes which are otherwise reflected in the error term of the wage equation.

When annual data are used, the direction of wage spillovers can be summarized as follows: the wage spillover models generally support the notion that spillovers are from union to nonunion wages. Also, there exists some evidence that feedback relationship exists between union and nonunion wages, and no evidence appears to contradict the existence of wage spillovers between the union and nonunion sectors.

<u>Using Quarterly Data</u>

Table 2.11 demonstrates the regression results of Wage Spillover Models (I) and (III) when quarterly data are used. Note that the time trend variable is used as proxy for union environment because quarterly union density data are not available.

First, we estimate the equations of Wage Spillover Model (III), and the results are reported in the first and second columns of Table 2.11. In the nonunion wage equation the coefficient of the wage differential D_{t-4} is positive but not significant, and the coefficient of the difference in the rate of wage change ΔW_{t-4} is positive and significant according to the t-test. This result implies that the difference in the rate of wage change is a more dominant reference in nonunion wage determination. On the other hand, neither of the references, D_{t-4} and ΔW_{t-4}, is significant in the union wage equation. Thus, Wage Spillover Model (III) suggests that the direction of spillovers is from union to nonunion wages but not vice versa.

The regression results of Wage Spillover Model (I), which includes only the wage differential variable D_{t-4}, are reported in the third and fourth columns of Table 2.11. In the nonunion wage equation the coefficient of D_{t-4} is positive and significant at the 5 percent level in a two-tailed test. In the union wage equation its coefficient is negative and significant at the 5 percent level in a one-tailed test.

Table 2.11

Regression Results of Wage Spillover Models: Quarterly Data with the Time Trend (t)

variable	Model (III)		Model (I)	
	ΔWN eq.	ΔWU eq.	ΔWN eq.	ΔWU eq.
U_t	-.540*	-.533*	-.664*	-.505*
	(3.87)	(2.16)	(4.88)	(2.18)
ΔP_{t-4}	.201*	.250*	.239*	.244*
	(4.38)	(3.64)	(5.27)	(3.73)
ΔP_{t-8}	.021	.194*	.064	.189*
	(0.45)	(2.72)	(1.36)	(2.72)
D_{t-4}	.116	-.203	.186*	-.221#
	(1.56)	(1.63)	(2.61)	(1.96)
ΔW_{t-4}	.261*	.047		
	(2.61)	(0.37)		
t	-.076*	-.145*	-.082*	-.143*
	(7.01)	(6.28)	(7.51)	(6.43)
constant	10.43*	10.66*	11.12*	10.44*
	(9.15)	(4.94)	(9.68)	(5.09)
$\bar{R}^2$	.968	.976	.963	.977
D.W.	2.16	1.89	2.05	1.92
SER	.330	.451	.355	.446
direction of spillovers	WU ⟶ WN		WU ⟷ WN	

Note: The results are after correcting for second-order serial correlation.
Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level in a two-tailed test.
# significant at the .05 level in a one-tailed test.

This result implies that union wages influence nonunion wages and vice versa, i.e., feedback relationship exists between them. Note that the coefficient of D_{t-4} becomes significant in the nonunion wage equation of Model (I), which is not significant in Model (III). This reflects the fact that the two references, D_{t-4} and ΔW_{t-4}, are not mutually exclusive.

In Table 2.11 the coefficient of the second-year lagged price change (ΔP_{t-8}) is positive and significant in the union wage equation while not significant in the nonunion wage equation. This is consistent with our expectations that the lag structure of price changes is supposed to be longer in the union sector due to the presence of multi-year contracts. And, the coefficient of the time trend is negative and significant in both wage equations. This seems to reflect the continuing decline in unionism since the mid-1970s.

In sum, the empirical results of wage spillover models, using both annual and quarterly data, provide strong evidence that spillovers exist from union to nonunion wages. And, the opposite spillovers from nonunion to union wages is also supported, although not so strongly, by the data. Therefore, the overall conclusion would be: (1) union wage behavior spills over into nonunion wages and vice versa, i.e., feedback relationship exists between union and nonunion wages; and (2) the degree of spillovers seems to be stronger from union to nonunion wages than from nonunion to union wages.

2.5 Conclusion

In this study we have examined whether union/nonunion wage spillovers exist and in which direction they run, relying on both causality tests and wage spillover models.

Based on the concept of Granger causal ordering, the results of the Granger, Sims, and modified Sims tests generally support the notion that causality runs both ways between union and nonunion wages. Also, they indicate that the degree of causation from union to nonunion wages is stronger than that from nonunion to union wages. These results seem to be robust with respect to annual data (1958-89) or quarterly data (1976 III-1990 I).

However, the causality test does not involve the theoretical framework of the spillover mechanism between union and nonunion wages. As a next step, the results of wage spillover models are examined using the same data set. The wage spillover models are constructed under the assumption that the wage determination in each sector is influenced not only by market forces but also by the wage comparison between the union and nonunion sectors.

Although the results of wage spillover models are, to some extent, sensitive to the choice of model specification, no evidence appears to contradict the results of the causality tests. They provide strong evidence that spillovers exist from union to nonunion wages and also provide some evidence that the opposite spillovers exists from nonunion to union wages.

Therefore, we can conclude as follows: (1) regarding the direction of wage spillovers, union wage behavior spills over into nonunion wages and vice versa over the past three decades in the United States. That is, the wage determination in each sector responds not only to market forces but also to the spillover effects from the other sector; (2) regarding the degree of wage spillovers, the time-series data generally support the notion that the spillovers from union to nonunion wages are stronger than those from nonunion to union wages.

CHAPTER 3

CROSS-SECTIONAL ANALYSIS OF UNION/NONUNION WAGE SPILLOVERS

The impact of unionism on wages is, as Lewis (1983) explained, likely to depend on the extent to which workers are organized in the work force. In other words, wage rates are likely to be affected not only by union membership, but also by the extent of union organization in an industry or a labor market. In this chapter, we will examine cross-section data on individual workers to determine the effects of industry and geographic area degree of unionization on the wages of union and nonunion workers.

3.1 Theoretical Implications of The Extent of Organization on Union and Nonunion Wages

3.1.1 The Effects on Union Wages

The wages of union workers tend to be higher, all else the same, the greater the extent of union organization in the relevant industry or labor market. The reason is that the ability of unions to achieve larger wage gains is likely to increase when there are few nonunion competitors in the product market, and as a consequence the elasticity of derived demand for the organized labor is low.[13] In this context,

[13] Freeman and Medoff (1982) report that the demand for labor tends to be less elastic in the union sector than in the nonunion sector.

Freeman and Medoff (1981a) argue that the percentage of workers organized can be regarded as an important determinant of union power.

3.1.2 The Effects on Nonunion Wages

The wages of nonunion workers may be influenced by the extent of labor organization. Most discussions of union effects on non-union wages have centered on what are referred to as the threat, the demand, and the crowding effects (Rosen, 1969; Kahn, 1978; Freeman and Medoff, 1981a; Moore et al., 1985; Podgursky, 1986).

First, union wage gains may compel nonunion firms to raise wages because of the threat of unionization, or because an increasing wage differential will negatively affect nonunion workers' efforts. This is the threat or morale effect. Note that this effect is far from market forces.

Secondly, if union wage gains result in product price differentials between union and nonunion firms, product demand will be shifted toward nonunion firms, and thus the demand for nonunion workers will increase. Unless the labor supply facing the nonunion sector is not perfectly elastic, this demand increase will put an upward pressure on wages in the nonunion labor market.

Finally, union wage gains are likely to lead to a reduction of employment in the union sector and thus a potential increase in the supply of labor to the nonunion sector. If labor shifts costlessly and completely from the union to the nonunion sector, this crowding effect (also called the labor-spillover effect) may substantially depress nonunion wages.

Therefore, the net effect of the extent of organization on nonunion wages is not clear a priori. Which of these effects dominates in a given economy is an empirical issue.

3.2 Review of The Literature

3.2.1 Empirical Results of Previous Studies

A number of studies have been conducted to analyze the effects of unionization on union and nonunion workers utilizing micro data. Freeman and Medoff (1981a) estimate separate union and non-union earnings equations with an explanatory variable for the extent of unionization in the three-digit industry, using 1973-75 Current Population Survey (CPS) data for manufacturing production workers. Their major findings are: (1) union density has a strong positive effect on the wages of union workers; (2) but it has either no association or positive and weak association with the wages of nonunion workers.

In contrast, Podgursky (1986) reports a large positive effect of industry unionization on the wages of both union and nonunion workers when the 1979 CPS data are used. Moore et al. (1985) also find a positive and significant effect of industry union density on both union and nonunion wages using aggregated CPS data for the period 1973-79.

Other studies have examined whether geographic area union density is positively related to the earnings of nonunion workers. Freeman and Medoff (1984) estimated the relationship between the percentage of blue-collar workers who are unionized in a Standard Metropolitan Statistical Area (SMSA) labor market and the wages of nonunion workers. The results, using the May 1973-78 CPS, show that the average wages of nonunion workers in an SMSA are positively associated with the percentage of workers unionized in an SMSA.

The impact of unions on the wages of nonunion workers may differ across individual groups. Detailed investigations of the characteristics of workers who gain and lose from working in a highly unionized geographic area have been made by Kahn (1978, 1980) and Holzer (1982). Kahn finds that non-union women and non-white men suffer from working in a highly unionized SMSA, while non-union white men benefit. Holzer's analysis shows that the wages and employment of young non-union blacks in highly unionized SMSAs are lower than in less unionized SMSAs.

On the other hand, some studies have analyzed the impact of unions on the functional distribution of income, that is, on the income share of union labor, nonunion labor, and capital. Using a general equilibrium model, the simulation results based on an assumed union wage differential of 15 percent show that virtually most of the gains by union labor come at the expense of nonunion labor, while income to capital is decreased only slightly (Johnson and Mieszkowski, 1970; Diewert, 1974; DeFina, 1985). This result implies that unionism is likely to decrease the wages of nonunion workers.

Therefore, it is not clear from these studies whether, or to what extent is the degree of unionization positively related to the wages of nonunion workers. There is no clear cross-section evidence of the union/nonunion wage spillovers in the literature.

3.2.2 Comparison of Model Specifications

The previous studies which analyzed the effects of industry union density differ in the functional form used in the regression and in the explanatory variables included in the wage equations.

Freeman and Medoff (1981a) estimate union and non-union wage equations separately. Their basic regression model is

$$\ln\ (WU_{ij}) = a_1X_{ij} + a_2IND_j + a_3T_j + e^u{}_{ij}$$
$$\ln\ (WN_{ij}) = b_1X_{ij} + b_2IND_j + b_3T_j + e^n{}_{ij}$$

where X_{ij} is a vector of the i-th worker's characteristics in industry j; IND_j is a vector of industry characteristics; T_j is the union density of the j-th industry; and e_{ij} is a random disturbance.

In contrast, Podgursky (1986) and Moore et al. (1985) use a single wage equation. The basic form is

$$\ln (W_{ij}) = c_1X_{ij} + c_2IND_j + c_3U_{ij} + c_4T_j + c_5U_{ij}T_j + e_{ij}$$

where U_{ij} is a dummy variable indicating whether a worker is covered by a collective bargaining agreement. In this model, the industry coverage effect for union workers can be estimated by coefficients, $c_4 + c_5$, while the effect on nonunion workers is captured by c_4.

Freeman and Medoff's two-equation model has some advantages because it does allow the coefficient of the earnings function to differ between the union and nonunion sectors. In general, unions appear to have the effect of flattening out the wage equation, that is, flat union-education and union-age profiles due to union seniority systems (Bloch and Kuskin, 1978). Podgursky's single equation model does not allow for these differences.

Turning to the control variables included in the regression, the Freeman-Medoff model specifies more detailed variables than the other models. Freeman and Medoff (1981a) includes three control variables for the industry (average firm size, four-firm concentration ratio, and injury rate), while Podgursky (1986) includes the four-firm concentration ratio only, and Moore et al. (1985) include establishment size only. The effects of industry union density are likely to be influenced by the choice of industry control variables. So, we will include all of the three industry variables.

In the next section, we will point out the weaknesses of previous studies and develop a revised strategy to analyze the impact of the extent of union organization on the wages of union and nonunion workers.

3.3 Econometric Specification and Data

3.3.1 A Revised Model Specification

The previous models which analyzed the effects of the extent of organization on union and nonunion wages involve some conceptual problems in the measurement of the union/nonunion wage spillovers. First, the industry union density data used in the previous studies are nationwide averages for three-digit industries. Hence, they are appropriate only when used for industries with nationwide product markets such as manufacturing (Holzer, 1982).

Second, the coefficient on the industry union density variable captures only the intra-industry wage spillovers within the three-digit industry classification. If there exist substantial threat effects across industries, that is, the wages in weakly unionized industries are influenced by those in heavily unionized industries, then the coefficient on the three-digit industry union density may not capture these inter-industry wage spillovers. Previous studies (Eckstein and Wilson, 1962; McGuire and Rapping, 1968) have shown that there exist significant inter-industry wage spillovers from the key group of industries such as "steel" and "automobile" industries in the United States.

Third, the use of geographic area (SMSA) union density can measure the effects of area-wide unionism on union and nonunion wages across industries. However, a possible bias may arise if the degree of unionism in one's industry-SMSA cell is correlated with the overall degree of unionism in the SMSA. In this case, if we include only the SMSA union density variable without the industry union density variable, the coefficient on the SMSA union density variable may reflect both intra-industry and inter-industry wage spillover effects.

Therefore, to investigate the cross-section evidence on the union/nonunion wage spillovers properly we should include both industry and geographic area (SMSA) union density in the wage determination model. The model we use is:

$$\ln\,(WU_{ijk}) = a_1\, X_{ijk} + a_2\, IND_j + a_3\, SMSA_k + a_4\, T_j + a_5\, T_k + \varepsilon^u{}_{ijk}$$

$$\ln\,(WN_{ijk}) = b_1\, X_{ijk} + b_2\, IND_j + b_3\, SMSA_k + b_4\, T_j + b_5\, T_k + \varepsilon^n{}_{ijk}$$

where X is a vector of individual worker's characteristics; IND_j is a vector of industry characteristics; $SMSA_k$ is a vector of SMSA characteristics; T_j, T_k are industry and SMSA union density, respectively; and ε^u, ε^n are random disturbances (i=individual, j=3-digit industry, k=SMSA).

The dependent variable is the natural log of hourly earnings for the individual worker. The semi-log form is derived from the standard human capital earnings model (Mincer, 1974), and it is consistent with all the earlier empirical work. The hourly earnings variable is weekly earnings divided by weekly hours. The union and nonunion samples are divided by whether a worker is covered by collective bargaining agreements rather than by the union membership status, since nonunion workers covered by collective bargaining agreements receive identical benefits as union members.[14]

The term X is a set of independent variables including demographic background (sex, race, marital status, veteran status, number of dependents), education, labor market experience (EXP= age-education-6) and its square, part-time dummy, 3 region dummies (Northeast, North-Central, and South), 4 occupation dummies[15], and 19 industry dummies[16]. These variables can capture variation in human capital investments and other worker and job characteristics influencing earnings.

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[14] Freeman and Medoff (1979) report that 91% of the covered workers are union members in the private sector. Also, the 1984-88 CPS data show that 93% of the covered workers are union members in manufacturing (Bureau of Labor Statistics).

[15] Four occupation dummies are craftsmen, operatives except transport, transport operatives, nonfarm laborers (the excluded group is service workers).

[16] We use two-digit Census Industry dummies. Among 20 two-digit industries, the ordnance industry is not included in the regression due to lack of data.

The explanatory variables of interest in this study are the industry and SMSA union densities which are measured as the proportion of production workers unionized in the three-digit Census industry (CIC) and the SMSA. In the union sample, the coefficient of industry union density reflects the degree of union power within the three-digit industry while the coefficient of SMSA union density represents the areawide inter-industry union power. In the nonunion sample, the coefficients of these variables, b_4 and b_5, capture the nationwide intra-industry threat effect and the areawide inter-industry threat effect, respectively.

As mentioned above, an important point in the estimation of the effects of industry and SMSA union density on union and non- union wages is the extent to which the effects of the relevant industry and SMSA factors are accounted for. The reason is that an omitted control variable which is partially correlated with both union density and wage rates will bias the estimated coefficients of the industry and SMSA union density variables. For this reason, variables measuring industry and SMSA characteristics that might be correlated with both wage rates and union density are included in the regression.

For industry-level control, four industry variables are included: average establishment size, the ratio of total wages to value added, the four-firm concentration ratio, and the injury and illness rate in three-digit industries (CIC).

First, the average establishment size accounts for the well-known positive effect of employer size on wages and the potential positive correlation between the extent of unionism and establishment size[17]. Numerous studies have reported that wage rates are higher in large plants (firms) than in small, which reflects at least in part the higher cost to large employers of judging labor quality (Mellow, 1982).

Second, the ratio of total wages to value added controls for the possible negative association of labor intensity (labor-to-capital ratio) with wage rates and the potential negative effect of labor intensity on industry union density. At a theoretical level, the marginal productivity of labor is positively related to capital intensity, and it is reported that wage rates are negatively associated with the ratio of total wages to value added (Masters, 1969). Also, there may be a positive correlation between industry union density and capital intensity because unions tend to be in a strong position when wages are a small percentage of unit costs in the industry.

Third, the four-firm concentration ratio captures the possible impact of industry concentration on wage rates and the likelihood that unionization is higher in more concentrated industries. Empirical studies report that concentrated industries tend to pay higher wage rates for

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[17] Establishment rather than firm size is chosen since most bargaining contracts are not firmwide in manufacturing.

given occupations (Weiss, 1966). Also, union organizers are likely to concentrate their efforts in industries which have a greater ability to pay higher wages.

Fourth, the injury and illness rate will control for the potential effect of dangerous work conditions on wages and their possible correlation with union density. We can find some partial evidence that workers are compensated by the market for dangerous job characteristics (Smith, 1979).

On the other hand, for SMSA-level control three SMSA variables are included: population size of the SMSA, the proportion of blacks (all non-white minorities) in the SMSA population, and the unemployment rate for the SMSA.

First, SMSA population size will control for the differences between the labor markets of larger and small urban areas (Holzer, 1982).

Second, the proportion of blacks accounts for its possible negative impact on wage rates and its potential positive correlation with SMSA union density. Kahn (1980) suggests that the proportion of blacks may approximate the degree of labor crowding in low-paying jobs within an SMSA. And, it is reported that workers are more likely to vote for union representation when they believe that they are treated unfairly (Farber and Saks, 1980).

Finally, the SMSA unemployment rate controls for the negative impact of unemployment on wage rates and its possible association with union density in local labor markets.

Inclusion of these control variables can minimize the possibility that the industry and SMSA union densities pick up their effects on wage rates. In addition, the inclusion of 19 two-digit CIC industry dummies reduce the chance that omitted industry factors may bias the estimated coefficients.

Here, we restrict our attention to production workers in manufacturing not only because unionized workers are primarily blue-collar labor but also because the threat effect of union organization may be diluted in heterogeneous samples which include white collar workers such as professional, technical, and managerial employees. The equations are estimated for workers in the 44 largest SMSAs on the CPS file.

3.3.2 The Data

This type of empirical study requires a large micro cross-section data set. We need not only the individual earnings and union-coverage but also detailed information on the characteristics of individuals, industries, and SMSAs.

Much of the information we need can be obtained from the CPS data file provided by the Bureau of Labor Statistics (BLS). The CPS is the monthly household survey used by the BLS to calculate the official unemployment rate. It contains many questions regarding the demographic and socioeconomic characteristics of household members. In addition to earnings data, the CPS file includes information on individual workers' characteristics such as age, sex, race, education, marital status, veteran status, union membership and coverage, residential location, etc.

Our regressions are based on the May 1979 CPS Supplement provided by the BLS. The total usable records of the May 1979 CPS are about 59 thousand employed workers. However, only 25 percent of the total CPS sample were asked the questions on earnings and hours of work in May, 1979. And, we restrict the samples to manufacturing production workers living in SMSAs. As a result, the samples in this study consist of 1,503 workers[18]: among them, 827 workers are the union sample and the others are the nonunion sample.

Data on the principal explanatory variables, proportion unionized figures for industries and for SMSAs, are taken from Freeman and Medoff (1979). Their estimates are based on the May 1973-75 CPS file, and their study is the only source from which estimates of the proportion of production workers unionized are available. This is the reason we do not use a more recent CPS file.

The CPS data file lacks information on industry and SMSA characteristics. The industry and SMSA variables are collected from other sources of statistics. The average establishment size is measured as the number of all employees divided by the number of establishments in each 3-digit CIC industry in 1977. Employee and establishment data by Standard

[18] In the Podgursky study (1986) the number of manufacturing production workers, regardless of their residence, is 3,093 using the May 1979 CPS.

Industrial Classification (SIC) are available from U.S. Department of Commerce, Bureau of the Census, *1977 Census of Manufactures*: Volume I, Table 1. These data are adjusted to 3-digit CIC industry classification of the CPS sample, using a conversion table provided by the Bureau of the Census.[19]

The ratio of total wages to value added are measured as the payroll for all employees divided by value added in each 3-digit CIC industry in 1977. Data source and the method of adjusting industry classification are the same as above.

The four-firm concentration ratio is measured as a weighted average of the fractions of shipments in 1977 accounted for by the four leading firms in the 4-digit SIC industries composing a 3-digit CIC industry. Here we use, as weights in the averaging process, value added in 4-digit SIC industries (see Weiss, 1963). Data on four-firm concentration ratios by 4-digit SIC industries are from U.S. Department of Commerce, Bureau of the Census, *1977 Census of Manufactures*: Volume I, Table 7.

The injury and illness rate is measured as the mean number of lost workdays due to injury and illness per full-time worker in each 3-digit CIC industry in 1978 and 1979. Data on injury and illness by SIC industries in 1978 and 1979 are from U.S. Department of Labor, Bureau of Labor Statistics, *Bulletin* (April, 1981).

Data on SMSA population size in 1980 and the proportion

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[19] See *CPS Technical Documentation*, May 1978, Appendix B.

of blacks in the SMSA population in 1980 are from U.S. Department of Commerce, Bureau of the Census, *State and Metropolitan Area Data Book* (1986). Unemployment rates for the SMSA in 1979 are from *State and Metropolitan Area Data Book* (1982).

To make an integrated data file which includes all information on individuals, industries, and SMSAs we merged the May 1979 CPS file with the industry data file including industry union density plus the four industry control variables and with the SMSA data file including SMSA union density plus the three SMSA control variables. Industry and SMSA data are matched to the CPS samples using the 3-digit CIC number and the SMSA code number.

3.4 Empirical Results

3.4.1 Data Summary

The sample we analyze is limited to manufacturing production workers who reside in the 44 largest SMSAs. Before reporting the regression results, we compare the sample characteristics of the union and nonunion sectors in our data.

Table 3.1 summarizes the data by providing means and standard deviations of the means for the characteristics of individual workers, industries, and SMSAs in the union and nonunion samples. Some significant differences exist between the two sectors. The mean value of hourly earnings is greater

Table 3.1

Sample Characteristics in the Union and Nonunion Samples

variable	union sample mean	union sample (S.D.)	nonunion sample mean	nonunion sample (S.D.)
hourly earnings ($)	6.93	(.080)	5.87	(.115)
years of education	12.1	(.086)	12.1	(.109)
labor market experience (EXP = age-education-6)	23.0	(.484)	18.9	(.555)
industry union density(%)	55.5	(.617)	44.7	(.621)
SMSA union density(%)	43.6	(.351)	38.0	(.443)
average establishment size (workers)	154	(6.65)	99	(4.68)
wages/value added(%)	44.4	(.277)	44.2	(.296)
concentration ratio(%)	40.3	(.656)	34.6	(.634)
injury & illness rate(%)	88.8	(1.52)	80.5	(1.78)
SMSA population size (in thousands)	3,423	(82.0)	3,395	(93.7)
proportion of blacks in SMSAs (%)	20.8	(.301)	21.2	(.325)
unemployment rate in SMSAs (%)	6.0	(.047)	5.6	(.049)
proportion of part-time workers(%)	2.9	(.584)	10.7	(1.19)
proportion of workers in Southern region(%)	11.4	(1.10)	20.0	(1.54)
sample size (N)	827		676	

Note: Data are from the May 1979 CPS bridged with the industry and SMSA characteristics.
Numbers in parentheses are standard deviations of the means, i.e., standard deviations of the variables divided by $N^{1/2}$.

and the standard deviation of the mean is smaller in the union sample than in the nonunion sample. This is consistent with the findings of previous studies that dispersion of wages is significantly narrower in unionized establishments than in nonunionized establishments due to unions' standard rate policies (Freeman, 1980, 1982).

Table 3.1 also demonstrates that workers employed in larger establishments and in more highly concentrated industries are more likely to be unionized, as reported by previous studies. This phenomenon can be explained by the fact that union organizers are likely to concentrate their organizing efforts in these establishments and industries because the costs of organization are relatively low in such establishments, and highly concentrated industries have a greater ability to pay higher wages.

There exists little difference in SMSA characteristics between the union and nonunion samples although the average population size and unemployment rate are somewhat higher in the union sample. Freeman and Medoff (1984) point out that the unemployment rate tends to be higher in a heavily unionized region than in a weakly unionized region.

Finally, the proportion of workers who live in Southern region is much smaller in the union sample than in the nonunion sample, and the same is true of the proportion of workers who are employed on the part-time basis. These are also consistent with the findings of previous studies.

3.4.2 Regression Results

Estimates of the wage equations for union and nonunion workers, based on the May 1979 CPS, are presented in Table 3.2. In this table, we report only the regression coefficients and t-statistics of the variables which are important to analyze the impact of industry and SMSA union density on earnings. The full regression results are reported in Appendix Table A.4. Note that establishment size is measured by the industry average in Table 3.2..

The principal results can be summarized as follows: (1) in the union sample both the industry and the SMSA union densities have positive and significant effects on the wages of union workers; and (2) in the nonunion sample the industry union density has a small and insignificant effect on nonunion wages, but the SMSA union density has a sizable and significant positive effect on the wages of nonunion workers.

Freeman and Medoff (1981) reported that using only industry union density as an explanatory variable, there is no spillover or a weak positive association between union density and the wages of nonunion workers (p.570). Our findings show that when both industry and regional (SMSA) union density variables are included, although industry union density has no sizable effect on nonunion wages there is a significant effect of geographic area (SMSA) union density on nonunion wages. This result implies that the threat effect is dominant in the inter-industry dimension within an SMSA rather than in the nationwide 3-digit intra-industry dimension. The inter-industry wage spillovers within an SMSA is also shown by the fact that SMSA union density has a sizable and significant positive effect on the wages of union workers.

Therefore, union environment in a geographic area should be considered as an important factor influencing both union wages and nonunion wages. The argument that there is no spillover from union density to the wages of nonunion workers is rejected.

In Table 3.2, estimated coefficients of the worker's personal characteristics variables are consistent with those obtained in many other studies. Education and labor market experience (=age-education-6) are positively and significantly associated with hourly earnings. When we compare the coefficient size of these variables between the union and nonunion sectors, we can also find that nonunion sector wages are more responsive to individual worker levels of education and experience, as reported by previous studies (Bloch and Kuskin, 1978; Duncan and Leigh, 1980).

Also, in both sectors there exists significant wage discrimination by sex and by race, although the degree of discrimination is smaller under unionism.

On the other hand, estimated coefficients on the industry and SMSA control variables are for the most part imprecise. Not only SMSA variables but also industry variables with the exception of the establishment size are not significant.

Table 3.2

Regression Estimates of the Wage Equations
with Industry Average Data on Establishment Size

variable	dependent variable: ln (hourly earnings)	
	union sample	nonunion sample
education	.027* (5.80)	.036* (6.95)
EXP(=age-education-6)	.012* (3.96)	.020* (5.39)
$EXP^2/100$	-.019* (3.18)	-.036* (4.97)
SEX (female=1)	-.233* (7.74)	-.280* (8.07)
RACE (white=1)	.080* (3.16)	.117* (2.98)
industry union density	.331* (2.57)	-.017 (0.12)
SMSA union density	.380* (2.75)	.608* (3.51)
establishment size/100	.000 (0.18)	.041* (2.23)
wages/value added	.139 (0.50)	-.535 (1.38)
concentration ratio	-.018 (0.13)	-.252 (1.36)
injury and illness rate	-.023 (0.51)	.038 (0.60)
SMSA population size (in logarithm)	.027 (1.28)	.006 (0.20)
SMSA proportion of blacks	-.192 (0.98)	-.159 (0.56)
SMSA unemployment rate(%)	-.013 (1.35)	-.010 (0.69)
R^2	.482	.549
SER	.263	.311
sample size	827	676

Note: Other variables included are a constant, the number of dependents, and dummy variables for marital status(3), veteran status, part-time, region(3), occupation(4), and industry(19).
SER denotes the standard error of the regression.
Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level.

Similar results have been reported in other empirical studies. Weiss (1966) pointed out that once the personal characteristics are introduced, the coefficient of the concentration ratio becomes non-significance. The implication seems to be that firms in concentrated industries do pay their employees more, but that they also get higher quality labor. Smith (1979) reported, in his review of the literature, that the effect of the injury and illness rate on wage rates is ambiguous in many cases. Concerning the SMSA variables, the coefficients of the unemployment rate and of the proportion of blacks are, as expected, negative but not significant in both samples. Similar results have been found in the work of Kahn (1980).

Although previous studies have generally found that large employers pay higher wages, Table 3.2 reveals that establishment size has a positive and significant effect on the wages of nonunion workers but no significant effect on the wages of union workers. This result seems consistent with the findings of previous studies that the estimated coefficient of establishment size is generally smaller under unionism (Freeman and Medoff, 1981a).

However, we have some reservations concerning data on establishment size because we have used industry average data on establishment size in Table 3.2. Ideally, employer size should be measured at the worker level, instead of being imputed on the basis of an industry average (Mellow, 1982).

The problem with the measurement of establishment size can be re-examined using the May 1979 CPS Supplement which includes an employer size question. The establishment size question is "How many persons are employed by your employer at the location where you work?". For this question, the survey allowed respondents to indicate one of five exclusive categories: less than 25, 25-99, 100-499, 500-999, and 1,000 and over.

After constructing four establishment size dummies (SIZE1= 25-99, SIZE2= 100-499, SIZE3= 500-999, SIZE4= 1,000+), we re-estimated the wage equations to investigate the effects of industry and SMSA union density on earnings. These results, using the micro establishment size data, are presented in Table 3.3. Note that the sample size in Table 3.3 is smaller than that of Table 3.2 because of missing values on establishment size for some of the observations.

The results in Table 3.3 are qualitatively similar to those in Table 3.2. Although there are small differences in the coefficients of the worker's personal characteristics, we get the same principal results: industry union density has a strong positive effect on union wages but not on nonunion wages, and SMSA union density has a sizable positive effect on both union and nonunion wages.

Turning to the establishment size dummies, we find significant effects only in the nonunion sector. The result that coefficients of the size dummies in the union sample are

Table 3.3

Regression Estimates of the Wage Equations with Micro Data on Establishment Size

variable	dependent variable: ln (hourly earnings)	
	union sample	nonunion sample
education	.030* (5.58)	.036* (6.36)
EXP(=age-education-6)	.009* (2.45)	.018* (4.33)
$EXP^2/100$	-.011 (1.63)	-.030* (3.77)
SEX (female=1)	-.225* (6.62)	-.296* (7.78)
RACE (white=1)	.044 (1.39)	.147* (3.25)
industry union density	.274* (2.06)	-.045 (0.28)
SMSA union density	.336* (2.19)	.493* (2.57)
SIZE1 (25-99)	-.035 (0.74)	.022 (0.55)
SIZE2 (100-499)	-.007 (0.16)	.057 (1.36)
SIZE3 (500-999)	.010 (0.20)	.155* (2.53)
SIZE4 (1,000+)	.069 (1.42)	.226* (4.45)
wages/value added	.297 (1.66)	.226 (0.58)
concentration ratio	-.121 (0.83)	.116 (0.62)
injury & illness rate	-.014 (0.35)	-.061 (1.09)
SMSA population (in log)	.038 (1.58)	-.006 (0.18)
SMSA proportion of blacks	-.177 (0.78)	-.097 (0.31)
SMSA unemployment rate(%)	-.001 (0.11)	-.017 (1.14)
R^2	.522	.584
SER	.255	.298
sample size	612	517

Note: Other variables included are the same as those in Table 3.2.
Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level.

smaller than those in the nonunion sample is consistent with the findings of previous studies that the impact of employer size on earnings is smaller under unionism. Again, the other industry and SMSA control variables are all insignificant in both union and nonunion samples.

In sum, the results presented here demonstrate that geographic area (SMSA) union density has a positive and significant effect on the wages of both union and nonunion workers, while industry union density has a strong positive effect only on the wages of union workers. These findings support the notion that there exist significant union-to-union and union-to-nonunion wage spillovers across industries within an SMSA. These results are robust with respect to the use of micro data or industry average data on establishment size. Thus, union environment in a geographic area (SMSA) should be regarded as an important factor influencing the earnings of both union and nonunion workers.

3.4.3 Correction for The Endogeneity of Union Status

Until now we abstracts from the problem of the endogeneity of union status in the wage determination model. Union status is likely to be endogenous in the wage equation in that workers' union membership decisions tend to be related to the anticipated wage differential in their best union/nonunion alternatives. It implies that the samples of union and nonunion workers are not drawn randomly from the population of workers, or union and nonunion workers with similar characteristics are systematically different. Then, ordinary least squares (OLS) estimates of the union and nonunion wage equations will be biased and inconsistent.

Several studies have attempted to deal with this issue using the sample selectivity and simultaneous-equations models.[20] Here, we re-examine the results obtained above using the sample selectivity model in which union status is treated as endogenous and selectivity bias resulting from unmeasured worker differences between the union and nonunion sectors is adjusted for. The model we use is based on those introduced by Lee (1978) and by Hirsch and Berger (1984). It consists of the following three equations:

$$\ln (WU_{ijk}) = a_1 X_{ijk} + a_2 IND_j + a_3 SMSA_k + a_4 T_j + a_5 T_k + \varepsilon^u{}_{ijk}$$

$$\ln (WN_{ijk}) = b_1 X_{ijk} + b_2 IND_j + b_3 SMSA_k + b_4 T_j + b_5 T_k + \varepsilon^n{}_{ijk}$$

$$U^*{}_{ijk} = c_1 X_{ijk} + c_2 IND_j + c_3 SMSA_k + c_4 T_j + c_5 T_k + c_6 (\ln WU_{ijk} - \ln WN_{ijk}) - \varepsilon_{ijk}$$

where i= individual, j= 3-digit industry, k= SMSA, and the disturbance terms are assumed to be normally distributed.

In the last equation let U^* be a latent union variable which is not observable but its dichotomous realization, U, is observed union status. This can be expressed as

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[20] For simultaneous equations models, see Ashenfelter and Johnson (1972) using aggregate industry data and Schmidt and Strauss (1976) using individual-worker data.

if $U^*_{ijk} > 0$, then individual i joins a union ($U_{ijk}=1$)

if $U^*_{ijk} \leq 0$, then he does not ($U_{ijk}=0$)

As long as the worker's decision to join a union depends on the wage differential, the error terms across equations will be correlated. In addition, the argument that union and nonunion samples are not drawn randomly implies that expected values of the error terms in the union and nonunion wage equations are not zero.

$$E(\varepsilon^u_{ijk} \mid U^*_{ijk} > 0) \neq 0$$

$$E(\varepsilon^n_{ijk} \mid U^*_{ijk} \leq 0) \neq 0$$

Heckman (1976) and Lee (1978) proposed a estimation method that the bias caused by sample selectivity is treated as a standard omitted variable problem and is controlled by inserting values of the "inverse Mill's ratio" (also called "selectivity variable") into the union and nonunion wage equations. In this model the truncation effect associated with sample selectivity in each sector may be written:

$$E(\varepsilon^u_i \mid U^*_i > 0) = \sigma_{\epsilon u} \, [-f(\hat{U}_i) \,/\, F(\hat{U}_i)]$$

$$E(\varepsilon^n_i \mid U^*_i \leq 0) = \sigma_{\epsilon n} \, [f(\hat{U}_i) \,/\, 1-F(\hat{U}_i)]$$

where f and F are, respectively, the density and distribution functions of a standard normal variable, and $\sigma_{\epsilon u}$, $\sigma_{\epsilon n}$ are covariances between the error term of the union status equation and the wage equation error terms (we drop the subscripts j, k for notational convenience). The estimate $\hat{U}_i$

is obtained using probit analysis of a reduced form union status equation which includes all exogenous variables, but not ($\ln WU_i - \ln WN_i$).

Then, the selectivity bias adjusted wage equations which are including the inverse Mill's ratios can be expressed:

$$\ln (WU_i) = a\ Z_{ui} + \lambda_u\ [f(\hat{U}_i)\ /\ F(\hat{U}_i)] + \varepsilon'_{ui}$$

$$\ln (WN_i) = b\ Z_{ni} + \lambda_n\ [-f(\hat{U}_i)\ /\ 1-F(\hat{U}_i)] + \varepsilon'_{ni}$$

where Z is a vector including all the independent variables (X, IND, SMSA, and T_j, T_k). Note that the adjusted error terms, ε'_{ui} and ε'_{ni}, will have expected values of zero.

The regression estimates of the wage equations using the sample selectivity approach are presented in Table 3.4.[21] In this model estimation of the coefficients of the inverse Mill's ratios provides a test of the null hypothesis that observed union and nonunion wages are randomly selected samples from the population wage distribution. The inverse Mill's ratio coefficients reported near the bottom of Table 3.4 show that there exists significant sample selectivity for the nonunion sector[22], while no significant selectivity is present in the union sector.

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[21] The regression results of the union status equation are reported in Table 4.1 of Chapter 4.

[22] A negative coefficient estimate, coupled with a selectivity variable (inverse Mill's ratio) defined to be negative, indicates positive selectivity for the nonunion sector.

Table 3.4

Regression Estimates of the Wage Equations using the Sample Selectivity Approach

variable	dependent variable: ln (hourly earnings)	
	union sample	nonunion sample
education	.028* (5.89)	.035* (5.03)
EXP	.010 (1.82)	.006 (0.87)
$EXP^2/100$	-.015* (1.99)	-.015 (1.24)
SEX (female=1)	-.219* (5.88)	-.209* (3.76)
RACE (white=1)	.106* (2.55)	.259* (3.30)
industry union density	.230 (1.22)	-.512 (1.76)
SMSA union density	.292 (1.57)	.117 (0.37)
establishment size/100	-.007 (0.51)	.000 (0.00)
wages/value added	.175 (0.63)	-.023 (0.04)
concentration ratio	.010 (0.75)	-.063 (0.25)
injury & illness rate	-.037 (0.74)	-.042 (0.48)
SMSA population(in log)	.035 (1.50)	.053 (1.16)
SMSA proportion of blacks	-.212 (1.08)	-.321 (0.87)
SMSA unemployment rate(%)	-.011 (1.19)	-.007 (0.42)
[inverse Mill's ratio]	-.104 (0.71)	-.492* (2.23)
sample size	827	676

Note: Estimation is done by the LIMDEP program (Selection Model). The inverse Mill's ratio is defined as $[f(\hat{U}_i)/F(\hat{U}_i)]$ for the union sector and as $[-f(\hat{U}_i)/1-F(\hat{U}_i)]$ for the nonunion sector.
Other variables included are the same as those in Table 3.2.
Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level.

As shown in Table 3.4, after correcting for the selectivity bias, the effects of union density become insignificant. Neither of industry and SMSA union density has a systematic impact on either sector wages. Note that the reduced form union status equation is used as a criterion of sample selection. In the reduced form union status equation, as seen in Table 4.1 of Chapter 4, both industry and SMSA union densities have significant effects on the likelihood of union membership.

In Table 3.4 the result that industry union density has no significant effect on union wages is not consistent with the findings of previous studies. Also, the result that the coefficients of EXP and establishment size are not significant in either sample of the union and nonunion sectors is not consistent with the general findings of earlier empirical work. Lee (1978), using the sample selectivity approach, reports that these variables have a positive and significant effect on wage rates.

In the literature, it is argued that the empirical results of the sample selectivity model are much more sensitive to changes in specification and choice of data set than those of the OLS model (Lewis, 1986; Freeman and Medoff, 1981b). Since we have some reservations on the empirical validity of the results in Table 3.4, we would not conclude that the results in Tables 3.2 and 3.2 are overturned by the sample selectivity approach.

3.5 Conclusion

Analysis of the May 1979 CPS data, with industry and SMSA union density as explanatory variables, yields cross-section evidence of the union-to-nonunion wage spillovers by supporting the notion that geographic area (SMSA) union density has a positive and significant effect not only on union wages but also on nonunion wages. We also find that industry union density has a positive effect only on the wages of union workers. Note that our sample is restricted to manufacturing production workers who reside in the 44 largest SMSAs.

The threat effect of unionization may exist for both the industry and the geographic area. However, the findings in this study imply that the threat effect for the nonunion sector is more dominant in the areawide inter-industry dimension rather than in nationwide intra-industry (3-digit) dimension, and thus geographic area (SMSA) union density should be regarded as an important factor in the wage determination model. To identify the threat effect, a case study approach might well be necessary. Foulkes' study (1980) of large nonunion firms also provides examples of such firms copying union wages within a geographic area (see also Freeman and Medoff, 1984).

A possible limitation on this conclusion is that the approach we use is a partial equilibrium analysis. In a general equilibrium model, when union wage gains lead to reductions of employment in the union sector, the nonunion sector must absorb both the displaced labor and capital. In such models, union wage gains are likely to lead in the long run to a downward pressure on nonunion wages through the so-called crowding effect.[23] However, we may reconcile the threat effect and the crowding effect on wages by considering the timing. The threat effect takes place in the short run, while the crowding effect takes longer to be realized if many employers reduce the quantity of labor demanded by not replacing retirees, rather than through direct layoffs. If this is the case, the cross-sectional analysis may not fully capture the crowding effect.

Another limitation to our conclusion is that the impact of the threat effect may change over time as the union environment changes or the tastes of unionism by employees and employers change. Similarly, the impact of the crowding effect may vary over the business cycle. It is suggested that the crowding effect would be stronger during periods of expanding employment than during recessions (Kahn, 1979). Thus, it may be argued that the effect of union density on nonunion wages is likely to change over time.

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[23] This argument may not hold if the union sector is more heavily capital intensive (Freeman and Medoff, 1984).

CHAPTER 4

THE ENDOGENEITY OF UNION STATUS IN WAGE DETERMINATION: A CROSS-SECTIONAL ANALYSIS

Most of U.S. studies of unionism have focused on the effects of unions on wage rates. For some time, it has been recognized that not only can unions procure wage gains but also workers' decisions to join a union are likely to be influenced by the wage differential. However, the choice of treatment of unionism as endogenous or exogenous in the wage equation is still an unsettled issue in the literature. In this chapter, the endogeneity of union status will be examined using cross-section, individual-worker data.

4.1 Two Different Views on The Endogeneity of Unionism

Most of empirical studies of unionism have been concerned with estimates of union/nonunion wage differentials. Usually, estimation of union wage effects is done by ordinary least squares (OLS) under the assumption that unionism is exogenous in wage determination. However, if union status is not determined independently of the wage rate, OLS estimation will lead to simultaneity bias.

Recently it is widely recognized that unionism and wage rates are determined simultaneously, but there still exist some controversies over the endogeneity of unionism in the wage model. In the literature, the following arguments have been presented in support of the endogeneity of union status.

First, individuals choose to work in the union sector depending on the wage differential between union and non-union, and the conditions of employment in each sector (Lee, 1978). Schmidt and Strauss (1976) also argue that employees voting in union representation elections consider the expected union and nonunion wages in casting their votes.

Secondly, union organizers tend to concentrate their organizing efforts in industries which have a greater ability to pay higher wages. Johnson (1975) suggests that the extent of unionization is likely to be greater in industries with higher wages because the costs of organization are relatively low in such industries due to the greater job specific training and lower turnover that may characterize higher wage industries. Moreover, Krueger and Summers (1987) argue that historically, high-wage industries already paid relatively high wages before the advent of wide-scale unionization in the U.S. For example, "the Big Three automobile manufacturers in the U.S. were wage leaders prior to successful union organization of General Motors and Chrysler in 1937 and Ford in 1941" (p.36).

Thirdly, union wage gains are likely to create an excess supply of workers to union jobs, and then some workers will queue for work in the union sector. If profit-maximizing firms select more productive workers from this queue, there would exist unmeasured differences in labor quality between union and nonunion workers, and thus the wage differential between truly similar workers would be smaller than the observed wage differential.

However, the alternative view that unionism should be treated as exogenous in the wage determination model is also presented in the literature. Mitchell (1980) argues that the issue is whether or not currently observed wage rates can be thought to have a significant influence on currently observed unionization. In the U.S. the pattern of unionization had been mainly established by the end of the 1940s, and industries that were heavily unionized then remain heavily unionized today. Thus, treating unionization as if it were a continuous function of today's wage rates creates a conceptual problem.

Bloch (1982) also contends that plausibility is not sufficient reason to infer the endogeneity of union status in a current wage determination model. First, movements from the nonunion into union sector may be costly. If occupation or residence is exogenous in a wage determination model, so is union membership. Secondly, if unions consider relative or absolute wages in choosing their organization targets, the appropriate wage in a union status equation is a wage existing prior to unionization, instead of the current wage. Finally, he points out that the basic statistical problem is whether the union membership and wage equations are identified. A necessary condition for identification of the wage equation is the existence of a variable that shifts the union membership but not the wage equation. If one attempts to identify the equation by arbitrarily excluding one of the variables, the estimates will not be reliable due to specification bias.

The points raised by Bloch and Mitchell are not necessarily for the exogeneity of unionism in the wage determination model. Rather, some of them may be considered as the arguments for more careful specification. However, it is not clear whether the simultaneous equations approach is empirically superior to the single equation approach in which unionism is treated as exogenous. In the next section empirical results of these approaches will be discussed.

4.2 Review of Previous Studies

4.2.1 Simultaneous Equations Approach

In the literature several studies have examined cross-section data to test for the endogeneity of unionism. Studies that use a simultaneous equations model of unionism and the wage level and find a negligible union wage effect include Ashenfelter and Johnson (1972), and Schmidt and Strauss (1976).

Ashenfelter and Johnson (1972) specify a three-equation model in which wage rates, unionization, and labor quality (education) are endogenous. They find, using aggregate cross-section data of manufacturing industries, that unionization has no significant impact on the wage rate, but the wage rate has a positive and significant effect on the extent of unionization. Schmidt and Strauss (1976), using individual-worker data, also reach similar results that unionism has no systematic impact on earnings while the earnings level has a positive and significant effect on the probability of union membership.

The positive unionism-wage relationship appears to support the argument of the endogeneity of unionism in wage determination. However, there is disagreement regarding the appropriateness of including a wage or earnings variable in the unionism equation. The use of the earnings level can be justified under the assumption that union services are a normal good, but this assumption seems not consistent with the fact that union membership is particularly unlikely for high-income workers. Instead, the union/nonunion wage differential seems to be a more appropriate variable because it can be regarded as an index measuring the benefit associated with union representation. In addition, as pointed out by Freeman and Medoff (1981b), the simultaneous equations approach have produced widely fluctuating results from arbitrary decisions over what to exclude in the process of identifying equations.[24]

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[24] Schmidt (1978), using the same data set used by Schmidt and Strauss (1976), reports that the earnings differential has no significant effect on the probability of union membership.

4.2.2 Sample Selectivity Approach

More recent studies (Lee, 1978; Duncan and Leigh, 1980; Hirsch and Berger, 1984) have viewed observed union status as the endogenous outcome of a choice process in which workers' decisions are related to the anticipated wage differential between their best union and nonunion alternatives. In other words, a worker will become a union member if the wage differential exceeds his or her reservation union wage premium which is regarded as a function of the characteristics of the person and the costs of becoming a union member (Lee, 1978).

These studies postulate a three-equation model determining union membership as well as union and nonunion wages and estimate the equations using the sample selectivity approach in which union status is treated as endogenous and selectivity bias resulting from unmeasured worker differences between the union and nonunion sectors is adjusted for. This type of empirical work can provide two kinds of criteria to test for the endogeneity of union status in the wage determination model.

The first is a test whether sample selectivity is present in the wage equation for each sector. If significant selectivity is present in either sector, we may conclude that workers' decisions of union status are associated with the wage comparisons between their best union and nonunion alternatives. No sample selectivity would suggest that union status is exogenously determined in the wage model.

Secondly, a more direct criterion is a test whether the worker's union status is influenced by his or her anticipated wage differential. If the wage differential has a systematic impact on the likelihood of union membership, it would support the argument that unionism should be treated as endogenous in the wage determination model. As long as the relative wage differential is regarded as an index measuring the benefit from union representation, the wage differential would have a positive effect on the likelihood of union membership.

Empirical results of the sample selectivity model are not consistent between studies. Some studies indicate that significant sample selectivity is present in either union or nonunion sector (Lee, 1978; Duncan and Leigh, 1980), but some studies do not (Duncan and Leigh, 1985). Also, some studies report that the union/nonunion wage differential has a positive and significant effect on the likelihood of union membership (Lee, 1978; Duncan and Leigh, 1980), but some studies do not (Hirsch and Berger, 1984). Thus, the empirical evidence on the hypothesis that unionism is endogenously determined in the wage model seems not clear.

In addition, the sample selectivity approach yields estimated union/nonunion wage differentials that appear to show substantial instability in the face of small changes in the model or the sample analyzed. Lee (1978), using a 1967 sample of full-time operatives, obtained estimates of the union/nonunion wage differential of 0.14 using OLS and 0.12

after adjusting for selectivity bias. In contrast, Duncan and Leigh (1980) obtained corresponding estimates of 0.13 and 0.30 with a sample of older white men from the 1969 National Longitudinal Survey. Hirsch and Berger (1984) reports corresponding estimates of 0.11 and 0.53, using a 1971 sample of manufacturing blue-collar workers from the Current Population Survey.[25] After a full survey of such studies, Lewis (1986) conclude that estimates obtained by the sample selectivity approach are so sensitive to sample and specification that estimates obtained by OLS estimation are more reliable in the support of union wage effects.

On the other hand, in a three-equation model determining union membership and union and nonunion wages, Duncan and Leigh (1985) explicitly test the null hypothesis of the exogeneity of union status using a test procedure proposed by Hausman (1978). In their study the estimator which is consistent and efficient under the null hypothesis is an ordinary least squares (OLS) estimator. This estimator is biased and inconsistent under the alternative hypothesis. When an instrumental variables (IV) estimator is chosen to compare with the OLS estimator, the Hausman test rejects the null hypothesis of the exogeneity of union status using 1969 National Longitudinal Survey data.

However, the inverse Mill's ratio procedure, reported in

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[25] For comparisons of estimates of wage differentials see Hirsch and Addison (1986).

the same paper by Duncan and Leigh (1985), indicates that no sample selectivity is present in either sector using the same sample. This result does not support the endogeneity of union status and contradicts that of the Hausman test. Duncan and Leigh prefer the Hausman test procedure to the inverse Mill's ratio procedure in that selectivity may be present but not revealed by the inverse Mill's ratio procedure because of its sensitivity to the underlying assumption that the error term is normally distributed.

Thus, in the literature there appears to be little consensus regarding the empirical validity of the endogeneity of unionism in the wage determination model.

In the next section, we will re-examine cross-section evidence of the endogeneity of union status using the sample selectivity model developed in Chapter 3. Our model differs from the earlier empirical work in that a more extensive set of explanatory variables are included and in that our sample is restricted to manufacturing production workers who reside in the 44 largest SMSAs.

4.3 The Sample Selectivity Model of Union Status

4.3.1 Econometric Specification and Data

The hypothesis of the endogeneity of union status is based on the argument that workers' decisions to join a union is likely to be related to the anticipated wage differential between their best union and nonunion alternatives. In other words, a worker will join a union if the wage differential exceeds his reservation union wage premium. This argument implies that the samples of union and nonunion workers are not drawn randomly from the population of workers, or union and nonunion workers with similar characteristics are likely to be systematically different. Following previous studies, we deal with this issue using the sample selectivity model in which selectivity bias is adjusted for by inserting the "inverse Mill's ratio" (see Heckman, 1976; Lee, 1978).

In this section the hypothesis of the endogeneity of union status will be tested using cross-section, individual-worker data. The sample selectivity model has already been described in Chapter 3, and will be outlined here with a minimum of discussion. As seen earlier, the model we use consists of the following three equations determining union and nonunion wage rates and union status:

$$\ln(WU_{ijk}) = a_1\,X_{ijk} + a_2\,IND_j + a_3\,SMSA_k + a_4\,T_j + a_5\,T_k + \varepsilon^u{}_{ijk}$$

$$\ln(WN_{ijk}) = b_1\,X_{ijk} + b_2\,IND_j + b_3\,SMSA_k + b_4\,T_j + b_5\,T_k + \varepsilon^n{}_{ijk}$$

$$U^*{}_{ijk} = c_1\,X_{ijk} + c_2\,IND_j + c_3\,SMSA_k + c_4\,T_j + c_5\,T_k + c_6\,(\ln WU_{ijk} - \ln WN_{ijk}) - \varepsilon_{ijk}$$

where X_{ijk}, IND_j, $SMSA_k$ are vectors of personal, industry, and SMSA characteristics, respectively (the variables included are the same as in Chapter 3); T_j, T_k are industry and SMSA union density; U^* is a latent union variable and its dichotomous realization, U, is observed union status; and the disturbance terms are assumed to be normally distributed.

The independent variables included in the union and nonunion wage equations are explained in detail in Chapter 3, and we will not discuss them again. Note that our model, unlike previous studies, includes SMSA union density and other SMSA characteristics variables in the wage equations.

Concerning the determinants of union status, a number of cross-sectional studies have examined the effects of personal characteristics on the likelihood of union membership. Union status has been found to be systematically related to the personal characteristics such as labor market experience (age), sex, race, region, and occupations.

Also, industry and SMSA characteristics can be regarded as important determinants of union status because they are likely to reflect the benefits and costs associated with unionism. Establishment size, industry concentration, capital intensity, and job hazards tend to be positively related to the likelihood of union membership.

Larger establishments are likely to involve lower organizing costs for a union, leading to a positive association between union membership and establishment size. Workers in highly concentrated industries are more likely to be unionized in that union organizers tend to focus on the industries with a greater ability to pay high wages. Greater capital intensity may lead to greater unionization because capital intensive firms are likely to have a more inelastic labor demand schedule which implies unions' ability to get higher wage gains. Job hazards may be positively related to unionism if workers believe that unions provide an effective collective voice for the job hazards and compensation mix (Hirsch and Addison, 1986).

Although the evidence of the effects of SMSA characteristics on unionism is sparse, it is reported that unionism is found to be more likely among workers in SMSAs than in non-urban areas. It seems to reflect differences in workers' attitudes toward or in organizing costs associated with union representation.

In this model the principal independent variable to test the endogeneity of union status is the union/nonunion wage differential for each worker. Other explanatory variables included in the union status equation are personal characteristics, four industry characteristics (establishment size, wages/value added, industry concentration ratio, injury and illness rate), three SMSA characteristics (population size, unemployment rate, proportion of blacks), and industry and SMSA union densities.[26]

The data used here is the cross-section, individual-worker data used in Chapter 3, that is, the May 1979 Current

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[26] Industry and SMSA union density variables are assumed to be exogenous to individual worker's union membership decision.

Population Survey (CPS) data, bridged with industry and SMSA characteristics. Our sample is restricted to production workers in manufacturing who reside in the 44 largest SMSAs, and the sample size is 1,503 workers (827 union workers and 676 nonunion workers).

4.3.2 Estimation Procedure

Actual estimation of the three-equation model is not straightforward because the error terms across equations are likely to be correlated. A statistical solution to this problem was developed by Heckman (1976) and by Lee (1978). It involves first estimating the union status equation by probit analysis, including all exogenous variables, but not ($\ln WU_{ijk} - \ln WN_{ijk}$). This is called a reduced-form union status equation which is obtained by substituting the wage equations into the union status equation. From this equation the inverse Mill's ratios are obtained to be included in the union and nonunion wage equations. The inclusion of the inverse Mill's ratio can yields selectivity-bias adjusted coefficient estimates of the union and nonunion wage equations.[27]

The estimated coefficients of the wage equations allow comparison of the anticipated wage differential between the union and nonunion sectors. From the estimated wage equations we can obtain the predicted value of ($\ln WU_{ijk} - \ln WN_{ijk}$) for

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[27] For detailed explanation of this procedure see Section 3.4.3 of Chapter 3.

each worker. Note that the predicted values can be calculated for all workers although coefficient estimates of the wage equations are obtained separately from the union and nonunion samples. The predicted value ($\ln \hat{W}U_{ijk} - \ln \hat{W}N_{ijk}$) serves as an estimate of the i-th worker's anticipated wage differential between his/her union and nonunion alternatives. Substituting ($\ln \hat{W}U_{ijk} - \ln \hat{W}N_{ijk}$) for ($\ln WU_{ijk} - \ln WN_{ijk}$), estimates of the "structural" parameters of the union status equation can be obtained using probit analysis. If the coefficient of the predicted wage differential is positive and significant, it can be regarded as evidence of the endogeneity of union status in the wage determination model.

4.4 Empirical Results

Probit (maximum likelihood) estimates of the coefficients in the reduced-form and structural-form union status equations are presented in Table 4.1. As Lee (1978) pointed out, one of the important advantages of the sample selectivity model is that we can estimate the structural form of the union status equation. The reduced-form union status equation is used to obtain the values of the inverse Mill's ratios included in the wage equations, and thus it is not much of our concern.

In the structural-form equation, the most powerful factor determining the worker's union status is the union/nonunion wage differential. As shown near the bottom of Table 4.1, the coefficient of the predicted wage differential is positive and significant, and its coefficient size is much greater than those of other variables. This result indicates that the worker's union status is positively associated with the anticipated wage differential between his or her union and nonunion alternatives. Therefore, the hypothesis of the endogeneity of union status cannot be rejected by the data.

Regarding the endogeneity hypothesis, it will be useful to examine the meaning of coefficient estimates of the inverse Mill's ratios included in the wage equations (these are reported in Table 3.4 of Chapter 3, p.87). As explained earlier, estimation of the coefficients of the inverse Mill's ratio provides a test of the null hypothesis that observed union and nonunion wages are randomly selected samples from the population wage distribution.[28] Rejecting the null hypothesis can be interpreted as alternative evidence of the endogeneity of union status.

The inverse Mill's ratio coefficients reported in Table 3.4 reveal that positive and significant sample selectivity is present for the nonunion sector, i.e., $E(\varepsilon^n_i \mid U_i=0) > 0$, but not for the union sector. The positive selectivity means that the individual chooses the relevant sector that pays him better than the average worker with the same characteristics,

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[28] In our model a worker is assumed to join a union if the wage differential exceeds his/her reservation union wage premium. Then, observed union and nonunion wages are likely to be non-randomly selected samples.

Table 4.1

Regression Estimates of the Union Status Equation: Probit Reduced-Form and Structural-Form Equations

variable	dependent variable: U_i			
	reduced-form		structural-form	
education	.003	(0.20)	.015	(0.93)
EXP	.042*	(3.78)	.036*	(3.27)
$EXP^2/100$	-.060*	(2.80)	-.061*	(2.88)
SEX (female=1)	-.271*	(2.53)	-.162	(1.59)
RACE (white=1)	-.462*	(4.44)	-.187	(1.64)
part-time	-.714*	(4.47)	-.551*	(3.51)
industry union density	.018*	(4.03)	.005	(1.25)
SMSA union density	.018*	(3.38)	.013*	(2.56)
establishment size/100	.119*	(2.47)	.075*	(2.23)
wages/value added	-.011	(1.01)	-.002	(0.33)
concentration ratio	-.006	(1.19)	-.003	(0.92)
injury & illness rate	.003	(1.69)	.001	(0.91)
SMSA population(log)	-.144	(1.62)	-.133	(1.30)
SMSA unemployment rate	-.002	(0.49)	-.000	(0.11)
SMSA % of blacks	.004	(0.48)	.003	(0.35)
[Region: base=West]				
Northeast	.378*	(2.95)	.196	(1.56)
North-Central	.262*	(2.27)	.057	(0.49)
South	-.089	(0.60)	-.010	(0.07)
[Occupation: base=service]				
craftsmen	.411	(1.89)	.448*	(2.08)
operatives	.461*	(2.17)	.353	(1.68)
transport operatives	.902*	(3.24)	1.009*	(3.66)
laborers	.848*	(3.23)	.373	(1.38)
$(\ln \hat{W}U_i - \ln \hat{W}N_i)$			1.906*	(5.70)
sample size	1,503		1,503	

Note: Other variables included are a constant, dummies for marital status (3), veteran status, and the number of dependents (19 industry dummies are also included in the reduced-form equation).
The numbers in parentheses are absolute values of asymptotic t-statistics.
* significant at the .05 level.

or more formally, only the upper section of the wage distribution is observed given worker characteristics (Lee, 1978). Note that the absence of positive selectivity in the union sector is not inconsistent with evidence of the strong impact of the predicted wage differential on union status. It is proved that coefficient estimates of the inverse Mill's ratio need not be positive or negative unless the error terms ε^u, ε^n, and ε are independent (Lee, 1978) and that they are sensitive to the underlying assumption that the error terms are normally distributed (Lee,1982).

Turning to the effects of other variables on union status, the results in Table 4.1 are generally consistent with earlier empirical work. Most studies do not find education to be a significant determinant of union status. This seems related to the fact that education tends to make workers more mobile and less dependent on internal union voice (Hirsch and Addison, 1986).

Table 4.1 also shows that there is a concave relationship between labor market experience (EXP= age-education-6) and union status, i.e., a positive coefficient on EXP and a negative coefficient on EXP^2. However, Farber and Saks (1980) find the likelihood of voting for a union to decrease with age using data from National Labor Relations Board (NLRB) elections. These different results may be interpreted as follows. Older workers may be less likely to vote for a union because they generally exhibit less positivity to changes in work environments, less mobility, and greater firm attachment. On the other hand, older workers may be more likely to be union members since union representation has been determined collectively by past workers and unions' seniority system is advantageous for them (Hirsch and Addison, 1986).

Previous studies generally find that women, whites, and part-time workers are less likely to be union members. It is reported that workers who dissatisfied with the current work environment are more likely to vote for a union (Farber and Saks, 1980). If there exists discrimination by race, non-whites may believe that they will be better in a union work environment with highly structured work rules and seniority provisions. It is also argued that union membership is less likely in job settings with less permanent labor force attachment or where part-time work schedules are present. Unionism of female and/or part-time workers would be less successful due probably to higher organizing costs. In Table 4.1 the coefficients of RACE (white=1) and SEX (female=1) are negative but not significant in the structural-form equation, after controlling for the predicted wage differential. This seems to imply that whites and females are more likely to be employed in the nonunion sector since it is more advantageous for them.

Concerning the impact of union density on the likelihood of union membership, industry union density has no significant effect, after controlling for the worker's predicted wage differential. SMSA union density, however, has a positive and significant effect on the worker's union membership. This implies that the workers in a heavily unionized region are more likely to join a union, all else the same.

Among the other industry and SMSA characteristics, our results show that the establishment size variable has a positive and significant effect on union status. It is consistent with the findings of earlier work that unionism is more likely in industries with larger establishments, reflecting lower organizing costs. However, other industry and SMSA variables have no systematic impact on the likelihood of union membership. Some studies have found that workers in highly concentrated industries are more likely to be unionized (Lee, 1978), but the results in Table 4.1 suggest that the concentration ratio has no systematic impact on union status after controlling for other factors. Similarly, Hirsch (1982), using industry data, finds no significant relationship between them after controlling for average firm sales.

As is well-known, regional differences in unionism are significant. Unionism is less prevalent in the southern and western regions than in the northern and central regions. Regional differences seem to be related to differences in attitudes toward unionism by employers and employees and differences in job structure not accounted for by other variables (Hirsch and Addison, 1986). It is also generally found that blue-collar workers are more likely to be union members. In Table 4.1, among those workers, craftsmen and transportation operatives are more likely to join a union after controlling for the predicted wage differential.

As a final exercise, we compare the magnitude of the wage differentials obtained from the sample selectivity and ordinary least squares (OLS) models. The approach taken here is to weight differences in estimated coefficients in the union and nonunion wage equations by mean values of worker characteristics. The mean logarithmic wage differential (d) may be written as:

$$d = (\hat{a}_u - \hat{b}_n)\ \bar{Z}$$

where $\ln WU_i = a_u Z_u + \varepsilon_{ui}$, $\ln WN_i = b_n Z_n + \varepsilon_{ni}$, and $\bar{Z}$ can represent either union means, nonunion means, or all-worker means. If we use all-worker means, it implies that "For a worker with average characteristics, what is the predicted wage differential between his working in the union and nonunion sectors?". The corresponding proportionate wage differential (D) can be written as D = exp(d)-1.

Using all-worker means, we obtain a proportionate wage differential estimate of 0.250 after correcting for sample selectivity bias, as compared to an OLS estimate of 0.105. Note that the selectivity-bias-adjusted estimate included the values of the inverse Mill's ratio in calculating the wage differential.

In the literature it is generally concluded that the reasonable size of average wage differentials is about

0.10-0.15, more probably near 0.10 when we consider the possibility of omitted explanatory variables (Hirsch and Addison, 1986; Lewis, 1986).[29] In our study the OLS estimate appears to be more reasonable as an average wage differential between otherwise similar union and nonunion workers.

In this context, we have also conducted a comparable probit analysis of union status determination using OLS estimates of the predicted wage differential. The results, reported in Appendix Table A.5, show a quite similar effect of the predicted wage differential on the likelihood of union membership (coefficient estimate= 1.685, t-ratio= 3.55).

Therefore, the qualitative results presented in Table 4.1 would generally not be overturned by differences in the magnitude of wage differential estimates.

4.5 Conclusion

This study extends recent investigations of the joint determination of union status and union and nonunion wage rates by including a more extensive set of explanatory variables and finds significant evidence of the endogeneity of union status in wage determination. The results of the sample selectivity model indicate that positive and significant sample selectivity is present for the nonunion sector and that

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[29] It is also reported that most estimates of the wage differential from the fixed-effect models using longitudinal data are close to 0.10.

the worker's predicted wage differential has a strong positive impact on the likelihood of union membership.

These findings do not refute the argument that workers' decisions to join a union are influenced by the wage comparison between their union and nonunion alternatives and that the potential union wage gain is an important determinant of union membership. Therefore, the hypothesis of the endogeneity of union status is supported by the data.

As pointed out by Hirsch and Berger (1984), union status is the outcome of not only private but also collective choice. If union representation has been determined collectively by the preferences of workers in the relevant industry or labor market, it would be more appropriate to include industry and SMSA characteristics in the model of union status determination because these characteristics are likely to reflect the benefits and costs associated with unionism. In this respect, our model can be regarded as more comprehensive than the previous ones which have focused mainly on personal characteristics.

Finally, note that the tastes of unionism by employees and employers may change over time. Changes in the union environment will shift not only the demand for union membership by workers but also the resistance of firms to unionism. In this study these effects are ignored since our analysis is cross-sectional and hence they are not expected to change in the short time period analyzed.

CHAPTER 5

SUMMARY AND CONCLUSIONS

Most of U.S. studies of unionism deal with the effects of unions on wage rates. The debate on the union wage effect concerns not only with what unions do for organized labor but also with what they do for unorganized labor. As explained earlier, there are two different views on the role of unions in wage determination.

The neo-classical position suggests that unions merely attempt to maintain a certain position relative to the nonunion sector and that nonunion wages are determined solely by market forces. Thus, unions are viewed as playing no significant role in aggregate wage determination. On the other hand, the institutional school argues that union wages are determined primarily by union power independently of the nonunion sector and that union wage behavior influences nonunion wage determination through the spillover mechanism. Hence, unions would have a significant impact on aggregate wage determination independently of market forces.

In Chapter 2, we have focused on whether union/nonunion wage spillovers exist and in which direction they run using time-series data of union and nonunion wages (1958-89). When we explicitly test wage interdependence between the union and nonunion sectors, the empirical results indicate that union wage behavior spills over into nonunion wages and vice versa, that is, feedback relationship exists between union and nonunion wages.

Therefore, these findings support the argument that the union and nonunion sectors emulate each other in the wage determination process, and thus the spillover effect, in addition to market forces, should be regarded as an important factor determining both union and nonunion wages.

Turning to the cross-sectional analysis, we have examined the effects of the extent of union organization on the wages of union and nonunion workers. In this part, our main concern is to determine whether or not unionism has a positive effect on the wages of nonunion workers because it can be regarded as cross-section evidence of the spillover effect from the union to nonunion sector.

The empirical results, based on the May 1979 CPS data, demonstrate that geographic area (SMSA) union density has a positive and significant effect not only on union wages but also on nonunion wages whereas industry union density has a positive impact only on union wages. These results imply that the threat effect for the nonunion sector is determined by area-wide unionism rather than nation-wide industry unionism.

These findings suggest that labor may not be tied to one specific industry (three-digit) and that movements between regions may be more costly than movements between industries. In other words, the labor market in the United States may be segmented more significantly by region than by industry.

Therefore, geographic area union density should be regarded as an important factor influencing wage rates, and the existence of the union-to-nonunion spillover effect is also supported, in part, by the cross-section data.

On the other hand, some authors have sought to take account of the reverse causality from wage rates to unionism. It means that not only can unions procure wage gains but also the worker's union status will be affected by the wage comparison between the union and nonunion sectors. If union status and wage rates are determined simultaneously, different estimation procedures will be required to determine a union/nonunion wage differential taking account of such endogeneity, and union wage effects will be interpreted as more limited than when unionism is regarded as exogenous.

The endogeneity of union status in wage determination is investigated in Chapter 4, using the sample selectivity model. The empirical results show that the workers in the nonunion sector are non-randomly (positively) selected samples from the population and that the worker's predicted wage differential has a strong positive impact on the likelihood of union membership. Here, the positive selectivity implies that the individual worker chooses the relevant sector that pays him better than the average worker with the same characteristics.

These findings do not refute the hypothesis that workers' decisions to join a union are influenced by the wage comparison between their union and nonunion alternatives, and hence the endogeneity of union status is supported by the data.

To summarize, both time-series and cross-sectional analyses support the notion that the role of unions in wage determination is not negligible in that union wage behavior spills over into nonunion wages and in that geographic unionism has a positive impact on both union and nonunion wages. Thus, the neo-classical hypothesis that nonunion wages are determined solely by market forces is rejected.

However, the role of unions in wage determination is limited within the market mechanism in that causality also runs from wage rates to unionism. Empirical results suggest that union wages are also influenced by the nonunion sector and that unionism occurs endogenously in the wage determination model. Thus, the institutionalist hypothesis that unionism plays an important role in wage determination independently of market forces is rejected.

In this context, the results obtained in this study support an intermediate position which lies between the neo-classical view and the institutional school.

APPENDIX

Table A.1

Annual Data Series of Wages and Union Density: 1958-89
(WU= union wage, WN= nonunion wage)

year	WU	WN	WU/WN	union density(%)
1958	100.0	100.0	100.0	30.3
1959	103.4	103.3	100.1	29.0
1960	106.9	105.9	100.9	28.6
1961	109.8	106.9	102.7	28.5
1962	112.7	108.7	103.7	30.4
1963	115.6	111.7	103.5	30.2
1964	118.1	113.9	103.7	30.2
1965	121.6	117.6	103.4	30.1
1966	125.4	122.2	102.6	29.6
1967	130.5	127.8	102.1	29.9
1968	137.0	134.2	102.1	29.5
1969	143.8	141.0	102.0	28.7
1970	152.0	148.2	102.6	29.6
1971	161.3	155.2	103.9	29.1
1972	169.7	162.9	104.2	28.8
1973	180.2	172.0	104.8	28.5
1974	194.6	185.8	104.8	28.3
1975	210.6	197.5	106.6	28.9
1976	226.8	210.4	107.8	27.9
1977	245.3	226.2	108.4	26.2
1978	266.9	244.1	109.3	25.1
1979	291.8	263.4	110.8	24.5
1980	323.9	284.3	113.9	23.2
1981	352.8	307.9	114.6	22.6
1982	373.4	325.1	114.9	21.9
1983	386.9	340.3	113.7	20.7
1984	402.7	356.1	113.1	19.4 (19.1)
1985	415.5	370.7	112.1	18.6 (18.2)
1986	422.9	385.9	109.6	18.1 (17.7)
1987	435.7	399.0	109.2	17.6 (17.3)
1988	445.2	413.1	107.8	17.3 (17.0)
1989	460.7	429.7	107.2	16.9 (16.6)

Note: The 1958-76 wage series are calculated from the BLS, Current Wage Developements, May, 1981, Table 3. The 1977-89 wage series are from the BLS, Current Wage Developments, February, each year. These wage series are year-end figures. The 1958-84 union density data are from Hirsch and Addison (1986). Union density figures in parentheses are from the BLS, Current Wage Developments, February, each year. The 1985-89 union density figures outside parentheses are calculated proportionately from the figures in parentheses.

Table A.2

Results of the Granger Test: 1960-89
Union Wage Equation

(i) $WU_t = \Sigma\, a_i\, WU_{t-i} + \Sigma\, b_i\, WN_{t-i} + \alpha + \beta\, t$ $(i = 1,2,..,m)$

H_0: $b_i = 0$ for all i

variable	m = 2	m = 3	m = 4
WU_{t-1}	1.276* (5.15)	1.296* (3.93)	1.062* (3.35)
WU_{t-2}	-.604* (2.31)	-.543 (1.33)	-1.009* (2.51)
WU_{t-3}		-.091 (0.28)	.471 (1.03)
WU_{t-4}			-.162 (0.41)
WN_{t-1}	1.169* (2.46)	1.154 (2.02)	1.743* (3.04)
WN_{t-2}	-.822 (1.80)	-.965 (1.41)	-1.207 (1.87)
WN_{t-3}		.175 (0.32)	1.119 (1.73)
WN_{t-4}			-.940 (1.58)
(F-statistic)	F_b= 4.65*	F_b= 2.45	F_b= 4.16*
t	.056 (0.17)	.046 (0.10)	-.729 (1.29)
constant	-2.725 (1.15)	-3.091 (1.05)	-4.367 (1.06)
$\bar{R}^2$	.9995	.9994	.9995
Durbin-h[a]	ρ=0	ρ=0	ρ=0

Note: [a] Durbin-h statistic cannot be computed. So, we use an alternative test (ρ=0) by creating a lagged residual variable. Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level.

Table A.3

Results of the Granger Test: 1960-89
Nonunion Wage Equation

(ii) $WN_t = \Sigma a'_i WN_{t-i} + \Sigma b'_i WU_{t-i} + \alpha' + \beta' t$ $(i=1,2,..,m)$

H_0: $b'_i = 0$ for all i

variable	m = 2	m = 3	m = 4
WN_{t-1}	1.177* (4.28)	1.029* (3.25)	1.167* (3.16)
WN_{t-2}	-.083 (0.32)	-.333 (0.88)	-.399 (0.96)
WN_{t-3}		.334 (1.11)	.617 (1.48)
WN_{t-4}			-.278 (1.48)
WU_{t-1}	.357* (2.49)	.502* (2.74)	.447* (2.19)
WU_{t-2}	-.453* (2.99)	-.481* (2.12)	-.603* (2.33)
WU_{t-3}		-.063 (0.35)	.070 (0.24)
WU_{t-4}			-.025 (0.10)
(F-statistic)	$F_{b'}$= 4.58*	$F_{b'}$= 3.46*	$F_{b'}$= 2.66
t	.485* (2.60)	.697* (2.81)	.524 (1.43)
constant	-1.468 (1.07)	-1.476 (0.91)	-1.672 (0.63)
$\bar{R}^2$	.9998	.9998	.9997
Durbin-h[a]	ρ=0	ρ=0	ρ=0

Note: [a] Durbin-h statistic cannot be computed. So, we use an alternative test (ρ=0) by creating a lagged residual variable. Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level.

Table A.4

Full Regression Results of the Wage Equations:
May 1979 CPS Data

variable	dependent variable: ln (hourly earnings)	
	union sample	nonunion sample
education	.027* (5.80)	.036* (6.95)
EXP(age-education-6)	.012* (3.96)	.020* (5.39)
$EXP^2/100$	-.019* (3.18)	-.036* (4.97)
SEX (female=1)	-.233* (7.74)	-.280* (8.07)
RACE (white=1)	.080* (3.16)	.117* (2.98)
married[a]	.097* (3.18)	.173* (4.64)
separated[a]	.010 (0.17)	.103 (1.48)
divorced,widowed[a]	.084* (2.02)	.213* (3.82)
no. of dependents	-.003 (0.36)	-.042* (3.89)
veterans	.041 (1.74)	.053 (1.47)
part-time	-.035 (0.61)	-.076 (1.79)
industry union density	.331* (2.57)	-.017 (0.12)
SMSA union density	.380* (2.75)	.608* (3.51)
[industry characteristics]		
establishment size/100	.000 (0.18)	.041* (2.23)
wages/value added	.139 (0.50)	-.535 (1.38)
concentration	-.018 (0.13)	-.252 (1.36)
injury & illness	-.023 (0.51)	.038 (0.60)
[SMSA characteristics]		
population size (log)	.027 (1.28)	.006 (0.20)
unemployment rate(%)	-.013 (1.35)	-.010 (0.69)
proportion of blacks	-.192 (0.98)	-.159 (0.56)
[Region: base=West]		
Northeast	-.111* (3.18)	-.135* (3.04)
North-Central	-.059 (1.83)	-.100* (2.51)
South	-.073 (1.70)	-.048 (0.97)
[Occupation: base=service workers]		
craftsmen	.215* (3.27)	.350* (5.04)
operatives	.091 (1.42)	.152* (2.24)
transport operatives	.091 (1.24)	.351* (3.37)
laborers	.065 (0.91)	.025 (0.28)

Table A.4 (Continued)

Full Regression Results of the Wage Equations:
May 1979 CPS Data

variable	dependent variable: ln (hourly earnings)	
	union sample	nonunion sample
[Industry Dummies]		
lumber	.041 (0.28)	-.053 (0.34)
furniture	-.138 (1.35)	-.236* (2.10)
stone,clay,glass	-.062 (0.65)	-.100 (0.71)
primary metals	-.031 (0.32)	.109 (0.85)
fabricated metals	-.057 (0.70)	-.065 (0.71)
machinery	.031 (0.40)	.072 (0.84)
electrical machine	-.060 (0.74)	.076 (0.82)
transport equipment	.045 (0.49)	.129 (1.12)
instruments	-.142 (0.94)	.129 (1.33)
food	.020 (0.23)	.036 (0.32)
tobacco	.326 (1.14)	- (-)[b]
textiles	-.072 (0.65)	-.140 (1.19)
apparel	-.151 (1.87)	-.049 (0.52)
paper	-.060 (0.66)	-.146 (1.09)
printing	.124 (1.52)	-.003 (0.03)
chemicals	.097 (1.02)	-.023 (0.19)
petroleum	.373* (2.32)	.120 (0.53)
rubber	-.069 (0.76)	-.077 (0.71)
leather	-.222 (1.45)	-.180 (1.34)
constant	.797* (3.19)	.907* (2.69)
R^2	.482	.549
SER	.263	.311
sample size	827	676

Note: [a] The base of marital status dummies is "Never married".
[b] There is no nonunion workers in the tobacco industry on the May 1979 CPS file.
The estimates shown in this table are full regression results of the wage equations estimated in Chapter 3 (see Table 3.2).
SER denotes the standard error of the regression.
Numbers in parentheses are absolute values of t-statistics.
* significant at the .05 level.

Table A.5

Comparisons of Estimates of the Union Status Equation: Probit Sample Selectivity vs. OLS Models

variable	dependent variable: U_i			
	sample selectivity		OLS[a]	
education	.015	(0.93)	.016	(1.03)
EXP	.036*	(3.27)	.054*	(4.71)
$EXP^2/100$	-.061*	(2.88)	-.085*	(3.84)
SEX (female=1)	-.162	(1.59)	-.249*	(2.47)
RACE (white=1)	-.187	(1.64)	-.402*	(3.86)
part-time	-.551*	(3.51)	-.735*	(4.62)
industry union density	.005	(1.25)	.015*	(3.97)
SMSA union density	.013*	(2.56)	.020*	(3.86)
establishment size/100	.075*	(2.23)	.084*	(2.24)
wages/value added	-.002	(0.33)	-.003	(0.50)
concentration ratio	-.003	(0.92)	-.004	(1.25)
injury & illness rate	.001	(0.91)	.001	(0.72)
SMSA population(log)	-.113	(1.30)	-.159	(1.80)
SMSA unemployment	-.000	(0.11)	.000	(0.13)
SMSA % of blacks	.003	(0.35)	.004	(0.50)
[Region: base=West]				
Northeast	.196	(1.56)	.334*	(2.71)
North-Central	.057	(0.49)	.159	(1.41)
South	-.010	(0.07)	-.032	(0.22)
[Occupation: base=service]				
craftsmen	.448*	(2.08)	.580*	(2.61)
operatives	.353	(1.68)	.515*	(2.44)
transport operatives	1.009*	(3.66)	1.408*	(4.59)
laborers	.373	(1.38)	.767*	(2.96)
$(\ln \hat{W}U_i - \ln \hat{W}N_i)$	1.906*	(5.70)	1.685*	(3.55)
sample size	1.503		1,503	

Note: [a] The predicted wage differential is calculated using OLS estimates of the union and nonunion wage equations.
Other variables included are the same as those in Table 4.1.
The numbers in parentheses are absolute values of asymptotic t-statistics.
* significant at the .05 level.

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