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TRADE UNIONISM, IMPLICIT CONTRACTING, AND THE RESPONSE TO DEMAND VARIATION IN U.S. MANUFACTURING*

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In the middle 1970's, economists studying the cyclical behavior of labor markets began to focus more attention on temporary layoffs. One factor stimulating interest in this topic was the recognition of the contribution of variation in layoff unemployment to variation in the aggregate unemployment rate. Between the second quarter of 1973 and the second quarter of 1975, the fraction of the unemployed on layoff jumped from 10 percent to 20 percent while the unemployment rate was rising from 4.9 to 8.7 percent. The apparent rigidity in wages and hours in the presence of large variations in layoff unemployment also contributed to interest in a theoretical explanation for layoffs. The newer theories address the questions raised by the prevalence of stable wage, variable employment practices by distinguishing contract markets from spot, or auction, markets and exploring the consequences of the durability of the typical employment relationship. The most frequently cited studies have investigated these issues from the perspective of risk shifting. The major alternative approach looks to transactions costs.[1]

The emphasis on understanding the process that generates layoffs has been well-placed. The contracting theories have provided some useful perspectives from which to view the operation of a labor market subject to frequent demand variation. There has been little new empirical evidence, however, and further progress will require more detailed knowledge about the manner in which these markets function. This paper presents an attempt to expand this knowledge. It reports an analysis of the effects of the decline in the demand for labor over the 1973-1975 recession on the wages, hours, and layoff unemployment of production workers in manufacturing.

A study of these manufacturing workers is relevant to the contract theories because they account for much of the average level of and cyclical variation in layoff unemployment. In May of 1973, 40 percent of those on layoff were manufacturing workers; two years later this figure had risen to more than 60 percent. In addition, this market exhibits the key attributes found throughout the contracting literature--low permanent separation rates, stable wages and hours, and frequent large shifts in demand. Some of these qualities can be found in the CPS. While layoff unemployment among hourly production workers was rising from less than 1 percent to more than 5.5 percent, their average workweek fell from 43.2 hours to 41.2, and real wages declined 3.5 percent.[2] Much of the fall in wages was a consequence of a 4 percentage point rise in the rate of inflation in the twelve months following May of 1973.

Concluding that the recession had no effect on nominal wage growth would be premature, however, for the results reported below reveal a surprisingly significant effect of industry demand shifts on hourly wages. The analysis combines data from the May Current Population Surveys (CPS) and the Bureau of Labor Statistics time series of employment by industry. The CPS provides observations on the wages, hours, and employment status of individuals in each year, and the latter source is used to construct measures of the shifts in each industry.

The CPS reveals that the experience of an individual worker depended heavily on whether he was a member of a labor union. The 1975 layoff unemployment rate for union members was 6.5 percent, over 2 percentage points higher than the rate for nonunion workers. The decline in union real wages of 2.5 percent was only half as large as the decline in nonunion wages. The paper investigates the possibility that this occurred because union workers are concentrated in particular types of industries and finds these differences too large to be accounted for by composition alone.[3] Other conclusions reached include:

(1) Variation in the fraction of employees working overtime accounted for the bulk of the variation in hours worked;

(2) Variation in this fraction was larger among union members than among nonunion workers; and

(3) A disproportionate number of the workers found on layoff had been employed in large establishments, while the employees who experienced adjustments in their hours worked per week were more likely to be working in small plants.

On the other hand, the analysis fails to uncover some relationships that one might have expected to find--neither wages nor industry unionization percentages display any association with the stochastic variability of industry employment.

Comparing the evidence with the details of competing theories of labor market contracting suggests that risk shifting plays a minor role in the response to cyclical demand shifts. The failure to establish a relationship between a stochastic variability in industry employment and any measures of contracting influence casts serious doubt on the relevance of risk shifting models. On the other hand, plant size does exhibit a significant relationship with variables that can be construed as measures of contracting strength, and this is interpreted as an indication that transactions costs are a more important determinant of the behavior of this labor market.

I. The Methods and Data

The introduction noted some differences in the experiences of union and nonunion workers. A primary objective here is determining the extent to which union-nonunion differences in the manifestations of cyclical activity are attributable to factors other than composition. Do the differences in the stability of wages and employment arise primarily from the conditions and characteristics of the industries in which union members are concentrated, or does the response to a given demand shock depend more on whether the establishment is unionized? There are environmental factors that predispose an establishment's employees to organize, and there is variation in the size and timing of cyclical demand shocks across firms, so answering this question requires one to have the ability to control for differences in these factors. Although, for this purpose, it would be desirable to use variables describing the environmental and demand conditions in each individual's place of work, the CPS does not identify employers so precisely. Therefore, each employer is assumed to be adequately described by average measures for the entire industry of which it is a part.

The primary analytical tool used in this study is multiple regression, in which the unit of observation is the individual worker. The dependent variables reflect the individual's wage, hours worked, and employment status in the reference week for the May CPS. The right side of the equation contains individual characteristics from the CPS record and industry variables obtained from other sources and attached to the CPS record. The measure of demand conditions at the time the individual was observed, one of the industry variables, is interacted with the union dummy to allow estimation of separate effects of demand shifts on union and non-union workers.

There are several possible directions in which one could proceed when devising a measure of cyclical demand shifts, and no one is clearly superior to the others. The measure used here is based on industry employment.[4] This choice is essentially dictated by the properties of the micro-data, which cover a short time interval and a broad cross section. The industry variable in the CPS is sufficiently detailed to identify 73 separate industries within manufacturing that can be matched with an industry (or group of industries) for which an employment series can be obtained from the BLS establishment data. From these series a measure of the state of labor demand in May of each of the three years was constructed for each of the industries.

A straightforward procedure was used to construct these measures. The log of employment was regressed on a set of seasonal dummy variables, time, and time squared using monthly observations over a seven and one-half year interval. The coefficients from that regression were then used to predict the log of employment for the twelfth month following the last observation in the interval. The difference between the actual and predicted values was used as the measure of the state of demand in that industry at that time. Because the observations on the individual workers were obtained in May, the regressions run to predict employment in May of year t used a sample period that ran from December of year t-9 to May of year t-1.

The principal variable used to capture differences in environments is an estimate of the number of production workers per establishment in each industry. The estimate was computed from data in the 1972 Census of Manufacturers. Industries with larger establishments are more heavily unionized, and the bureaucracy characteristic of larger plants should have some effect on the response to demand variation. Also, workers in large plants may be more interdependent and may have skills that are more firm specific.

The establishment size measure is the only environmental control used in the analysis of adjustments in hours and employment. The wage analysis contains two others. A measure of stochastic employment variability is included to see if wages behave differently in industries with historically unstable employment. Its value is the sum of squared residuals from a quadratic trend regression (that included seasonal dummies) of the log of industry employment over the period January 1958 to June 1975. The other variable in the wage analysis, estimated from the May CPS, is the percentage of male, full-time production workers in each industry who are union members. It captures the effect of other factors correlated with unionism that might affect the level or flexibility of wages.

Table 1 contains some statistics indicating how the industry measures vary across CPS industry groups. The demand shift variables, labeled prediction errors, show a very large decline in the demand for labor over the period. Employment was above the level predicted from recent trends by an average of 13.5 percent in May of 1973, but by 1975 it had fallen to 11 percent below trend. A comparison of the means of these shift and volatility measures in the durable and nondurable sectors indicates that employment has been more stable in nondurables by a factor of about three. Establishments in nondurables are also slightly smaller and less heavily unionized.

Table 2 shows the distributions of the industry variables after their attachment to the records of the CPS. The unit of observation here is the individual worker. The means of the shift measures for union and nonunion workers are interesting, for they reveal that the concentrations of workers in industries that experienced large reductions in employment were about the same for both union members and nonunion employees. Thus, the 1975 union-nonunion difference in layoff unemployment does not appear to be a consequence of greater union penetration in the more volatile durable goods sector. The table does indicate a marked difference in the concentrations of the two types of workers in industries characterized by large plants, however.

Section II discusses regression analysis of the behavior of real wages. The May CPS reports the straight time hourly wage for workers who were paid by the hour. The log of this value deflated by the consumer price index is the dependent variable in the regression equations. Estimates are reported based on a sample that pooled the data from all three years and also from regressions run on data from each year separately.

Section III discusses two regressions, one for employment status and the other for hours worked. Both use a qualitative dependent variable with values ranging from 1 to 3. The estimates reported are partial derivatives obtained from maximum likelihood estimation of conditional logit models.

The employment status equation analyzes the determinants of the probability that an individual will be disemployed. The dependent variable distinguishes people on temporary layoff from those who have been discharged permanently.[5] Those unemployed or not working for other reasons, such as illness or strikes, are not included in the sample. Table 3 shows the percentages of workers in this sample unemployed due to job loss. As one might expect, union members are less likely to be discharged, and this raises the possibility that the larger rise in union layoff unemployment over this period might have been caused by a larger proportion of disemployed nonunion workers being discharged rather than laid off. The logit model allows the investigation of this possiblity.

The dependent variable used in the analysis of variation in the length of individuals' workweeks is somewhat unconventional. The May CPS reports the number of hours each respondent works in a normal week as well as the number of hours he actually worked in the survey reference week. Table 4 contains statistics on a variable designated "excess hours," which is the difference between actual and normal weekly hours expressed as a percentage of normal hours. The table reveals that the small and stable mean for this variable conceals considerable hours variation for individual workers. Over 30 percent of manufacturing employees did not work their usual number of hours during the reference weeks of these surveys, and the average deviation from the normal level was not trivial.

The stability across years of the means of the positive and negative values of this variable is quite remarkable. Presumably, it reflects the presence of upper and lower bounds on hours worked.[6] When demand falls temporarily, employers lay off workers and do not reduce hours further once they have shortened the workweek to a certain point. When demand rises, firms prefer to add to the number of workers rather than extend the workweek beyond some limit. This property of the variable complicates the analysis; a regression directly employing excess hours as the dependent variable yields estimates that are difficult to interpret, because they are subject to truncation and selectivity biases. The variable does exhibit some year-to-year variation, however, and the table indicates it is primarily due to variation in the proportions of employees working more or fewer hours than normal. Therefore, the hours equation reported in Section III uses a qualitative dependent variable whose value is determined by the sign of excess hours.

II. Analysis of Real Wage Rates

The wage equations discussed in this section contain a mixture of standard and uncommon variables. In addition to the industry measures described in Section I, the equations include age and its square, years of schooling, and sets of dummies capturing variance attributable to differences in marital and union status, area population density, region, occupation, and state laws regarding the establishment of the union shop.[7] The industry variables and the right-to-work dummy are interacted with the union membership dummy. Table A-1, appended, contains the coefficients of the variables that are not of primary concern in this paper. Space does not permit a thorough discussion of these estimates, but a few are sufficiently important to note briefly.

Some of the coefficients suggest the existence of union threat effects of the type sought by Rosen (1969). The nonunion establishments most susceptible to union organization are large, in heavily unionized industries, and in states that allow the union shop. To make the overtures of union organizers less appealing, nonunion plants in these environments appear to have an incentive to pay above-market wages. The estimates in

Table A-1 are consistent with this scenario. Nonunion wages are lower in right-to-work states and positively related to establishment size and the industry unionization rate, while the two former variables have no effect on union wages. The coefficient of the latter interacted with the union dummy indicates the union wage premium is a positive function of the extent of union organization in the industry.

Table 5 displays the coefficients of primary interest estimated using a sample that pooled the observations from all years. The figures in the first column refer to an equation containing only a set of year dummies to capture between-year wage variation. Their coefficients show an approximately 5-percent decline in real nonunion wages. The interaction of the year dummies with the union dummy shows an upward trend in the union wage premium and, therefore, a smaller decline in real wages for union members.

The second column reports an equation containing both the set of year dummies and the prediction error. The positive coefficient for the prediction error indicates procyclic variation in nonunion wages, and the negative coefficient of the interaction between the prediction error and the union dummy indicates the response of union wages to be significantly smaller than the response of nonunion wages. The sum of the two coefficients, labeled a_1+a_2 in the table, reveals procyclic variation in real union wages. The year dummies in this specification show that nonunion wages, after adjustment for year-to-year shifts in demand, did not decline at all during 1973-74 and rose about 3 percent during 1974-75. The adjusted union wage premium declined over the period, as adjusted union wages fell in the first year and remained constant the second.

The figures in the third column refer to an equation containing the prediction error without the year dummies. The coefficients indicate significant procyclic wage variation for both groups. The estimates do not, however, indicate that wage sensitivity to cyclical shifts was high. A shift in labor demand causing employment to drop from a level consistent with a prediction error of 15 percent (the mean of the positive prediction errors in 1973) to -15 percent (the mean of the negatives in 1975) would cause nonunion real wages to fall 6 percent, or 0.15 1967 dollars. The effect on union wages would be about one-fourth as large.

If short-run demand shifts and long run productivity trends were the only influences on real wages, the dummy variable coefficients in column 2 would show steady year-to-year growth. They do not, and the history of the period contains at least part of the explanation. Consumer prices rose 6 percent in the year preceeding May 1973. In the subsequent twelve months the increase was 10 percent, and it fell slightly to 9 percent in the year prior to May 1975. The jump in the rate of growth of prices probably accounts for much of the absence of growth during the first year in adjusted nonunion wages and during both years for adjusted union wages.

Assume for the moment that the prediction errors fully capture the effect of short-run shifts and that productivity growth and the change in the path of prices were the only other factors influencing wages. The effect of inflation on real union wages implied by these estimates is then very large, and much higher and longer-lived than the effect on nonunion real wages. For example, if trend growth is assumed to be 3 percent, then the inflation change was responsible for 5.1 percentage points of the 5.4 percentage point reduction in real union wage growth during 1973-74 and 3.0 of the 3.7-percentage point drop the following year. Inflation's effect on nonunion wage growth was limited to 3 percentage points of the 5.5-percentage point reduction in 1973-74.

These are rough approximations, but the estimates of the effect of the price level change on union wage growth are a little large for comfort. If nominal union wages are less sensitive to external shocks than nonunion wages, then real union wages should show a relatively large response to inflation and a relatively small response to short-run demand shifts. Many union contracts contain clauses providing automatic cost-of-living adjustment, however, so one could expect to see a smaller impact far from unanticipated price level movements than is suggested here.

Table 6 displays the results from estimating the coefficients using data for each year separately, and it reveals one reason for the small union wage response to demand shifts estimated with the pooled sample. The coefficients in Table 5 reflect both within-year and between-year wage variation, and those in Table 6 indicate the prediction errors capture much more within-year union wage variance in 1974 and 1975 than in 1973. Thus, the estimates from the pooled sample may understate the actual union wage response, so the effect of the inflation increase implied by the dummy variable coefficients in column 2 may be too large.

The pattern of the response estimates found in Table 6 is interesting for its causes as well as its effects. It may be at least partly a consequence of the Nixon administration price control program. Comprehensive government intervention into the wage determination process began with the freeze on all wages and prices on August 15, 1971. In the following November the freeze was replaced with a system of supervised wage and price increases. During this "Phase II," which lasted through early 1973, the wage increases of approximately 70 percent of the nation's workers were subject to the approval of the Pay Board. Wage and price controls lasted through April of 1974, but after Phase II the proportion of workers covered declined steadily, and the supervision of wage setting grew less formal. Thus, most of the wages observed in 1973 and some of those observed in 1974 were probably influenced by the price control program.

Traditional analyses of the effect of controls have focused on the behavior of average wages. Most studies of the Nixon program using this approach have found little effect.[8] The dummy variable coefficients in Table 5 support this conclusion. Controls could alter wage distributions without affecting economy-wide averages, however, by dampening the response to short-run movements in demand. Workers in industries experiencing unusually high demand would be denied the larger raises market conditions would otherwise bring about, and employers in low-demand industries might feel that awarding an increase much below the maximum allowable would cause morale problems. Such considerations would account for the relatively small coefficients for the prediction errors computed using the 1973 and 1974 samples, when employment in most industries was above trend. Students of the subject contend that wage controls have a larger impact on the union sector, and that hypothesis is supported by the estimates in Table 6; the 1975 response is 2.6 times the 1973 response for union members and 1.4 times larger in the nonunion sector. But the period covered is short and highly volatile, so any conclusions must be regarded as extremely tentative.

In any case, the cyclic response of union wages is considerably smaller than the response of nonunion wages. These estimates do not identify the source of the response differential, however. Does the lower sensitivity of union wages arise from unionism directly, or is it a consequence of the environment in which union workers are most commonly found?

To answer this question, the prediction error variables were interacted with the unionization rate, employment variability, and establishment size variables in an equation similar to the one reported in column 3 of Table 5. The coefficients of interest are in Table 7. The estimates attribute little of the response differential to the environment. The estimated response of nonunion wages declines only 20 percent when the means of the industry variables for union members (from Table 2B) are substituted for the means for nonunion workers (from Table 2A). A similiar exercise reduces the union wage response by 10 percent. Establishment size is the most important variable.

A surprising result emerging from the wage analysis is the failure to find evidence that historic employment variability influences wage levels. Theory predicts risk averse workers will demand a premium for accepting jobs in firms with histories of cyclically sensitive employment. The coefficients of the measure of stochastic employment variability in Table 5 do not show this to be the case, however. The elasticities are small and positive for nonunion workers and small and negative for union workers. In only one case is an estimated elasticity larger than its standard error.

III. Analysis of Adjustment in Employment and Hours

The logit models reported in this section contain fewer explanatory variables than the wage equations. These regressions contain only age and its square, years of schooling, establishment size, the prediction errors, and dummies for union membership and occupation. A lagged prediction error measure is also included. This variable, which is the average of the prediction errors for December through May of the previous year, is included to investigate the possibility that an unusual level of employment in one year might affect an individual's probability of being unemployed or working more than usual when he is observed in the subsequent year. For example, one would expect to find a smaller fraction of employees in an industry working overtime if their employers did an exceptionally large amount of hiring in the previous year.

The tables in this section report derivatives of the probabilities with respect to the independent variables. The tables show separate estimates for union and nonunion workers. The calculation of these derivatives is explained in the Appendix. The procedure requires the assignment of values to each of the probabilities; the values used here were obtained by evaluating the logit functions when the prediction errors are 0 and the other independent variables are set equal to their means conditional on union membership, and are included at the bottom of the tables containing the derivatives. This provides estimates of the effect of the variables in a steady state situation, i.e., a world that has experienced no recent stochastic variation in employment. The logit coefficient estimates are in the Appendix.

The employment status model indicates that cyclical variation in union layoff unemployment exceeds variation in nonunion layoff unemployment and that much of the difference cannot be attributed to either environmental differences or the more frequent use of discharges by nonunion employers. Table 8 contains the derivatives computed from the logit coefficients in Table A-2. The contemporaneous prediction error derivatives imply that, if union and nonunion workers were alike and equally numerous, changes in the employment status of union members would account for about 60 percent of the year-to-year variation in total job loss unemployment. Variation in this unemployment among nonunion workers is almost equally divided between layoffs and discharges, but 70 percent of the variation in union job loss unemployment arises from layoffs. Variation in union layoff unemployment alone is almost as great as variation in unemployment from both sources combined for nonunion workers.

The other two industry variables also have a significant effect on the employment status probabilities. The lagged prediction error derivatives support the hypothesis that, within an industry, employment in one year and unemployment in the subsequent year are positively correlated. Firms that are well stocked with workers from the previous year's hiring will have a larger fraction of employees on layoff and will be less reluctant to discharge those who are redundant or poorly qualified. Establishment size has little effect on an industry's job loss unemployment rate, but it does influence the proportion of unemployed job losers who expect to be recalled. Layoff unemployment is an increasing function of plant size, but discharge unemployment declines as this variable rises. Establishment size and the union membership dummy are responsible for most of the union-nonunion differences in the steady-state probabilities shown at the bottom of Table 8.

The remaining variables in the model performed much as one would expect. Unemployment declines with age, and the magnitude of the effect of this variable is smaller for older workers. Craftsmen, the most highly skilled of the workers included in the analysis, experience slightly lower unemployment rates. The schooling derivatives are somewhat more interesting. They reveal that education has a stronger effect on discharge unemployment than on unemployment from layoffs. If investment in firm specific skills is an important determinant of an employer's choice to use layoffs rather than discharges when reducing his workforce, then this finding supports the proposition, advanced by Oi (1962), that investment in general skills is positively correlated with investment in specific skills.

In the hours regression, as in the employment status model, the industry variables provide the most interesting information. Table 9 reports the derivatives computed from the coefficients listed in Table A-3. The estimates reveal that cyclical variation in the fraction of employees working overtime is large relative to variation is in the fraction working short weeks. The lagged and contemporaneous prediction errors have a much greater effect on overtime than on short time for both union and nonunion workers. Another striking feature is the importance of the lagged prediction error. The derivatives indicate that the influence of the previous year's hiring on the probability of working the usual number of hours is about as great as that of current demand conditions. The sign of the relationship corresponds with what one would expect: high employment in one year reduces the following year's percentage of employees working overtime and raises the percentage that will be working less than usual.

The hours model also indicates that employees in industries characterized by large plants are less likely to experience hours variation in a steady state situation. Both derivatives with respect to establishment size in Table 9 are negative and statistically significant, so seasonal and other very short-lived shifts affect the hours of fewer workers in larger establishments. A comparison of Tables 8 and 9 suggests that this variable has more influence on the adjustment of hours than employment.

When stochastic shifts occur, union members are more likely to experience hours adjustment than nonunion workers. The contemporaneous prediction error derivatives are more than twice as large for union members, and the difference is particularly large for the probability of working overtime. The remaining derivatives in this model reveal nothing exceptional. Older workers, those with more schooling, and craftsmen are more likely to be observed working overtime and/or less likely to report a short workweek. Thus, these variables indicate a positive relationship between skill and excess hours.

IV. Interpreting the Results

Other studies of the effect of unionism on the response to demand variation have reported findings similar to those discussed above. Several studies, beginning with Lewis (1963), analyzed aggregate time series data and found that union wages respond less to cyclical shifts than nonunion wages. Medoff (1979) found a positive relationship between the unionization and layoff rates in state-industry cells. Raisian analyzed data from the Michigan Panel Study of Income Dynamics using a technique similar to the one here. The only major discrepancy between his results and those discussed above is his finding of countercyclical movement in union wages. Raisian's sample included workers from all industry groups, however.

The major advantage of this study is that the evidence presents a stronger argument for the case that unionism itself affects the firm's response to demand variation. The previous studies are unlikely to convince the skeptical that the effects attributed to unionism are not a consequence of a correlation between union organization and demand variability. Of course, the possibility exists that such a correlation produced the results obtained in this study also, but it is smaller.

If one accepts the conclusion that the union-nonunion differences are due to unionism rather than the environment in which union members are concentrated, then the explanation for this becomes an interesting question. Medoff calls attention to the political aspects of union decision making and to Freeman's (1978) application to the labor market of Hirshman's exit-voice approach. The key point in this view is that the preferences of senior workers have more weight, relative to those of junior workers, in the determination of conditions of employment under collective bargaining than when the workforce is not organized. Because senior workers are the last laid off and the first rehired, they will be affected less by demand variation if the firm relies heavily on employment adjustment when demand shifts. One of the consequences of unionization is, therefore, that the employer will perceive greater desire among workers for layoffs and rehires and less desire for flexible wages.

Raisian attempts to interpret his results within the context of the recent literature on implicit contracting, which examines the circumstances surrounding the durable associations between many employers and their employees. The labor market is dichotomized into auction and contracting segments, and the behavior of the two is compared. Contract market employers must be conscious of their reputations, for their treatment of current employees affects their ability to recruit workers in the future. This concern shapes the contract employer's response to demand variation. Raisian contends that his evidence suggests that unionized employers behave as if they were in the contracting segment, while nonunion employers are better characterized as auction market participants.

Raisian's interpretive efforts show promise, but they do not go far enough. There are two prominent approaches to the study of the contrasts between contractual labor markets and auction labor markets. The first, which will be referred to as the transactions cost model, is based on the idea that the gains from trade in some activities are larger when the transactors maintain a continuous long-term association with a restricted set of trading partners. This approach has evolved from the Becker (1962) and Oi (1962) discussions of firm specific human capital to the broader treatments of contracting by Wachter and Williamson (1978) and Klein, Crawford, and Alchian (1978). The second approach views long-term associations as the result of implicit arrangements in which the parties on one side of the transaction are induced to bear a disproportionate share of the burden imposed by instability in market supply and demand conditions. This risk shifting model captured considerable attention a a result of the nearly simultaneous but independent efforts of D. Gordon (1974), Baily (1974), and Azariadis (1975). Rather than simply associating unionism with contracting, the evidence in this paper can be more profitably used to distinguish which approach to contracting the data favor.

The cornerstone of the transactions cost approach is the observation that investment in physical assets or acquisition of specialized information or knowledge facilitates the exchange and production of many types of goods. Undertaking such an investment has two effects on exchange. The first, to which the literature on specific human capital is primarily addressed, is that the markets in which these goods are traded will be characterized by practices that tend to prolong the associations between trading partners. Mobility is costly in these markets, because the returns on the transaction specific assets are collected only as long as the association is maintained. The second consequence of transaction specific investments is the behavior arising from the presence of the quasi-rents that exist in activities where impediments to mobility are found. The mobility costs bestow a degree of short-run market power on at least one of the parties, and market participants can be expected to anticipate that their trading partners might attempt to exploit this power once a specialized investment has been made. The potential for such opportunistic behavior encourages the adoption of practices that limit the opportunities for exploitation and stimulates the development of institutions that regulate behavior.

Changes in supply and demand often generate confrontations concerning such exploitation, particularly when information on market conditions is not equally accessible to buyers and sellers. Those faced with higher costs of learning the true state of the market will tend to suspect that any effort to alter the terms of trade to their disadvantage is an attempt to seize a larger share of the specific quasi-rents. Transactions cost theorists perceive this atmosphere of distrust to be an important source of the price rigidity observed in contracting markets, and they assert it causes shifts in supply or demand to be met initially with adjustments in the quantity exchanged instead of changes in price. This priority is preferred because it appears to distribute more equitably the burden of the adjustment across both buyers and sellers rather than forcing one side to shoulder it alone. Changes in the terms of trade are made only at regularly specified intervals, and the establishment of a new price is in some cases even overseen by a third party, such as a government regulatory body or a union representative.

This general view of the effects of mobility costs can be applied to labor markets in a straightforward manner. The aspects of the model of immediate interest here are the consequences arising from the existence of appropriable quasi-rents. This segment of the theory provides a fresh perspective on the sluggish response of wages to changes in demand. Feldstein (1975) pointed out that the superior information on product demand possessed by managers of firms combined with workers' immobility may account for the use of adjustments in employment rather than adjustments in wages immediately after a change in market conditions. Employees resist wage reductions, even when they are justified by assertions that the value of labor services has fallen, for the workers' relatively incomplete information leads them to suspect that their employer is taking advantage of the short-run monopsony power arising from obstacles to their mobility. The employer can establish the credibility of his claim by first reducing output and employment, and thereby absorbing more of the effect of the demand shock himself. Reductions in wages would be permitted only after the employees had been convinced that they were not being exploited.

The risk shifting theory of contracting approaches the analysis of long-term relationships between transactors from a quite different angle. The fundamental assumption is that business firms are less averse to risk than the households that supply their labor services or purchase their products. The proponents of the approach defend this assumption by appealing to the ability of stockholders to diversify and to Knight's (1921) conjecture that entrepreneurs have a higher tolerance for risk than people who are employed by others. Auction markets in this environment would leave some profit opportunities unexploited, for the firm could obtain more favorable terms of trade by doing business on a regular basis with a small subset of potential transactors and shielding them from the risk to which the auction market would subject them. He can pay (charge) a lower (higher) average price by smoothing short-run price and quantity variation over some extended interval. Here, then, it is the transfer of risk that motivates the long-term relationships observed in contracting markets.

Risk shifting contracts will dampen the variation in labor earnings by reducing the flexibility of both wages and employment relative to what would occur in an auction market.[9] The most interesting product of research on this question is the implication that contracts will amost entirely eliminate variation in the wages of covered workers, but will not eliminate layoffs. This asymmetry arises from two assumptions: that contract labor is a wholly fixed factor, and that the employees value their time spent away from work. The first assumption means that the firm, after acquiring a stock of employees of the desired size, is free to vary the utilization of this stock but cannot change its size until the contracting period has ended. It implies that, within the contracting interval, the time paths of wage rates and employment will be independent. Wages will, therefore, be held fixed, because greater wage flexibility will increase income variability but will not reduce the incentive to adjust employment.[10] The second assumption guarantees this incentive will exist in firms that experience large fluctuations in demand, for they can increase profits by laying off workers when the value of the marginal worker's product falls below the value of his time in an alternative use.

The two approaches have in common a possible role for third parties--the provision of monitoring services. Officials of labor unions may be regarded as policing agents; the explicit contracts negotiated and administered by these specialists will be more costly to breach than implicit nonunion agreements. This presumes that the rules governing the employment relationship are similar in organized and unorganized establishments, but union workers enjoy greater protection under these rules because a third party ensures that they are followed.

Although unions serve the same basic function in the two approaches, they provide their services through different channels. In the risk shifting model, the contracting process may break down through persistent default by either side. The employer may attempt to take advantage of low spot market wages in periods of low demand by reducing the wages of his own workers, or the workers may desert contract market employers for auction market firms in periods when demand is high. Unions reduce the gains to default by both parties. Strikes raise the cost of failure to maintain historic rates of real wage growth, and greater reliance on seniority in the allocation of promotions, raises, and layoffs discourages job hopping.[11] Union services will, therefore, be more valuable in cyclically sensitive industries.

The transactions cost approach focuses on the complexities associated with production in large establishments, where the distance between those who direct the enterprise and those who execute the operations is bridged by a sometimes intricate bureauacracy. Unions can help shape this bureaucracy and oversee its operation. The performance of these functions reduces wage flexibility by reducing the frequency and magnitude of wage adjustments. The costs of strikes make firms seek long-term contracts, and the negotiating process may encourage particpants to resist giving up gains won in previous settlements. Transactions cost theorists perceive the introduction of formal grievance procedures to be the crucial feature of industrial relations contributed by unionism. These procedures are obviously of greater value in large, relatively impersonal plants.[12]

Determining which of the two approaches implies labor market behavior that is more consistent with what is actually observed in manufacturing requires identifying variables that indicate where the influences of contracting are strongest and then comparing the effects of increasing influence with the implications of the theories. Both claim contracting will reduce wage variation and increase the use of layoffs and rehires relative to discharges and new hires, and unionization can also be construed as an indicator of the strength of contracting influences. Thus, the extent to which contracting affects market behavior can be monitored through variables affecting the layoff-discharge ratio and the unionization rate as well as through unionism itself. The two approaches differ primarily in their treatment of employment variation. The transactions cost theorists view variation in wages and employment as substitutes--employment adjustments are larger and more frequent because they are less costly. Less wage flexibility will be accompanied by more employment variation. Risk shifting implies less variation in both employment and wages.

The evidence presented in the previous sections favors the transactions cost approach. Establishment size is a better indicator of the strength of contracting influence than cyclical employment variability. The contribution of layoffs to total job loss unemployment rises as establishment size increases, the correlation of the industry unionization rate with the average number of production workers in the industry is .30 (significant at the 5-percent level), and the flexibility of nonunion wages is negatively related to this variable. On the other hand, the measure of cyclical employment variability shows no significant relationship with any of these factors. The evidence also supports the transactions cost implication that contracting is associated with more employment variation, for union workers were observed to experience greater cyclical swings in job loss unemployment than nonunion workers. The response of wages to changes in the rate of inflation suggests that contracts stabilize nominal wages, also an implication of transactions costs, rather than stabilizing real wages as implied by risk shifting.

V. Conclusion

This study has had two purposes--to examine the response of the market for manufacturing production workers to between-year shifts in demand and to interpret that response in the context of the implicit contracting literature. The investigation of the response followed the path broken by Smith and Welch and Raisian who devised measures of industry specific demand movements constructed from aggregate data and employed these measures in an analysis of microdata. The evidence obtained was then used to examine which of the variants of labor market contracting yields more accurate implications about the sources of the now familiar manifestations of non-auction market behavior.

The empirical technique is not highly refined, and the reader has doubtlessly found points at which the analysis could be improved. There are, however, two solid conclusions that can be drawn:

(1) Wages and hours worked per week in manufacturing respond to the same cyclical forces that put pressure on employment, and these forces cause real wages and hours to adjust in the same direction as employment; and

(2) Implicit and explicit contracting affects the extent to which wages are adjusted relative to adjustments in hours and employment.

But the analysis covers a short span of time that includes only the contraction phase of one business cycle. It would be overly optimistic to presume that the simple specifications in Sections II and III would provide consistent explanations of movements in wages and labor utilization over a longer period, for the dynamics of a recovery may differ substantially from those of a recession.

The case on the source of contracting in the labor market is also far from closed. Many of the relationships cited in support of the transactions cost view may be attributed to other causes. For example, unionism may be more prevalent in large plants because they offer economies of scale in the acquisition of members. Thus, the most that can be claimed here is that the evidence cited is more consistent with the implications of transactions costs than risk shifting. There is much room for further research on this question, also.

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FOOTNOTES

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[1]The relevant topics in the contracting literature are discussed in Section IV.

[2]These figures and those that follow are for full time workers.

[3]Medoff (1979) and Raisian (1979) reached similar conclusions. Their work will be discussed in Section IV.

[4]Forerunners of this procedure are Smith and Welch (1978) and Raisian (1979). The former used deviations from trend in employment for state group by major industry cells. Raisian used major industry unemployment rates.

[5]The CPS identifies people who have been laid off and expect to be recalled within 30 days, people who have been laid off and expect to be recalled but do not know when (or expect more than 30 days to pass before they receive recall notice), and people who have been released by their employers and do not anticipate recall. The BLS refers to all

three as "job losers," and this convention is followed here. Workers in the two former classes will be referred to as "on layoff," and those in the latter category will be referred to as "discharged."

[6]For formal models of this phenomenon, see Baily (1977) or Pearce (1980).

[7]If the respondent did not reside in one of the most populous states, the geographic information in the CPS placed him in a state group rather than an individual state. If the group contained both right-to-work and other states, then the value assigned the right-to-work variable is the probability he lived in a right-to-work state, based on his two-digit industry.

[8]See Weber and Mitchell (1978, p. 314) for a list of studies and their conclusions.

[9]Actually, there is some disagreement on this issue. Baily (1974) asserts risk shifting will lead to what was formerly called "labor hoarding" and reduce employment variation. Grossman (1978) claims the reverse, but his concept of risk shifting appears to run more along the lines of reducing variation in workers' consumption streams than reducing risk.

[10]Baily (1977) stresses this point.

[11]Grossman attributes the popularity of seniority systems to the notion that senior workers have demonstrated that they will uphold a labor contract.

[12]See Feller (1973) and Vollmer (1960) for a fuller discussion. Note that this is the reverse of the exit-voice view of unions, which holds that union grievance procedures induce workers to change employers less freauently. The transactions cost model postulates the existence of mobility costs that deter frequent job changes, and workers facing these costs will have a strong incentive to seek implementation of procedures that will enable them to obtain better working conditions without changing jobs.

TABLE 1

STATISTICS ON INDUSTRY VARIABLES

A. All Manufacturing (73 industries)

	Mean	Standard Deviation	Minimum	Maximum	Number Positive	Mean of Positives	Mean of Negatives
Prediction Errors:							
1973	.135	.125	-.075	.612	68	.149	-.048
1974	.070	.117	-.449	.464	59	.100	-.056
1975	-.111	.111	-.345	.194	11	.069	-.143
Index of Stochastic Employment Variability	.660	.882	.012	5.30			
Production Workers per Establishment (1,000's)	.116	.169	.006	.965			
Fraction Union Members	.519	.162	.095	.859			

B. Major Subsectors

	Durables (38 Industries)		Nondurables (35 Industries)	
	Mean	Standard Deviation	Mean	Standard Deviation
Prediction Errors:				
1973	.194	.135	.072	.074
1974	.092	.148	.047	.064
1975	-.153	.111	-.065	.092
Index of Stochastic Employment Variability	.981	1.07	.312	.407
Production Workers per Establishment (1,000's)	.128	.182	.102	.156
Fraction Union Members	.539	.180	.498	.139

SOURCES: Employment and Earnings
Census of Manufacturers
May Current Population Surveys

TABLE 2

STATISTICS ON INDUSTRY VARIABLES AFTER ATTACHMENT TO CPS RECORDS

A. Nonunion Workers

	Mean	Standard Deviation	Percent Positive	Mean of Positives	Mean of Negatives
Prediction Errors:					
1973	.165	.129	95	.176	-.052
1974	.071	.114	78	.106	-.056
1975	-.137	.108	11	.063	-.163
Index of Stochastic Employment Variability	.665	.856			
Production Workers per Establishment (1,000's)	.085	.122			
Fraction Union Members (for industry)	.484	.151			

B. Union Workers

	Mean	Standard Deviation	Percent Positive	Mean of Positives	Mean of Negatives
Prediction Errors:					
1973	.151	.122	95	.162	-.045
1974	.077	.105	78	.108	-.028
1975	-.125	.105	12	.060	-.152
Index of Stochastic Employment Variability	.728	.785			
Production Workers per Establishment (1,000's)	.168	.239			
Fraction Union Members (for industry)	.613	.155			

SOURCES: Employment and Earnings
Census of Manufacturers
May Current Population Surveys

TABLE 3

LAYOFF AND DISCHARGE UNEMPLOYMENT IN MANUFACTURING

	Percent Unemployed Due to:	
	Temporary Layoff	Permanent Discharge
Nonunion		
1973	0.7	1.7
1974	1.0	1.4
1975	4.6	6.2
Union		
1973	0.7	1.1
1974	1.6	0.9
1975	7.0	3.2

NOTES: Population restricted to hourly white male production workers who were either employed or job losers.

SOURCE: May Current Population Surveys.

TABLE 4

EXCESS HOURS IN MANUFACTURING

	Mean	Standard Deviation	Percent Positive	Mean of Positives	Percent Negative	Mean of Negatives
Nonunion						
1973	1.51	14.1	23.0	18.7	13.2	-21.1
1974	0.18	14.2	19.4	17.4	13.5	-23.7
1975	0.34	14.5	17.6	19.8	13.5	-23.3
Union						
1973	2.87	15.2	25.3	21.1	10.9	-22.7
1974	2.04	13.7	20.8	20.7	10.5	-21.6
1975	0.02	13.0	14.2	20.0	13.2	-21.6

NOTES: All figures are percentages. Population restricted to white male hourly production workers.

SOURCE: May Current Population Surveys.

TABLE 5

ESTIMATES FROM WAGE EQUATIONS:
POOLED SAMPLES FOR 1973-75

	1	2	3
Prediction Error (a_1)	--	.255 (7.4)	.198 (8.6)
(Prediction Error) x D(u) (a_2)	--	-.215 (4.8)	-.144 (4.8)
$a_1 + a_2$	--	.040 (1.4)	-.054 (2.8)
D(1974)	-.025 (2.7)	-.002 (0.2)	--
D(1975)	-.049 (5.1)	.028 (2.0)	--
D(1973) x D(U)	.045 (2.1)	.076 (3.3)	--
D(1974) x D(U)	.056 (2.6)	.067 (3.0)	--
D(1975) x D(U)	.073 (3.4)	.038 (1.7)	--
D(U) (Union Dummy)	--	--	.057 (2.8)
Ln(Demand Variability Index) (b_1)	.010 (2.4)	.005 (1.1)	.006 (1.5)
[Ln(Demand Volatility Index)] x D(U) (b_2)	-.013 (2.6)	-.008 (1.7)	-.010 (2.0)
$b_1 + b_2$	-.003 (0.9)	-.004 (1.2)	-.004 (1.4)
R^2	.3800	.3837	.3831
Standard Error of Estimate	.2320	.2313	.2314

Number of Observations: 9,417
Mean of Dependent Variable: 1.07

NOTES: Absolute t-ratios in parentheses. Remaining coefficients in Table A-1.

SOURCES: <u>Employment and Earnings</u>
<u>Census of Manufacturers</u>
May Current Population Surveys

TABLE 6

ESTIMATES OF COEFFICIENTS OF
PREDICTION ERROR IN WAGE EQUATIONS:
SEPARATE SAMPLES FROM 1973-75

	1973	1974	1975
Prediction Error (a_1)	.295 (4.2)	.272 (4.5)	.418 (5.6)
(Prediction Error) x D(U) (a_2)	-.250 (2.7)	-.174 (2.1)	-.301 (3.2)
$a_1 + a_2$	.045 (0.7)	.098 (1.8)	.117 (2.0)
R^2	.3679	.3940	.3953
Partial R^2 (both variables)	.0056	.0074	.0118
Standard Error of Estimate	.2331	.2264	.2344
Mean of Dependent Variable	1.10	1.07	1.06
Number of Observations	3,247	3,177	2,993

NOTES: Absolute t-ratios in parentheses.

SOURCES: Same as Table 5

TABLE 7

REGRESSION COEFFICIENTS FOR COMPARING WAGE RESPONSE IN DIFFERENT INDUSTRY ENVIRONMENTS[a]

Prediction Error	.221 (5.3)
(Prediction Error) x D(U)	-.195 (3.6)
(Prediction Error) x (% Union)	-.014 (0.1)
(Prediction Error) x (% Union) x D(U)	.074 (0.4)
(Prediction Error) x PWE[b]	-.567 (2.1)
(Prediction Error) x PWE x D(U)	.318 (1.0)
(Prediction Error) x Var[c]	.015 (0.8)
(Prediction Error) x Var x D(U)	.014 (0.6)

NOTES: [a]The equation also contained all variables reported in the equation reported in column 3 of Table 5 and Table A-1.

[b]PWE is production workers per establishment.

[c]Var is the measure of employment variability.

SOURCES: Same as Table 5

TABLE 8

DERIVATIVES WITH RESPECT TO VARIABLES
IN EMPLOYMENT STATUS MODEL

	Nonunion		Union	
	Layoff	Discharge	Layoff	Discharge
Age	-.043 (6.7)	-.077 (7.3)	-.059 (6.7)	-.046 (7.3)
Schooling	-.003 (0.1)	-.160 (3.3)	-.005 (0.1)	-.096 (3.3)
D(Craftsman)	-.460 (2.8)	-.573 (2.2)	-.622 (2.8)	-.343 (2.2)
Ln(PWE)	.262 (3.7)	-.233 (1.9)	.351 (3.7)	-.141 (2.0)
Lagged Prediction Error	1.59 (2.3)	2.29 (2.1)	2.15 (2.3)	1.37 (2.1)
Contemporaneous Prediction Error	-6.07 (7.8)	-6.63 (6.9)	-11.5 (14.2)	-4.05 (6.2)
Steady State Probability	.012	.018	.016	.011

NOTES: Derivatives are multiplied by 100.
PWE is production workers per establishment.
Absolute t-ratios in parentheses.

SOURCES: See Table A-2.

TABLE 9

DERIVATIVES WITH RESPECT TO VARIABLES
IN HOURS ADJUSTMENT MODEL

	Nonunion		Union	
	Overtime	Short time	Overtime	Short time
Age	-.004 (0.1)	-1.81 (5.9)	-.001 (0.3)	-.161 (5.9)
Schooling	.567 (2.9)	-.420 (2.8)	.540 (2.9)	-.369 (2.8)
D(Craftsman)	.320 (0.4)	-1.41 (1.9)	.263 (0.3)	1.25 (1.9)
Ln(PWE)	-1.28 (3.1)	-.849 (2.5)	-1.28 (3.2)	-.767 (2.6)
Lagged Prediction Error	-15.8 (4.1)	4.47 (1.4)	-15.4 (4.1)	3.85 (1.4)
Contemporaneous Prediction Error	11.5 (2.9)	-1.59 (0.5)	24.4 (7.5)	-5.24 (2.1)
Steady State Probability	.223	.129	.215	.113

NOTES: Derivatives are multiplied by 100.
PWE is production workers per establishment.
Absolute t-ratios in parentheses.

SOURCES: See Table A-3.

APPENDIX

The multinomial logit model, introduced by Theil (1969), is of the form

$$(1)\quad \ln\left(\frac{P_i}{P_k}\right) = B_iX + \varepsilon \qquad (i = 1, \ldots, k-1),$$

where the dependent variable assumes integer values from 1 through k. X is the vector of independent variables, P_i is the probability that the dependent variable is equal to i, and the B_i are vectors of coefficients with elements b_{ij}. The probabilities must sum to 1, so the individual P_i and the independent variables are related by

$$(2)\quad P_i = \begin{cases} 1/(1+\sum_1^{k-1} S_m) & \text{if } i = k \\ P_k S_i & \text{if } i < k, \end{cases}$$

where

$$S_i = \exp\,(B_iX) \qquad (i = 1, \ldots, k-1).$$

The derivative of P_i with respect to the j^{th} independent variable (denoted P_{ij}) is then

$$(3)\quad P_{ij} = P_i(b_{ij} - \sum_1^{k-1} P_i b_{ij}).$$

Sample frequencies are commonly used for the probabilities P_i. A set of steady-state probabilities is desired here, however, and these can be obtained from equations (2) by assigning suitable values to the independent varaibles and using the logit coefficients reported in Tables A-2 and A-3. To approximate the steady-state, the prediction error variables are set to 0 and the remaining variables are set to their means conditional on union membership. This allows computation of a steady state value for each group. The establishment size means are in Table 2, the nonunion and union mean ages are 35 and 39, and the mean years of schooling for both groups is 12. (The craftsman dummy was arbitrarily set to 0.)

The age derivative for both groups and the contemporaneous prediction error derivative for union workers required computation of linear combinations. The former depends on the value of age, and the derivatives reported in Section III are for a worker 38 years old. The contemporaneous prediction error derivatives for union workers are computed using the sums of the logit coefficients of this variable and the logit coefficients of the interaction with the union dummy.

TABLE A-1

COEFFICIENTS OF REMAINING VARIABLES FROM WAGE REGRESSIONS[a]

Schooling	.024 (21.8)
Age	.030 (18.2)
Age Squared	3.31×10^{-4} (16.2)
D(Never Married)	-.044 (3.5)
D(Married, Spouse Present)	.042 (4.2)
D(SMSA > 1 million)	.031 (4.6)
D(SMSA < .25 million)	-.065 (10.5)
D(North East)	-.095 (11.7)
D(North Central)	-.013 (1.7)
D(South)	-.083 (8.5)
D(Laborer)	-.151 (15.3)
D(Operative)	-.118 (22.8)
D(RTW)[b]	-.066 (6.0)
D(RTW)xD(U)	.058 (4.9)
% Union	.151 (5.5)
(% Union) x D(U)	.160 (4.4)

TABLE A-1--continued

Production Workers per Establishment	.128 (3.6)
(Production Workers per establishment) x D(U)	-1.58 (4.0)

NOTES: [a]These are the coefficients of the remaining variables from the regression reported in column 3 of Table 5. Absolute t-ratios in parentheses.

[b]D(RTW) is a dummy whose value is 1 if the individual resides in a state that outlaws the union shop.

TABLE A-2

LOGIT COEFFICIENTS FOR EMPLOYMENT STATUS MODEL

	Layoff	Discharge	Chi Square
Constant	.184 (2.3)	1.20 (1.4)	--
Age	-.171 (4.5)	-.198 (4.9)	42.6
Age Squared	1.76×10^{-3} (3.6)	2.03×10^{-3} (3.8)	26.6
Schooling	-4.26×20^{-3} (0.1)	-.088 (3.3)	11.2
D(Craftsman)	-.385 (2.8)	-.321 (2.2)	12.4
D(U)	.294 (1.9)	-.315 (2.3)	9.3
Ln(PWE)	.214 (3.7)	-.126 (1.9)	17.8
Lagged Prediction Error	1.34 (2.3)	1.28 (2.1)	9.3
Contemporaneous Prediction Error	-5.08 (8.0)	-3.72 (7.0)	106.8
(Contemporaneous Prediction Error) x D(U)	-2.03 (2.5)	-.110 (0.1)	6.4
Sample Proportions	.026	.022	--

Number of Observations: 12,044
Overall Chi Square: 598.5

NOTES: PWE is Production Workers per Establishment. Figures in parentheses are absolute t-ratios.

TABLE A-3

LOGIT COEFFICIENTS FOR HOURS ADJUSTMENT MODEL

	Overtime	Short Time	Chi Square
Constant	-2.31 (6.6)	-.701 (1.7)	--
Age	.037 (2.3)	-.029 (1.5)	8.9
Age Squared	-5.26×10^{-4} (2.6)	1.60×10^{-4} (0.6)	7.8
Schooling	.028 (2.4)	-.030 (2.2)	13.7
D(Craftsman)	-.002 (0.0)	-.125 (1.9)	3.8
D(U)	-.016 (0.3)	-.068 (1.0)	1.1
Ln(PWE)	-.090 (3.8)	-.098 (3.3)	21.2
Lagged Prediction Error	-.887 (3.9)	.169 (0.6)	17.4
Contemporaneous Prediction Error	.668 (2.8)	.030 (0.1)	8.3
(Contemporaneous Prediction Error) x D(U)	.749 (2.4)	-.209 (0.5)	7.2
Sample Proportions	.204	.122	--

Number of Observations: 10,700
Overall Chi Square: 170.2

NOTES: PWE is Production Workers per Establishment. Figures in parentheses are absolute t-ratios.