WAGES AND THE RISK OF PLANT CLOSINGS*

by

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Abstract

This paper examines the empirical relationship between the probablity a plant closes and the compensation paid to employees in the plant. The paper uses data on over 6500 manufacturing plants from the LRD to estimate the market hedonic wage locus and the probability of plant failure. The empirical results reported in this paper indicate that the probability of plant failure is systematically related to the plant's market share, age, recent growth, and variable cost to revenue ratio. The market hedonic wage regression indicates that workers employed by multiplant firms earn a positive compensating wage differential for the risk of plant closing but workers employed in single-plant firms do not. Additionally, the paper provides evidence on the general pattern of wage variation across heterogeneous employers. Establishment wage rates are significantly affected by plant size, age, geographic location, industry, capital intensity, and value added per worker.

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I. Introduction

The extent of plant closing and job loss resulting from the sectoral reallocation of U.S. production has been of great concern in recent years. While sectoral demand shifts certainly affect the rate of plant closing other forces are also at work. Evans (1987) and Dunne, Roberts, and Samuelson (1988, 1989b) find that patterns of plant and firm failure are consistent with a process of within-industry selection in which relatively high-cost producers fail. These selection forces lead to substantial variation in the probability of failure across producers within an industry. These findings suggest that the layoff risk faced by workers as a result of plant closing varies not only with the demand for their industry's output but also with the efficiency of their employer relative to competing producers in the same industry.

This paper examines the empirical relationship between the probability a plant closes and the compensation paid to employees in the plant. Two linkages are possible. First, workers may demand higher wages from failure-prone employers in order to compensate them for the higher risk of layoff they face. This force would produce a positive correlation between wages and the probability of plant closing that is present both across and within industries. Second, wage increases by a single plant affect the plant's profitability and competitive position within its output market. If an increase in wages simply raises variable costs without increasing plant revenue then a positive within-industry correlation between wages and failure probability should result. However, if the increase in wages is compensated by increased worker quality or raises worker productivity, as suggested in the efficiency wage literature (Katz, 1986), there may no correlation. This paper develops a two-equation empirical model of plant failure and wage determination and estimates both the compensating differential employees require to work in a failure-prone plant and the effect of wage changes on the probability of plant survival.

Previous studies of the relationship between unemployment risk and wages have utilized survey data on workers. Empirical estimates of the compensating differential for unemployment risk are generally positive but the wage premium is often small and statistically weak.[1] Hamermesh (1988, 1989) examines the effect of wage cuts on the probability of plant survival and finds that large wage concessions are needed to effect small changes in the probability of survival. One limitation of these studies is that, because they utilize worker data, the measures of unemployment risk they develop depend only on worker and industry characteristics. They are unable to capture the substantial variation in unemployment risk which workers in the same industry face as a result of the efficiency of their employer.

In this paper we utilize a micro data set on over 6500 U.S. manufacturing plants for the period 1974-1978. The longitudinal elements in the data allow us to estimate a model of plant failure, derive estimates of the probability of plant closing that vary with plant and industry characteristics, and quantify the effect of wage increases on failure probabilities. The market hedonic wage locus is estimated using the cross-sectional variation in plant characteristics, including the probability of failure, and the average hourly wage paid to production workers in the plant. From this the compensating wage differential is estimated.

The empirical model used in this paper also provides insights into the general pattern of wage variation across heterogeneous employers. This complements recent research by Brown and Medoff (1989) and Groshen (1988), that identifies an important role for employer size, and by Krueger and Summers (1988), that emphasizes the importance of across-industry wage differences. The data used in this paper allow us to control for a more detailed set of producer characteristics, including age and capital intensity, than has been previously possible. This research also complements recent work by Leonard (1987), Dunne, Roberts, and Samuelson (1989a), and Davis and Haltiwanger (1989) on the importance of the intrasectoral and intraregional turnover of employers in generating gross employment flows. All of these empirical studies indicate a significant, but still largely unexplored, role for employer heterogeneity in the process of wage and employment determination.

The empirical results reported in this paper indicate that the probability of plant failure is systematically related to observable plant characteristics including the plant's market share, age, and recent growth experience. A ten percent increase in production worker wages is estimated to cause a .15 percentage point increase in the probability of plant failure. This is similar to Hamermesh's (1988) finding that large wage cuts are necessary to offset the effects of adverse output market shocks on plant closing. This small effect could arise either because plant-level variation in efficiency is large relative to expenditure on production workers or because higher wages are, at least partially, offset by higher worker quality or productivity. Finally, the growth rate of industry output is not related to the probability of plant closing but higher industry-region unionization rates are positively related to plant closing.

The hedonic wage regression indicates that workers employed by multiplant firms earn a positive compensating differential for the risk of plant closing. For plants owned by single-plant firms no significant compensating differential is found. Other plant characteristics which are found to be significantly related to production worker wages are size, age, geographic location, industry, capital intensity and value added per worker. Finally, the presence of unemployment insurance reduces the compensating differential.

The next section of this paper develops a simultaneous equation model of plant failure and wages. The third section describes the plant-level data. The fourth section provides the empirical estimates and the fifth section examines the robustness of the findings to alternative specifications of the model, particularly the inclusion of unemployment insurance variables.

II. An Empirical Model of Plant Failure and Wages

The data used in this paper are drawn from the Annual Survey of Manufactures panel of plants for the years 1974-1978. The main goal of this empirical model is to examine the covariation between the average wage paid to production workers in a plant in 1974 and the probability the plant closes over the four-year period. The basic model consists of two equations. The first describes the probability that a plant which is in operation in 1974 fails over the 1975-78 period. The failure probability is modeled as a function of plant and output market characteristics. Of particular importance is a plant's ratio of variable cost to revenue because changes in wage rates, other factors held fixed, will alter this ratio. The second equation is the hedonic wage locus. The average wage paid in a plant is modeled as a function of plant and local labor market characteristics.

The model of failure for a single plant i is

$$Y_i^* = \theta' X_i + \alpha\, g(W_i) + u_i \tag{1}$$

where Y_i^* is a latent variable reflecting the future profitability of the plant, X_i is a k-element vector of explanatory variables, and θ is a k-element vector of parameters to be estimated. To control for the effect of wages on plant profitability we include $g(W_i)$ which is a variable that will depend on the wage rate paid by the plant. Finally, α is a parameter to be estimated and u_i is a normally distributed random variable with zero mean and constant variance. In practice Y_i^* is unobservable but instead we observe a dummy variable

$$\begin{aligned} Y_i &= 1 \text{ if } Y_i^* < 0 \\ &= 0 \text{ otherwise.} \end{aligned} \tag{2}$$

The dummy variable Y_i equals 1 for each plant that closes over the 1975-1978 time period and equals 0 for each plant that remains in operation for the entire period. The probability that plant i closes over the 1975-1978 period is defined as $P_i \equiv Pr(Y_i^* < 0)$.

The vector of explanatory variables X_i contains both plant and output market characteristics in order to control for the state of demand in the plant's output market and the relative efficiency of the plant within its output market. The plant characteristics included are: a dummy variable to distinguish whether the plant is owned by a single-plant or multiplant firm, a dummy variable to identify whether multiplant owners have other plants in the same four-digit industry, the plant's market share in its four-digit output market, a set of categorical variables to measure the plant's age, age-market share interaction terms, the share of energy in total material cost, and the past rate of output growth in the plant. The average wage is assumed to affect a plant's failure rate through its effect on the ratio of variable cost to plant revenue. The variable $g(W_i)$ is measured as the plant's total expenditure on labor and materials relative to its total revenue. Variables that are not plant specific are the annual average rate of real output growth for the plant's four-digit industry for the years 1974-1978, two-digit industry dummy variables, and the unionization rate for the two-digit industry and region in which the plant is located.[2] All variables, with the exception of the industry growth rate, are measured in 1974.

The justification for the variables is as follows: ownership type, market share, age, and age-market share interactions are all predicted to affect failure rates in the model of producer growth and failure developed by Jovanovic (1982).[3] The share of energy in material cost is included to control for possible adverse effects of the energy price increases in 1974. The past growth variable is intended to control for the recent history of the plant. If plants face adjustment costs when changing their scale of operation, the relatively efficient plants will tend to expand gradually over time while inefficient plants would tend to contract prior to failure. The industry growth rate controls for fluctuations in demand over the relevant time period. The industry-region unionization rate is included to measure a possible effect of union status on the decision to close a plant.[4] Finally, the variable cost-revenue ratio is included to control for cost heterogeneity across producers, including the effect of wage differences on failure probabilities. In one case, plants that have higher variable costs, holding revenue fixed, are less likely to cover their fixed costs in the long run and are thus more likely to shut down. Alternatively, if higher wages are correlated with higher worker-productivity there may be no effect on failure rates.

The second equation of the empirical model is the market hedonic wage equation. It relates the plant's average wage to a set of plant and local labor market characteristics including the plant's probability of failure. The basic estimating equation is

$$\ell n W_i = \beta' Z_i + \gamma f(P_i, S_i) + \varepsilon_i \qquad (3)$$

where $\ell n W_i$ is the natural logarithm of the average hourly wage paid to production workers in plant i in 1974. Z_i is a m-element vector of plant and labor market characteristics, and β is a m-element vector of parameters. The variables included in Z_i are: the dummy variables for ownership type described above, plant size measured as total employment, dummy variables for plant age, the log of the capital-labor ratio, log of value-added per worker, the industry-region unionization rate, the local unemployment rate, and sets of two-digit industry dummies and regional dummies. Finally, $f(P_i, S_i)$ is an n-element vector of explanatory variables that are all functions of the probability of plant failure P_i and a set of variables S_i that capture aspects of the local labor market. These include the local unemployment rate and aspects of the unemployment insurance program that may affect the compensating differential by altering the cost of a plant failure to workers.[5]

Virtually all of these variables have been found to be important in hedonic wage regressions although the reasons are not always clear. Brown and Medoff provide a detailed analysis of the employer size-wage relationship and speculate on the possible importance of plant age. The capital-labor ratio controls for differences in marginal products. Several papers including Krueger and Summers (1987), Dickens and Katz (1987), Katz and Summers (1989), and Blanchflower, Oswald and Garrett (forthcoming) have argued that the wage process may reflect bargaining over firm or industry rents. They have found measures of profit or value-added per worker to be important in explaining firm or industry wage variation. The local unemployment rate is included to reflect local labor market conditions. Unionization rates, industry, and region have been found to be important in most wage studies.

III. Data and Estimation

The data analyzed in this paper are drawn from information collected in the 1974 through 1978 Annual Survey of Manufactures. The ASM is a yearly survey of approximately 70,000 U.S. manufacturing plants that is taken in all noncensus years. The yearly survey covers approximately 20 percent of all manufacturing plants. The group of plants to be included in an ASM panel are chosen every five years and plants contained in an ASM panel are surveyed each year for five years if they remain in operation. The panel of plants used in this study was surveyed each year from 1974 through 1978.[6]

This paper examines a sample of plants from the 1974-78 ASM panel representing fifty four-digit SIC industries or ten two-digit SIC industries.[7] Three criteria were used in constructing the final sample. First, because it will be important to control for differences in local labor market conditions when estimating the compensating differentials the sample was limited to plants located in one of the 28 states for which unemployment rates were available for 1974.[8] These states account for over ninety percent of U.S. manufacturing employment. Second, because we want to control for the plant's recent history when estimating the model of plant failure we limit the sample to plants in operation in 1974 that were also in operation in 1972.[9] Third, all plants that were deleted from the ASM panel for administrative reasons unrelated to plant closing were removed.[10] The final dataset contains 6683 separate manufacturing plants.

An important measurement issue concerns the identification of plants which close over the 1975-1978 period. All plants that stop reporting data during one of the ASM survey years and then do not report data in all subsequent years and do not exist in the next Census of Manufactures (taken in 1982) are classified as closed.[11] Using this definition 904 plants that were in operation in 1974 closed between 1975 and 1978. This gives an overall sample closure rate of 13.53 percent.

Several other variables used in the analysis require further explanation. The basic wage rate used is the average hourly wage earned by production workers in a given year. It is measured as total expenditure on production worker wages divided by production worker hours.[12] Plant age is measured as categorical variables based on the year of entry of the plant.[13] The base category are the oldest plants in the sample; they began operation before 1964. Two additional dummy variables are defined as:

Age1 = 1 if the plant began operation in 1968 through 1972

= 0 otherwise

Age2 = 1 if the plant began operation in 1964 through 1967

= 0 otherwise.

Three categories of plant ownership are also distinguished with a set of dummy variables. The base category includes plants that are owned by single-plant firms. A dummy variable is used to identify plants owned by multiplant firms regardless of the industry in which the firm's other plants operate. A second dummy variable identifies plants owned by multiplant firms which operate other plants in the same four-digit industry.

The empirical model given by equations (1) - (3) treats both the probability of plant failure and the plant's wage rate as endogenous. Both equations are estimated using instrumental variables.[14] In the failure equation the endogeneity of the plant wage rate results in correlation between the error term and the variable cost-revenue ratio which is used as a regressor. The variable cost-revenue ratio is replaced with an instrumental variable constructed as the fitted value from a reduced form regression of the ratio on all exogenous variables in the model X_i, Z_i, and S_i. Equation (1) is then estimated with maximum likelihood.

Estimation of the wage equation (3) requires an instrumental variable for the probability of plant failure. This is constructed from a reduced form probit regression of Y_i on all exogenous variables in the model X_i, Z_i, and S_i. Denoting $\hat{Y}_i$ as the fitted value of Y_i from the regression, an instrumental variable for the probability of failure is constructed as $\hat{P}_i = \Phi(-\hat{Y}_i)$ where Φ is the cumulative density function for the standard normal random variable.

In this model the output market variables identify the hedonic wage locus and the local labor market variables identify the probit model of plant failure. More specifically, the plant's market share, age-market share interactions, energy share, past plant employment growth, and industry-region growth rates reflect the plant's relative efficiency and shifts in market demand for the plant's output and are used to identify the wage locus.[15] In the structural probit model of plant failure, total plant employment, the plant's capital-labor ratio, value-added per worker, the local unemployment rate, regional dummy variables, and the state-level unemployment insurance variables are omitted. These variables reflect differences in the marginal product of labor, potential rents, and local labor market conditions that could affect wages.

IV. Empirical Results

<u>Plant Closing</u>: The estimation results for the probit model of plant closing are reported in Table 1. They indicate that the plant's variable cost-revenue ratio, market share, age, age-market share interactions, past employment growth, and the industry-region unionization rate are significantly correlated

with failure rates. An increase in the ratio of variable cost to plant revenue acts to increase the probability of plant closing. The coefficient of .442 implies that a ten percent increase in the plant's expenditure on production worker wages, holding revenue fixed, would result in a .15 percentage point increase in the probability of plant failure over the four-year period.[16] This suggests that substantial wage changes have little effect on the probability of plant closing. There are at least two reasons this could occur. In the sample production worker wages account for 16 percent of a plant's variable cost on average. If plants differ substantially in efficiency, so that total costs vary substantially across plants even when output levels and factor prices are equal, then across-plant differences in production worker wages may simply be too small to overcome other sources of cost heterogeneity. Alternatively, if worker wage increases are fully or partially offset by productivity improvements, as suggested in the efficiency wage literature, then wage increases would be expected to have little or no effect on failure rates.

Larger plants have significantly lower rates of closing. The age dummy variables, age 1 and age 2, indicate successively older plants and the omitted category is the oldest group of plants. The coefficients are positive and decline with age which indicates that the failure rate is higher for younger plants. One of the age-size interactions is also statistically significant and the pattern indicates that, while larger plants within each age group have lower failure rates, this reduction in failure rates is most significant for the younger plants.[17]

The plant's past employment growth has a significant negative effect on the failure probability. This can reflect that the process of plant investment or disinvestment is a gradual one and that plant's tend to shrink prior to closing. The coefficient on the industry-region unionization rate indicates higher plant closing rates in areas and industries with higher unionization rates. The coefficient of .340 implies that a ten percent increase in the unionization rate increases the probability of plant closing by .255 percentage points. This implies an elasticity of failure with respect to the unionization rate of .210.[18]

The coefficients on the industry dummy variables indicate that, relative to the food processing industry (SIC 20), the primary metals (SIC 33) industry has a lower rate of plant closing and the instrument (SIC 38) industry has a higher rate. Finally, the plant's ownership dummies, energy share, and industry growth rate are not significantly related to plant closing.

Overall the results indicate a very significant role for plant characteristics in the process of plant closing. From the estimates in Table 1 we construct a predicted probability of failure for each plant in the sample. A variance decomposition of this data indicates that 92.6 percent of the total variation occurs across plants within the same two-digit industry while 7.4 percent occurs across industries. This suggests that the majority of the variation in the risk of job loss which workers face as a result of plant closing occurs across employers rather than industries. This is precisely the variation in risk that is not possible to quantify using micro data on workers. This finding also suggests that there may be substantial within-industry variation in the wage premiums which plants must pay to compensate workers for the risk of plant closing. We now turn to estimation of this wage premium.

Wage Rates: Coefficient estimates for the hedonic wage equation are reported in Table 2. In this simplest specification the compensating differential for plant failure is treated as a constant, rather than a function of the exogenous variables. Thus $f(P_i, S_i)$ in equation (3) is initially set equal to P_i.

The first column of Table 2 reports the results for the full sample of 6683 plants treating the probability of plant failure as endogenous. The probability of plant failure is positively and significantly correlated with the average production worker wage. Converting the coefficient of .519 to an elasticity results in a value of .070 evaluated at the mean failure probability in the sample. The empirical results indicate a positive, statistically significant compensating differential for workers employed in plants with a higher probability of failing. Relative to workers in a plant which has no probability of failing, those working in a plant with the mean probability of failure (.135 over four years) have a wage premium of 7.3 percent.[19]

The remaining coefficients in the first column indicate that plant characteristics, geographic location, and industry of operation contribute significantly to variation in wages across plants. We first discuss the role of plant characteristics. The coefficients on the two ownership dummy variables indicate that, relative to plant's owned by single-plant firms, plants operated by multiplant firms pay lower wages. Wages are 5.7 percent lower when the firm operates plants in other manufacturing industries and 3.7 percent lower when the firm operates other plants in the same industry. The coefficient on plant size, measured as the log of total employment, equals .059 which is very similar to the estimates from the wide range of alternative sources discussed by Brown and Medoff (1989). In addition to these variables, plant age is also significantly correlated with wages. Average production worker wages increase with plant age. Relative to the oldest plants in the sample (those that began operation prior to 1964), the youngest plants (began operation between 1968 and 1972) have average wages that are 6.9 percent lower. This difference occurs even after controlling for plant size and the probability of failure, both of which are correlated with age.[20] Two additional plant characteristics that are positively correlated with wages are the plant's capital-labor ratio and the level of value-added per employee. The former can reflect differences in the marginal product of labor across plants while the latter can reflect across-plant differences in output prices. Plants with higher output prices will have higher value-added per employee other factors held fixed. If this price variability results in plant-specific rents that are, at least partially, shared with workers then a positive correlation between wages and value-added per employee would result.

The final variables in column 1 reflect geographic and industry wage variation. Relative to the food industry (SIC 21) all industries except apparel (SIC 23) have significantly higher average wages. The highest average wages appear in primary metals (SIC 33), fabricated metals (SIC 34), and nonelectrical machinery (SIC 35). The regional differences indicate that, relative to the northeast, production worker wages are significantly lower in the South Atlantic and South Central U.S. Finally, wages are higher in industries and regions with high unionization rates and in markets with high local unemployment rates. The last result is surprising because high local unemployment rates should tend to reduce worker's reservation wages and result in lower average wages. Alternatively, if high local unemployment is correlated with layoffs of low-wage employees in the plant or if the unemployment rate controls for measures of unemployment risk not captured in the plant's probability of failure then the positive correlation would result.[21]

The second and third columns of Table 2 disaggregate the plants into two groups: plants owned by single-plant firms and those owned by multiplant firms. The results indicate a very different set of correlations for the two types of plants. For the single-unit plants only value-added per employee among all the plant characteristics is statistically significant. The probability of failure, plant size, age, and capital intensity are not statistically significant. In particular, plants with higher failure probabilities have lower wages although the coefficient is not significantly different from zero.

In contrast the plant characteristics, including the probability of failure, are important determinants of wages in multiunit plants. The signs and magnitudes are similar to those reported in the regressions for all plants. This is not surprising given the fact that multiunit plants account for 80.2 percent of the sample. The coefficient on the probability of plant failure increases slightly to .670. This indicates that a plant with the mean probability of failure pays wages that are 9.5 percent higher than a plant which has zero probability of failing. Overall the decomposition into single and multiunit plants indicates a very different role for both plant characteristics and the risk of plant closing between the two groups.

V. Alternative Empirical Specifications

This section discusses a number of alternative specifications for the model and checks for the sensitivity of the results to changes in specification. The first case we examine involves a change in the measurement of production worker wages. Hourly wages are now defined to include both direct payments as well as supplemental payments which include social security and fringe benefits. In the wage regression for multiunit plants, the coefficient on the probability of failure increases slightly to .688 (from .670) reported above. The other regression coefficients are unchanged. For the single-unit plants the coefficient on the failure probability remains negative and statistically insignificant. The point estimate is -.500 compared with the value of -.441 reported in Table 2. With the exception of the youngest age category, the other plant characteristics remain statistically insignificant. Overall, the patterns reported in Table 2 are not sensitive to this change in the definition of worker compensation.[22]

The second sensitivity check involves a change in the definition of a failing plant. Defining failure as those plants that report their closing to the Census bureau reduces the sample failure rate from 13.53 to 8.31 percent. We believe that this definition underestimates the amount of plant closing, particularly for small, young plants. The only results of the probit regression that are affected by this change in definition are that the age coefficients and unionization rate become statistically insignificant.

The wage regressions discussed until this point have included the probability of failure as a single regressor. Abowd and Ashenfelter (1981), Adams (1985), and Topel (1984, 1986) find that either the wage rate or the compensating differential for unemployment risk are sensitive to the presence of unemployment insurance programs. In this section we introduce variables that allow the compensating differential to vary across plants depending on differences in unemployment insurance programs and local unemployment rates.

The theoretical reasons for including these interaction effects are straightforward. The presence of generous unemployment insurance benefits reduces the cost to the worker of being laid off due to a plant closing. This in turn reduces the compensation needed to induce the worker to accept employment in a risky plant. In contrast, high local unemployment rates can indicate that loss of a job due to plant closing may result in a longer spell of unemployment than would be the case if unemployment rates were low. It will be necessary for failure-prone plants in these high unemployment areas to offer a larger wage premium to compensate for this increase in the length of the expected unemployment spell.

In order to empirically examine these effects we introduce two new variables to quantify aspects of the state unemployment insurance programs. In general terms these programs replace a certain percentage of a worker's annual income up to a maximum yearly payment. The percentage of income replaced and the maximum payment vary across states. For our purposes it is important to know both of these state-level UI variables as well as whether plant wages are high enough that workers are bound by the maximum yearly UI payment.[23] The latter is important because it affects the rate of insurance coverage in high wage plants, with higher wages leading to a lower rate of insurance protection. This will in turn affect the extent to which the market, through the compensating wage differential, must provide insurance.

In order to allow the compensating differential to vary with the UI system and local unemployment rates the market hedonic wage equation (3) is specified as:

$$\ell nW_i = \beta' Z_i + \gamma_1 P_i + \gamma_2 P_i \cdot U_i + \gamma_3 (P_i \cdot R_i \cdot D_{1i}) + \gamma_4 (P_i \cdot \ln M_i \cdot D_{2i}) + \varepsilon_i \qquad (4)$$

where U_i, R_i and M_i are, respectively, the local unemployment rate, the state UI replacement rate, and the stale maximum annual UI payment.[24] D_{1i} and D_{2i} are dummy variables that classify plants into the correct UI regime based on whether workers in a plant are constrained by the maximum allowable UI payment. When classifying plants into UI regine it is not appropriate to simply compare the observed plant-level wage with the state-specific wage that would subject workers to the maximum payment because random disturbances to plant i wages could then affect the classification of the plant making the dummy variables D_{1i} and D_{2i} endogenous. Instead, a reduced form wage equation is estimated by regressing plant wages on all exogenous variables in the model, Z_i, X_i and S_i, and basing the classification on the predicted wage $\hat{W}_i$ from the reduced form regression. Denoting W_i^M as the annual wage that would qualify workers in plant i for the maximum annual UI payment in their state, the dummy variables are defined as:

$$D_{1i} = 1 \text{ if } \hat{W}_i < W_i^M$$
$$= 0 \text{ otherwise}$$

$$D_{2i} = 1 - D_{1i}.$$

In the sample 51.6 percent of the plants have D_{1i} equal to 1 and the remaining 48.4 percent have D_{2i} equal to 1.

The results of instrumental variable estimation of equation (4) are reported in Table 3. When pooling both the single and multiunit plants the results indicate that the presence of unemployment insurance reduces the magnitude of the compensating differential required to work in a risky plant. Both of the interaction terms between the probability of failure and the UI variables are negative and statistically significant. The interaction term between the failure probability and the local unemployment rate is not statistically significant.

One way to evaluate these coefficients is to calculate the change in the compensating differential as the replacement rate or maximum annual payment changes. For plants whose workers would not qualify for the maximum annual payment (D_{1i} = 1) there is an estimated 10.0 percent wage premium for working in a plant with the mean failure probability rather than working in a plant with a zero probability of failure. Reducing the annual replacement rate by 25 percent from its average sample value, would raise the wage premium to 11.3 percent. For plants whose workers are constrained by the maximum annual UI payment ($D_{2i} = 1$) the implied wage premium for working in a plant with the mean failure probability is estimated to be 5.5 percent. A 25 percent decline in the maximum annual UI payment level would raise the compensating differential to 5.9 percent. In both cases the compensating differential increases as the UI benefits are decreased and the wage premium appears more sensitive to the UI program among lower wage plants.

When the sample is disaggregated into single-unit and multiunit plants the presence of unemployment insurance continues to reduce the compensating differential for both groups of plants although neither coefficient is statistically significant for the single-unit plants. Overall there is evidence of a positive effect of plant failure probabilities on wages with the effect being reduced by unemployment insurance programs. The effect is stronger for the multiunit plants.

VI. Conclusions

This paper uses micro data on over 6500 U.S. manufacturing plants to estimate a model of plant failure and wage determination. Estimates of the compensating differential which employees require to work in a failure-prone plant and the effect of wage changes on the probability of plant survival are constructed. The results indicate that workers in a plant which has the average probability of failure in the sample, .135 over four years, earn 7.3 percent higher wages than workers in a plant with a zero probability of failure. In addition, wage increases for production workers have little effect on the probability of plant failure. A ten percent increase in wages, holding plant revenue fixed, increases the probability of failure by .15 percentage points.

The empirical results also indicate a substantial role for plant characteristics in explaining variation in plant failure and wages. In the case of plant failure, increases in the plant's market share, age, and past employment growth or increases in the unionization rate in the plant's industry and geographic region increase failure rates. The hedonic wage regressions indicate that wages also increase with plant size, age, capital intensity, and value-added per worker. Plant characteristics are substantially more important for plants owned by multiplants firms.

Ideally an empirical study of wages, whether measured for individual workers or as an average wage for an employer, should control for characteristics of both the worker and employer. Because data on the characteristics of employers has been difficult to obtain, the demand side of the labor market is frequently either ignored or controls are limited to industry-level characteristics. While this study controls for a detailed set of employer characteristics, it is unable to control for differences in the characteristics of workers across plants. If the average quality of workers differs across manufacturing plants then this could contribute to inter-plant variation in average wages.[26] However, the effect of this omitted variable bias on the estimated compensating differential for plant failure is not likely to be large. The estimated wage regressions have controlled for plant size, age, capital intensity, value added, and industry, all of which are more likely to be correlated with average worker quality than is the probability of failure. In addition, this study has been limited to production worker wages in order to reduce the importance the inter-plant variation in worker skills.

Controlling for employer heterogeneity is especially important when quantifying the compensating differential for plant closing because that decision depends on the efficiency of the plant relative to competitors in its output market. Empirical studies which must rely on inter-industry differences in plant closing or layoff rates are unable to account for the substantial variation in unemployment risk faced by different workers that results from the within-industry heterogeneity of the employers. In our data this within-industry variation accounts for 92.6 percent of the total sample variation in the predicted probability of plant failure. In terms of the empirical model developed here it is this within-industry employer heterogeneity which allows identification of the compensating differential for unemployment risk.

Footnotes

[1]Hall (1972), Marsten (1985), and Topel (1986) examine regional variation in average wages and unemployment rates. Adams (1985) and Bronars (1983) explain variation in worker wages with measures of industry employment turnover including the variance of industry employment and the mean and variance of industry layoffs. Abowd and Ashenfelter (1981) use the time-series variation in the hours worked by an individual as a measure of the unemployment risk which the individual faces. Topel (1984) estimates individual-specific measures of unemployment risk which vary with the worker's characteristics and his industry of employment. Hamermesh (1989) concludes that "there is no evidence that [the costs of displacement to workers] are perfectly offset by prior compensating wage differentials." Rosen (1986) summarizes the theoretical framework underlying compensating differential studies.

[2]Output markets are defined at the four-digit level. Industry dummy variables are included at the two-digit level to make reporting and comparison with other studies easier. The model was also estimated with four-digit industry dummies and any results that were sensitive to this change are discussed below.

[3]Jovanovic's model assumes that producers have varying efficiency levels and that each producer learns his relative efficiency over time. The model predicts that larger plants and older plants will have lower probabilities of failure. Size, age, and ownership type were all found to affect plant failure in U.S. manufacturing in the study by Dunne, Roberts, and Samuelson (1989b).

[4]The actual union status of the plants is unknown. The industry-region unionization rate is constructed as the proportion of workers that report they are unionized in the 1973, 1974, 1975, and 1976 CPS.

[5]In general the variables in S_i should enter multiplicatively with P_i because they alter the effect of P_i on wages. Biddle and Zarkin (1988) and Viscusi and Moore (1987) discuss the need to use interaction effects to quantify the effect of worker's compensation on the compensating differential workers demand to accept jobs with a high risk of injury. Topel (1984) also models the effect of unemployment insurance using interaction terms.

[6]This particular <u>ASM</u> panel was chosen for analysis because we judged the quality of the time-series linkages for the plants to be superior to those of other <u>ASM</u> panels. In addition, only a single <u>ASM</u> panel was chosen for analysis because it allows us to construct plant failure rates that are not distorted by changes in the sample of plants being surveyed. For a full discussion of the <u>ASM</u> panels and Census Bureau plant-level data see McGuckin and Pascoe (1988).

[7]The industries were chosen to reflect the mix of high and low-wage industries in the U.S. manufacturing sector.

[8]Unemployment rates by state and SMSA were constructed from data reported in <u>Geographic Profile of Employment and Unemployment 1974</u>, U.S. Department of

Labor, Bureau of Labor Statistics, Report 452, 1976. Data was available for plants in 28 states and 34 SMSA's.

[9]A complete census of the manufacturing sector was taken in 1972. This criteria allows us to construct growth rates over the two year period 1972-74 for each plant in the sample and thus account for the fact that plants that closed over the 1975-78 period may have undergone gradual contraction prior to closing. Because of this we are unable to include plants that are very young (began operation in 1973 or 1974) in the sample. Alternatively we included these youngest plants in the sample and deleted the past growth variable. This resulted in a slight increase in the failure rate in the sample, from .135 to .156, because the youngest group of plants fail more often, but had no appreciable effect on the compensating differential estimates reported below.

[10]Additionally, plants that reported production worker wages less than .50 dollars per hour or more than 35.00 dollars were deleted as well as plants that had zero values for expenditure on materials or value of shipments.

[11]A second more stringent definition of plant closing was also employed. It requires that the plant actually report that it had closed its operation. This definition provides a conservative lower bound on failure rates because many plants close without actually reporting the fact to the Census Bureau. Under this definition 8.31 percent of the plants in the sample fail. Results using this definition are reported below.

[12]We also used a measure of worker compensation which included supplemental labor costs. The <u>ASM</u> collects information on total supplemental labor costs for each plant. These costs apply to both production and nonproduction workers and include required payments such as social security and unemployment insurance as well as fringe benefits. We divided these supplemental labor costs between production and nonproduction workers using the proportion of total plant wages paid to the two groups. An average hourly wage for production workers which includes supplements was then constructed by dividing total production worker wages plus imputed supplements by production worker hours. Results using this alternative wage variable are discussed below.

[13]See Dunne, Roberts, and Samuelson (1989b) for a more detailed description of the methodology used to assign the plant age categories. Measurement error will tend to result in plant's being assigned to age categories that are too young. This will tend to reduce any effect of age in the regressions.

[14]In both equations the standard errors are corrected for the fact that at least one of the regressors in each equation is a predicted value from a reduced-form regression and is thus subject to sampling variation. The methodology developed by Murphy and Topel (1985) is used.

[15]This hedonic wage model implicitly assumes that workers are identical in the preferences toward risk of job loss and wages. Variations in these plant characteristics and wages thus represent different points on the worker's indifference curve between risk and wages. Studies using worker surveys to quantify the compensating differential assume that the only producer characteristic that affects the probability of job loss is the

industry of occupation. As a result these studies rely on industry dummy variables to identify the compensating differential.

[16]This effect is constructed by calculating the estimated probability of failure evaluated at the means of all the explanatory variables. The expenditure on labor is then increased by ten percent and the estimated probability of failure recalculated holding all other variables fixed.

[17]These age and size patterns are similar to those reported in Dunne, Roberts, and Samuelson (1989b). Their findings were based on a very different data set which included over 250,000 plants observed at five year intervals over the 1967-1982 period.

[18]Blanchflower and Oswald (1989a) estimate a probit model for plant failure using establishment data from England. They find that an increase in the two-digit industry unionization rate decreases the plant closure rate. This effect appears strongest in the highly unionized sectors. They argue this negative correlation may arise because unions raise plant shutdown costs. In their study they also find that plant size and age are negatively correlated with the closure rate and that the average industry wage has no effect on plant closing.

[19]If $\hat{w}_1$ is the predicted wage in the plant with a failure probability of .135 and $\hat{w}_0$ is the predicted wage in the riskless plant then relative wages are calculated as $\hat{w}_1/\hat{w}_0 = \exp((.519)(.135)) = 1.073$.

[20]Brown and Medoff (1989) suggest that the robust relationship between employer size and wages may result from failure to control for plant age. While our results indicate that age is an important factor in explaining wage variation, the size effect still remains.

[21]Blanchflower and Oswald (1989b) review the literature on the relationship between unemployment and wages. They report that most studies using micro data on workers find a small negative wage elasticity with respect to the unemployment rate. They also find a negative relationship in their empirical work. Blanchflower, Oswald, and Garrett (forthcoming) also report that higher county unemployment rates are correlated with lower wages using establishment data. In contrast studies that use regional data, including Hall (1972), Marsten (1985), and Topel (1986), find a positive correlation between unemployment rates and average wages across different geographic areas.

[22]The model was also estimated using four-digit rather than two-digit industry dummy variables. The only variable to change in statistical significance was the unionization rate in the wage regression. The coefficient fell from .072 to .051 with a standard error of (.027). Including four-digit industry dummies reduced the coefficient on the probability of failure slightly from .519 to .472.

[23]Most states also have minimum UI payments which are made to workers in low-wage jobs. In our sample all plants had average wage rates that would qualify a worker for more than the minimum UI payment. As a result we treat plants as falling above or below the statutory maximum.

[24]All information on unemployment insurance coverage is drawn from "Significant Provisions of State Unemployment Insurance Laws, July 8, 1974," United States Department of Labor, Manpower Administration, Unemployment Insurance Service.

[25]See Brown (1980) for discussion of omitted variables bias in compensating differential models.

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