WORKING PAPER
ALFRED P. SLOAN SCHOOL OF MANAGEMENT

UNION WAGE GAINS UNDER REGULATION: EVIDENCE FROM THE TRUCKING INDUSTRY

Nancy L. Rose
MIT
June 1985

MIT Sloan School of Management Working Paper: #1683-85

MASSACHUSETTS
INSTITUTE OF TECHNOLOGY
50 MEMORIAL DRIVE
CAMBRIDGE, MASSACHUSETTS 02139

UNION WAGE GAINS UNDER REGULATION: EVIDENCE FROM THE TRUCKING INDUSTRY

Nancy L. Rose
MIT
June 1985

MIT Sloan School of Management Working Paper: #1683-85

I am grateful to Hank Farber, Paul Joskow, Frank Fisher, and Jim Poterba for helpful conversations and comments. Wayne Gray and the National Bureau of Economic Research provided data assistance. Financial support was provided by a National Science Foundation fellowship and a Sloan Foundation grant to the MIT economics department.

This paper uses wage evidence to investigate the ability of the Teamsters Union to capture a share of regulatory rents in the trucking industry. Previous work has documented the existence of monopoly rents in the motor carrier industry, and has linked these to economic regulation of the industry by the Interstate Commerce Commission (ICC).[1] However, there has been relatively little economic analysis of the Teamsters' role in the trucking industry.[2] This is a potentially serious oversight; the Teamsters have been a powerful force in the industry, and been perhaps as influential as ICC regulation in determining the performance of trucking firms. In addition, an investigation of union behavior in this industry may yield insights relevant to other regulated markets. The interaction of unionization and regulation was explored briefly by Hendricks (1975), with respect to electric utilities. The trucking industry is a fertile ground for evaluating potential interactions, due to the strength of both the union and regulatory protection.

The study analyzes labor market data to determine union losses resulting from motor carrier deregulation. I estimate union wage premia over the 1973 through 1984 period, based on data from the Current Population Survey. Because the data span the period before and after trucking deregulation, they can be used to examine the response of union wage differentials to regulatory reform. This provides evidence on

---

[1]See Moore (1978, 1983), Frew (1981), and Rose (1984), for example.

[2]The published empirical work in this area appears to consist of studies by Annable (1973), Moore (1978), and Garnel (1972). Unpublished theses by Arnold (1970) and Hayden (1977) also address aspects of this issue.

union rent losses resulting from increased competition in the trucking industry. The estimated losses will, in general, be a lower bound on the total rents captured by the union under regulation.[3]

The paper also extends earlier studies that have attempted to determine Teamster Union rent gains from estimates of the level of the union wage premium under regulation (see Annable (1973), Moore (1978), and Hayden (1977)). Most previous calculations rely heavily on aggregate data on average annual compensation; micro data estimates have been made only for 1967 in the published literature (Moore, 1978), and for 1973-75 in an unpublished study (Hayden, 1977).

The results suggest union wage premia in the neighborhood of 55 percent during the early to mid-1970s. These are considerably higher than average union differentials for the labor force as a whole, although they are comparable to the estimates obtained by Moore and Hayden for the trucking industry. The union premium dropped sharply with the onset of motor carrier deregulation in the late 1970s, to roughly 25 percent of nonunion wages, and remained substantially below earlier levels through 1984. The premium in real dollars per hour for a "representative" worker declined by one-third over the decade studied.

The study is divided into two sections. The first describes the data and methodology. The second section presents the results and discusses their implications for the interaction of unionization and regulation.

---

[3]A number of factors may create a wedge between the level of the union premium as estimated here, and total union rents. These include possible productivity differences between union and nonunion workers, differential non-wage compensation, and biases in cross-sectional estimates of union wage differentials.

I. Methodology and Data

This study estimates wage equations from a cross-section of individual workers' wages and characteristics. The data are from the Current Population Surveys (CPS) conducted by the Bureau of the Census. Wages are likely to be a function of four factors: workers' characteristics including union status, firm characteristics, occupation and industry specific effects, and geographic wage levels. The May Current Population Surveys (CPS) provide information on most of these factors. Available data on worker qualities include union status, education, age, sex, race, and marital status. I control for industry and occupation effects by limiting the study to truck drivers employed in the for-hire trucking industry. Data on worker location are used to control for geographical wage variation via regional fixed effects.

Unfortunately, the May CPS typically does not provide information on the firms employing the responding individuals.[4] The inability to control for firm characteristics may bias the estimated union wage differential. For example, if small firms typically pay lower wages, ceteris paribus, because nonunion drivers in the trucking industry are more likely to work for smaller firms, omitting firm size may cause the estimates to overstate the union wage premium.[5] This suggests that one should use caution interpreting the <u>level</u> of the estimated premium,

---

[4]Supplements occasionally gather data on workers' establishments; see, for example, the May 1979 Pension Plan Supplement.

[5]Mellow (1983) investigates the firm size effect on wages, and finds that estimated union wage differentials are smaller in the presence of controls for firm characteristics, notably size. In addition to firm size, differences in firm markets may affect estimated differentials.

although this potential bias should not invalidate tests of changes in union premia through time.[6] The Current Population Survey is otherwise an excellent source of data for this project, as it provides information on large samples of workers (necessary to obtain adequate subsamples of truck drivers) and is available over a long period of time, permitting analysis of union premiums before and after trucking deregulation. Therefore, with the caveats described above, I rely on the CPS data to estimate my wage equations.

The investigation covers the period 1973 through 1984. The CPS survey design is such that half the respondents in one year are reinterviewed the following year. Because of this overlap, and to conserve computational resources, I use every other year's survey during the 1970s (1973, 1975, 1977, and 1979). This yields an essentially complete, though non-overlapping, set of CPS respondents from 1973 through 1979. Beginning in 1980, the Census Bureau sharply reduced the number of respondents that were asked both union status and wage questions, resulting in substantially smaller potential sample sizes during the 1980s. This, combined with my interest in the pattern of wage behavior after deregulation, leads me to use every year's survey

---

For example, the inability to distinguish workers employed in the more profitable less-than-truckload (LTL) sector from those employed in the more competitive, less unionized, and less profitable specialized commodity sector may distort the results. The inability to exclude United Parcel Service (UPS) employees, who negotiate a separate union contract and who are unlikely to have been affected by deregulation of the for-hire trucking industry, may bias estimates against a decline in the post-deregulation union premium.

[6]Freeman and Medoff (1984) discuss other potential sources of bias in cross-sectional estimates. However, even if the level of the differential were overstated by the use of cross-sectional data without adequate controls for all worker and firm characteristics, there is no

from 1980 to 1984 (with the exception of May 1982, for which union status information is missing from my CPS tape).[7]

My sub-sample is restricted to full-time truck drivers working in the trucking industry. This excludes non-driver employees in the for-hire trucking industry, as well as truck drivers who work in private carriage. This narrow sample definition should control for most industry and occupation specific effects on wage. Because women comprise a very small fraction of truck drivers (frequently 0 in the CPS samples, and never more than 2 percent), they are excluded from the sample. No self-employed drivers (owner-operators) satisfied criteria for inclusion in the sample.

I estimate a conventional semi-log wage equation of the form:[8]

$$(1)\quad \text{LHWAGE} = \beta0 + \beta1\cdot\text{UNION} + \beta2\cdot\text{EDUC} + \beta3\cdot\text{EXP} + \beta4\cdot\text{EXP}^2 + \beta5\cdot\text{NONWHITE} + \beta6\cdot\text{SINGLE} + \beta7\cdot\text{NE} + \beta8\cdot\text{SOUTH} + \beta9\cdot\text{WEST}$$

where: LHWAGE = natural log of the hourly wage rate

UNION = 1 if a union member, 0 otherwise. If union membership data were missing, UNION was coded as 1 if the worker reported being covered by a union contract, 0

---

reason to expect the bias to alter in such a way as to substantially <u>reduce</u> estimated union wage differentials over time. Indeed, the rapid growth of new, smaller nonunion firms in the 1980s, and the maintentance of higher wages by drivers for UPS (see note 2, supra) might exacerbate the potential for overstatement of union wage differentials post-deregulation.

[7]Estimates for 1979-1980-1981 and for 1983-1984 may not be statistically independent, since the samples for contiguous years will include about half of the same individuals, given the survey design.

[8]This equation is standard in much of the labor economics literature. See Bloch and Kuskin (1978) for a discussion of this type of equation versus separate union and nonunion equations.

otherwise. Since the Teamsters' contracts typically impose union shops, this seems an appropriate substitution.

EDUC = number of years of schooling completed

EXP = experience, defined as (Age - EDUC - 6)

NONWHITE = 1 if race is non-white, 0 otherwise

SINGLE = 1 if marital status is other than married with spouse present, 0 if status is married, spouse present.

NE = 1 if region is Northeast, 0 otherwise

SOUTH = 1 if region is Southern, 0 otherwise

WEST = 1 if region is Western, 0 otherwise

The North Central region dummy variable is omitted, so that estimates are relative to this region. The results presented below are robust to variations in this specification.

II. Results

Table 1 presents estimated union wage coefficients and sample mean wages (in logs) and unionization rates for each year of data. The estimates span four union contract (National Master Freight Agreement) periods: 1973-75, 1976-78, 1979-81, and 1982-84. The full set of estimated coefficients for each equation are reported in appendix table 1A. The remaining coefficients reported in the appendix table generally are of the expected sign, although their effects are often quite small and imprecisely estimated. This is not particularly surprising, given how narrowly the sample is defined. One might expect that wages of full-time truck drivers in the for-hire trucking industry would not vary much across workers with different education or experience levels.

Several aspects of Table 1 deserve mention. Note first that the estimated union coefficients prior to deregulation--in 1973, 1975, and 1977--are all tightly clustered at .44. This implies union premia of 55 percent above nonunion wages. These findings comport well with Moore's (1978) estimate of a 48 percent premium in 1967 and Hayden's (1977) estimated 50 percent premium in 1973-75.[9] However, these differentials are nearly twice the average union differential for cross-sections of industries estimated on similar data sets. For example, Freeman and Medoff (1984, p.46) report an average union premium of 21 percent for a cross-industry sample of respondents from the 1979 CPS. This suggests that the Teamsters may have been able to capture a larger

[9]The data I use overlap with Hayden's data, although the specification of the wage equation differs slightly from his. Because of this, the similarity in results is to be expected at least in the 1973 and 1975 equations.

TABLE 1

Estimated Wage Premiums and Sample Characteristics:

Annual Data

Year	Union Coefficient	Mean LHWAGE Union	Mean LHWAGE Nonunion	Proportion UNION	Sample Size
1973	.440 (.050)	1.692	1.210	.63	231
1975	.448 (.056)	1.834	1.322	.55	245
1977	.441 (.040)	1.990	1.513	.60	303
1979	.225 (.055)	2.061	1.793	.58	168
1980	.199 (.089)	2.070	1.819	.58	91
1981	.231 (.089)	2.229	1.970	.61	82
1983	.350 (.084)	2.268	1.915	.55	84
1984	.321 (.081)	2.342	1.974	.34	89

share of rents for truck drivers than the average union captured for its members.[10]

The second interesting feature of this table is that the union differential drops sharply in 1979. From 1979 through 1981, the union coefficient is about .22--half its level in 1973-1977--implying premia of only 25 percent above nonunion wages. Although the estimated union coefficient increases somewhat in 1983 and 1984, to an average of .33, it remains substantially below the pre-deregulation level.

Finally, there appears to be a reduction in the percent of drivers with union affiliation at the end of the sample period. In 1984, only 34 percent of the sample reported belonging to a union. Whether this is an aberration of the 1984 sample or indicative of a decline in the union presence in trucking remains an open question, and will not be addressed in this paper. Although much of the growth in trucking firms has been through entry of nonunion carriers, additional evidence is required to ascertain whether union representation in the industry has diminished by any substantial amount.

To test the statistical significance of the decrease in union wage differentials, I estimate wage equations pooling data across years.[11] These results are reported in Table 2. Separate intercepts are estimated

---

[10]This conclusion holds up against industry-specific union differentials estimated by Freeman and Medoff from 1973 CPS data. Freeman and Medoff (p. 50) report that half of the 62 industries in their sample had union premia greater than 15 percent, but only 8 of the 62 had union premia greater than 35 percent.

[11]I have estimated pooled equations omitting contiguous years (that is, eliminating 1980 and 1984 observations), to ensure independence of observations across time. Neither the coefficients nor the hypothesis tests for homogeneity across time are materially affected by excluding these years. The standard errors rise somewhat, but this is

TABLE 2

Pooled Wage Equations

Variable	1973-1977	1979-1984	1973-1984
UNION	.446 (.042)	.263 (.033)	---
UNION1 (1973-77)	---	---	.445 (.027)
UNION2 (1979-84)	---	---	.262 (.032)
EDUC	.018 (.007)	.012 (.009)	.016 (.005)
EXP	.015 (.004)	.018 (.005)	.016 (.003)
EXP^2	-.0003 (.0001)	-.0003 (.0001)	-.0003 (.0001)
SINGLE	-.028 (.042)	-.076 (.040)	-.057 (.029)
NONWHITE	-.127 (.048)	-.050 (.048)	-.087 (.034)
NE	-.042 (.035)	-.115 (.045)	-.070 (.028)
SOUTH	-.081 (.033)	-.070 (.041)	-.078 (.026)
WEST	.060 (.036)	.118 (.046)	.084 (.028)
Mean Intercept	1.070	1.637	1.421
NOB	779	514	1293
R^2	.41	.40	.45
SSR	88.76	61.30	151.08

These equations all include time effects for each sample year. The mean of the estimated time effects is reported as the mean intercept.

for each year, but all other coefficients are constrained to be the same within the pooled sample. The first column pools data from the pre-deregulation period, 1973 through 1977. Column 2 pools data over the deregulatory period, 1979 through 1984. These two samples are combined in column 3, which reports results for the full sample, 1973-1984, allowing the union coefficient to differ across regulatory regimes.

As anticipated from Table 1, the constrained union coefficient for the pre-deregulation period is quite precisely estimated at .446 with a standard error of .027.[12] By contrast, the union wage coefficient over the pooled deregulation sample in column 2 is .263 (standard error, .033). This is less than 60 percent of the size of the union wage effect measured in the column 1 results. Further, we cannot reject at conventional levels of statistical significance the hypothesis that wages over the five deregulation years are generated by a common process. The F-statistic to test the null hypothesis (H0) of homogeneous coefficients for 1979 through 1984 is .811, which is distributed as F(36,464) under H0.

The difference in the union coefficient pre- and post-deregulation is highlighted by the full sample results reported in column 3. The null hypothesis of homogeneous coefficients across samples 1 and 2, excepting the intercepts and the union coefficients, cannot be rejected

---

expected from the smaller sample size.

[12]The restrictions implied by pooling the regulatory years cannot be rejected at conventional levels of significance. The F-statistic to test these restrictions is 1.029, which is distributed as F(18,749) under the null hypothesis of homogeneous coefficients.

at conventional levels of significance.[13] The restriction of equal union coefficients for the two samples can be rejected at the .005 level.[14] This suggests that the dominant change in the wage behavior of truck drivers over the period studied is a decline in the ability of the union to maintain its wage advantage.

The decline in union wage differentials is coincident with the onset of ICC deregulation of the trucking industry, which began in earnest in the fall of 1978. However, although a number of indicators show the effects of increased competition in the industry as early as 1979,[15] the impact of competition became much more pronounced in 1980 and beyond. It therefore seems surprising that the relative union wage would respond to deregulation as early as 1979. To try to understand better the nature of the decline in the union differential, I calculate the predicted union and nonunion wage rates over time for a driver with fixed characteristics.

My "representative" driver is a married white male with 12 years of education, 20 years of experience, living in the North Central region. Wages are predicted for this "representative" driver from a single equation pooling data from all sample years. Intercepts and union

---

[13]The F-statistic to test this hypothesis is .922, which is distributed as F(62, 1213) under the null hypothesis of homogeneous coefficients.

[14]The t-statistic to test the hypothesis of equal union coefficients pre- and post-deregulation is 4.54.

[15]Moore (1983) looks at a number of industry indicators, including a sample of truckload and less-than-truckload rates, average revenues per ton-mile, return on transportation investment, entry applications and number of regulated carriers, average employee compensation, and many others. A number of these series show changes beginning in 1978 or

coefficients are permitted to vary across years; all other coefficients are constrained to be the same over time. The complete set of results for this equation are reported in appendix table 2A. Table 3 presents predicted wages in both current dollars (columns 1 and 2) and May 1984 constant dollars (columns 3 and 4).[16] The union differential in 1984 dollars per hour is presented in column 5.

Two aspects of this table are noteworthy. First, the declining percentage union differential observed in Tables 1 and 2 is associated with a decline in the growth rate of nominal union wages from 1977 to 1979, and an increase in the growth rate of nonunion wages between these years. Predicted union wages decline in real terms from 1979 on, with real wages in 1983-84 falling nearly 20 percent below the average predicted real wages in 1973-77. Non-union wages do not exhibit this pattern. Real nonunion wages remain high through 1981 (compared to their early 1970s levels), and experience a decline of only 8 percent in 1983-84 relative to 1973-77.

Second, the reduction in the dollar union wage premium for this representative worker is substantial. The union wage differential declines in <u>nominal</u> dollars in 1979-1981; in real terms the differential falls by one-half relative to its pre-deregulation levels. Despite an

---

1979; however, the larger effects are usually found in 1980 or 1981. It is also difficult to disentangle the role of deregulation from that of the 1979-80 recession in assessing the cause of the price, revenue, and rate-of-return reductions that Moore's series show in 1979.

[16]The constant dollar figures were obtained using the Urban Worker Consumer Price Index for May of each year.

TABLE 3

Predicted Union and Nonunion Hourly Wages
For a Representative Driver
(Pooled Sample Estimates)

	Nominal Dollars		Constant 1984 Dollars		
Year	Predicted Union Wage	Predicted Nonunion Wage	Predicted Union Wage	Predicted Nonunion Wage	Real Differential
1973	\$6.24 (.17)	\$4.01 (.20)	\$14.70	\$9.45	\$5.25
1975	7.04 (.22)	4.51 (.21)	13.70	8.77	4.93
1977	8.36 (.21)	5.36 (.22)	14.35	9.20	5.15
1979	9.04 (.32)	7.17 (.32)	13.08	10.37	2.71
1980	9.22 (.52)	7.39 (.40)	11.67	9.35	2.32
1981	10.73 (.61)	8.56 (.52)	12.35	9.86	2.50
1983	10.91 (.63)	7.93 (.45)	11.38	8.27	3.11
1984	12.10 (.75)	8.60 (.41)	12.10	8.60	3.50

Predicted values are computed as $\exp(X\beta + \frac{1}{2}\sigma^2)$, where $X\beta$ is the predicted log wage and σ^2 is the error variance from a single wage equation pooling all sample years, allowing for separate intercepts and union coefficients for each year. Coefficients are reported in table 2A. Standard errors of predicted values are in parentheses. The representative driver is a married, white male with 12 years of education and 20 years of experience, living in the North Central region.

increase in the constant dollar differential in 1983-84, it remains only two-thirds of the 1973-77 level. Thus, the union is losing not only in its percentage premium relative to nonunion wages, but also in the dollar amount of the union wage premium. The union premium in 1975 was $2.53 per hour in nominal dollars, or about $5700 per year. In 1983, the current dollar premium was $2.98 per hour, or $6700 per year. This represents a real decline of nearly $4100 per year in 1984 dollars.

This wage pattern is unlikely to be explained by an institutional failure of the union contract to compensate adequately for inflation. The major union contract in the trucking industry is the National Master Freight Agreement (NMFA), estimated to cover 300,000 to 500,000 trucking employees during the 1970s. The initial decline in the union premium comes between 1977 and 1979, despite an uncapped cost-of-living adjustment clause in the NMFA in force from 1976 through March 1979. Second, the 1979 data are from May, one month after the new National Master Freight Agreement for 1979-1981 was signed. Presumably, the wage data should reflect the newly negotiated union wages. Finally, the decline in the the differential persists through the 1979 contract and into the 1982 contract. Explaining this pattern seems to require a more fundamental cause than unanticipated inflation levels in the mid-1970s.

Nor does the declining union differential seem to be part of a general trend in union wages. Freeman and Medoff (1984, pp. 52-54) argue that the overall union differential <u>widened</u> during the late 1970s. This is counter to the pattern observed in my results for the trucking industry, which indicate a dramatic and lasting reduction in the union premium after 1977. While industry observers are quick to attribute wage concessions in the 1982 contract to the effects of deregulation, it appears that deregulation reduced the union's relative wage advantage long before the 1982 negotiations.

The magnitude of the loss implied by these findings is substantial. A conservative estimate of the reduction in union rents is calculated as follows: First, I assume that deregulation reduced the 1973-77 union premium of 55 percent to the 1983-84 average premium of 39 percent. This understates the change implied by the 1979-84 average premium of 30 percent reported in column 3 of Table 2. Second, I assume that non-union wages are unaffected by deregulation. If nonunion wages were bid up under regulation (e.g., due to the union "threat" effect), then part of the decline in nonunion wages should be attributed to deregulation. My calculation will omit any loss due to a decline in the overall wage level. Finally, I exclude potential losses from a reduction in union employment. There is substantial anecdotal evidence that Teamster unemployment rates in the trucking industry have increased substantially since deregulation,[17] and that nonunion operations have captured an increasing share of the market. My computation of rent reductions will exclude the union's loss on jobs that are no longer held by union employees.

Given this set of assumptions, the loss to the union can be calculated as 11.5 percent of union compensation, which is the decline

---

[17]The Wall Street Journal (October 9, 1984, p.1) reports that the Teamsters Union estimates 100,000 union jobs have been lost since 1980. Business Week (September 26, 1983, p. 147N) reports that layoffs have left one-third of the Teamsters' trucking members without union jobs. It is difficult to disentangle the role of regulatory reforms from that

relative to what compensation would have been had the 55 percent union differential been maintained.[18] I apply this to the 1983 aggregate labor bill for 886 Class I motor carriers. The result suggests an annual reduction of $950 million in union rents for employees of these firms.[19]

---

of the 1981-82 recession; however, the persistence of such high union unemployment rates appears to be considered abnormal by industry observers.

[18]The loss (Δ) is given by: $\Delta = (1.39-1.55)\cdot$nonunion wage $= -.16\cdot$ nonunion wage. The percent loss is Δ/union earnings $= -.16/1.39 = .115$.

[19]This calculation assumes that 60 percent of the employees are unionized, and that union earnings are 1.39 times nonunion earnings. Applying this to 1983 total employee compensation of $12.24 billion for the 886 Class I motor carriers reported by the ICC yields total union compensation of $8.27 billion, and a rent loss of $951 million (11.5 percent). Compensation data are from U. S. Interstate Commerce Commission Bureau of Accounts, Transport Statistics in the United States. Motor Carriers. Part 2, 1983. Class I motor carriers are regulated firms with more than $5 million in annual gross revenues. Although this calculation excludes most trucking firms (which are quite small), these are the very largest companies in the industry.

Conclusion

This study indicates substantial losses to union employees in the trucking industry as a consequence of motor carrier deregulation. Union wages in the early to mid-1970s were estimated to be 55 percent greater than nonunion wages, holding constant worker characteristics. This is consistent with the results of earlier studies of union wage differentials in the regulated trucking industry. However, this differential declined sharply, beginning in 1979. The percentage union wage premium drops by one-third to one-half with the regulatory reforms of the late 1970s and early 1980s. Annual earnings for a "representative" union driver in 1983-1984 were $3000 (11 percent of earnings) less than they would have been had the pre-deregulation union wage differential been maintained. The annual loss in union rents is computed as $950 million for employees of the 886 largest regulated trucking firms. It appears that deregulation, which I have shown in earlier work (Rose (1984)) dramatically decreased the expected profitability of trucking firms, also substantially reduced the rents captured by unionized labor in the trucking industry.

TABLE 1A

Annual Wage Equations

Variable	1973	1975	1977	1979	1980	1981	1983	1984
Constant	.832	1.071	1.225	1.761	1.432	1.487	1.688	1.560
	(.185)	(.184)	(.130)	(.202)	(.331)	(.354)	(.363)	(.325)
UNION	.440	.448	.441	.225	.199	.231	.350	.321
	(.050)	(.056)	(.043)	(.055)	(.089)	(.089)	(.084)	(.081)
EDUC	.025	.014	.020	.005	.018	.029	.019	.019
	(.013)	(.013)	(.009)	(.014)	(.023)	(.024)	(.025)	(.025)
EXP	.017	.013	.018	.007	.023	.021	.0008	.031
	(.008)	(.008)	(.006)	(.008)	(.012)	(.012)	(.015)	(.011)
EXP^2	-.0004	-.0002	-.0004	-.0001	-.0003	-.0004	-.0001	-.0007
	(.0002)	(.0002)	(.0001)	(.0002)	(.0003)	(.0002)	(.0003)	(.0002)
SINGLE	.020	.012	-.079	-.182	-.031	.078	-.181	.001
	(.085)	(.090)	(.054)	(.072)	(.015)	(.111)	(.109)	(.081)
NONWHITE	-.198	-.159	-.062	-.150	-.023	-.144	.077	.044
	(.089)	(.107)	(.065)	(.080)	(.120)	(.133)	(.136)	(.112)
NE	.090	-.078	-.109	-.156	-.161	-.202	.086	-.172
	(.065)	(.073)	(.050)	(.077)	(.124)	(.116)	(.112)	(.106)
SOUTH	-.075	-.058	-.102	-.110	-.085	-.034	.004	-.120
	(.056)	(.065)	(.051)	(.067)	(.110)	(.118)	(.114)	(.096)
WEST	.043	.164	-.012	.121	-.036	.121	.201	.463
	(.072)	(.072)	(.049)	(.080)	(.117)	(.124)	(.115)	(.111)
NOB	231	245	303	168	91	82	84	89
R^2	.38	.35	.42	.26	.19	.24	.33	.26
SSR	25.02	34.41	27.19	17.26	11.49	10.01	9.62	9.29

TABLE 2A

RESULTS FOR POOLED SAMPLE WAGE EQUATION

Year	Constant	Union Coefficient	Variable	Coefficient
1973	.944 (.084)	.444 (.052)	EDUC	.016 (.006)
1975	1.063 (.089)	.445 (.050)	EXP	.015 (.003)
1977	1.235 (.078)	.445 (.040)	EXP^2	-.0003 (.0001)
1979	1.526 (.089)	.232 (.053)	SINGLE	-.059 (.029)
1980	1.556 (.094)	.222 (.075)	NONWHITE	-.089 (.032)
1981	1.703 (.098)	.227 (.079)	NE	-.069 (.027)
1983	1.627 (.096)	.319 (.077)	SOUTH	-.078 (.026)
1984	1.708 (.088)	.342 (.075)	WEST	-.084 (.027)

Number of Observations: 1293
R^2 .46
SSR 150.79

REFERENCES

Annable, James E. "The ICC, the IBT, and the Cartelization of the American Trucking Industry." QUARTERLY REVIEW OF ECONOMICS AND BUSINESS 13:2 (1973) 33-47.

Arnold, T. R. "The Teamsters Union as a Determinant of the Structure of the Trucking Industry." Unpublished PhD. dissertation. Department of Economics, Syracuse Univeristy. 1970.

Bloch, Farrell E. and Kuskin, Mark S. "Wage Determination in the Union and Nonunion Sectors." INDUSTRIAL AND LABOR RELATIONS REVIEW 31 (1978) 183-204.

Freeman, Richard and Medoff, James. WHAT DO UNIONS DO? New York: Basic Books, Inc. 1984.

Frew, James R. "The Existence of Monopoly Profits in the Motor Carrier Industry." JOURNAL OF LAW AND ECONOMICS 24 (1981) 289-315.

Garnel, Donald. THE RISE OF TEAMSTER POWER IN THE WEST. Berkeley: University of California Press. 1972.

Hayden, James F. "Collective Bargaining and Cartelization: An Analysis of Teamster Power in the Regulated Trucking Industry." Senior Honors Thesis. Harvard College. 1977.

Hendricks, Wallace. "The Effect of Regulation on Collective Bargaining in Electric Utilities." THE BELL JOURNAL OF ECONOMICS 6 (1975) 451-465.

Mellow, Wesley. "Employer Size, Unionism, and Wages." In NEW APPROACHES TO LABOR UNIONS, RESEARCH IN LABOR ECONOMICS, SUPPLEMENT 2, pp. 253-282. Edited by Joseph Reid, Jr. Greenwich, CN: JAI Press, Inc. 1983.

Moore, Thomas G. "The Beneficiaries of Trucking Regulation." JOURNAL OF LAW AND ECONOMICS 21 (1978) 327-343.

__________. "Deregulating Rail and Truck Transportation." Hoover Institution Working Paper in Economics E-83-13. July 1983.

Rose, Nancy L. "The Incidence of Regulatory Rents in the Motor Carrier Industry." Mimeo. MIT Department of Economics. 1984.