WORKING PAPER
ALFRED P. SLOAN SCHOOL OF MANAGEMENT

LABOR RENT-SHARING AND REGULATION: EVIDENCE FROM THE TRUCKING INDUSTRY

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Revised September 1986

Sloan School of Management Working Paper: #1828-86

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This paper is a substantially revised version of an earlier working paper, number 1683-85 (July, 1985). I am grateful to Orley Ashenfelter, Henry Farber, Wallace Hendricks, James Poterba, Lawrence Summers, and seminar participants at Chicago, Columbia, MIT, Pennsylvania and Yale for helpful conversations and comments. I thank Henry Farber, Wayne Gray, and the National Bureau of Economic Research for assistance with the CPS data. Lars Bespolka and Anne Grøn provided excellent research assistance. This research was begun with financial support from a National Science Foundation fellowship and a Sloan foundation grant to the MIT economics department.

LABOR RENT-SHARING AND REGULATION: EVIDENCE FROM THE TRUCKING INDUSTRY

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An increasing body of evidence suggests that labor may be an important claimant to firms' profits. While there is a well-developed literature on union rent-sharing, recent empirical and theoretical work suggests that nonunion workers also may capture a share of rents. This study uses wage responses to an exogenous reduction in trucking industry rents to estimate the extent of union and nonunion rent-sharing. The results should be of interest to both labor economists interested in noncompetitive theories of wage determination and regulatory economists interested in assessing the magnitude and distribution of regulatory rents.

The paper evaluates wage responses to motor carrier deregulation in the late 1970s and early 1980s. The results indicate substantial declines in union premia over nonunion wages as a consequence of regulatory reforms. Union premia dropped from an average of 50 percent over nonunion wages during the 1973-78 period, to 30 percent over nonunion wages during 1979-84. Data for nonunion workers in the industry indicate wage declines relative to economy-wide average wages, although the magnitude of nonunion rents is not precisely quantified.

The 1983 annual loss in aggregate union rents is estimated at $657 to $1.2 billion for union employees of the 886 largest regulated trucking firms. These results suggest that union workers captured 60 to 70 percent of total industry rents, and provide strong support for labor rent-sharing hypotheses.

Increasing evidence suggests that labor may be an important claimant to firms' profits. Most studies in this area have analyzed the ability of labor unions to capture a share of rents. For example, Clark (1984), Ruback and Zimmerman (1984), and Salinger (1984) use cross-sectional data on firms or lines of business to estimate the effect of unionization on profitability. More recent theoretical and empirical studies by Dickens (1986), Dickens and Katz (1986a,b), Krueger and Summers (1986a,b), and Rotemberg and Saloner (1986) provide support for noncompetitive models of wage determination in the nonunion, as well as union, sector. These analyses suggest that rent-sharing may extend to nonunion workers, although they do not provide direct tests of this effect.[1]

The importance of labor rent-sharing extends well beyond the labor economics literature. For example, understanding rent-sharing is essential to analyzing government regulation. Regulatory protectionism can create rents over which workers and firms may negotiate; regulatory profit constraints may distort firms' labor input decisions or alter firms' relative bargaining strength vis-à-vis unions; the political nature of regulatory agencies can expand the scope of potential games between firms and workers. The extent of union rent-sharing and collective bargaining patterns in regulated industries have been explored by a number of authors, including Hendricks (1975, 1977) and Ehrenberg (1979), for

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[1] Dickens and Katz (1986a,b) and Krueger and Summers (1986a,b) report substantial dispersion in wages across industries, holding constant occupation and worker characteristics. They find high correlatations of industry wage differentials across occupations, countries, and union and nonunion workers. Theoretical models that predict nonunion rent-sharing include Dicken's (1986) model of union threat effects and Rotemberg and Saloner's (1986) game-theoretic model of wage determination for unorganized workers.

example.[2] Failure to account for these effects may lead to underestimates of regulatory rents and distortions.

Examining wage responses to exogenous reductions in rents, such as those caused by regulatory reforms, can provide a strong test of both union and nonunion rent-sharing.[3] My analysis focuses on the trucking industry, which for several reasons provides a fertile ground for evaluating rent-sharing hypotheses. First, previous work has documented the existence of monopoly rents in the motor carrier industry, and linked these to economic regulation of the industry by the Interstate Commerce Commission (ICC).[4] Second, regulatory reforms in the late 1970s and early 1980s provide a natural experiment that allows us to observe wage responses to reductions in these rents. Finally, a single, very powerful union--the International Brotherhood of Teamsters--represents almost all unionized workers in the heavily unionized for-hire trucking industry. This tends to increase the bargaining power of organized labor in trucking, and suggests that, at the very least, the industry should provide a strong test of union rent-sharing.

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[2] Hendricks (1975) analyzes electric utilities' relative bargaining strengths in different regulatory environments; Hendricks (1977) investigates occupational wage patterns across regulated and unregulated industries; and Ehrenberg (1979) discusses the effect of regulatory constraints and political influences on collective bargaining, focusing on telecommunications in New York. In addition, Hendricks, Feuille, and Szerszen (1980) and Card (1985) analyze interactions between regulation and collective bargaining in the airline industry; Arnold (1970), Annable (1973), Hayden (1977), and Moore (1978) discuss potential Teamster Union rents in their investigations of trucking regulation.

[3] Card (1985) and Hendricks et al. (1980) explore the impact of airline deregulation on unionized labor in that industry. They find only limited effects from deregulation, although Hendricks et al. note that this may be a result of substantial union power combined with a bargaining structure that is impervious to regulatory changes.

[4] See Moore (1978, 1986), Frew (1981), and Rose (1985a), for example.

This paper uses wage responses to motor carrier deregulation to estimate the extent of labor rent-sharing. The first two sections of the paper concentrate on union wage responses to motor carrier deregulation. Section I examines union contracts for indications of changes in bargaining outcomes after deregulation. The contracts reveal substantial differences between settlements before and after the Motor Carrier Act of 1980, with contracts in 1982 and 1985 specifying considerably less favorable terms for union workers relative to historical patterns. However, it is difficult to interpret this evidence without reference to some benchmark wage behavior. Aggregate earnings data indicate a decline in average trucking wages, but cannot distinguish a decline in the wage level from a decline in unionization rates. I therefore evaluate union wage behavior against a nonunion trucking wage benchmark.

Section II uses data from the Bureau of the Census's Current Population Surveys (CPS) to estimate union wage premia over nonunion wages for the 1973 through 1984 period. Because these data span the period before and after trucking deregulation, they can be used to examine the response of union wage differentials to regulatory reform. The results indicate declines in the average union wage premia from 50 percent of nonunion wages during the early to mid-1970s, to 30 percent after deregulation. Annual earnings for a "representative" union driver in 1983-84 were at least $2200, or 8 percent, less than they would have been had the regulatory union wage differential been maintained.

Nonunion rent-sharing is evaluated in section III. The behavior of nonunion wages in the trucking industry is compared to that of average wages for the economy as a whole. This analysis provides evidence on the impact of deregulation on nonunion wages, and hence on possible nonunion rent-sharing. The results indicate some decline in nonunion trucking wages relative to economy-wide average wage levels. Although this effect is not precisely quantified, it provides some support for models of nonunion rent-sharing.

The implications of these results for aggregate labor rents in the trucking industry are developed in section IV.

I. Union Rent-sharing: Contract Evidence

Contract agreements provide a natural starting point for evaluating the effect of regulatory reforms on union-firm relations. The Teamsters Union has negotiated national agreements in the trucking industry since 1964. These "National Master Freight Agreements" (NMFAs) specify wage increases and benefits changes, as well as a broad range of work rules, over each three-year contract period. Supplements to the NMFA allow for local and regional variations in wage levels. The union also negotiates a number of separate specialized commodities agreements, which apply to firms in the least unionized sectors of the motor carrier industry.[5] The National Master Freight Agreement is the most extensive of these various agreements: about 75 percent of the Teamsters' 375,000 to 400,000 freight division members in the mid-1970s were covered by the NMFA.

Identifying the timing of regulatory reforms is critical to evaluating their effect on union wages and contract terms. Deregulation of the trucking industry began in earnest during the fall and winter of 1978, with a series of ICC administrative reforms (see Rose, 1985a). The movement toward administrative "deregulation" intensified over 1979; congressional action in this area culminated in June 1980 with approval of the Motor Carrier Act of 1980. Although a number of industry indicators show the effects of increased competition as early as 1979,

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[5]These include four conference-wide "Iron and Steel and Special Commodity Riders," a national "Automotive Transporters Agreement," and a number of contracts governing tank truck carriage and the household good industry. These agreements were a response to the increasing use of nonunion owner-operators by specialized commodity carriers in the 1950s and 1960s.

the bulk of the impact appears in 1980 and beyond.[6] Based on this chronology of regulatory reform, the NMFAs can be separated into "regulation" and "deregulation" periods. The regulation period covers contracts signed through 1976. The deregulation period includes the 1979, 1982, and 1985 agreements, with one caveat. Because the 1979 NMFA was signed before the outcome of deregulation was clear, it may look more like the regulation-era contracts preceding it than it looks like the post-1980 agreements. For this reason, the 1979-82 contract may be considered "transitional".

Table 1 highlights the provisions of the six NMFAs signed from 1970 to 1985.[7] As the table indicates, the 1982 and 1985 agreements represent dramatic departures from the earlier pattern of contracts. The 1982 contract called for no general wage increase over the life of the contract, provided for cost-of-living (COLA) adjustments under the contract to be diverted to pension and health/welfare (benefits) funds, and provided for no other increases in employer contributions to benefits funds. The only wage change over the contract life was a 47 cents per hour increase specified by the 1979 NMFA. This was part of a 72 cents per hour COLA that the 1979 contract deferred to 1982; the remaining 25 cents per hour was diverted to benefits funds. Employers agreed to provisions

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[6]Moore (1986) 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 1979; however, the larger effects are usually found in 1980 or 1981. It is, however, difficult to disentangle the role of deregulation from general economic conditions in assessing the determinants of the price, revenue, and rate-of-return reductions that Moore's series show in 1979.

[7] The table focuses on changes in hourly wages to facilitate interpretation. The NMFAs also specify increases in mileage rates for approximately 100,000 over-the-road drivers that are paid on a mileage rather than hourly basis.

TABLE 1

National Master Freight Agreements:
Highlights of Contract Provisions, 1970-85

Contract Years	Average Base Contract Wage	General Wage Increase	Cost-of-Living Provisions	Employer Contributions to Pensions, Benefits
1970-73	\$ 4.50 (est.)	\$1.85/hour	\$.01/hr per .3 ΔCPI; .08 max.	\$ 8.00 per week increase
1973-76	\$ 6.32	\$ .95/hour	\$.01/hr per .3 ΔCPI; .06 min., .08 max.	\$16.00 per week increase
1976-79	\$ 7.55	\$1.65/hour	\$.01/hr per .4 ΔCPI 4/77 \$.01/hr per .3 ΔCPI 4/78 uncapped	\$17.00 per week increase
1979-82	\$ 9.60	\$1.50/hour	\$.01/hr per .3 ΔCPI, defer 3d yr. increase	\$30.00 per week increase
1982-85	\$12.80	none	\$.01/hr per .3 ΔCPI, diverted to benefits. deferred COLA: \$.47 to wages, \$.25 to benefits.	No general increase
1985-88	\$13.26	\$1.50/hour; pay decreases for part-time, new employees	\$.31 of each year's .50 increase considered COLA	\$ .30 per hour increase

Sources: Bureau of National Affairs, Collective Bargaining Negotiations and Contracts, "Wage Patterns: Trucking," 1984. Bureau of National Affairs, Labor Report, April 1985.

protecting union jobs from being transferred to nonunion workers, and Teamsters agreed to relaxation of some work rules.[8]

The 1985 contract introduced a number of additional concessions. Although it provided for contract wage increases for current full-time employees ($1.50/hour over the life of the contract, of which 93 cents was considered to be a cost-of-living adjustment), the agreement sharply reduced wages for part-time employees and new hires. The base rate for casual and part-time employees was cut 17 percent, to $11.00 per hour. The wage structure for new hires provided for lower wages during the first three years of employment (at 70%, 80%, and 90% of the full-time rate during the first, second, and third years, respectively). Finally, the NMFA contract terms *overstate* the union's compensation package in the industry. National bargaining was substantially weakened in the 1980s, as increasing numbers of employers sought--and obtained--individual concessions from their local or regional union. A substantial number of these smaller companies refused to sign the 1982 or 1985 NMFA.[9]

Negotiators for both sides attributed these radical departures from historical contract terms to the effects of deregulation on the industry. Deregulation brought with it increased entry and price competition. This led to a surge in the number of nonunion companies, and coupled with the 1981 recession, resulted in deteriorating financial conditions for large numbers of unionized carriers and unemployment for as many as 20 to 30 percent of the Teamsters' freight division

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[8] The most significant work rule concession was to permit over-the-road drivers to make local deliveries. This was part of various supplemental agreements, and was not included in the NMFA itself.

[9] Lieb (1984) reports that the number of companies participating in the 1982 agreement had fallen to 284, from almost 500 in the early 1970s, and that the number of employees covered by the NMFA had declined by 30 percent.

members.[10] These conditions induced the union and trucking firms to agree in November 1981 to reopen the 1979 contract, even though it had five months left to run. The resulting contract was approved and in place by March 1982. Continued deterioration of industry conditions led trucking management and union leadership to agree on further concessions in the summer of 1983. This agreement would have reduced wages for part-time employees and laid-off employees recalled to work. Although the union membership rejected the proposal in fall 1983, quite similar provisions eventually were embodied in the 1985 contract.[11]

The 1982 and 1985 National Master Freight Agreements suggest that deregulation disrupted the historical pattern of union wage-setting in the trucking industry, and led to substantial concessions in union wage levels. Unfortunately, this evidence is not decisive. Although recent union contracts appear to have been strongly concessionary, this conclusion depends on what settlements in the counterfactual state would have been. Some sort of benchmark against which to measure NMFA wages is necessary. For example, if economy-wide wages exhibited the same pattern as NMFA wages, one would not wish to conclude that deregulation accounted for the wage reductions in trucking.

One possible benchmark would be to measure wage movements in the trucking industry relative to wage movements in comparable, but non-regulated industries. I examine wages for three reference groups: manufacturing, mining,

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[10] These figures are based on Teamster estimates, which may be subject to substantial error. See the U.S. General Accounting Office's (1982) report (hereafter, GAO (1982)).

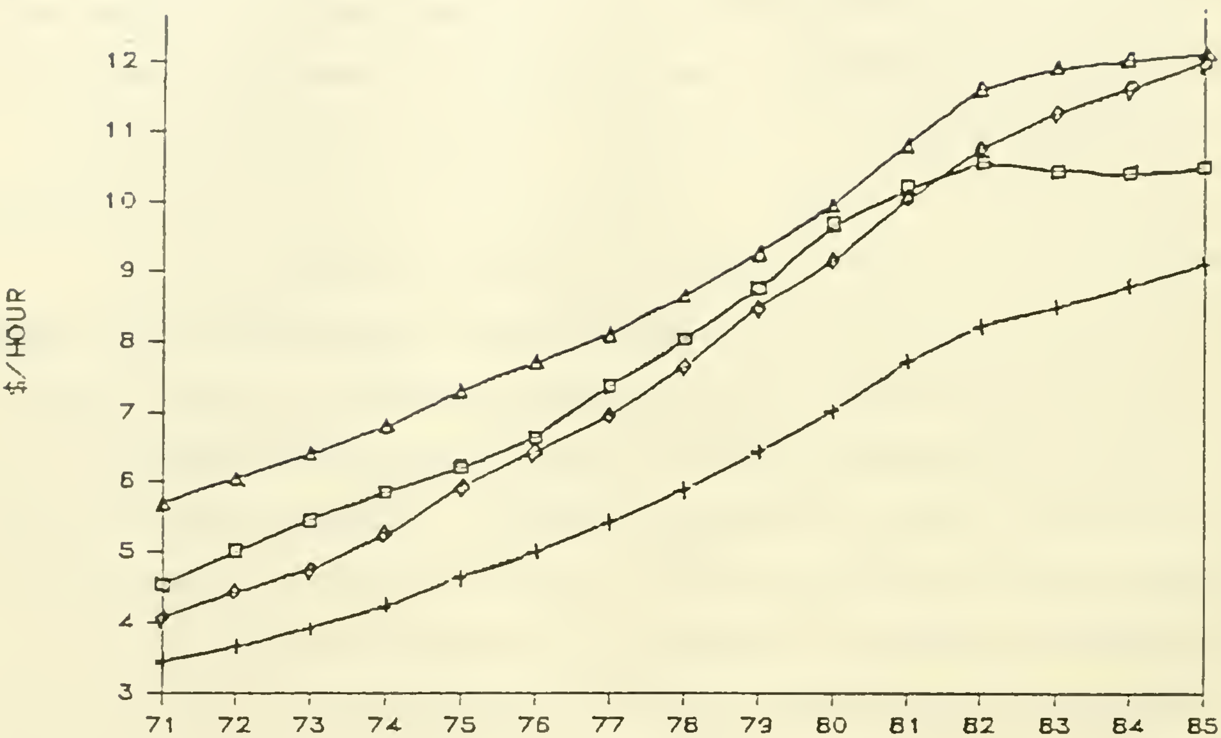
[11] As noted above, the 1985 contract lowered wages for part-time employees and new hires, but did not reduce wages for recalled workers. The 1985 contract was challenged on the grounds that part-time workers were excluded from voting, but was settled and in force by June 1985.

and construction. Aggregate data on average hourly earnings for these groups and the trucking industry generally support the conclusions suggested by the union contract terms. Figure 1 plots the movement of average hourly earnings for the four groups from 1971 through 1984. Wages in the trucking industry move roughly together with mining and construction earnings until 1980. After 1980, trucking wages decline, both relative to mining and construction wages, and absolutely in 1983. Of course, these findings do not necessarily imply that wages in the trucking industry have declined. If union workers earn more than nonunion workers, a decline in the proportion of unionized workers could reduce average earnings even if union wage levels were unaffected.

An alternative approach is to measure union wages against nonunion wages in the trucking industry. This comparison is the focus of the next section. It has two main advantages. First, this reference group is quite similar to the union truck drivers I am most interested in analyzing. Second, looking at union wage behavior relative to nonunion wages in the industry proves a straightforward way to quantify the effects of deregulation. One caution is urged, however. If deregulation reduced nonunion wages, then this reference point may understate the ability of the Teamsters Union to capture regulatory rents, and consequently understate the effect of deregulation on union wages.

FIGURE 1

COMPARISON OF AVERGE HOURLY EARNINGS



□ TRUCKING + MANUFACTURING ◊ MINING △ CONSTRUCTION

Source: U. S. Department of Labor, Bureau of Labor Statistics. Employment and Earnings. Annual average hourly earnings reported for 1971-83. 1984 earnings are from July; 1985 earnings are from June. Manufacturing, mining, and construction data are from September 1985 issue, Table C-1, and August 1985 issue, Table C-1. Trucking data are from Table C-2, various issues.

II. Union Rent-sharing: CPS Evidence

This section uses microdata from the Current Population Survey (CPS) to quantify changes in union wage behavior over time. The CPS provides information on large samples of workers (necessary to obtain adequate subsamples of truck drivers) and is available over a sufficiently long period of time to permit analysis of union premia before and after trucking deregulation. Wage and union status information are available on the May CPS for each year from 1973 through 1984, with the exception of the 1982 survey, for which union status information was not collected. Although the data set has numerous shortcomings (see Freeman (1986) for a discussion), these features make it attractive for the present investigation.

A. Methodology

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 CPS provides information on most of these. Available data on worker qualities include union status, education, age, sex, race, and marital status. 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. I control for industry and occupation effects by restricting the sample to full-time truck drivers employed in the for-hire trucking industry. This excludes non-driver employees in the trucking industry, as well as truck drivers who work in private carriage and self-employed drivers (owner-operators). Data on worker location are used to control for geographical wage variation via regional fixed effects.

Unfortunately, the CPS typically does not provide information on the firms employing the responding individuals.[12] 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.[13] This suggests that one should use caution interpreting the *level* of the estimated premium, although this potential bias should not invalidate tests of *changes* in union premia through time.[14]

The CPS data are used to estimate a conventional semi-log wage equation of the form:

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

[13] 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. I used data from the May 1979 CPS Pension Plan Supplement to test the significance of this effect in the trucking industry. Although the point estimates suggest that larger firms tend to pay both higher nonunion wages and higher union premia over nonunion wages, the small sample size makes statistical inference difficult. Due to the large standard errors of the estimates, statistical tests would not reject the restriction that wages levels and union premia are constant across firm size in the trucking industry.

In addition to firm size, differences in firm markets may affects estimated differentials. 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 (UPS) employees, who negotiate a separate contract and who realized slightly better terms than the NMFA in their 1982 contract, may bias estimates against a decline in the deregulation union premium.

[14] 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 reason to expect the bias to change in such a way as to substantially *reduce* estimated union wage differentials over time. Indeed, the rapid growth of new, smaller nonunion firms in the 1980s, and the maintenance of higher wages by drivers for UPS might exacerbate the potential for overstatement of post-deregulation union wage differentials.

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

where: LHWAGE = natural log of the hourly wage rate. Wages less than $1.00 per hour are assumed to be coding errors and the observations are deleted.

UNION = 1 if a union member, 0 otherwise.

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.

B. CPS Results

Table 2 presents sample mean union and nonunion wages, estimated union wage coefficients from equation (1), and unionization rates for each year. The estimates span four NMFA contract periods: 1973-76, 1976-79, 1979-82, and 1982-85. The full set of estimated coefficients for each equation is reported in

TABLE 2

Estimated Wage Premiums and Sample Characteristics:

Annual CPS Data

Year	Mean Wage Union	Mean Wage Nonunion	Union Coefficient	Proportion Union	Sample Size
1973	$5.65	$3.71	.440 (.050)	.63	231
1974	5.75	4.14	.345 (.052)	.62	258
1975	6.65	4.21	.448 (.056)	.55	246
1976	6.85	4.41	.391 (.050)	.61	236
1977	7.54	4.92	.441 (.040)	.60	304
1978	8.00	5.74	.346 (.046)	.57	291
1979	8.20	6.41	.212 (.055)	.58	170
1980	8.51	6.50	.199 (.089)	.58	91
1981	9.82	7.67	.208 (.089)	.61	83
1983	10.22	7.27	.337 (.083)	.54	85
1984	10.94	7.68	.321 (.081)	.34	89

Standard errors in parentheses.

appendix table 1A.[15] Several aspects of Table 2 deserve mention. First, the estimated union coefficients during the regulation period cluster around .4, implying a union premium of 50 percent above nonunion wages.[16] This finding comports well with Moore's (1978) estimate of a 48 percent premium in 1967 and Hayden's (1977) estimated 50 percent premium in 1973-75.[17] However, these differentials are substantially above the average union differential for cross-sections of industries estimated on similar data sets. For example, Freeman (1986) reports an average union premium for all blue collar workers of .25, or 36 percent, during the 1973-78 period. This suggests that the Teamsters may have been able to capture higher rents for truck drivers than the average union captured for its members.[18]

The second interesting feature of this table is that the union differential drops sharply between 1977 and 1979. From 1979 through 1981, the union coefficient averages about .21--*half* its level in 1973-1978--implying premia of only 23 percent above nonunion wages. The estimated union coefficient increases

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[15] 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.

[16] The 1974 coefficient is a notable exception. The low estimate in this year is driven by a sharp decline in the estimated premium for the outgoing rotation groups in the CPS. The union coefficient for these groups is .260 (.066 standard error) in 1974, as compared to .451 (.068) in 1973.

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

[18] This conclusion holds up against industry-specific union differentials estimated by Freeman and Medoff (1984) 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.

somewhat in 1983 and 1984, to an average of .33, although it remains substantially below the average regulation level. The larger standard errors on the post-1979 estimates suggest that the 1979 through 1984 estimates are likely to be statistically indistinguishable.[19]

Finally, there may be a reduction in the proportion 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 such a substantial amount.

To test the statistical significance of the decrease in union wage differentials, I estimate wage equations pooling data across years.[20] These results are reported in Table 3. Separate intercepts are estimated for each year, but all other coefficients are constrained to be the same within the pooled sample. The

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[19] The higher estimated union premia in 1983 and 1984 could also be real phenomena, linked to our inability to control for job tenure and firm characteristics. The average seniority of unionized workers probably rose during the 1980s, as union unemployment and layoff rates rose. In addition, the average size of nonunion firms declined during the 1980s in response to substantial entry into the industry. Both of these factors could tend to increase estimated union premia in equations that omit controls for tenure and firm size. See notes 13 and 14, *supra*.

[20] The CPS survey design calls for half the respondents in a given year to be reinterviewed the following year. This design creates an overlap in respondents for adjacent years' surveys. Because of this, estimates from contiguous years may not be statistically independent. I therefore have estimated pooled equations omitting contiguous years (eliminating all even year observations) to ensure independence of observations across time. The hypothesis tests for homogeneity are not materially affected by excluding these years. The coefficients' standard errors rise somewhat, but this is expected from the smaller sample size.

TABLE 3

Pooled Wage Equations

Variable	1973-1978	1979-1984	1973-1984
UNION	.402 (.020)	.254 (.033)	---
UNION-REG (1973-78)	---	---	.403 (.019)
UNION-DEREG (1979-84)	---	---	.251 (.031)
EDUC	.021 (.005)	.012 (.009)	.019 (.004)
EXP	.021 (.003)	.021 (.005)	.020 (.003)
EXP^2	-.0004 (.0001)	-.0003 (.0001)	-.0004 (.0001)
SINGLE	.025 (.028)	-.084 (.039)	-.015 (.023)
NONWHITE	-.135 (.033)	-.051 (.048)	-.107 (.027)
NE	-.053 (.025)	-.111 (.045)	-.067 (.022)
SOUTH	-.083 (.023)	-.066 (.041)	-.079 (.020)
WEST	.035 (.026)	.119 (.045)	.058 (.022)
Mean Intercept	1.010	1.593	1.226
NOB	1566	518	2084
R^2	.39	.25	.43
SSR	182.79	62.24	246.89

These equations include time effects for each year. The mean of the estimated time effects is reported as the mean intercept. Standard errors are in parentheses.

first column pools data over the regulatory period, 1973 through 1978. 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.

The constrained union coefficient for the regulatory period is estimated at .402 (standard error, .020). By contrast, the union wage coefficient over the pooled deregulation sample in column 2 is .254 (.033), or roughly 60 percent of the size of the union wage premium measured in column 1. Further, we cannot reject the hypotheses that wages *within* each of these periods are generated by a common process.[21]

The full sample results reported in column 3 reject the restriction of equal union coefficients across the regulation and deregulation periods at the .001 level.[22] The null hypothesis of homogeneous coefficients across the entire sample period, excepting the time intercepts and the union coefficients, cannot be rejected at conventional levels of significance.[23] 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 relative wage advantage.

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[21] The F-statistic to test the restrictions implied by pooling the regulatory years, 1973-1978, is 1.09, which is distributed as F(45,1506) under the null hypothesis (HO) of homogeneous coefficients. The critical value at a 10 percent significance level is 1.24. The F-statistic to test the null hypothesis of homogeneous coefficients for the deregulation period, 1979-1984, is .75. This test statistic is distributed as F(36,468) under HO, with a critical value of 1.29 at the 10 percent significance level.

[22] The t-statistic to test the hypothesis of equal union coefficients over regulation and deregulation periods is 4.67.

[23] The F-statistic to test this hypothesis is 1.038, which is distributed as F(89, 1974) under the null hypothesis of homogeneous coefficients. The critical value at a 10 percent significance level is 1.17. Note, however, that the hypothesis that the coefficients in columns 1 and 2 are the same, excepting the union coefficients, will be rejected at the 5 percent significant level. The F-statistic for this test is 1.95, which is distributed as F(8,2055).

A puzzling difference between these results and the contract evidence is in the timing of the union relative wage decline. Analysis of the Teamsters' contracts in Section I suggested that major union concessions were first made in the 1982 NMFA, while the CPS wage equations show an initial decline in the union premium possibly as early as 1978, with a sharp reduction by 1979. The aggregate wage evidence falls somewhere between the two, with average trucking wages beginning to decline relative to the reference group wages after 1980. While these are not necessarily contradictory (for example, non-contract wage concessions, allegedly made by some Teamster locals for many smaller firms, or an increase in nonunion wages in the late 1970s are consistent with both the NMFA and CPS findings), they do warrant further investigation.

Evaluating Wages For a "Representative" Driver

A first step toward understanding the nature of the decline in the union differential is to look at wages over time for a worker of given characteristics. I use as a "representative" driver a married white male with 12 years of education, 20 years experience, living in the North Central region.[24] Predicted union and nonunion wages for this worker are calculated from a single wage equation pooling data from all sample years. Intercepts and union coefficients are permitted to vary across years; all other coefficients are constrained to be the same over time.[25]

Table 4 reports predicted wages for the representative driver in both current

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[24] These characteristics are close to the 1979 sample means.

[25] The complete set of results for this equation are reported in appendix table 2A. The constant terms reported in the table are estimates of LHWAGE for a nonunion driver with representative characteristics.

TABLE 4

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

Year	Nominal Dollars: Predicted Union Wage	Nominal Dollars: Predicted Nonunion Wage	Constant 1984 Dollars: Predicted Union Wage	Constant 1984 Dollars: Predicted Nonunion Wage	Constant 1984 Dollars: Real Differential
1973	$6.26 (0.19)	$4.06 (0.16)	$14.64	$9.49	$5.15
1974	6.36 (0.19)	4.53 (0.18)	13.39	9.55	3.85
1975	7.09 (0.23)	4.56 (0.17)	13.68	8.81	4.87
1976	7.48 (0.24)	4.88 (0.19)	13.65	8.91	4.74
1977	8.38 (0.24)	5.45 (0.19)	14.37	9.34	5.03
1978	8.88 (0.27)	6.24 (0.22)	14.14	9.94	4.20
1979	9.00 (0.33)	7.23 (0.32)	12.88	10.34	2.54
1980	9.27 (0.46)	7.39 (0.43)	11.69	9.32	2.36
1981	10.62 (0.54)	8.64 (0.55)	12.13	9.86	2.26
1983	10.88 (0.57)	8.03 (0.46)	11.34	8.37	2.97
1984	12.01 (0.78)	8.63 (0.41)	12.01	8.63	3.38

The representative driver is a married, white male with 12 years of education and 20 years of experience, living in the North Central region. Because the wage equation assumes a lognormal error distribution, the predicted wage is calculated as $\exp(X\beta + \sigma^2/2)$, where $X\beta$ is the predicted log wage and σ^2 is the error variance from the pooled equation reported in appendix table 2A. Standard errors, in parentheses, are computed as $\exp(X\beta + \sigma^2/2) \cdot \mathrm{var}(X\beta)^{1/2}$.

dollars (columns 1 and 2) and 1984 constant dollars (columns 3 and 4).[26] 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 2 and 3 is associated with a decline in the growth of nominal union wages from 1977 to 1979 relative to trend and higher growth rates for nonunion wages between these years. This is illustrated in figures 2a and 2b which plot predicted union and nonunion wages in current dollars and 1984 constant dollars, respectively. Predicted union wages are lower in real terms during the 1980s, with real wages in 1983-84 falling 16 percent below the average real wages in 1973-78. In contrast, real nonunion wages remain relatively high through 1981 (compared to their early 1970s levels), and decline only 9 percent in 1983-84 relative to their average over 1973-78.

Second, the reduction in the dollar union wage premium is substantial. The union wage differential declines in *nominal* dollars between 1977 and 1979; in real terms the differential falls by almost one-half relative to its level under regulation. Despite an increase in the real differential in 1983-84, it remains only two-thirds of the 1973-78 level. 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 $2.85 per hour, or $6400 per year. This represents a real decline of nearly $4200 per year in 1984 dollars.

This wage pattern does not appear attributable to the NMFA's failure to compensate adequately for inflation during the contract's life. The initial decline in the union premium comes between 1977 and 1979, despite contract wage

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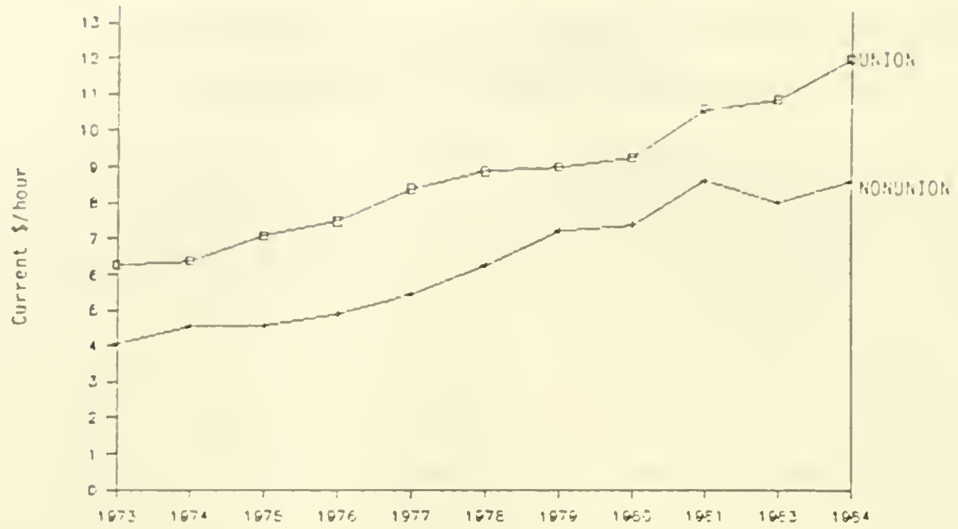
[26] The constant dollar figures were obtained using the average urban Consumer Price Index for each year.

FIGURE 2

PREDICTED HOURLY WAGES FOR UNION AND NONUNION REPRESENTATIVE DRIVER

2A CURRENT DOLLARS



2B: CONSTANT 1984 DOLLARS

increases that basically matched CPI increases over this period.[27] In addition, the decline in the differential persists through the 1979 contract period and into the 1982 contract period. This suggests a more fundamental cause for the relative union decline than unanticipated inflation in the late 1970s.

Comparison of Trucking Wages with Alternative Benchmarks

Disentangling changes in union premia in the trucking industry from general economy-wide movements or idiosyncracies of the CPS data set is more problematic. To evaluate these possiblilities, I compare the CPS results for the trucking industry to those for four alternative benchmark groups. Union coefficients for these groups are presented in Table 5. For ease of reference, union coefficients for the trucking industry are reproduced in column 1. Column 2 reports Freeman's (1986) results for a cross-industry sample of all blue collar workers. Columns 3-5 report my results for workers in the construction, motor vehicle manufacturing, and printing industries.[28] These industries were selected because they employed large numbers of workers (necessary to obtain sufficient sample sizes from the CPS tapes), were heavily unionized, and were unregulated. Both the construction and motor vehicle manufacturing industries also underwent

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[27] Scheduled contract wage increases (including COLAs) between April 1977 and April 1979 amounted to roughly 18%; this compares to a 19.7% increase in the CPI between May 1977 and May 1979.

[28] The union coefficients are estimated from year-by-year equations similar to (1), incorporating occupational fixed effects within each industry. Roughly similar results are obtained by looking at particular occupational categories within each of the three benchmark industries.

TABLE 5

Union Coefficients for Benchmark Industry Groups

Year	Trucking	Average Blue Collar	Construction	Motor Vehicle Mfg.	Printing
1973	0.440	0.23	0.420	0.124	0.289
	(.050)	(.01)	(.016)	(.042)	(.036)
1974	0.345	0.22	0.391	0.122	0.207
	(.052)	(.01)	(.016)	(.037)	(.037)
1975	0.448	0.22	0.410	0.067	0.243
	(.056)	(.01)	(.016)	(.036)	(.039)
1976	0.391	0.25	0.416	0.063	0.239
	(.050)	(.01)	(.017)	(.043)	(.037)
1977	0.441	0.28	0.438	0.200	0.263
	(.040)	(.01)	(.015)	(.037)	(.038)
1978	0.346	0.27	0.417	0.201	0.288
	(.046)	(.01)	(.015)	(.039)	(.034)
1979	0.212	0.22	0.319	0.253	0.156
	(.055)	(.01)	(.023)	(.046)	(.048)
1980	0.199	0.21	0.335	0.268	0.280
	(.089)	(.01)	(.026)	(.079)	(.095)
1981	0.208	0.21	0.352	0.292	0.192
	(.089)	(.01)	(.029)	(.082)	(.077)
1983	0.337	0.22	0.291	0.148	0.258
	(.083)	(.01)	(.041)	(.085)	(.096)
1984	0.321	0.25	0.369	0.204	0.219
	(.081)	(.01)	(.034)	(.099)	(.075)

Standard errors in parentheses. Blue collar estimates are from Freeman (1986). Remaining estimates are from annual CPS wage equations for workers within each industry, controlling for occupational fixed effects, sex, and the variables described in equation (1) above.

changes that potentially reduced industry rents.[29] Use of these benchmarks may tend to understate the extent of regulation-induced changes in the trucking industry.

These results are somewhat mixed. Several aspects of the benchmark results are qualitatively similar to those I obtain for the trucking industry. For example, the union coefficient declines between 1978 and 1979 in three of the four benchmark groups; the exception is motor vehicle manufacturing. This decline is reversed in 1980 in the printing industry, but the average union premia for all blue collar and construction workers remain near their 1979 levels for several years, before increasing in 1984. Despite these apparent similarities, however, important distinctions between the benchmark results and those for the trucking industry suggest that movements in trucking union wage premia are not merely reflective of general trends in the economy as a whole or in the CPS data.

Consider first the average blue collar union premium, as estimated by Freeman (1986). This union coefficient rises from .23 in 1973 to a peak of .28 in 1977, then falls back to .22 in 1979, a decline in 10 percent from its average level over 1973-77. In contrast, the differential in trucking hovers around .4 from 1973-77, then drops to *one-half* of its 1973-77 average level by 1979. The decline in the trucking differential brings it into the range of the average blue collar differential--a significant change from its pre-1979 relative position. Moreover, Freeman notes that by 1984 the blue collar union differential had returned to its 1973-78 average level. For the trucking industry, the differential remained 20 percent below its average 1973-78 level. The patterns of union premia for the other benchmark industries are varied. The premia in the

[29] For example, the oil price shock and subsequent demand shift to imported cars adversely affected motor vehicle manufacturers during this period.

construction industry exhibit behavior similar to that of the trucking premia. The 1973-77 and 1983-84 average premia are essentially identical for the two industries, although the average construction premium during 1979-84 is substantially higher than the trucking premium during this time period. This similarity may be a result of the decline in union influence in the construction industry during the 1970s, as discussed by Allen (1985). This pattern is not repeated in the other benchmark industries. The union premium in motor vehicle manufacturing appears to have increased during the late 1970s and early 1980s and declined somewhat in 1983-84, in sharp contrast to the trucking pattern. No pattern in the union premia in the printing industry emerges from the data. The estimated premia vary considerably from year to year, and the post-1978 average level is only slightly below the pre-1978 average level.

Although it is difficult to draw decisive conclusions from these comparisons, it does not appear that the decline in the union wage premium in the trucking industry is simply an artifact of the particular data set used. The pattern observed in the trucking industry indicates a fair amount of stability in the union premium over the regulatory period, followed by a dramatic and lasting reduction in the premium in 1979 and thereafter. This decline is consistent with the contract evidence and aggregate wage data evaluated in section I, although the CPS data suggest that deregulation may have reduced the union's *relative* wage advantage before the effects show up in contract negotiations and aggregate earnings indices.

III. Nonunion Rent-Sharing

The final possibility I investigate is nonunion rent-sharing. The analysis thus far has measured the decline in union wages relative to nonunion wages in the trucking industry. If nonunion workers also shared in the rents available in the regulated trucking industry, this comparison will tend to understate both union rents and total labor rents. An emerging literature in labor economics suggests that nonunion rent-sharing is a distinct possibility.

A number of recent papers have investigated the existence of inter-industry wage differentials and the role of rent-sharing in explaining their persistence. Krueger and Summers (1986a,b) and Dickens and Katz (1986a,b) provide empirical evidence on the substantial dispersion in wages across industries. These studies report high correlations of industry wage differentials across occupations, countries, and union and nonunion sectors. The authors consider a variety of possible explanations for their findings, including efficiency wage models and possible union and nonunion rent-sharing. Their findings are generally consistent with noncompetitive models of wage determination, although they do not provide clear evidence supporting a single theory.

Recent theoretical models also predict nonunion rent-sharing. Dickens (1986) provides a model of union threat effects, in which unorganized workers are able to capture higher wages as the firm tries to deter workers from organizing a union. In Dickens's model, as long as nonunion wages are near union wage levels, workers will have no incentive to bear the cost of organization. Firms avoid the costs associated with having a unionized workforce by paying workers enough to eliminate workers' net benefits from organizing a union. An alternative model, which does not rely on the presence of a unionized sector in the industry, is proposed by Rotemberg and Saloner (1986). They construct a game-theoretic model in which incumbent workers are more valuable than new hires, technology is characterized by putty-clay features (so that the capital-labor ratio is fixed *ex post*), and workers have some bargaining power vis-a-vis firms. In this model, unorganized workers are able to capture a share of potential rents and quasi-rents in their wages. The model predicts that wages will be a function of profitability and the capital-labor ratio.

The studies suggest the desirability of investigating nonunion rent-sharing in the trucking industry. An analysis of nonunion wage responses to deregulation is important not only for understanding the distribution of rents in the trucking industry, but also for evaluating the nonunion rent-sharing hypotheses proposed in these recent labor studies. To evaluate the extent of nonunion rent-sharing, we need a benchmark against which to measure nonunion wages; I use an economy-wide average wage for this purpose. These are estimated from separate union and nonunion equations for all private sector employees reported in the May CPS for each year. The equations are similar to (1), with occupational, regional, and one-digit industry fixed effects.[30] The "all-industry average wage" is computed as the wage for a blue collar worker with the characteristics described for the representative driver in table 4, based on the average of the industry fixed effects. Estimated union wages are included for comparison.

These wages are reported in table 6, along with the predicted wages for union and nonunion trucking drivers from table 4.[31] The results suggest declines

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[30] I am grateful to Henry Farber for making his moment matrices of these data available to me. Because of minor differences in the data construction, the wage equation for the all-industry regressions differs slightly from (1).

[31] The conclusions discussed below are not materially affected if the table 4 trucking wage estimates are replaced by estimates from separate union and nonunion trucking wage equations for each year This robustness is not

TABLE 6

COMPARISON OF TRUCKING WAGES TO AVERAGE ECONOMY WAGES

YEAR	NONUNION ALL-INDUSTRY AVERAGE	NONUNION TRUCKING	UNION ALL-INDUSTRY AVERAGE	UNION TRUCKING
1973	$5.34	$4.06	$5.68	$6.26
1974	5.70	4.53	6.06	6.36
1975	6.17	4.56	6.23	7.09
1976	6.39	4.88	6.89	7.48
1977	6.72	5.45	7.49	8.38
1978	7.04	6.24	8.33	8.88
1979	7.68	7.23	7.93	9.00
1980	8.21	7.39	8.81	9.27
1981	9.50	8.64	9.59	10.62
1983	10.71	8.03	11.65	10.88
1984	10.38	8.63	13.30	12.01

Trucking wage estimates reproduced from Table 4. All-industry average wage estimates are for a married white male with 12 years of education and 20 years of experience, and are estimated from separate union and nonunion wage equations including all private sector employees. Equations control for sex, education, experience, marital status, and race, and include occupational, regional, and one-digit industry fixed effects. All-industry averages are based on average of industry fixed effects, and are evaluated for blue collar workers.

in both union and nonunion trucking wages relative to the all-industry averages during the deregulation period. These patterns are clearly evident in figures 3a and 3b, which plot trucking and all-industry wages for nonunion and union workers, respectively. Until 1979, nonunion trucking wages rise steadily relative to average nonunion wages. After 1979, trucking wages fall increasingly behind average nonunion wages. All-industry average nonunion wages increase 35 percent between 1979 and 1984, as compared to only 19 percent for the trucking industry. The nonunion wage gap rises from an average of $.63 per hour in 1978-79 (8 percent of industry average nonunion wages) to an average of $2.21 per hour in 1983-84 (21 percent of industry average nonunion wages).[32] Similar patterns are evident in the union wage comparisons.[33]

These results should be interpreted with some caution. The sample sizes of nonunion trucking drivers are quite small, averaging only 35 to 45 workers in the 1980s. These suggest quite substantial standard errors on the wage estimates. In addition, the magnitude of the trucking wage decline depends critically upon the period over which changes are measured. Because trucking wages gain relative to average wages during the 1970s, using 1978 or 1979 as the regulation base year and comparing changes through 1983 or 1984 will lead to larger estimates of nonunion rents than would be predicted using earlier years as the regulation base.

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surprising, given the pattern of sample mean wages reported in table 2.

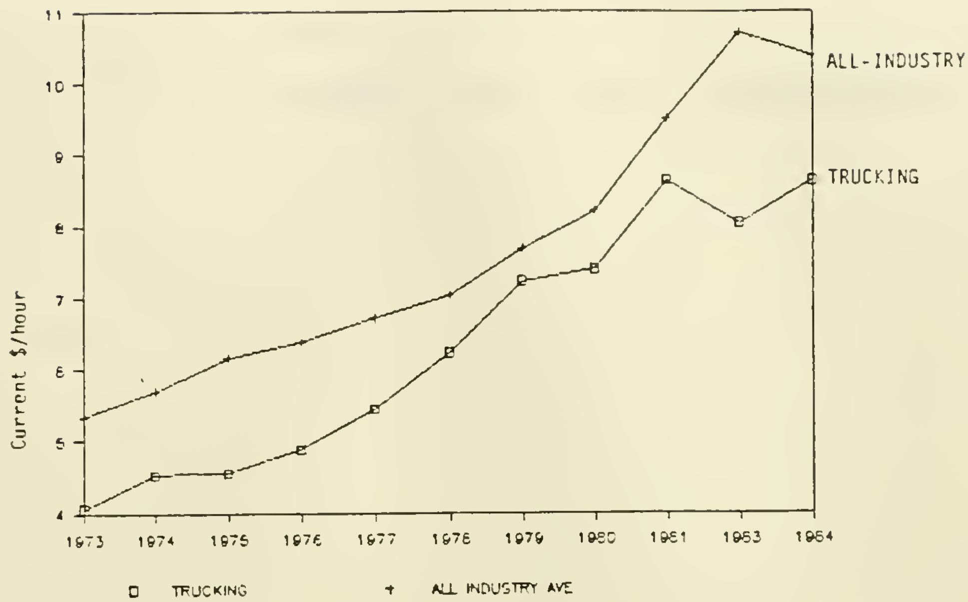
[32] Nonunion trucking wages are roughly the same percentage of all-industry average nonunion wages in 1983-84 that they are in 1973-74. This may somewhat complicate the interpretation of the post-1979 increase.

[33] The rate of increase in union trucking wages is dramatically lower than that for all-industry average union wages after 1979. There is a 67 percent increase in all-industry average union wages between 1979 and 1984, compared to a 33 percent increase in trucking wages over the same period. This disparity is sufficient to reverse the historic superiority of trucking wages; by 1983, union trucking wages have for the first time in the sample fallen below all-industry average union wages.

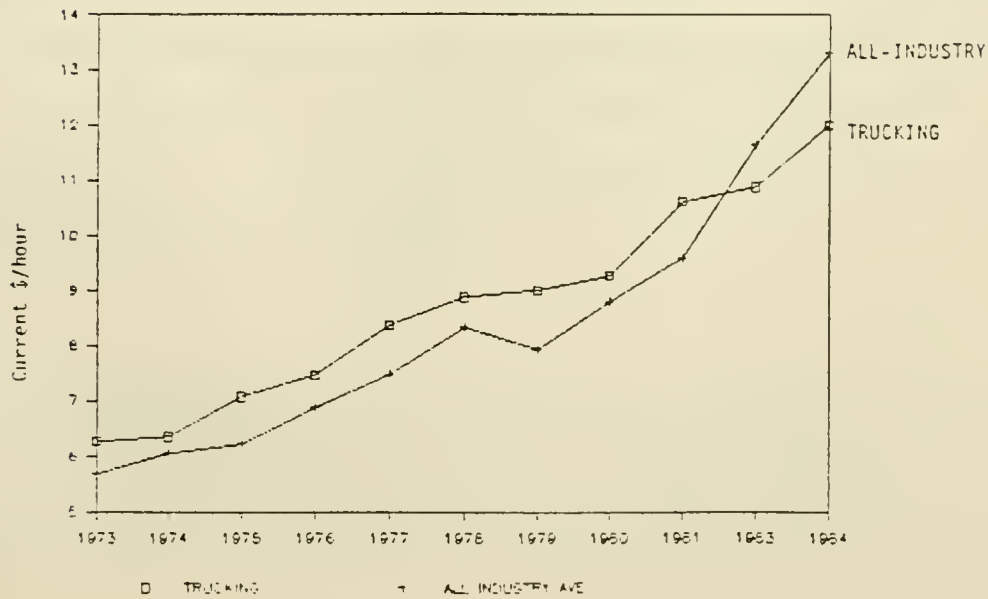
FIGURE 3

COMPARISON OF TRUCKING WAGES WITH ALL-INDUSTRY AVERAGE WAGES

3A: NONUNION WORKERS



3B: UNION WORKERS



With these caveats in mind, however, the findings nevertheless appear to support the hypothesis that nonunion, as well as union, trucking drivers shared in the rents available under motor carrier regulation.

IV. Aggregate Union Losses

The magnitude of the rents implied by these results is substantial. To put the findings into context, I calculate a rough measure of the total losses resulting from deregulation. I restrict this calculation to rents captured by the Teamsters Union, because of the difficulty quantifying nonunion rents. The calculation is summarized in table 7. A number of assumptions are required to compute aggregate union rents. First, I assume that deregulation reduced the 1973-78 union premium of 50 percent to the 1983-84 average premium of 39 percent. The sensitivity of the estimate to this assumption is checked by recalculating losses assuming a decline to the 1979-84 average premium of 30 percent. Second, I assume that nonunion wages are unaffected by deregulation. This will understate both union and nonunion rents to the extent that nonunion wages were bid up under regulation. My calculation attributes the entire change in the union premium to a reduction in rents, but omits any change due to a decline in the overall trucking 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, and that nonunion carriers have captured an increasing share of the market.[34] My computation of rent reductions excludes the union's loss from jobs that are no longer held by union employees.

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[34] The Wall Street Journal (October 9, 1984, p.1) reports that the Teamsters Union estimates 100,000 jobs have been lost since 1980. Lieb (1984) and Hendricks (forthcoming) discuss the effects of increased nonunion competition on union employment. Although it is difficult to disentangle the role of regulatory reforms from that of the 1981-82 recession, the persistence of such high union unemployment rates appears to be considered abnormal by industry observers.

TABLE 7

Aggregate Losses and Division of Rents

1. Aggregate losses for union

A. Assumptions:
- union premia declines from 50% to 39%
  sensitivity check: union premium declines from 50% to 30%
- nonunion wages are unaffected by deregulation
- unionization rate is constant at 60%

B. Result for Class I intercity motor freight carriers in 1983; 886 firms (revenues greater than $5 million)

Aggregate employee compensation:	$12.24 billion
Implied union compensation:	8.27 billion
Implied union loss:	657 million
Implied union loss at 30%:	1.24 billion

2. Division of rents between firms and union: general freight carriers

A. Assumptions:
- above assumptions 1.A. for union rent share
- present discounted value (pdv) of firm rents equal to 8.8% of 1978 aggregate revenue for general freight carriers (from Rose, 1985a)
- 10 percent discount rate, 50 percent corporate income tax rate
- rents are perpetuity

B. Result for Class I intercity general freight common carriers in 1978, 345 firms (revenues greater than $3 million)

Aggregate revenues:	$17.50 billion
Implied pdv of firms' rents:	1.54 billion
Implied annual pre-tax firm rents:	308 million
Aggregate employee compensation:	8.82 billion
Implied union compensation:	6.11 billion
Implied annual union rents at 39%:	448 million
Implied annual union rents at 30%:	814 million

C. Comparison of rent shares, 1984 dollars

Annual pre-tax firm rents:	$490 million
Annual pre-tax union rents:	$713 million - $1.3 billion
Union share:	59 - 73 percent
Total rents/Total revenues:	4.3 - 6.4 percent

The loss to an individual union driver is estimated in the results above at 8 to 15 percent of his current compensation, which is the decline relative to what his compensation would have been had the 50 percent union differential been maintained.[35] The total union loss is estimated by aggregating over all Class I motor carriers, which are regulated firms with more than $5 million in annual gross revenues. Although this group excludes a high proportion of trucking firms (which are quite small), it comprises the very largest companies in the industry, accounting for half or more of total revenues. Applying the assumptions described above to 1983 total employee compensation for Class I motor carriers implies union compensation of $8.27 billion, and an estimated annual reduction of $657 million in rents for unionized employees of these firms.[36] If the average deregulation union premium of 30 percent is used in the calculations, the estimated union loss rises to $1.24 billion.

Finally, the reduction in union rents can be compared to the losses incurred by owners of trucking firms, to estimate the relative shares of rents for unionized labor and capital. Rose (1985a) estimates the deregulation-related loss in market value for a sample of general freight trucking firms at 8.8 percent of 1978 revenues. Applying this estimate to revenues for all 1978 Class I general freight common carriers yields an estimated after-tax loss of $1.54 billion in present discounted value. This corresponds to roughly $308 million in annual pre-tax rents in 1978, or $490 million in 1984 dollars. Applying the calculation

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[35] The loss, Δ, for a 39% premium is given by: $\Delta = (1.39 - 1.50) \cdot$ nonunion wage $= -.11$ nonunion wage. The percent loss is Δ/union earnings $= -.11/1.39 = -.08$.

[36] Total employee compensation of $12.24 billion for the 886 Class I motor carriers reporting to the ICC is taken from the U.S. Interstate Commerce Commission Bureau of Accounts, Transport Statistics in the United States. Motor Carriers. Part 2, 1983.

described above to employee compensation for these firms yields an estimate of $448 to $814 million in pre-tax union rents in 1978, or $713 million to $1.3 billion in 1984 dollars. These calculations, while crude, suggest that the Teamsters Union may have been the dominant beneficiary trucking regulation capturing 60% or more of the total rents in the industry.[37]

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[37] Total rents are measured as union rents plus firm rents. This excludes rents captured by nonunion labor and other factors of production. This finding is consistent with the estimated union share obtained in a Tobin's q type of model of trucking firms; see Rose (1985b). The union share is strikingly similar to Salinger's (1984) estimate of a 77 percent average union rent share in his cross-industry studying employing Tobin's q.

Conclusion

This study examines industry wage responses to motor carrier deregulation. Union contract evidence and aggregate data on industry average earnings suggest substantial changes in historical wage determination patterns after motor carrier deregulation. Microdata estimates of union premia over nonunion wages in the trucking industry indicate declines of 30 to 40 percent in the size of the union differential beginning in 1979. Annual earnings for a "representative" union driver in 1983-1984 were at least $2200, or 8 percent, less than they would have been had the regulatory union wage differential been maintained. These results suggest considerable rent-sharing by union workers. The annual loss in union rents is estimated at $657 million to $1.24 billion for employees of the 886 largest regulated trucking firms in 1983. The results also indicate a decline in nonunion trucking wages relative to economy-wide wage levels. Although the paper does not precisely quantify this effect, the result provides some support for models of nonunion rent-sharing. Further research in this area seems desirable.

Comparison of the results in this paper with estimates of rents accruing to owners of capital indicate that the Teamsters Union may have been the primary beneficiary of regulatory rents in the trucking industry, capturing as much as 60 to 70 percent of total rents. The findings suggest that the assumption of competitive factor prices can be quite misleading, and provide strong support for labor rent-sharing hypotheses.

APPENDIX TABLE 1A

Annual Wage Equations

Variables	1973	1974	1975	1976	1977	1978	1979	1980	1981	1983	1984
Constant	.815	.667	1.058	.970	1.207	1.280	1.668	1.408	1.421	1.756	1.528
	(.190)	(.173)	(.188)	(.161)	(.133)	(.163)	(.204)	(.336)	(.364)	(.366)	(.329)
UNION	.440	.345	.446	.391	.441	.346	.212	.199	.208	.337	.321
	(.050)	(.052)	(.056)	(.050)	(.040)	(.046)	(.055)	(.089)	(.089)	(.083)	(.081)
EDUC	.025	.047	.014	.017	.020	.011	.008	.018	.029	.010	.019
	(.013)	(.013)	(.013)	(.011)	(.009)	(.012)	(.014)	(.023)	(.025)	(.024)	(.025)
EXP	.018	.016	.013	.034	.019	.028	.012	.024	.025	.004	.033
	(.009)	(.008)	(.009)	(.007)	(.006)	(.007)	(.008)	(.013)	(.013)	(.015)	(.011)
EXP^2	-.0004	-.0002	-.0002	-.0006	-.0004	-.0005	-.0002	-.0003	-.0004	-.0001	-.0007
	(.0002)	(.0002)	(.0002)	(.0002)	(.0001)	(.0002)	(.0002)	(.0003)	(.0002)	(.0003)	(.0002)
SINGLE	.020	.08	.012	.035	-.078	.082	-.207	-.031	.042	-.151	.001
	(.085)	(.07)	(.090)	(.074)	(.054)	(.060)	(.072)	(.105)	(.110)	(.106)	(.081)
NONWHITE	-.198	-.13	-.158	-.015	-.060	-.199	-.153	-.023	-.131	.056	.044
	(.089)	(.07)	(.103)	(.097)	(.064)	(.075)	(.081)	(.120)	(.135)	(.135)	(.113)
NE	.090	-.084	-.079	-.116	-.109	-.011	-.146	-.161	-.198	.084	-.172
	(.065)	(.065)	(.072)	(.063)	(.049)	(.062)	(.078)	(.124)	(.117)	(.113)	(.106)
SOUTH	-.075	-.060	-.058	-.176	-.103	-.048	-.098	-.085	-.033	-.002	-.120
	(.056)	(.055)	(.065)	(.057)	(.051)	(.056)	(.067)	(.106)	(.120)	(.114)	(.096)
WEST	.043	.052	.164	.014	-.012	-.025	.131	-.035	.097	.229	.046
	(.072)	(.073)	(.072)	(.065)	(.049)	(.058)	(.080)	(.117)	(.124)	(.112)	(.111)
NOB	231	258	246	236	304	291	170	91	83	85	89
R^2	.38	.30	.35	.41	.42	.30	.26	.19	.22	.27	.33
SSR	25.02	30.87	34.41	24.90	27.19	34.63	17.84	11.49	10.43	9.78	9.29

Standard errors in parentheses.

TABLE 2A

Results for Pooled Sample Wage Equation

Year	Constant	Union Coefficient	Variable	Coefficient
1973	1.341	.434	(EDUC-12)	.019
	(.040)	(.048)		(.004)
1974	1.452	.338	(EXP-20)	.006
	(.039)	(.045)		(.001)
1975	1.458	.440	$(EXP-20)^2$	-.0004
	(.037)	(.045)		(.0001)
1976	1.526	.427	SINGLE	-.017
	(.040)	(.047)		(.023)
1977	1.635	.431	NONWHITE	-.108
	(.035)	(.041)		(.027)
1978	1.771	.353	NE	-.066
	(.035)	(.042)		(.022)
1979	1.918	.219	SOUTH	-.078
	(.044)	(.054)		(.020)
1980	1.941	.226	WEST	.057
	(.059)	(.074)		(.022)
1981	2.096	.207		
	(.064)	(.078)		
1983	2.024	.303		
	(.057)	(.075)		
1984	2.095	.330		
	(.048)	(.078)		

Number of Observations	2084
R^2	.429
SSR	245.97

Standard errors in parentheses.
* Constant terms are estimates LHWAGE for nonunion "representative" drivers in each year (see text for description of characteristics).

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