Labor Rent Sharing and Regulation: Evidence from the Trucking Industry

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Labor is likely to be an important claimant to firms' rents, particularly in a regulated environment. This study analyzes wage responses to trucking deregulation to test labor rent-sharing hypotheses. The results indicate substantial declines in union wages as a consequence of reduced regulatory rents. Union premia over nonunion wages fell from 50 percent to less than 30 percent, implying aggregate annual losses of $950 million to $1.6 billion. Rent spillovers to nonunion drivers and truck drivers outside the regulated trucking industry appear insignificant. The results suggest that union workers captured more than two-thirds of total industry rents and provide strong support for union rent-sharing hypotheses.

I. Introduction

Labor may be an important claimant to firms' economic profits. A number of empirical studies have analyzed the ability of labor unions to capture a share of rents. Clark (1984), Ruback and Zimmerman (1984), and Salinger (1984), for example, used cross-sectional data on firms or lines of business to estimate the effect of unionization on...
profitability. All three found evidence of significant union rent sharing. Recent studies by Dickens (1986), Rotemberg and Saloner (1986), Dickens and Katz (1987a, 1987b), and Krueger and Summers (1987, in press) provide theoretical and empirical 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.

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, regulatory barriers to entry may enhance unions’ power, and the political nature of regulatory agencies can expand the scope of potential games between firms and workers. The extent of union rent sharing and the pattern of collective bargaining in regulated industries have been explored by a number of authors, including Hendricks (1975, 1977) and Ehrenberg (1979). These studies indicate that regulation may have important effects on industry wages; failing to account for these effects may lead to underestimates of regulatory rents and distortions.

Examining wage responses to reductions in rents, such as those caused by regulatory reforms, can provide a powerful test of both union and nonunion rent sharing. My analysis focuses on the truck-

1 Hendricks (1975) analyzed electric utilities’ relative bargaining strengths in different regulatory environments; Hendricks (1977) investigated occupational wage patterns across regulated and unregulated industries; and Ehrenberg (1979) discussed the effect of regulatory constraints and political influences on collective bargaining, focusing on telecommunications in New York. In addition, Hendricks, Feuille, and Szerszen (1980), Cappelli (1985), Card (1986), and Hendricks (in press) analyzed interactions between regulation and collective bargaining in the airline industry; Arnold (1970), Garnel (1972), Annable (1973), Hayden (1977), and Moore (1978) discussed potential Teamster Union rents in their investigations of trucking regulation. Hirsch (1986) examined the pattern of drivers’ wages through time in an analysis similar to the present study. His results are substantively quite similar to the ones reported here, although he used a slightly different wage equation specification. Hirsch’s interpretation of his results does, however, differ somewhat from the conclusions reached below. See Joskow and Rose (in press) for a discussion of the literature on labor rent sharing in regulated industries.

2 Most studies of regulation assume competitive factor markets and treat factor costs as exogenous to the regulatory system. If factor prices include a share of regulatory rents, this assumption will underestimate the effect of regulation on product prices and associated efficiency losses.

3 Hendricks et al. (1980), Cappelli (1985), and Card (1986) explored the impact of airline deregulation on unionized labor in that industry. They found 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.
ing 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) (see, e.g., Moore 1978, 1986; Frew 1981; Rose 1985a). 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 should tend 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.

The remainder of the paper is organized into five sections. Section II outlines trucking regulatory reforms and discusses their predicted effects on wages under labor rent-sharing hypotheses. Section III investigates rent sharing by unionized trucking industry drivers, using information on union contracts and microdata from the Bureau of the Census’s Current Population Surveys (CPS). These data provide strong evidence of union rent sharing. The average union premium over nonunion wages declines from 50 percent during the regulation period to less than 30 percent after deregulation. This corresponds to annual earnings losses of $3,800, or 14 percent, for a “representative” union driver in 1983–85. Possible rent spillovers to other drivers—in particular, nonunion trucking industry drivers and unregulated (private carriage) drivers—are explored in Section IV. Microdata from the CPS are used to compare nonunion trucking industry wages and wages for private carriage drivers (both union and nonunion) to wages for other blue-collar workers outside the trucking industry. The relative wage responses over time provide little evidence of significant rent sharing by nonunion and private carriage drivers. These suggest that rent spillovers were minimal and that unionized trucking industry drivers were the primary labor beneficiary of regulatory rents. Estimates of the aggregate labor rents in the trucking industry are developed in Section V, and a brief conclusion follows in Section VI.

II. The Regulatory Experiment

Trucking regulatory reforms of the late 1970s and early 1980s constitute an almost ideal natural experiment for evaluating rent-sharing hypotheses. From 1935 through the mid-1970s, ICC regulation of the trucking industry included stringent entry controls, restrictions on
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partially regulated and exempt carriers, and collective rate making. This system raised trucking rates above competitive levels, ensuring high economic profits for regulated trucking firms (see Moore 1978, 1986; Frew 1981; Rose 1985a). The transformation of regulation in the late 1970s led to substantial entry of new firms, expansion of existing firms, and enhanced price competition. These reforms created a considerable—and, from the standpoint of existing trucking firms and their workers, largely exogenous—shock to potential industry rents. Wage responses to this reduction in rents are used to measure the extent of labor rent sharing in the trucking industry.

A. The Timing of Regulatory Reforms

Identifying the timing of regulatory reforms is a prerequisite to evaluating their effect on wages. Regulatory reforms of the ICC were concentrated in the late 1978–79 period, although the commission had hinted at the possibility of internal reform as early as 1975 or 1976 (see Rose 1985a; Robyn 1987). These reforms fundamentally altered the regulatory environment and contributed to substantial changes in industry structure and performance. By 1979, entry applications from existing carriers reached record highs, average rates of return fell sharply, and average revenue per ton-mile declined.

Congressional action on trucking regulation lagged ICC initiatives; not until late 1979 were the key congressional actors willing to consider reform legislation, and the Motor Carrier Act of 1980 was not approved until June 1980. The act did not extend deregulation and in some cases rescinded or limited additional ICC reforms. By resolving uncertainty over the legality and permanence of the ICC’s reforms, however, the act did contribute to substantial industry reorganization. Entry by new firms, for example, exploded after the act was signed, leading to significant increases in competition and further reductions in average trucking rates and returns.

4 Rothenberg (1987) argues that motor carrier deregulation was largely independent of interest group lobbying activities. Deregulation followed no major change in the distribution or magnitude of regulatory rents, nor were there changes in the relevant congressional committees that should have reduced the influence of proreregulation interests. Deregulation appears to have been primarily an executive branch phenomenon that developed in spite of intense lobbying campaigns by both the American Trucking Associations (the industry trade association) and the International Brotherhood of Teamsters.

5 See Moore (1986). The return and revenue impacts must be at least in part attributable to the 1979–80 oil price shock and recession, however. In general, it is difficult to disentangle the relative effects of deregulation and economic conditions in these aggregate series.

6 Most important, the act did not deregulate trucking, but only reduced the intrusiveness of regulation. The ICC maintains regulatory jurisdiction over the motor carrier industry.
This chronology suggests that the precise date of "deregulation" is ambiguous, although the effects of regulatory reforms should begin to show up sometime between 1978 and 1981. I adopt the convention of dating deregulation from 1979, while recognizing that the boundaries are somewhat blurry.

B. Predicted Wage Responses to Deregulation

The implications of labor rent sharing for wage responses to deregulation are outlined below. I consider potential rent sharing by three groups of workers: union workers in the regulated trucking industry, nonunion workers in the regulated trucking industry, and truck drivers employed outside the regulated sector.

1. Union Workers in the Trucking Industry

Although economists have long acknowledged the potential role of labor unions as rent-seeking organizations (see, e.g., Lewis's [1959] discussion of "monopoly unionism"), until recently there have been few direct empirical tests. Studies by Clark (1984), Ruback and Zimmerman (1984), and Salinger (1984) found evidence that unions reduce firms' profitability in a manner consistent with rent sharing. This suggests that unions may be important coclaimants to firms' monopoly rents.\(^7\)

Regulation is likely to have augmented the potential for union rent sharing in the trucking industry.\(^8\) Regulatory operating ratio constraints (which require that the average ratio of operating expenses to operating revenues for groups of trucking firms exceed some threshold level, set at 93 percent during the 1970s) may induce excessive use of labor or reduce the resistance of firms to higher labor costs (see Daughety 1984). Hendricks (1975) also found that regulatory rate-making processes can alter firms' incentives for holding down labor costs. Moreover, regulatory barriers to entry and restrictions on

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\(^7\) Labor unions may also seek to create rents by restricting output and raising product prices; see Lewis (1959) for a discussion. Although Arnold (1970) and Annable (1973) argue that the Teamsters cartelized the trucking industry, creating monopoly rents for its members, the presence of substantial firm rents and regulatory mechanisms that encouraged cooperative price setting by firms independent of union wage linkages suggests that rent sharing rather than rent creation best describes the union's role in trucking.

\(^8\) This paper cannot determine whether rent sharing is a general phenomenon or one restricted to regulated industries. As regulation seems to enhance, but not create, union bargaining power, it seems plausible that regulation will affect the degree but not the existence of rent sharing. This conclusion is reinforced by the Clark (1984), Ruback and Zimmerman (1984), and Salinger (1984) cross-industry studies.
independent rate setting enhanced the Teamsters Union’s strength. Entry restrictions virtually eliminated the threat of new nonunion entry and curtailed the potential expansion of existing nonunion carriers into new markets. Uniform rates reduced the ability of nonunion carriers to attract business from union carriers by limiting price competition.

If the union rent-sharing hypothesis is correct, union wages should decline in response to trucking deregulation. To the extent that union workers gained more than nonunion workers did, we would expect union premia over nonunion wages also to decline. Wage responses should be inversely correlated with the magnitude of regulatory rents. This suggests that wages should decline most in the general freight, less-than-truckload (LTL) sector, which earned substantial regulatory rents (Rose 1985a). Specialized commodity common and contract carriers were less protected and less profitable under regulation, implying smaller, but still negative, union wage responses.

2. Nonunion Workers in the Trucking Industry

Nonunion workers in the trucking industry also may share in the rents created by trucking regulation. Lewis (1963) argues that high union wages may spill over to nonunion wages as nonunion employers raise wages to reduce the probability of successful organization of their work forces; Dickens (1986) has developed a formal model of this phenomenon. Potential empirical support is provided by Dickens and Katz (1987a, 1987b) and Krueger and Summers (1987, in press), who found high correlations of industry wage differentials across union and nonunion workers. These results are consistent with rent sharing by both union and nonunion workers, although they do not provide a direct test of this hypothesis.

For union “threat” effects to raise nonunion wages, the threat of organization must be credible, and raising nonunion wages must be a cost-effective way of avoiding higher union wages (Lewis 1963; Rosen 1969). While threat effects may have been important during the 1940s and 1950s, their recent significance in trucking is dubious. The probability of successful organization of nonunion firms during the 1970s and 1980s appears to have been low; indeed, Teamster trucking membership has declined continuously since the mid-1960s.9 This

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9 Teamsters' organizational drives relied heavily on the use of secondary boycotts and "hot cargo" clauses, both of which were declared illegal during the 1950s. The Teamsters have engaged in little subsequent organizational activity in the trucking industry and have met with frequent failure in these few attempts. Although low organizational success could be evidence of successful nonunion use of higher wages, this seems implausible given the magnitude of average union/nonunion wage differentials during the 1960s and 1970s.
suggests that spillovers to nonunion wages via threat effects are likely to have been small over the last 15 years and implies little response of nonunion wages to deregulation.

Nonunion wages could be affected by deregulation even in the absence of rent sharing if the supply of nonunion labor is less than perfectly elastic. In this case, regulation might lower the nonunion wage by reducing the demand for labor (operating through the output effects of high regulated prices).\textsuperscript{10} Deregulation would have ambiguous effects on nonunion wages: higher nonunion labor demand (through output effects of lower prices and increased nonunion price competition) would tend to raise nonunion wages, while further shrinkage of the union sector would tend to raise labor supply to the nonunion sector, lowering nonunion wages.\textsuperscript{11}

It seems unlikely, however, that the labor supply curve for nonunion truck drivers has a substantial positive slope. Truck driving is a low-skill occupation with considerable turnover. In addition, there is a large pool of drivers outside the regulated interstate trucking industry who perform essentially the same job; this group includes owner-operators (self-employed workers who own and operate their own trucks), private carriage drivers (employees of companies shipping their own goods), and delivery drivers.\textsuperscript{12} It seems reasonable to assume a quite elastic supply of labor to each industry segment and probably to the occupation as a whole.

The predicted response of nonunion wages to deregulation is therefore ambiguous. If there is nonunion rent sharing, wages should tend to fall after deregulation. These declines should be smaller in magnitude than union wage responses, however, since nonunion wage premia are derivative of union premia. If the supply of nonunion labor is less than perfectly elastic, wages may increase with deregulation. In either circumstance, however, the wage response is likely to be small.

3. Truck Drivers outside the Regulated Trucking Industry

Labor and product market linkages may also have induced rent spillovers to drivers outside the regulated trucking industry. The two

\textsuperscript{10} Regulatory commodity and route restrictions that reduced the efficiency of truck, and therefore driver, use may have had somewhat offsetting effects on labor demand.

\textsuperscript{11} Aggregate trucking industry employment declined after 1979 and remained below 1979 levels through 1985, primarily as a result of the general economic recession. Shifts in the union/nonunion composition of employment are likely to have increased nonunion employment levels despite the aggregate decline.

\textsuperscript{12} Of the 2 million truck drivers in the United States in 1980, 800,000 were employed in the for-hire trucking industry (Perry 1986, p. 49), and many of these are outside the federally regulated trucking sector.
largest groups in this category are owner-operators and drivers employed by private carriers. Data on owner-operators' wage responses to deregulation are not available, although rent spillovers to this group are likely to have been negligible. For this reason, I focus on private carriage drivers.

Under regulation, shippers could choose between regulated carriage and providing their own transportation. Because these are substitutes, higher regulated prices tend to create quasi rents for private carriage (by raising the cost of its closest substitute). Notice, however, that while regulation raised trucking industry rates above competitive levels, ICC restrictions also raised private carriers' costs. Private carriers could haul only their own goods, could not hire owner-operators, and could not sublease their trucks and drivers to regulated carriers for return trips. These restrictions substantially increased private carriage costs (Moore 1978) and constrained the size of regulation-induced quasi rents.

Private carriage employees—especially truck drivers, who are most directly associated with the rent-producing activity—are potential claimants to the pool of regulatory quasi rents. The earlier analysis suggests that unionized drivers are likely to have captured a larger share of potential rents than did nonunion drivers. Moreover, the level of bargaining power for private carriage drivers appears to have been much lower than that of regulated carrier drivers. This is due in part to frequent National Labor Relations Board (NLRB) refusals to recognize a firm's truck drivers as a separate collective bargaining unit. These made it more difficult to organize private carriage truck drivers and eliminated the possibility of uniform national wage agreements, reducing the Teamsters' influence. This suggests that rent spillovers to union private carriage drivers were smaller than those to regulated drivers. For nonunion drivers, the analysis of nonunion trucking industry drivers can be applied. This implies that wage responses of private carriage truck drivers to deregulation are likely to be small in magnitude but should be negative if rent spillover models are correct.

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13 This results from a variety of factors, including regulatory restrictions that prohibited owner-operators from substituting for regulated carriers, legal restrictions on owner-operators' ability to bargain collectively, and the ease of entry and exit into the owner-operator sector.

14 This was reinforced by the fact that much private carriage is truckload traffic since regulated truckload rates tended to be closer to competitive levels than were regulated LTL rates.

15 The Teamsters Union had a particularly strong incentive to organize private carriage drivers and raise their wages, in an effort to impede their substitution for regulated Teamster drivers. The NLRB rulings tended to preclude this. As a result, both unionization rates and Teamster coverage of unionized drivers were much lower for private carriage drivers than for regulated trucking drivers (Perry 1986).
III. Union Rent Sharing

This section reviews evidence on union rent sharing within the trucking industry. Subsection A analyzes union collective bargaining agreements over the 1973-85 period. Subsection B uses microdata from the CPS to estimate the wage differential between union and non-union workers in the trucking industry.

A. Contract Evidence

Union contracts provide a natural starting point for evaluating the effect of regulatory reforms on labor rent sharing. The Teamsters Union has negotiated national contracts in the trucking industry since 1964. These National Master Freight Agreements (NMFAs) specify wage increases and benefit changes, as well as a broad range of work rules, over each 3-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. The NMFA is the most extensive of these various agreements: about 75 percent of the Teamsters' 375,000-400,000 freight division members in the mid-1970s were covered by the NMFA.

On the basis of the chronology of regulatory reform in Section II, the NMFAs can be separated into "regulation" and "deregulation" periods. The regulation period covers contracts signed through 1976; the deregulation period clearly includes the 1982 and 1985 agreements. The 1979 contract is more difficult to place. The 1979 NMFA was signed in April 1979, before the outcome of deregulation was clear, so its terms may resemble those of the regulatory period. The ICC's reform program and the political debate over deregulation may, however, have induced the Teamsters to moderate their wage demands in an effort to win political support for their proregulation position (see Rothenberg 1987, p. 324). This suggests that the 1979 NMFA may resemble the post-1980 agreements. Given these competing influences, the 1979-82 contract may be considered "transitional."

16 These include four conferencewide 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 (Levinson 1980).

17 The administration had promised "more modest" deregulation legislation if the Teamsters accepted a wage settlement within President Carter's inflation guidelines (Robyn 1987, pp. 35-36; Rothenberg 1987). The magnitude of the 1979 settlement and the fact that it occurred after only a 2-week strike/lockout suggest that the union was less than completely conciliatory.
Table 1 highlights the wage and benefits provisions of the six NMFAs signed from 1970 to 1985. 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, allowed cost-of-living adjustments (COLAs) under the contract to be diverted to pension and health/welfare (benefit) funds, and provided for no other increases in employer contributions to benefit funds. The only wage change over the contract life was part of a 72 cents per hour COLA that the 1979 agreement had deferred to 1982; the remaining 25 cents per hour was diverted to benefit funds. Employers agreed to provisions protecting union jobs from being transferred to nonunion workers, and Teamsters agreed to relaxation of some work rules.

Negotiators for both sides attributed these radical departures from historical contract terms to the effects of deregulation on the industry. The postderegulation surge in nonunion entry and price competition, coupled with the 1981 recession, resulted in deteriorating financial conditions for large numbers of unionized carriers and unemployment for as many as 20–30 percent of the Teamsters’ freight division members. These conditions induced the union and trucking firms to agree in November 1981 to open contract negotiations 5 months early. 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, rejected by the union membership, would have reduced wages for part-time employees and laid-off employees recalled to work. Similar provisions eventually were embodied in the 1985 contract.

Although the 1985 NMFA provided moderate wage increases for current full-time employees ($1.50 per hour over the life of the contract, of which 93 cents was considered to be a COLA), it introduced substantial new concessions. The base rate for casual and part-time employees was cut 17 percent, to $11 per hour. The wage structure

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18 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 who are paid on a mileage rather than hourly basis.

19 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.

20 These figures are based on Teamster estimates that may be subject to considerable error. See the U.S. General Accounting Office’s (1982) report.

21 Although the 1985 contract lowered wages for part-time employees and new hires, it did not reduce wages for recalled workers. The 1985 contract was challenged on the grounds that part-time workers were excluded from voting, but it was settled and in force by June 1985.
| 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/hr.             | 1¢/hr. per .3 CPI; 8¢ max./yr. | $8/wk. increase                          |
| 1973–76        | 6.32                           | $0.95/hr.             | 1¢/hr. per .3 CPI; 6¢ min., 8¢ max./yr. | $16/wk. increase                         |
| 1976–79        | 7.55                           | $1.65/hr.             | 1¢/hr. per .4 CPI (4/77) | $17/wk. increase                         |
| 1979–82        | 9.60                           | $1.50/hr.             | 1¢/hr. per .3 CPI, defer 3d yr. increase | $30/wk. increase                         |
| 1982–85        | 12.80                          | None                  | 1¢/hr. per .3 CPI, diverted to benefits; deferred COLA: 47¢ to wages, 25¢ to benefits | No general increase                       |
| 1985–88        | 13.26                          | $1.50/hr.; pay decreases for part-time, new employees | $1¢ of each year’s 50¢ increase considered COLA | $0.30/hr. increase                       |

Source.—Bureau of National Affairs 1984, 1985.
for new hires provided for lower wages during the first 3 years of employment (at 70 percent, 80 percent, and 90 percent of the full-time rate during the first, second, and third years, respectively). Moreover, the 1982 and 1985 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 union. A large number of smaller companies refused to sign the 1982 or 1985 NMFA. Substantial unemployment among union members further reduced contract coverage.

The 1982 and 1985 NMFAs suggest that deregulation disrupted the historical pattern of union wage setting in the trucking industry and led to substantial wage concessions. Unfortunately, this evidence is not decisive; it depends on what settlements in the counterfactual state would have been. If, for example, economywide 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 solution is to measure wage movements in the trucking industry relative to wage movements in comparable unregulated industries. Figure 1 compares 1971–85 trucking wages to those for three reference groups: manufacturing, mining, and construction. These aggregate data on average hourly earnings generally support the conclusions suggested by the union contract terms. Earnings in the trucking industry moved roughly in tandem with mining and construction earnings until 1980. After 1980, average trucking wages declined, both relatively, with respect to mining and construction wages, and absolutely, in 1983. Of course, these findings do not necessarily imply that actual 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.

B. Microdata Estimates

An alternative approach, which accounts for shifts in the union-nonunion mix, compares union wages to wages for nonunion workers in the trucking industry. This controls for occupational and industry-specific effects and provides a straightforward quantification of the effects of deregulation on union workers. The analysis below uses data on individual workers’ wages, obtained from the CPS, to mea-

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22 Lieb (1984) reported 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. Perry (1986) estimated that roughly 150 employers participated in the 1985 NMFA wage agreements.
sure changes in union wage behavior over time. Although the CPS has numerous shortcomings (see Freeman 1986; Lewis 1986), its large samples of workers (necessary to provide adequate subsamples of truck drivers) and its availability over a long time period make it attractive for the present investigation. One caution is urged, however. If deregulation reduced nonunion wages, then this benchmark may understate the ability of the Teamsters Union to capture regulatory rents and consequently understate the effect of deregulation on union wages. This concern is addressed in Section IV.

1. Methodology and Data

Wages would ideally be modeled as a function of an employee’s characteristics and union status, the firm’s characteristics, occupation- and industry-specific effects, and geographic wage levels. This suggests a conventional human capital semilog wage equation of the form

$$\ln(\text{wage}_{ijk}) = \beta_0 + \beta_1 \cdot U_{ijk} + \mathbf{X}_{ijk} \mathbf{\beta} + \mathbf{Z}_{ijk} \mathbf{\Gamma} + \mathbf{R}_{ijk} \mathbf{\Phi} + \delta_j + \eta_k + \epsilon_{ijk},$$

(1)

where $U_{ijk}$ is the union status of individual $i$ in occupation $j$ and industry $k$, $\mathbf{X}_{ijk}$ is a vector of worker characteristics, $\mathbf{Z}$ is a vector of firm characteristics for worker $ijk$, $\mathbf{R}_{ijk}$ is a vector of regional dummy variables, $\delta_j$ is an occupational fixed effect, $\eta_k$ is an industry fixed effect, and $\epsilon_{ijk}$ is the error term for individual $ijk$. 
The May CPS contains information on most of these factors for the years 1973–85, except for 1982, in which union status information was not collected. Data on workers’ characteristics include union status, education, age, sex, race, and marital status. Regional fixed effects are used to control for geographical wage variation. 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 nondriver employees in the trucking industry, as well as truck drivers who work in private carriage and self-employed drivers (owner-operators). Private carriage drivers are analyzed in Section IV; earnings information unfortunately is not collected for self-employed workers.

Three CPS shortcomings are particularly troublesome for this study. First, the CPS typically does not provide information on the firms employing respondents. 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, then because nonunion drivers in the trucking industry are more likely to work for smaller firms, omitting firm size may cause the cross-sectional estimates to overstate the union wage premium. In addition, differences in firms’ markets may affect estimated premia. For example, regulatory protectionism generally is thought to have conferred the largest benefits on firms in the LTL sector; those in the specialized commodity, truckload sector operated in a more competitive environment, even under regulation (see Rose [1985a] and the references cited therein). The inability to distinguish workers employed in the LTL sector from those employed in the less profitable and less unionized specialized commodity sector may distort the results.

Second, union status should be modeled as an endogenous variable. If unions raise wages relative to nonunion wages, one would expect workers to queue for union jobs. This might allow firms to ration union jobs by selecting higher-quality workers. Given imperfect con-
tROLS FOR WORKER QUALITY IN WAGE EQUATIONS, THIS SELECTION PROCESS MAY LEAD ORDINARY LEAST SQUARES (OLS) ESTIMATES TO OVERSTATE THE LEVEL OF THE UNION PREMIUM. WHILE THIS PROBLEM CAN IN PRINCIPLE BE TREATED, THE PRACTICAL PROSPECTS FOR DOING SO ARE BLEAK. LEWIS (1986) SURVEYED THE LITERATURE ON SELECTIVITY BIASES IN WAGE EQUATIONS AND CONCLUDED THAT THE RESULTS ARE TREMENDOUSLY SENSITIVE TO MINOR VARIATIONS IN SPECIFICATION, ECONOMETRIC TECHNIQUE, ERROR DISTRIBUTION ASSUMPTIONS, AND SAMPLE COVERAGE. THE MEAN OLS BIAS ESTIMATED BY THESE STUDIES IS RELATIVELY SMALL AND NOT ALWAYS POSITIVE (E.G., −0.07 FOR A SET OF 58 ESTIMATES ON NARROWLY DEFINED OCCUPATION-INDUSTRY GROUPS), THOUGH THE STANDARD DEVIATION IS QUITE LARGE. THESE FINDINGS, COMBINED WITH THE CPS'S Dearth OF INFORMATION ON WHICH TO BASE A UNION SELECTION EQUATION, LED ME TO ESTIMATE A SINGLE WAGE EQUATION TREATING UNION STATUS AS UNCORRELATED WITH THE WAGE RESIDUALS, $\epsilon_{ijk}$. THESE TWO SHORTCOMINGS SUGGEST CAUTION IN INTERPRETING THE LEVELS OF THE ESTIMATED UNION PREMIA. They are likely to bias the results against finding rent sharing, however. In particular, as the trucking industry union sector shrank during the early 1980s, its concentration of experienced, high-quality workers was likely to have increased. This suggests a tendency for estimated premia to increase through time, contrary to the prediction of the rent-sharing model.

The third problem with CPS data is their exclusive focus on wage compensation, to the exclusion of employee benefits. The analysis ideally would compare the total compensation package for union and nonunion workers. While a 1976 Council on Wage and Price Stability report indicated that benefit costs were substantially higher in the union than in the nonunion sector (see perry 1986, p. 56), time-series data on nonunion benefits do not appear to exist. We therefore are unable to infer how our results would change if we could substitute total compensation for wage compensation.

With these caveats in mind, the CPS data were used to estimate a semilog wage equation of the form

26 Freeman and Medoff (1984), Hirsch and Addison (1986), and Lewis (1986) 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, the analysis in this paper would follow as long as the bias did not systematically reduce estimated union wage differentials over time. The rapid growth of new, smaller nonunion firms in the 1980s and the mainten- nance of higher wages by United Parcel Service drivers suggest that the opposite tendency may be most plausible.

27 Sherwin Rosen has pointed out that more able workers may have better alternative opportunities, leading them to exit the union sector more quickly after deregulation and reversing the bias argument. If union trucking wages remain above opportunity wages elsewhere (even though they decline in relative terms), this effect may not be substantial. Its significance is an empirical question that cannot be answered with CPS data.
LHWAGE = \beta_0 + \beta_1 \cdot UNION + \beta_2 \cdot EDUC + \beta_3 \cdot EXP \\
+ \beta_4 \cdot EXP^2 + \beta_5 \cdot NONWHITE + \beta_6 \cdot SINGLE \\
+ \beta_7 \cdot NORTHEAST + \beta_8 \cdot SOUTH + \beta_9 \cdot WEST,
\tag{2}

where LHWAGE = natural log of the hourly wage rate, in nominal dollars;\textsuperscript{28} UNION = 1 if a union member, 0 otherwise; EDUC = number of years of schooling completed minus 12; EXP = years of experience, proxied by (age - number of years of schooling - 6) - 20; NONWHITE = 1 if race is nonwhite, 0 otherwise; SINGLE = 1 if marital status is other than "married with spouse present," 0 if status is "married, spouse present"; and NORTHEAST, SOUTH, and WEST are regional dummy variables.

Education and experience variables were normalized so that the intercept, \beta_0, measures the predicted wage for a married white male with a high school education and 20 years of work experience, living in the North Central region. A worker with these characteristics will be referred to as a "representative driver." The results presented below are robust to variations in the specification of equation (2) such as measuring union status by contract coverage rather than membership, including interaction terms between education and experience, controlling for usual weekly hours, and estimating separate union and nonunion wage equations.

2. Basic CPS Results

Table 2 reports sample characteristics and estimated union premia for each year of CPS data, 1973–85 (except for 1982, as noted earlier). The table presents mean wages for union and nonunion drivers, sample unionization (coverage) rates, estimated union wage coefficients (\beta_1 from eq. [2]), and sample sizes for each year. (The sharp decline in sample sizes after 1978 results from changes in the CPS survey design and does not reflect employment attrition in the trucking industry.) The data span four NMFA contract periods: 1973–76, 1976–79, 1979–82, and 1982–85. Several aspects of the data are noteworthy.

First, the CPS samples track the Bureau of Labor Statistics (BLS) average trucking wage reported in figure 1 reasonably well. The average trucking wage calculated from CPS data lies below the BLS aver-

\textsuperscript{28} Hourly wages were used where reported. For workers who were not paid by the hour, the hourly wage was imputed as (weekly earnings/weekly hours). Wages less than $1.00 per hour were assumed to be coding errors, and the observations were deleted. This treatment did not materially affect the results.
Table 2
Sample Characteristics and Estimated Wage Premia (Annual CPS Data)

| Year | Mean Wage ($) | Union Coefficient | Proportion Union | Sample Size |
|------|---------------|-------------------|------------------|-------------|
|      | Union (1)     | Nonunion (2)      | (3)              | (4)         | (5)         |
| 1973 | 5.66          | 3.71              | .440 (.050)      | .63         | 230         |
| 1974 | 5.75          | 4.14              | .345 (.052)      | .62         | 258         |
| 1975 | 6.65          | 4.21              | .448 (.056)      | .55         | 245         |
| 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          | 6.37              | .213 (.056)      | .58         | 169         |
| 1980 | 8.51          | 6.50              | .199 (.089)      | .58         | 91          |
| 1981 | 9.82          | 7.67              | .208 (.089)      | .61         | 83          |
| 1983 | 9.99          | 7.42              | .292 (.083)      | .53         | 79          |
| 1984 | 10.97         | 7.78              | .346 (.091)      | .31         | 78          |
| 1985 | 10.05         | 8.36              | .177 (.092)      | .28         | 108         |

Note.—Union coefficients estimated from semilog wage equations following (2). Standard errors are in parentheses.

age wage by 10–15 percent, but the movements of the two series through time are quite close.29 Given this similarity, the disaggregated CPS data on union and nonunion wage movements and union coverage rates can be used to analyze the wage patterns noted in the discussion of figure 1. Table 2 suggests that shifts in union-nonunion composition are not the primary cause of the decline in BLS average aggregate trucking wages. Instead, a major factor in the relative trucking wage decline appears to be the slow rate of growth of union wages relative to nonunion wages during the late 1970s and early 1980s and the low growth of both union and nonunion wages during the mid-1980s. This conclusion is supported by the relatively minor changes in unionization rates through the early 1980s. Not until 1984

29 The CPS wage increase between 1979 and 1980 is much smaller, and the CPS increase between 1980 and 1981 much larger, than the corresponding changes in the BLS wages. The 1979–81 movements between the two data sets are comparable, however.
does the union coverage rate substantially decay, although the extent of the decline by the end of the period is enormous. In 1984 and 1985, the union coverage rate is half its former level, at roughly 30 percent of the sample.

This apparent decline in union coverage is the second striking feature of table 2. The finding supports the Teamsters’ claim that the union has been decimated by deregulation as unionized firms have exited and nonunion entrants (including nonunion subsidiaries of unionized companies) have captured increasing traffic shares. The small sample sizes by the end of the period and the overlap between the 1984 and 1985 samples (both of which are a result of CPS survey design) limit our confidence in any particular CPS point estimate of union coverage, however. Nor are there other reliable time-series estimates of industry union coverage with which to compare the CPS results. Some independent confirmation is provided by Perry’s (1986, p. 110) report that the number of workers covered by the NMFA declined by 30–40 percent between 1979 and 1985, although NMFA coverage and union coverage are not identical concepts. While this provides some support for the CPS findings, further evidence seems desirable to ascertain the precise extent of the decline in union representation.

Finally, the pattern of estimated union premia provides strong support for union rent-sharing hypotheses. The union coefficients reported in column 3 of table 2 were estimated from annual wage equations following equation (2); the full set of results for each equation is reported in Appendix table A1. The estimated union coefficients during the regulation period, 1973–78, cluster around 0.40, implying a union premium of 50 percent above nonunion wages. 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 1967 and Hay...
cent premium in 1973–75. The results, particularly when combined with Moore’s 1967 estimate, suggest a great deal of stability in the union premium over the regulation period. Further, these differentials are substantially above the average industry union differential estimated from similar data sets. For example, Freeman (1986) reported an average union premium for all blue-collar workers of 0.25, or 28 percent, during the 1973–78 period. Freeman and Medoff (1984, p. 50) estimated industry-specific union differentials from 1973 CPS data. They reported that half of the 62 industries in their sample had union premia greater than 15 percent but that only eight of the 62 had union premia greater than 35 percent. This suggests that the Teamsters may have been able to capture higher rents per member than most unions did.

The sharp decline of the union premium after deregulation provides additional support for the rent-sharing hypothesis. From 1979 through 1985, the union coefficient averages about 0.24—slightly more than half its level in 1973–78—implying premia of 27 percent above nonunion wages. The average deregulation level of the premium is almost identical to the average union premium for all blue-collar workers reported by Freeman (1986). Although the estimated union coefficient is somewhat higher in 1983 and 1984 than in the other deregulation years, the large standard errors on the post-1979 estimates suggest that the 1979–85 estimates are statistically indistinguishable.

The statistical significance of the decrease in the union wage differential was tested by estimating wage equations pooling data across years.34 These results are reported in table 3. Separate intercepts were estimated for each year, but all other coefficients were constrained to be the same within the pooled sample. The first column pools data over the regulatory period, 1973–78. Column 2 pools data over the deregulatory period, 1979–85. These two samples are combined in column 3, which reports results for the full sample, 1973–85, allowing the union coefficient to differ across regulatory regimes.

The constrained union coefficient for the regulatory period is estimated at 0.402 (standard error, 0.020). By contrast, the union wage coefficient over the pooled deregulation sample in column 2 is 0.237 (0.033), or roughly 60 percent of the size of the union wage premium measured in column 1. Further, we cannot reject the hypothesis that

34 Because of the overlap in respondents discussed in n. 30, estimates from contiguous years may not be statistically independent. Hypothesis tests are not materially affected by excluding contiguous years (eliminating all even-year observations) from the pooled equations. The remainder of the paper treats observations from different years as independent.
TABLE 3
WAGE EQUATIONS ESTIMATED FROM POOLED CPS DATA

| Variable           | 1973–78 (1) | 1979–85 (2) | 1973–85 (3) |
|--------------------|-------------|-------------|-------------|
| UNION              | .402        | .237        | ...         |
|                    | (.020)      | (.033)      |             |
| UNION-REG (1973–78)| ...         | ...         | .404        |
|                    |             |             | (.019)      |
| UNION-DEREG (1979–85)| ...     | ...         | .235        |
|                    |             |             | (.030)      |
| EDUC               | .021        | .008        | .018        |
|                    | (.005)      | (.008)      | (.004)      |
| EXP                | .006        | .006        | .006        |
|                    | (.001)      | (.002)      | (.001)      |
| EXP²               | -.0004      | -.0003      | -.0003      |
|                    | (.0001)     | (.0001)     | (.0001)     |
| SINGLE             | .025        | -.094       | -.023       |
|                    | (.028)      | (.037)      | (.022)      |
| NONWHITE           | -.135       | -.089       | -.117       |
|                    | (.033)      | (.047)      | (.027)      |
| NORTHEAST          | -.053       | -.085       | -.061       |
|                    | (.025)      | (.043)      | (.022)      |
| SOUTH              | -.083       | -.030       | -.068       |
|                    | (.023)      | (.038)      | (.020)      |
| WEST               | .035        | .121        | .060        |
|                    | (.026)      | (.043)      | (.022)      |
| Mean intercept     | 1.536       | 2.018       | 1.779       |
| No. observations   | 1,564       | 608         | 2,172       |
| SSR                | 182.79      | 76.19       | 261.18      |

Note.—All 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.

wages within each of these periods are generated by a common process. 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. The hypothesis of homogeneous coefficients across the entire sample period, excepting the time intercepts and the union coefficients, cannot be rejected at conventional levels of

35 The F-statistic to test the restrictions implied by pooling the regulatory years, 1973–78, is 1.09, which is distributed as $F(45, 1,503)$ under the null hypothesis ($H_0$) of homogeneous coefficients. The F-statistic to test the null hypothesis of homogeneous coefficients for the deregulation period, 1979–85, is 0.72, distributed as $F(45, 548)$ under $H_0$. The critical value for both test statistics is approximately 1.24 at the 10 percent significance level.

36 The t-statistic to test the hypothesis of equal union coefficients over regulation and deregulation periods is 4.87.
significance. This suggests that the dominant change in the wage behavior of truck drivers over the period studied may be a decline in the union's ability to maintain its relative wage advantage.

3. Evaluating Wages for a Representative Driver

The union wage decline in the CPS results predates the declines apparent in the union contracts and the aggregate BLS data. While these findings are not necessarily contradictory (e.g., noncontract 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 timing), they do warrant further investigation.

To explore the decline in the union differential, I examined wages over time for a worker of given characteristics. I used the representative driver defined earlier: a married white male with 12 years of education and 20 years of experience, living in the North Central region. Predicted union and nonunion wages for this worker were calculated from a single wage equation pooling data from all sample years. Intercepts and union coefficients were permitted to vary across years; all other coefficients were held constant over time.

The results suggest that the lower union premium 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 may suggest some erosion of union power prior to deregulation, consistent with the arguments advanced by Levinson (1980). The effect on real wages is illustrated in figure 2, which plots predicted union and nonunion 1985 constant dollar wages for a representative driver over time. Predicted union wages declined steadily after 1978, with real wages in 1983–85 averaging $11.63, or 20 percent less than the average real wage of $14.39 in 1973–78. In contrast, real nonunion wages remained relatively high through 1981 (compared to their early 1970s levels) and declined only 7 percent between 1973–78 and 1983–85, from $9.60 to $8.91.

The reduction in the dollar union wage premium was substantial. The union wage differential declined in nominal dollars between 1977 and 1979; in real terms the differential fell by one-half relative to its

37 The $F$-statistic to test this hypothesis is 1.026, which is distributed as $F(98, 2,051)$ under the null hypothesis of homogeneous coefficients. The critical value at the 10 percent significance level is 1.19. Note, however, that the hypothesis that the coefficients in cols. 1 and 2 are the same, excepting the union coefficients, will be rejected at the 2.5 percent significance level. The $F$-statistic for this test is 2.29, which is distributed as $F(8, 2,157)$. 
level under regulation. Despite an increase in the union premium in 1983–84, the 1983–85 average real differential was $2.71—less than 60 percent of its 1973–78 average real level of $4.78 (expressed in 1985 constant dollars). For example, the union premium in 1975 was $2.51 per hour in nominal dollars, or about $5,600 per year (based on a 45-hour week, 50-week year). In 1983, the nominal dollar premium was $2.38 per hour, or $5,300 per year. This represents a real decline of nearly $5,500 in annual earnings in 1985 dollars.

4. The Influence of Economic Conditions

A major impediment to using time-series data to estimate regulatory effects is the difficulty of disentangling the impact of regulatory and other changes in the industry’s operating environment. The use of nonunion trucking labor as a benchmark should control for most industry-specific shocks. Two sources of economic variation—inflation and employment fluctuations—may affect union premia independently of regulatory changes. Oil price shocks also may affect the premium, but these seem likely to operate only through their effects on industry output and employment.

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38 Oil price shocks also may affect the premium, but these seem likely to operate only through their effects on industry output and employment.
such as the late 1970s (depending, of course, on the structure of COLA clauses built into the contracts). Union wage rigidity also implies that the premium should move countercyclically: contract wages will remain high even as unemployment reduces more flexible nonunion wages. These factors seem unlikely to explain the substantial drop in the union premium after 1979, however, given the stability of the union premium in response to the 1973–74 OPEC price shock, the 1975 recession, and the relatively high inflation rates of the mid-1970s.

I explore the implications of these factors for the union premium results by modeling the union coefficient as a function of economic conditions and a deregulation shift parameter (DEREG). Three measures of economic conditions are considered: (1) INFL, the inflation rate as measured by the consumer price index; (2) UNEMP, the average civilian unemployment rate; and (3) DTRKEMP, the change in total trucking industry employment. To estimate this relation, I replaced \( \beta_1 \) in equation (2) with

\[
\beta_1 = \alpha_0 + \alpha_1 \cdot UNEMP + \alpha_2 \cdot INFL + \alpha_3 \cdot DTRKEMP + \alpha_4 \cdot DEREG
\]

and used the pooled 1973–85 sample to estimate this modification of equation (2) (including separate intercepts for each year). Variations of these estimates are reported in table 4. They suggest that union premia vary inversely with the inflation rate and directly with unemployment. The change in trucking industry employment appears to

\[\text{TABLE 4}\]

**Influence of Economic Conditions on Union Wage Premia: Estimates from a Pooled CPS Wage Equation, 1973–85**

\[
\beta_1 \cdot UNION = (\alpha_0 + \alpha_1 \cdot UNEMP + \alpha_2 \cdot INFL + \alpha_3 \cdot DTRKEMP + \alpha_4 \cdot DEREG) \cdot UNION
\]

|      | (1)  | (2)  | (3)  | (4)  | (5)  |
|------|------|------|------|------|------|
| \( \alpha_0 \): constant | .402 | .402 | .402 | .402 | .403 |
|      | (.019)| (.019)| (.020)| (.020)| (.019)|
| \( \alpha_1 \): UNEMP | ... | .016 | .012 | -.004 | ... |
|      | (.013)| (.014)| (.020)| ...| ...|
| \( \alpha_2 \): INFL | ... | ... | -.006 | -.011 | -.010 |
|      | (.006)| (.008)| (.006)| ...| ...|
| \( \alpha_3 \): DTRKEMP | ... | ... | ... | -.004 | -.0033 |
|      | ... | ... | ... | (.004)| (.0025)|
| \( \alpha_4 \): DEREG | -.169 | -.177 | -.172 | -.160 | -.163 |
|      | (.035)| (.035)| (.036)| (.037)| (.035)|
| SSR  | 261.18 | 260.99 | 260.88 | 260.74 | 260.74 |

**Note.**—Estimated from a modified eq. (2), incorporating time effects for each year and replacing \( \beta_1 \) with the parameterization above. Standard errors are in parentheses.
have slightly more explanatory power than does the aggregate unemployment rate, although neither is statistically significant. Most important, controlling for variations in economic conditions does nothing to weaken the conclusion that deregulation has substantially reduced union premia.

5. Comparing Trucking Union Premia with Economywide Union Premia

I finally examine whether the changes in the trucking union premium reflect either general economywide movements in union premia or idiosyncrasies of the CPS data set rather than regulation-specific responses. This is accomplished by comparing the CPS results for the trucking industry with those for a cross-industry sample of all private industry blue-collar workers. Freeman (1986) estimated average blue-collar premia from CPS data for the 1973–84 period. His analysis suggests that the 1979 decline in the union coefficient may reflect an aberration of the CPS data set for that year. He found that the average blue-collar union coefficient rose from 0.23 (0.01) in 1973 to a peak of 0.28 (0.01) in 1977, then fell back to 0.22 (0.01) in 1979, a decline of 10 percent from its average level over 1973–78.\textsuperscript{39} Despite the common 1979 union premium decline, however, the behavior of the two series suggests that the trucking results are not merely reflective of general trends in the economy as a whole or in the CPS data. The trucking union differential averaged a fairly stable 0.40 (0.02) from 1973–78 then dropped to one-half of its 1973–78 average level in 1979. This decline brought the trucking union differential into the range of the average blue-collar differential—a significant change from its pre-1979 relative position. Over the 1979–84 period, the average trucking premium was almost identical to the average blue-collar premium: 0.25 (0.04) versus 0.22 (0.004), respectively.

The decline in the union wage premium in the trucking industry was neither a general economywide phenomenon nor a mere artifact of the particular data set used. The pattern of wages observed in the trucking industry indicates considerable stability in the union premium over the regulatory period, followed by a dramatic and sustained reduction in the premium in 1979 and thereafter. This decline is consistent with the contract evidence and aggregate wage data eval-

\textsuperscript{39} In earlier work, I found similar declines in the 1979 union premium for a number of industries. The results in Sec. IV, based on separate union and nonunion equations for workers in blue-collar occupations, indicate a lower premium in 1979 but suggest that the average union premium for blue-collar workers excluding truck drivers is constant (or increasing) over the sample period. These results imply an average union coefficient of 0.233 (0.005) for 1973–78 and 0.240 (0.010) for 1979–85.
uted earlier, although the CPS data suggest that deregulation may have reduced the union's relative wage advantage before the effects showed up in contract negotiations and aggregate earnings indices.

IV. Rent Spillovers

If nonunion workers in the trucking industry shared in regulatory rents, the decline in union wages relative to nonunion wages will tend to understate both union rents and total labor rents. The analysis in Section II also suggested that regulation may have raised wages of owner-operators and drivers in exempt trucking firms, such as private carriers. To evaluate these possibilities, we need a benchmark against which to measure the wage behavior of these two groups of drivers. The ideal benchmark would be a measure of these workers' opportunity wages. As an approximation to this, I used the average wage for blue-collar workers.40

Unfortunately, wage data for nonunion, owner-operator, and private carriage drivers are quite sparse. The CPS appears to be the only source of consistent time-series wage data for these workers, and since it does not collect wage information from self-employed workers, owner-operators are excluded. This restricts the analysis to nonunion trucking industry drivers and private carriage drivers.

To estimate the pattern of truck driver wage premia through time, I extracted samples of all private, nonagricultural blue-collar workers reporting complete wage equation information from the May CPS for each year. These were divided into union and nonunion subsamples; separate wage equations were estimated for each subsample and each year. The equations modified equation (2) to include fixed effects for each occupation class and a separate intercept for trucking industry workers.41

The estimates were used to construct (1) LBASEWAGE, the average estimated log hourly wage (LHWAGE) for a non–truck driver,

40 The choice of a benchmark or reference group should satisfy two conditions: First, mobility from drivers to the benchmark occupations (and vice versa) should be high; i.e., the benchmark should represent true alternative employment opportunities. Second, the benchmark should not be so narrowly defined that changes in trucking wages have a significant effect on wages in the benchmark occupations; otherwise, the measure will understate the effect of deregulation. This suggests a trade-off: broader reference groups will do less well on the first criterion and better on the second, and conversely for narrower reference groups. The blue-collar workers I chose represent a fairly broad benchmark. However, Hirsch's (1986) results for narrower reference groups suggest that the results are robust to the definition of the benchmark.

41 The equations included a FEMALE dummy variable since women are represented in these larger samples. Only full-time workers were included. The results are qualitatively similar to those obtained from equations that estimated a separate fixed effect for each industry group.
non–trucking industry representative worker (computed from the sample weighted average of occupational fixed effects for all occupations except truck driver, for a married white male with 12 years of education and 20 years of experience, living in the North Central region); (2) DRIVE PREM, the occupational premium for truck drivers over the BASEWAGE; (3) TRKDRV PREM, the premium over the BASEWAGE for truck drivers in the trucking industry (computed as DRIVE PREM plus the trucking industry intercept less the average intercept for all other industries); and (4) BASE UNPREM, the average union premium for non–truck drivers, non–trucking industry workers (computed as the difference between the LBASEWAGEs for union and nonunion workers). These measures are reported in table 5 for union and nonunion workers.

A. Nonunion Trucking Industry Drivers

Column 2 of table 5 describes the wage behavior of nonunion trucking industry drivers relative to the average nonunion blue-collar wage benchmark (LBASEWAGE, in col. 1). Nonunion trucking industry drivers earn less than the LBASEWAGE (as evidenced by the typically negative coefficients), although the difference is seldom statistically significant. Conclusions about wage behavior through time are hampered by the substantial year-to-year variation in estimated premia, due in part to the small number of trucking industry drivers in the post-1978 samples. This variance through time, combined with large standard errors on the point estimates, cautions against relying on premia estimates for any single year.

The last three rows of table 5 report average premia over the regulated (1973–78) and deregulated (1979–85) sample periods and the difference between the two periods. These suggest that nonunion trucking industry drivers were essentially unaffected by trucking deregulation. The average premium during the regulated period was $-0.030 (0.011)$; this increased slightly to $-0.014 (0.018)$ after deregulation, but the difference (1.7 percentage points) is neither substantively nor statistically significant. This result can be contrasted to the change in wages for union drivers in the trucking industry, relative to the union LBASEWAGE benchmark, reported in column 5. The premium over LBASEWAGE for union trucking industry drivers declined by $12 (2.0)$ percentage points after deregulation—an order of magnitude larger than the nonunion effect. These results indicate that, despite substantial union gains, nonunion trucking industry drivers were unable to capture a significant share of regulatory rents. This is consistent with insignificant union threat effects in the industry and suggests a reasonably elastic supply of nonunion labor to the industry.
| Year | Nonunion | | | Union | | |
|------|----------|----------|----------|----------|----------|----------|
|      | LBASEWAGE | TRKDRV PREM | DRIVE PREM | BASE UNPREM | TRKDRV PREM | DRIVE PREM |
|      | (1)       | (2)      | (3)       | (4)       | (5)       | (6)       |
| 1973 | 1.411     | -0.037   | -0.085    | 0.209      | 0.171      | -0.007    |
|      | (0.010)   | (0.028)  | (0.020)   | (0.012)    | (0.021)    | (0.020)   |
| 1974 | 1.486     | -0.013   | -0.104    | 0.205      | 0.120      | -0.014    |
|      | (0.010)   | (0.027)  | (0.018)   | (0.012)    | (0.019)    | (0.019)   |
| 1975 | 1.553     | -0.065   | -0.100    | 0.224      | 0.119      | -0.035    |
|      | (0.010)   | (0.027)  | (0.019)   | (0.012)    | (0.021)    | (0.020)   |
| 1976 | 1.629     | -0.072   | -0.128    | 0.224      | 0.100      | -0.046    |
|      | (0.010)   | (0.027)  | (0.018)   | (0.013)    | (0.021)    | (0.021)   |
| 1977 | 1.678     | -0.012   | -0.091    | 0.264      | 0.144      | -0.031    |
|      | (0.009)   | (0.025)  | (0.016)   | (0.012)    | (0.019)    | (0.019)   |
| 1978 | 1.771     | 0.019    | -0.091    | 0.267      | 0.123      | -0.059    |
|      | (0.009)   | (0.025)  | (0.017)   | (0.012)    | (0.020)    | (0.019)   |
| 1979 | 1.890     | 0.020    | -0.071    | 0.215      | 0.060      | -0.040    |
|      | (0.013)   | (0.034)  | (0.024)   | (0.017)    | (0.027)    | (0.026)   |
| 1980 | 1.955     | -0.049   | -0.114    | 0.213      | 0.020      | -0.069    |
|      | (0.016)   | (0.044)  | (0.031)   | (0.022)    | (0.038)    | (0.033)   |
| 1981 | 2.013     | 0.068    | -0.060    | 0.249      | 0.056      | -0.045    |
|      | (0.018)   | (0.051)  | (0.033)   | (0.024)    | (0.041)    | (0.040)   |
| 1983 | 2.127     | -0.095   | -0.065    | 0.220      | -0.028     | -0.130    |
|      | (0.020)   | (0.052)  | (0.038)   | (0.028)    | (0.048)    | (0.049)   |
| 1984 | 2.157     | -0.020   | -0.058    | 0.263      | 0.023      | -0.046    |
|      | (0.018)   | (0.041)  | (0.034)   | (0.025)    | (0.052)    | (0.047)   |
| 1985 | 2.169     | -0.005   | -0.121    | 0.283      | -0.059     | -0.046    |
|      | (0.018)   | (0.037)  | (0.035)   | (0.026)    | (0.050)    | (0.049)   |
| Average, 1973–78 | ... | -0.030 | -0.100 | ... | 0.232 | 0.129 | -0.032 |
|      | ... | (0.011) | (0.007) | ... | (0.005) | (0.008) | (0.008) |
| Average, 1979–85 | ... | -0.014 | -0.082 | ... | 0.241 | 0.012 | -0.062 |
|      | ... | (0.018) | (0.013) | ... | (0.010) | (0.018) | (0.017) |
| Difference | ... | 0.017 | 0.018 | ... | 0.009 | -0.117 | -0.030 |
|      | ... | (0.021) | (0.008) | ... | (0.011) | (0.020) | (0.009) |

Note.—Samples include all private, nonagricultural blue-collar workers. Estimated from a modified eq. (2). The dependent variable is ln(hourly wage). See text for a description of the equation and reported coefficients. Standard errors are in parentheses.
B. *Truck Drivers outside the Trucking Industry*

The behavior of wages for private carriage drivers, relative to the average blue-collar LBASEWAGE, is measured in column 3 for nonunion workers and column 6 for union workers. The wage behavior for nonunion drivers was similar to that for nonunion trucking industry drivers, although both the wage level and the year-to-year variation in premia were lower outside the trucking industry. The regulation period average premium was $-0.100 (0.007)$, implying that truck driver wages were 10 percent below average blue-collar wages. There was a small gain of $0.018 (0.008)$ after deregulation, to an average premium level of $-0.082 (0.013)$. The effect is statistically significant but substantively unimportant. While the increase in the premia may be consistent with a slightly upward-sloping supply curve for truck drivers, the magnitude of the effect suggests a fairly large elasticity.

The premia for unionized private carriage drivers were uniformly negative (though often not statistically different from zero) and reasonably stable from year to year. The data suggest a slight decline in wages for these workers after deregulation: the average premium of $-0.032 (.008)$ during the regulated period fell to $-0.062 (0.017)$ during the deregulation period. At $-3.0 (0.9)$ percentage points, this decline is small in absolute terms but large relative to the change in the corresponding premium for nonunion workers. These results can be restated in terms of union wage premia. For private carriage truck drivers, the union premia over nonunion workers averaged $0.301 (0.011)$, or 35 percent over 1973–78. The union wage premium fell to $0.260 (0.023)$, or 30 percent, over 1979–85.\(^{42}\) This compares to a roughly constant average union premium of $0.240 (0.010)$ for blue-collar workers excluding truck drivers (BASE UNPREM, in col. 4). These findings may suggest some rent spillovers into union wages outside the trucking industry, but this effect is small.

V. *Aggregate Labor Rents*

The wage response results imply substantial labor rents. To put the findings into context, I calculated a rough measure of the total labor losses resulting from deregulation. Because the rent spillovers to nonunion workers and drivers outside the for-hire trucking industry appear to be small or zero, I restricted this calculation to rents cap-

\(^{42}\) This average deregulation union premium is almost identical to the average 1979–85 trucking industry union wage premium of $0.266 (0.026)$ estimated from these equations, and is only slightly above the 1979–85 average blue-collar union wage premium of $0.240 (0.010)$.\)
tured by the Teamsters Union. A number of assumptions were required to compute aggregate union rents. First, I assumed that deregulation reduced the 1973–78 union premium of 50 percent to 35 percent (this is somewhat above the 1983–85 average premium of 31 percent). The sensitivity of the estimate to this assumption was checked by recalculating losses assuming that the premium declined to 25 percent (slightly below the 1979–85 average premium of 27 percent). Second, I assumed that nonunion wages included no regulatory rents. To the extent that nonunion wages were bid up under regulation, this assumption would understate both union and nonunion rents. My calculation attributed the entire change in the union premium to a reduction in rents but omitted any effect from changes in the overall trucking wage level. Finally, I excluded potential losses from a reduction in union employment. This ignored the effect of a decline in union coverage by excluding the union’s loss from jobs that were no longer held by union employees.

Given these assumptions, the loss to an individual union driver was estimated at 10–20 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. The total union loss was 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 regulated trucking firms (which are quite small), it accounts for at least half of total regulated revenues. Applying the assumptions described above to 1983 total employee compensation for Class I motor carriers implies union compensation of $8.2 billion and an estimated annual reduction of $900 million in rents for unionized employees of these firms. If the average deregulation union premium is assumed to be 25 percent, the estimated union loss rises to $1.6 billion.

Finally, the reduction in union rents can be compared with the losses incurred by owners of trucking firms to estimate the relative shares of rents for unionized labor and capital. Rose (1985a) estimated the deregulation-related loss in market value of general freight trucking firms at 8.8 percent of 1978 revenues. Applying this estimate to 1978 revenues for all Class I general freight common carriers yields an after-tax loss of $1.5 billion in present discounted value. This corresponds to annual pretax rents of roughly $300 million in

\[\Delta = (1.35 - 1.50) \cdot \text{nonunion wage} = -0.15 \cdot \text{nonunion wage}.\]

\[\text{The percentage loss is } \frac{\Delta}{\text{union earnings}} \cdot 100 = \frac{-15}{1.35} = -11.1 \text{ percent.}\]

\[\text{Total employee compensation of } $12.24 \text{ billion for the 886 Class I motor carriers reporting to the ICC was taken from the U.S. Interstate Commerce Commission Bureau of Accounts, } \textit{Transport Statistics in the United States: Motor Carriers}, \textit{pt. 2}, 1983.\]
1978 ($500 million in 1985 dollars). Applying the assumptions described above to calculate 1978 annual pretax union rents for these firms yields an estimate of $600 million to $1.0 billion, or $950 million to $1.6 billion in 1985 dollars. These calculations, while crude, suggest that the Teamsters Union may have been the dominant beneficiary of trucking regulation, capturing 65–76 percent of the total rents in the industry. The estimated total annual rents are 5–9 percent of industry revenues.

VI. Conclusion

This study examined industry wage responses to motor carrier deregulation. Union contract evidence and aggregate data on industry average earnings suggest substantial changes in wage determination patterns after motor carrier deregulation. Microdata estimates of union premia over nonunion wages in the trucking industry indicate declines of roughly 40 percent in the size of the union differential, beginning in 1979. Annual earnings for a representative union driver in 1983–85 were $3,800, or 14 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 was estimated at $900 million to $1.6 billion for employees of the 886 largest regulated trucking firms in 1983. Contrary to recent speculation on nonunion rent sharing (Dickens and Katz 1987a, 1987b; Krueger and Summers 1987, in press), the results provide little evidence of rent spillovers to either nonunion trucking industry drivers or truck drivers outside the regulated trucking industry.

Comparison of the results in this paper with estimates of rents accruing to owners of capital indicates that the Teamsters Union was the primary beneficiary of regulatory rents in the trucking industry, capturing as much as 75 percent of total rents. The findings suggest that the assumption of competitive factor prices can be quite misleading, at least for regulated markets, and provide strong support for union rent-sharing hypotheses.

The conversion from after-tax present discounted value to annual pretax rents is based on a 50 percent corporate income tax rate, a 10 percent discount rate, and perpetuity annual rents.

Total rents were measured as union rents plus firm rents. Because regulatory rents are inferred from deregulation losses, this measure implicitly assumes that firms and unions share equally in the decline. If unions are less able to retain rents than firms are, my measure will overstate the union's share, and conversely if unions are better able to resist wage reductions. The findings of this paper are broadly consistent with the estimated union share obtained in 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 study employing Tobin's q.
### Appendix

**TABLE A1**

**Annual CPS Wage Equations**

| Variables | 1973 | 1974 | 1975 | 1976 | 1977 | 1978 | 1979 | 1980 | 1981 | 1983 | 1984 | 1985 |
|-----------|------|------|------|------|------|------|------|------|------|------|------|------|
| Constant  | 1.329| 1.458| 1.405| 1.611| 1.670| 1.775| 1.933| 1.972| 2.110| 1.928| 2.121| 2.053|
| UNION     | .440 | .345 | .448 | .391 | .441 | .346 | .213 | .199 | .208 | .292 | .346 | .177 |
| EDUC      | .025 | .047 | .014 | .017 | .020 | .011 | .007 | .018 | .029 | .005 | .014 | .005 |
| EXP       | .003 | .006 | .005 | .009 | .004 | .008 | .005 | .011 | .008 | .001 | .003 | .007 |
| EXP²      | -.0004 | -.0002 | -.0002 | -.0006 | -.0004 | -.0005 | -.0002 | -.0003 | -.0004 | -.0001 | -.0005 | -.0003 |
| SINGLE    | -.085 | -.073 | -.090 | -.074 | -.054 | -.060 | -.072 | -.105 | -.110 | -.092 | -.100 | -.118 |
| NONWHITE  | -.198 | -.133 | -.158 | -.015 | -.060 | -.199 | -.153 | -.023 | -.131 | .047 | .075 | -.282 |
| NORTHEAST | -.090 | -.084 | -.079 | -.116 | -.109 | -.011 | -.146 | -.161 | -.198 | .088 | -.121 | .039 |
| SOUTH     | -.075 | -.060 | -.058 | -.176 | -.103 | -.048 | -.098 | -.085 | -.033 | -.008 | -.131 | .110 |
| WEST      | .043 | .052 | .164 | .014 | -.012 | -.025 | .129 | -.035 | .097 | .225 | .065 | .141 |
| No. observations | 230 | 258 | 245 | 236 | 304 | 291 | 169 | 91 | 83 | 79 | 78 | 108 |
| R²        | .38 | .30 | .35 | .41 | .42 | .30 | .26 | .19 | .22 | .23 | .30 | .17 |
| SSR       | 25.02 | 30.87 | 34.41 | 24.90 | 27.19 | 34.63 | 17.84 | 11.49 | 10.43 | 9.18 | 8.49 | 14.52 |

**Note.**—Dependent variable is ln(hourly wage). See eq. (2) for a description of the equation and variables. Standard errors are in parentheses.
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