Trucking Deregulation and Labor Earnings: Is the Union Premium a Compensating Differential?

Barry T. Hirsch  
Department of Economics  
Florida State University  
Tallahassee, Florida 32306-2045  
(904)644-7207

Abstract

This paper examines wage determination among union and nonunion truck drivers using the 96 monthly Current Population Surveys for 1983-90. Union density in the previously regulated for-hire sector of the trucking industry fell from about 60 percent during the regulatory period of the 1970s to about 25 percent by 1990. Union log wage premiums fell from 0.40 in the 1970s to 0.30 or below in the 1980s. Longitudinal estimates from multiple panels for 1983-84 through 1989-90 suggest far smaller union premiums, supporting the thesis that part of the wage advantage following deregulation is a compensating differential for driver quality.

The author appreciates the comments and suggestions of David Card, Jack Fiorito, William Linde, an anonymous referee, and seminar participants at Florida State University. David Macpherson assisted in creation of the data set, as well as providing comments.
Introduction

Regulation and subsequent deregulation of the trucking industry by the Interstate Commerce Commission (ICC) has provided a natural experiment by which researchers can study, among other things, regulatory effects on labor earnings, union coverage, and union and nonunion rents. Recent studies by Rose (1987) and Hirsch (1988) utilize the May Current Population Survey (CPS) public use tapes for 1973-85 to examine these issues. Based on their analyses of truck drivers in the previously regulated for-hire sector of the industry, they conclude that both union density (i.e., the proportion of drivers who are union members) and the union-nonunion wage differential decreased markedly in the transition from a regulated to a largely deregulated environment. Moreover, they find little evidence for nonunion rent sharing during the regulatory period. The wages of nonunion drivers in the covered sector, as well as both union and nonunion drivers in the unregulated private carrier sector, largely mirrored economy-wide wage levels both during and following regulation of the industry.1

The purpose of this paper is twofold. First, evidence on unionism and wages among truck drivers is examined using a data set that updates and overcomes the serious sample size limitations of previous studies. Second, longitudinal analysis using seven panels of truck drivers allows us to estimate the extent to which union-nonunion differentials among drivers represent compensating premiums for driver quality.2

Deregulation and the Wages of Union and Nonunion Drivers

Regulatory authority over motor freight carriage was initially provided to the Interstate Commerce Commission (ICC) by the Motor Carrier Act of 1935. As it evolved, the ICC severely restrained both price and entry competition among for-hire common carriers engaged in intercity and interstate cartage. Largely exempt were private and local carriers, and carriers of unprocessed agricultural products.

Regulatory constraint on entry and price competition, along with rapid growth in demand for trucking services, improvements in truck technology, and expansion of the U.S. highway system, combined to create large rents in the trucking industry. Rents took the form of higher prices and costs than would have

---

1 References to earlier studies are provided by Rose (1987) and Hirsch (1988). Ying and Keeler (1991) provide an analysis of deregulation's effects on pricing in the motor carrier industry. For an analysis of the effect of deregulation on wages and employment in the airline industry, see Card (1989). In contrast to evidence from the trucking industry, Card finds that wage losses have been small for pilots, flight attendants, mechanics, and airline workers in total.

2 Although not the focus of the paper, the size of the data set also allows for the first time an examination of gender differences in labor earnings among drivers.
existed in a competitive industry. These rents were subject to capture by capital owners (or, more precisely, initial owners of ICC route certificates) and by labor (Rose 1985, 1987). Within this regulated environment, the International Brotherhood of Teamsters (IBT) gained power and captured for union drivers a substantial share of industry rents. Union power was concentrated in the less-than-truckload (LTL) sector of the industry, including the terminals through which LTL trucks and cargo passed. Although estimates are not precise, Rose (1987) calculates that labor captured roughly two-thirds to three-fourths of the total rents in the regulated sector of the industry during the 1970s.

Administrative deregulation within the ICC during the late 1970s and the subsequent Motor Carrier Act of 1980 effectively eliminated most price and entry restrictions in the industry. As expected, deregulation facilitated the entry and expansion of nonunion trucking operations and produced high rates of business failure among trucking companies, particularly high-cost union operations. Whereas traffic had shifted to the unregulated private carrier sector (i.e., firms transporting their own cargo) during the regulatory period, deregulation generated a shift of traffic back to the now deregulated for-hire sector.

The effects of deregulation on union density and union wage premiums were predictable. No longer able to maintain costly union operations in the face of lower-cost nonunion operations, union density (i.e., the proportion of truck drivers who are union members) and union-nonunion wage differentials decreased following deregulation. Rose and Hirsch find that union density in the previously regulated for-hire sector fell from about 60 percent during the 1970s and early 1980s to about 30 percent by 1984-85, while the union-nonunion wage differential fell from about 50 percent during 1973-78 to an average 30 percent during 1979-85.

Although for any given labor market environment, sustainable union premiums and union density are inversely related (Lazear, 1983), deregulation represented a dramatic change in the environment, leading to both lower coverage and wage premiums. Stated alternatively, the Teamsters would have had to have agreed to far more substantial decreases in union wages and benefits if they were to preserve union coverage among drivers at regulatory period levels. The effects of deregulation on the wages and employment of nonunion drivers in the previously regulated sector, and both union and nonunion drivers in the unregulated private carrier sector, depends on labor demand and supply elasticities, threat effects, and rent sharing.
These issues are examined in some depth by both Rose and Hirsch. They conclude that there was little evidence of nonunion rent sharing during the regulatory period and that changes in nonunion driver wages following deregulation largely mirrored wage movements elsewhere in the economy (implying a highly elastic labor supply curve).

A primary limitation of the Rose and Hirsch studies was the small size of the truck driver samples in the previously regulated for-hire sector after 1980 (about 100 drivers per year, most of whom are nonunion). Associated with their small samples was large variability in estimates of union wage effects. For example, the Motor Carrier Act (MCA) of 1980 was passed in June 1980. Yet Rose and Hirsch find sharp decreases in union premiums for May 1979 and 1980 prior to its passage, as well as in 1981 following passage, but a return to historically higher levels for 1983-84 (the May 1982 CPS survey did not include the earnings supplement). The 1985 data indicated a very sharp drop in the union premium, to a level less than half as large as the 1984 estimate.

Rose and Hirsch date the deregulation period from 1979-85, using as justification the administrative deregulation prior to Congressional passage of the MCA. Their conclusions are based in large part on a comparison of averages for the 1973-78 and 1979-85 periods. Their conclusions would have differed considerably had comparisons been based instead on the periods 1973-80 and 1981-85, corresponding to passage of the June 1980 MCA. What is not clear is the extent to which the substantial year-to-year variability reported in their studies reflected small sample sizes, or actual changes over time in union density and the earnings of union and nonunion drivers.

Data and Evidence

The data base for this study comprises 96 monthly CPS surveys conducted between January 1983 and December 1990. Use of the 1983-90 CPS surveys, besides increasing annual sample sizes approximately twelve-fold, provides more recent evidence on wage determination in the trucking industry. Evidence on union density and union wage premiums thus allows us to reexamine the conclusions reached by Rose and Hirsch. In addition, design of the CPS allows construction of two-year panels in which data on

---

3 Beginning in January 1983, union status questions were asked of the outgoing rotation groups (ORG) in each monthly survey rather than only in the May public use surveys. The ORG “earnings microdata files” are made available by the Data Services Group at the Bureau of Labor Statistics.
a portion of drivers are observed for 1983-84 through 1989-90 (see the Data Appendix). Based on longitudinal evidence, changes in wages associated with individual-specific changes in union status are examined. Such evidence allows us to address the question of what portion of the sizable union-nonunion wage differential is in fact a rent and what portion results from unobserved quality differences between union and nonunion drivers.

In order to compare results to those for May 1973-85, selection of the sample and modeling follow that presented in Hirsch (1988). The 1983-90 CPS sample includes all male truck drivers in the labor force, ages 16 to 64, who have data provided on usual weekly earnings, usual hours worked per week, union status, and whose hours worked are at least 30 hours. We distinguish between truck drivers in the for-hire or common carrier sector of the industry (this includes both general commodity and contract carriers), defined as drivers who designate their industry of employment as the trucking service industry, and drivers in the private carrier sector, defined as drivers who designate their industry of employment as something other than the trucking service industry. Owner operators are not included in the sample since self-employed workers are not asked the CPS earnings supplement questions. Drivers employed in the for-hire sector sample generally were covered by ICC regulation and thus affected directly by deregulation, whereas drivers in the largely unregulated private carrier sector were affected only indirectly by deregulation.

Table 1 provides descriptive data for the years 1973-90 on sample sizes, union density, union and nonunion wages for truck drivers, and the unadjusted and adjusted (i.e., regression based) union-nonunion log wage differential. The samples are separated into the for-hire and private carrier sectors. In the top of the table, 1973-85 information is based on calculations from May CPS surveys presented in Hirsch (1988, Tables 1-3, fn. 17, and unpublished data), with weighted means presented for the “regulatory period” 1973-78, and May data annually thereafter. Data then are presented for 1983-90 based on the twelve monthly CPS surveys in each year. Wages are expressed in constant fourth-quarter 1990 dollars.4

Union density fell sharply in the previously regulated for-hire sector, from an average 60% during the regulatory period to less than 40% by 1983-85, and to 24% by 1990. Union density in the private carrier sector

---

4 Monthly CPS wages are deflated by the quarterly personal consumption expenditure (PCE) component of the GNP implicit price deflator, taken from the mid-year 1991 CITIBASE tape. Data for May 1973-85 taken from Hirsch (1988) are converted from constant 1985 to 1990:4 dollars. Nominal weekly earnings are topcoded at $999 through 1988, and at $1,927 beginning in 1989. Because so few truck drivers are topcoded, no adjustment to wages is made.
sector, which had been about 36% during the 1973-78 period, had fallen by 1983-85 to about 30%, and to 25% by 1990. Estimates of union coverage densities (not shown), calculated by adding to the number of union members those nonmember drivers who say they are covered by a collective bargaining agreement, are about .015 (1.5 percentage points) higher than the membership densities reported in Table 1. The monthly CPS surveys for 1983-90 largely reinforce earlier conclusions about union density reached by Rose and Hirsch, although year-to-year changes are far more stable than suggested by the use of data only from the May surveys. Union density among for-hire drivers has fallen sharply following deregulation, while changes among drivers outside the trucking industry have been modest.

Worth noting in Table 1 is the relative growth in employment in the for-hire sector following deregulation. Whereas the ratio of the private carrier sample size to the for-hire sample size is 2.1 during the 1973-78 regulatory period, it had fallen to 1.2 in the 1990 sample of drivers. This change reflects the distortionary effects of ICC regulation. During the regulatory period, many companies conducted their own trucking operations rather than ship goods by regulated common carriers. Following deregulation, firms have lessened use of their private trucking operations, with traffic being shifted toward the now more competitive for-hire sector.

Average real wage rates among union drivers in the for-hire sector (unadjusted for worker characteristics) have clearly declined from the regulatory period, but have shown relatively slow change since the mid-1980s. Wages among nonunion for-hire drivers remained relatively stable over the entire 1973-90 period, declining only modestly from their level in the 1970s. Union and nonunion drivers in the private carrier sector show little change in wages between the regulation and deregulation years. The maintenance of real wages in the private carrier sector, despite a shift of traffic and employment toward the for-hire sector, is consistent with a highly elastic labor supply curve and job skills that are primarily occupation- rather than firm-specific. Alternative (non-trucking) real wages during this period for workers with low levels of schooling tended to show little change or moderate deterioration. The descriptive evidence presented for 1983-90 also suggests that much of the year-to-year variability found by Rose and Hirsch following deregulation was due to the small sample sizes of drivers in the single-month surveys utilized through 1985. Month-to-month variability is examined explicitly below.
Table 1 also presents regression estimates of the union-nonunion differential, averaged for 1973-78, and annually for 1979-90.\(^5\) Estimates are based on the coefficients, \(\theta_{ky}\), attached to union dummy variables in semilogarithmic wage equations. That is, we estimate:

\[
\ln W_{iky} = \alpha_{ky} + \beta_{ky}X_{iky} + \theta_{ky} UN_{iky} + \epsilon_{iky},
\]

where \(\ln W_{iky}\) represents the natural logarithm of hourly earnings (usual weekly earnings divided by usual hours worked per week) of driver \(i\), in sector \(k\) (for-hire or private carrier), in year \(y\). The vector \(X\) includes variables measuring years of schooling, years of potential experience and experience squared, race, marital status, veteran status, and region (8 dummies).

Coefficient estimates (\(\theta\)) on the union membership variable (UN) are provided separately for truck drivers in the for-hire and private carrier sectors. In the previously regulated for-hire sector, the logarithmic union-nonunion wage differential fell from .39 during the regulatory period, to about .30 during the late 1980s.\(^6\) These latter estimates are substantially higher than previous estimates for May 1980-81 and May 1985, but similar to the May 1979-85 deregulation average. Thus, the results provide support for the general conclusions reached by Rose and Hirsch, while at the same time illustrating the need for substantially larger sample sizes than are available in a single month’s survey. Estimated union premiums in the previously unregulated private carrier sector are similar or only moderately lower following deregulation than during the regulatory period, in the range .25-.31 during 1983-90 compared to an estimate of .30 during 1973-78.

Finally, Table 1 allows a comparison of the union coefficient \(\theta\), measuring union-nonunion wage differentials conditional on characteristics included in vector \(X\), with the mean union-nonunion log wage differential unadjusted for characteristics. Accounting for differences in worker characteristics explains little of the observed union premium, as seen by the fact that the conditional differentials are only moderately lower than the unadjusted differentials. In part this reflects relatively small differences between union and nonunion drivers in schooling and other characteristics, and in part the relatively low returns on some of these characteristics. For example, the within-occupation (trucking) rate of return to schooling

\(^5\) The 1973-78 averages are presented in Hirsch (1988, Table 2); annual estimates for May 1979-85 are from Hirsch (1988, p. 315n) for the for-hire sector and from unpublished results for the private carrier sector. The annual 1983-90 estimates are estimated for the 12 combined monthly CPS surveys for each year. For discussion of econometric issues in the estimation of union-nonunion wage gaps, see Lewis (1986) and Hirsch and Addison (1986).

\(^6\) Logarithmic wage differentials, \(\theta\), can be converted to approximate percentage differences, \(D\), by \(D = [\exp(\theta)-1]100\). More precise approximations are compared in Giles (1982).
averages .018 across years in the for-hire sector. While differences in schooling between union and nonunion drivers are small, unionized drivers tend to be older (by about 6 years) and somewhat more likely to be married and white than are nonunion drivers.

Previous analysis by Hirsch (1988) for the 1973-85 period indicated that real wage movements among nonunion truck drivers in the for-hire sector, and among both union and nonunion drivers in the private carrier sector, largely mirrored wage movements among a control group of non-truck operatives, selected to approximate long-run reservation wage and employment opportunities. Rose (1987) utilizes broader control groups and reaches a similar conclusion. Table 2 provides information for 1983-90 for two samples of male workers – a group of non-truck operatives with occupations equivalent to those chosen by Hirsch (1988) and a sample of all blue-collar workers in manufacturing. Other selection criteria are identical to those used to select the truck driver samples. Information presented for the two control groups parallels that found in Table 1.

The descriptive data provided in Table 2 provides measures of economy-wide wage movements and union-nonunion differentials during the 1983-90 period. The pattern previously found for truck drivers in both the for-hire and private carrier sectors corresponds reasonably closely to economy-wide patterns among the control groups. There is relatively little change in union premiums and average real wage rates during the 1980s, although the control groups do display some deterioration in real wages during the late 1980s (fringe benefits, of course, are not included in this measure of hourly earnings). The similarities in Tables 1 and 2 help reinforce the previous conclusion that apart from the wages of unionized drivers in the for-hire sector during and immediately after deregulation, the wage movements of both union and nonunion drivers have largely mirrored economy-wide wage movements. The similarity in wages between truck drivers and other occupations, even in the face of a restructuring of the trucking industry and substantial shifts in traffic associated with deregulation, strongly suggests a highly elastic labor supply curve. Apart from the substantial decrease in wages among union drivers in the for-hire sector following deregulation, real wages of drivers have proven to be relatively invariant with respect to demand and employment shifts.

In order to illustrate more directly the gains associated with larger sample sizes, separate monthly regressions in the for-hire sector are estimated for January 1983 - December 1990. Table 3 presents the
means (unweighted) of the monthly estimates, as well as the minimum and maximum estimates. Month-to-month variability is found to be substantial. For example, in 1985 union premium estimates range from a low of .106 for the May sample to a high of .525 for the December sample.\(^7\) No obvious seasonal pattern is evident in the month-to-month variability. Rather, the results reported in Table 3 imply rather clearly that reliable inferences cannot be drawn from regression results based on single-month samples of approximately 100 drivers, about 30 of whom are union members. Annual estimates of the union wage premium based only on the May survey do not permit inferences as to the exact timing of the labor market responses to deregulation. The month-to-month variability in estimates demonstrated in Table 3 does help explain the rather weak correspondence found between contract terms in the National Master Freight Agreements and May CPS estimates of union wage premiums during 1973-85 (see Rose, 1987; for a recent summary of NMFA contract terms, see Bureau of National Affairs, 1991).

Finally, use of the 96 monthly CPS surveys permits analysis of male-female wage differentials, not possible previously owing to the small number of female drivers (Rose 1987, p. 1159n). There are 444 female drivers in the combined samples for 1983-90 (and 22,496 males). Female drivers tend to be younger (35.1 versus 38.6 years), more educated (11.7 versus 11.0 years), less likely to be married (50.2% vs. 73.0%), and less likely to be union members (15.5% vs. 29.2%) than their male counterparts. Estimation of a pooled male-female wage equation for 1983-90 (\(N = 22,940\)) produces a coefficient (\(|t|\) of -0.195 (11.01) on a female dummy variable, indicating female drivers have hourly earnings 17.7% lower than similar males (all variables are the same as in previous regressions). The unadjusted mean log wage differential is 0.244. Estimation of the gender differential using separate male and female wage equations, and weighting alternatively by male and female coefficients, produces virtually identical estimates. Is the estimated gender wage gap large or small? It is low relative to economy-wide gender gaps. But much of the economy-wide gap reflects inter-occupational differences, while gaps within narrowly-defined occupations are generally lower than the gap among truck drivers. The CPS provides neither sufficiently detailed information on worker and job characteristics, nor large enough samples of female drivers to permit a detailed analysis of the determinants of the gender gap. In fact, the severely limited employment of female drivers may say

\(^7\) Note that the rather unrepresentative May 1985 survey formed the end-point used in both the Rose (1987) and Hirsch (1988) studies.
more about discriminatory barriers facing women than do estimates of wage differentials among existing drivers.⁸

**Longitudinal Analysis: Is the Union Premium a Compensating Differential?**

A potential problem in previous analyses has been the inability to control for unobservable determinants of truck driver quality. Specifically, employee and employer sorting make it likely that part of the observed union-nonunion wage differential among drivers reflects a compensating quality premium rather than a rent. Quality differences are most likely to arise because employers paying union wages can select higher quality drivers from among drivers in the job queue. Quality differences might also arise owing to lower quit rates or greater occupation- and firm-specific skills among more highly paid drivers. Higher quality drivers also might have a greater preference for union representation and thus can be organized at lower cost. Measured differences in years of schooling are unlikely to provide a good measure of differences in driver quality.

In order to address the issue of unmeasured driver quality, we exploit the panel construction of the CPS. Individuals are included in the CPS for 8 months – 4 consecutive months in the survey, followed by 8 months out, followed by 4 months in. Earnings supplement questions (weekly earnings, hours, and union status) are asked only of the outgoing rotations (surveys 4 and 8); thus individuals provide usable observations for the same month in consecutive years. For example, the 1989-90 panel includes male drivers observed in their 4th survey in 1989 and 8th survey in 1990, who report their occupation as truck driver during both years. Excluded from the 1989-90 panel are drivers in their 8th survey in 1989, their 4th survey in 1990, workers who are drivers in 1989 or 1990 but not both, and females. Estimates in this section are provided for 7 panels of drivers – the periods 1983-84 through 1989-90. In order to construct the panels, drivers are identified based on household identification numbers, as well as age, survey rotation, month, and year information.⁹

---

⁸ The 1979, 1983, and 1988 May/June CPS Pension Supplements, containing information on tenure with current employer, include only 13 female drivers. These women have an average ratio of years of tenure to years of potential experience of .182, as compared to .314 for male drivers (N = 1,167).

⁹ The Data Appendix describes construction of the longitudinal data set. Sample selection criteria are the same as previously, except drivers are required to work at least 30 hours per week in at least one rather than both years. Although correcting for some econometric problems, the subsequent analysis may introduce additional problems owing to measurement error in change variables and because changes in union status may not be independent of
The longitudinal samples consist of 3,659 male drivers in the for-hire and private sectors, each with multiple observations in adjoining years. Since wage determination and union-nonunion wage differentials are found to be broadly similar in the for-hire and private sectors following deregulation, I opt for the efficiency gains provided by a pooled analysis of the two sectors and multiple years (a sector variable is included). Four union states are defined: nonunion stayers (UN00), union stayers (UN11), union joiners (UN01), and union leavers (UN10), where the first number refers to union status in year 1 and the second to union status in year 2. Over the period approximately 9 percent of our sample changes union status during the year, roughly half being union joiners and half leavers (measurement error in the union status variable is discussed subsequently).

Longitudinal estimates of union wage premiums are estimated as follows. Let the wage equation in levels, pooled across years, be:

\[
\ln W_{iy} = \alpha_y + \beta X_{iy} + \tau UN_{iy} + \Phi_i + \epsilon_{iy},
\]

where \( \ln W_{iy} \) represents the log wage of driver \( i \) in year \( y \), vector \( X \) includes control variables and its coefficient vector \( \beta \) (assumed fixed over time), \( \alpha_y \) represents regression constants allowed to vary by year, \( \tau \) represents the logarithmic union-nonunion wage premium assumed fixed over time, \( \Phi_i \) is a driver-specific fixed-effect constant over the 2 periods for which each driver is observed, and \( \epsilon_{iy} \) is a random error term with zero mean and constant variance. If the omitted fixed effect (\( \Phi_i \)) capturing unmeasured driver skill differences is correlated with union status, estimation of (2) in levels form will lead to biased estimates of \( \tau \).

In difference form, letting \( t \) equal panel period \( y \) minus \( y-1 \), equation (2) becomes:

\[
\Delta \ln W_{it} = \Delta \alpha_t + \beta \Delta X_{it} + \tau \Delta UN_{it} + \Delta \epsilon_{it}.
\]

In difference form, fixed effect \( \Phi_i \) falls out, therefore purging correlation between UN and unmeasured driver quality differences. Panel specific intercepts \( \Delta \alpha_t \) capture both wage changes among drivers who do not change union status and the effect of one additional year of experience (\( \Delta EXP = 1 \) for all observations, whereas \( \Delta EXP SQ \) varies with the level of EXP). The union wage premium is now estimated based on wage changes among union status changers, where union joiners have \( \Delta UN = 1 \), union leavers \( \Delta UN = -1 \), and nonunion and union stayers \( \Delta UN = 0 \). The control variables in \( X \) vary little or remain fixed over a two-year unobservables correlated with wages. For previous estimation of a wage change equation using matched CPS data, see Mellow (1981).
period. Nonwhite, veteran status, and region (by design of the CPS) remain fixed and fall out, while changes in experience equal one and are reflected in the period intercepts. In addition to the change in union status and period dummies, included explanatory variables are change variables for schooling, experience squared, marital status, and sector (for-hire or private), and dummies for region.

Equation (3) restricts estimates of to be symmetrical such that wage gains for union joiners are equivalent to wage losses for leavers, and assumes panel period wage changes are equivalent for nonunion and union stayers. The following less restrictive specification relaxes these assumptions, allowing wage changes to differ between nonunion and union stayers and different estimates of for union joiners and leavers.

\[
\Delta \ln W_{it} = \Delta \alpha_t + \beta \Delta X_{it} + \tau_{01} UN01_{it} + \tau_{10} UN10_{it} + \tau_{11} UN11_{it} + \Delta \varepsilon_{it}.
\]

UN01, UN10, and UN11 are dummy variables set equal to 1 if the driver is a union joiner, leaver, or stayer, respectively. The parameter \( \tau_{01} \) provides a longitudinal estimate of the union wage premium based on the sample of union joiners, compared to the base group of nonunion stayers; \(- (\tau_{10} - \tau_{11})\) provides an alternative estimate of the union premium based on the sample of union leavers, compared to the group of union stayers. Separate annual estimates of \( \tau_{01} \), \( \tau_{10} \), and \( \tau_{11} \) are possible, but entail considerable loss in efficiency and a high variability in estimates owing to the small number of union joiners and leavers in each panel.

Table 4 provides regression results for the longitudinal sample of 3,654 drivers, each observed in consecutive years (the periods 1983-84 through 1989-90). Three sets of union premium estimates are presented. The first column presents a 1984-90 premium estimate based on a wage levels equation whose observations are the second year for each driver; the second column presents a longitudinal estimate based on the coefficient of \( \Delta UN \) in a wage change equation (eq. 3), thus restricting the estimated premium to be equivalent for joiners and leavers; and the third column presents alternative longitudinal estimates based separately on union joiners and leavers (eq. 4). The levels equation includes the same set of control variables included previously, plus year dummies. The wage change equations include variables measuring changes in schooling, experience squared, marital status, and sector, as well as dummies for period and region (the union coefficients are invariant to inclusion of the dummies). The control variables provide
virtually no explanatory power. And as seen by the low $R^2$s in the change equations, union status change (not surprisingly) explains a tiny portion of the year-to-year change in individual wages.

The estimate of the 1984-90 union premium for this sample of drivers, based on the levels equation shown in the first column, is 0.279 ($|t| = 22.05$), consistent with evidence presented in Table 1. Longitudinal estimates of the union premium are substantially lower. The wage change results presented in the second column provide a union premium estimate of only 0.049 (2.51), about a sixth as large as the levels estimate. The less restrictive specification in column three permitting alternative estimates of the union premium based on union joiners and leavers, $\tau_{01}$ and $-(\tau_{10} - \tau_{11})$, respectively, produces estimates of 0.051 (1.74) for joiners and .053 (1.82) for leavers.

In levels analysis not shown, union joiners are found to earn more than nonunion drivers even in the initial year when both are nonunion, and union leavers are found to earn more than nonunion drivers in the second year when both are nonunion. These results, along with the longitudinal estimates presented in Table 4, are consistent with the existence of a queue of qualified drivers for union jobs, with those selected being more able than the average nonunion driver.

How much of the union wage advantage represents a rent, and how much is a compensating premium for labor quality? There is no unique decomposition of the union-nonunion wage differential into quality and rent components. But if it is assumed that there is no measurement error and that union status change is exogenous, or more specifically, that drivers who are union members for either or both of the years have homogeneous labor quality for given measured characteristics (other than union status), such a division can be approximated by comparing the union premiums estimated above for joiners and leavers, to the observed differential in the levels equation.$^{10}$ Let 0.053 be our best estimate of the longitudinal union premium, based on the average of the estimated wage premium for union joiners and leavers in equation 3. This is only 19% of the observed levels premium (.053/.279=.190). Thus, the estimates imply that much of

$^{10}$The direction of bias is indeterminate if union status change is endogenous. For example, if union leavers voluntarily leave union for nonunion jobs, there will be a downward bias in the union premium estimate for leavers since unusually high wage opportunities increase the probability of voluntary job change. However, involuntary union job loss is likely among drivers in companies with particularly high wage scales, leading to estimates of bias in the opposite direction since these drivers would experience unusually large wage decreases.
the observed union-nonunion wage differential observed for truck drivers is not a rent but, rather, reflects a compensating premium for driver quality.

As is widely recognized (e.g., Freeman 1984; Griliches and Hausman 1986; Card 1991), measurement error in the union variable can cause substantial downward bias in longitudinal estimates of the union premium; the bias be larger the smaller the number of union changers and the larger the proportion of the sample with incorrectly classified union status. There are two principal sources of measurement error in the union variable – inaccurate answers given by survey respondents or recording errors by the Census, and incorrect allocations by the Census among workers without responses to the union question. The proportion of inaccurate responses or recording errors on the union question is reasonably small (2.5%-3.0%) and has little effect on levels estimates of the union premium (Mellow and Sider 1983; Card 1991). Its effect on the magnitude of longitudinal estimates is less certain but can be substantial. Freeman (1984) provides a range of estimates, but assumes that matched firm records are accurate (see Card, 1991, on this issue) and does not account for the fact that serially correlated measurement error lessens the downward bias (for analysis of this issue, see Bound and Krueger, 1991). Measurement error is increased here by focusing on a single occupation, since occupational mobility is a source for some of the true variation in union status, and by use of brief two-year periods, since the noise-to-signal ratio is higher the shorter the time period. On the other hand, driving skills are largely occupation rather than firm specific, so drivers can change employers and union status without changing occupation. Unknown is whether truck drivers are more or less likely than are average workers to report incorrectly their union status.

Approximately 2% of workers have their union status allocated by the Census through a matching of non-respondents to responding workers with similar characteristics. Among the sample of union status changers, a higher proportion will have had their union status allocated, since an incorrect allocation is likely to produce a measured changer. Unfortunately, the allocation flag indicating whether union status is allocated is not contained on the 1983-88 microdata earnings file tapes provided to the BLS by the Census; they are available beginning in 1989. How serious is the measurement error and bias resulting from incorrect Census allocations? We are not sure. Examining the 1989-90 panel, 16% (8 out of 50) of the union joiners and leavers had their union status allocated in either 1989 or 1990. Longitudinal union
premium estimates for the 1989-90 panel, however, are similar in change equations with and without the workers whose status is allocated in at least one of the years (the coefficients on $\Delta UN$ are .118 and .114, respectively). Separate estimates for joiners and leavers are not symmetric in the 1989-90 sample and display greater sensitivity to exclusion of drivers with allocated union status.\textsuperscript{11}

Although the seriousness of measurement error in the union change variable cannot be known with certainty, the large difference in estimates between the longitudinal and levels equations (.053 vs. .279) makes us confident that a substantial portion (half may be a good guessimate) of the union-nonunion differential for truck drivers is a compensating quality premium. For example, suppose that measurement error is so severe that longitudinal estimates of the quality-adjusted union premium are biased downward to a third of their true value. Even then, over 40% of the estimated cross-section union-nonunion wage differential (already adjusted for measurable wage determinants) can be attributed to unmeasured quality differences ($1 - [(3 \times .053)/.279] = .430$). The results support the expectation that employers offset some of the union-nonunion wage differential through selective hiring; both driver quality and wages are higher than they would be absent unionization. Although the pure rents accruing to unionized truck drivers cannot be estimated with accuracy, the continuing decline in union density during 1983-90 suggests that these rents remain substantial.

The previous longitudinal analysis examined wage changes among workers who are truck drivers in adjacent years. The data set also allows an examination of wage changes among workers changing occupations – from truck driver in year 1 to another occupation, or from another occupation in year 1 to driver in year 2. Such a sample is constructed using the same sampling criteria used to construct the previous panel of drivers. Estimates from this sample allow measurement of the gains and losses associated with changes into and out of both truck driving and union membership.

The CPS earnings files for 1983-90 allow construction of a panel of 1,473 workers who are truck drivers in year 1 (survey rotation 4) and not in year 2 (rotation 8), and of 1,580 workers who are not drivers

\textsuperscript{11} Prior to going to press, I examined the 1991 CPS earnings files and constructed a 1990-91 panel of drivers. The panel contains 57 union changers, 11 of whom had union status allocated in at least one of the years. Estimates from change equations including and excluding drivers with allocated status, using the combined 1989-90 and 1990-91 panels, produce coefficients on $\Delta UN$ of .122 and .104, respectively. This limited evidence does not indicate any downward bias in longitudinal estimates arising from the Census allocation of union status.
in year 1 and are in year 2. In both samples, about 12% have changed union status, with roughly an even split between joiners and leavers. Preliminary analysis (not shown) indicated that for both samples of workers, union premium estimates based alternatively on union joiners and leavers are highly similar. For ease of exposition, results therefore are presented using a specification with the $\Delta \text{UN}$ variable and for a panel ($N = 6,712$) pooling the three longitudinal samples: truck drivers in adjacent years ($TT = 1$, $N = 3,659$), drivers in year 1 and not in year 2 ($TN = 1$, $N = 1,473$), and not a driver in year 1 and a driver in year 2 ($NT = 1$, $N = 1,580$). The following regression results are obtained from a wage change equation for the combined 1983-84 through 1989-90 panels ($|t|$ in parentheses):

$$\Delta \ln W_t = .001TN - .005NT + .048TT \Delta \text{UN} + .131TNDUN + .116NT \Delta \text{UN}, \ N = 6,712 \ R^2 = 0.016$$

(0.10) (0.43) (2.34) (4.59) (4.25)

The wage change regression also includes variables measuring changes in schooling, experience squared, marital status, and presence in the for-hire trucking sector, as well as dummies for region and period. The omitted reference group represents individuals who are truck drivers in both years ($TT = 1$). The absence of wage gains or losses associated with leaving or entering truck driving (the coefficients on $TN$ and $NT$, respectively) is of interest. Small gains or losses associated with occupational change could reflect relatively low (i.e., quickly learned) investment in training by drivers. Alternatively, the loss in human capital resulting from occupational change could be substantial since driving skills are to a considerable degree occupation specific, but this loss may be offset by gains associated with voluntary (i.e., non-random) job change.

The estimated longitudinal union premium for the trucker-trucker sample of drivers is about 5%, as seen previously in Table 4. But for the samples of drivers changing occupation, estimated wage changes associated with the change in union status are 12%-14%. Although higher than the estimate from the TT sample (perhaps due in part to lower measurement error bias among occupation changers), the longitudinal premiums remain far lower than estimates of the differential in levels, which are approximately .30 for all three groups of workers (levels results for the occupational changers are not shown). Although the exact magnitude of the union premium associated with union status change is again in doubt, the large difference
in premium estimates between the levels and change equations make us confident that much of the observed union-nonunion wage differential in fact reflects labor quality differences in drivers.

**Interpretation and Conclusions**

Deregulation in the trucking industry provides an unusually rich opportunity to study the creation and dissolution of economic rents associated with entry, price, and operating restrictions. Past studies have provided strong evidence that rents arising from regulation were reflected not only in higher profitability and in the market value of operating rights, but also in the form of higher labor costs. Previous analyses, however, were limited by small sample sizes of truck drivers, and the concomitant high variability of estimates, during the post-regulation period. This paper examines wage determination among truck drivers using the 96 monthly samples from the Current Population Surveys for 1983-90, thus facilitating more current and precise estimation of union density and union-nonunion wage differentials during the deregulatory period, as well as the use of longitudinal analysis to examine wage changes for union joiners and leavers.

The evidence is broadly supportive of the conclusions reached previously by Rose and Hirsch. In the previously regulated for-hire sector of the trucking industry, union density fell continuously from about 60% during the regulatory period to about 25% by 1990. Union-nonunion wage differentials (in log differences) fell from about 0.40 during 1973-78 to a range of .26-.33 during 1983-90. Union-nonunion wage differentials among drivers in the unregulated private carrier sector were about 0.30 during the earlier period, but have fallen little during the 1980s and are roughly similar in magnitude to union premiums in the for-hire sector.

Longitudinal analysis is performed on seven panels of truck drivers for whom data are available for 1983-84 through 1989-90. Wage changes among union joiners and leavers allow inferences to be made about unmeasured quality differences between union and nonunion drivers. Union premium estimates based on longitudinal data are considerably lower than are union-nonunion differentials estimated from levels data. Based on results from the panel data, we conclude that a substantial portion of the union-nonunion wage differential is a compensating premium for unmeasured driver quality, consistent with the hypothesis that union employers select and retain relatively high-quality drivers.
Data from supplemental sources also is examined to shed further light on this paper's results. Corroborative evidence on unionization and wages among truck drivers is available in private surveys of long-haul truckload (TL) drivers in 1980 and 1987.\(^{12}\) The National Motor Transportation Data Base contains survey data collected from drivers at 20 U.S. truckstop sites, with approximately 17,000 observations in 1980 and 19,000 in 1987. These surveys do not sample a representative cross-section of drivers but, rather, a fairly narrow segment of the total universe of drivers. The survey primarily includes long-haul truckload drivers along major long-haul (rail competitive) corridors, a large proportion of whom are involved in irregular-route common carriage. The survey largely excludes short-haul TL drivers who generally run on a straight through nonstop schedule, and drivers in the highly unionized less-than-truckload (LTL) segment who drive terminal to terminal.

Union density among those drivers surveyed (primarily for-hire TL) fell sharply during the deregulation period, from 19.4% in 1980 to only 2.6% in 1987 (coverage among the tiny regular-route LTL common carriage component of this sample fell from 74.4% to 56.4% during this period). The cost differential between union and nonunion trucking has exhibited little change during the period. Estimated costs per mile were 22.3% higher among union than nonunion trucking in 1987, as compared to 20.5% in 1980. Consistent with our finding that part of the union wage premium reflects a quality differential, the survey finds that union drivers are older, have more trucking experience, and have greater tenure with their present firm than do nonunion drivers. Moreover, the gap between union and nonunion drivers in these three characteristics widened between 1980 and 1987.

The relationship between truck drivers’ union status and job tenure also is explored more fully using the CPS Pension Supplements attaching to the 1979, 1983, and 1988 May/June public use tapes. The Supplements include a question asking number of years with the current employer. Comparison of union and nonunion years of tenure may be a reasonable proxy for relative years of driver experience.\(^{13}\) Taking the ratio of years of tenure on current job to years of potential experience (age-schooling-5), we obtain a

\(^{12}\) Information on these surveys has been kindly provided and summarized by William Linde, Intermodal Policy Division, American Association of Railroads.

\(^{13}\) Union workers generally have lower permanent turnover rates than do nonunion workers. If, on the one hand, this generalization is also correct for truck drivers, the ratio of tenure to years of total driver experience will be higher for union than for nonunion drivers. On the other hand, the business failure of numerous union trucking operations during the 1980s would lead to a bias in the opposite direction.
value of .383 for male union drivers in the combined 1979, 1983, and 1988 Pension Supplements, as compared to only .277 for the sample of nonunion drivers. In 1988 union drivers averaged 10.95 years of tenure on the current job and 25.76 years of potential experience. The comparable figures for nonunion drivers are 5.17 years of tenure and 20.44 years of potential experience. The tenure-experience ratio exhibited a small increase among union drivers between 1979 and 1988, while showing little change among nonunion drivers. The tenure results reinforce and help explain the earlier finding that a substantial portion of the union-nonunion wage differential appears to be a compensating premium for higher quality. Union drivers possess greater tenure and (by inference) driving experience than do nonunion drivers of similar age.

Deregulation in the trucking industry has limited the ability of firms to continue maintenance of costly unionized trucking operations, except where high wages can be offset by productivity advantages or where union operations are partially protected from nonunion competition as is the case in many LTL markets. Evidence from the CPS indicates that the cost pressures brought to bear upon union wages have produced a narrowing in what once was a very large union-nonunion wage differential in the for-hire sector of the industry. For the Teamsters to have maintained union density at levels obtained during the regulatory period, however, the union would have had to agree to wage and benefit concessions far greater than they did. Evidence clearly indicates that the rank-and-file were unwilling to support many of the concessions actually recommended by the Teamster leadership, while the even greater concessions that were necessary to protect employment of union drivers were not politically feasible. Currently, the trucking industry is characterized by a union sector much smaller than during the regulatory period and in which union drivers continue to realize a sizable albeit reduced wage premium, a portion of which is offset by higher quality among union drivers.
Data Appendix: Construction of the Longitudinal Sample from the CPS Outgoing Rotation Files

Households are included in the CPS for 8 months – 4 consecutive months in the survey, followed by 8 months out, followed by 4 months in. The CPS contains household identification numbers (ID), but not individual identifiers. Since January 1983, the CPS has included the earnings supplement questions (weekly earnings, hours, union status, etc.) each month, rather than just in the May surveys. But only the outgoing rotation groups 4 and 8 are asked these questions. Individuals potentially can be identified for the same month in consecutive years; that is, individuals in rotation 4 in year 1 can be matched to individuals in rotation 8 in year 2. The CPS ran a test sample from July-September 1985 in order to implement new population weights and area identifiers. As a result, rotation 4 households interviewed in July 1984 through September 1985 were not reinterviewed a year later in 1985 and 1986. Therefore, the longitudinal sample for 1984-85 is approximately one-half, and the sample for 1985-86 approximately one-quarter, the normal size.

The longitudinal file was created in the following manner. Separate data files were created separately for males and females, and for pairs of years (rotation 4 1983, and rotation 8, 1984; rotation 4, 1984, and rotation 8, 1985; etc.). Within each file, individuals were sorted as appropriate on the basis of ascending and descending household ID, year, and age. To be considered an acceptable matched pair, a rotation 8 individual had to be matched with a rotation 4 individual with identical household ID, identical survey month, and an age difference between 0 and 2 (since surveys can occur on different days of the month, age change need not equal 1). Several passes were necessary because a single household may contain more than one male or female pair. Checks were provided to insure that only unique matches were selected. For each rotation 8 individual, the search was made through all rotation 4 individuals with the same ID to make sure there was only 1 possible match; the file was resorted in reverse order and each selected rotation 4 individual was checked to insure a unique rotation 8 match. As uniquely matched pairs were identified they were removed from the work file. Incorrect changes in the variables marital status, veteran status, race, and education (e.g., a change in schooling other than 0 or 1, a change from married to never married, etc.) were used to delete “bad” observations in households where there were multiple observations and ages too close to separate matched pairs. Several passes at the data were made. In
households where two pairs of individuals could be separated based on a 1 year but not the 0 to 2 year age change, a 1 year criterion was used. If a unique pair could not be identified based on these criteria, they were not included in the data set (e.g., four observations with two identical pairs, or three individuals with two possible matches using the 0 to 2 age change criterion).

The match rate in the longitudinal analysis is 64%. A match rate of 64.3% is obtained based on the percentage of the 7,883 rotation 4 truckers for 1983-89 included in the levels analysis (this excludes observations during July 1984 - September 1985 for reasons stated above) who also are included in the subsequent longitudinal analysis – 5,068 rotation 8 workers in the trucker-trucker and trucker-nontrucker samples during 1984-90 (the combined TT and TN samples shown in the text were $N = 5,132$ and included workers with 30 or more hours in at least one of the years; 64 of these had fewer than 30 hours in year 1 and were not included in the levels analysis or the calculation of the match rate). Alternatively, a match rate of 64.4% is obtained based on the percentage of the 8,066 rotation 8 truckers for 1984-90 included in the levels analysis who also are included in the longitudinal analysis – 5,191 rotation 4 workers in the trucker-trucker and nontrucker-trucker samples during 1983-89. The principal reasons that matches cannot be made are if a household moves (thus changing the household ID), if an individual moves out of a household (young workers are less likely to be matched), if an individual drops out of the labor market or fails to meet other sample selection criteria, or if the Census is unable to reinterview a household and/or receive information on the individual. Note that the match rate reported here is similar to the match rate of 68.8% for the 1987-88 CPS reported by Card (1991) using a broader-based sample and a less stringent probabilistic matching algorithm obtained from the Bureau of Labor Statistics.


Table 1. Mean Real Wages, Union Density, and Union Wage Premium Estimates for Male Truck Drivers, 1973-90

<table>
<thead>
<tr>
<th>Year</th>
<th>Drivers in For-Hire Sector</th>
<th>Drivers in Private Carrier Sector</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>N</td>
<td>DEN</td>
</tr>
<tr>
<td>1973-78</td>
<td>1,533</td>
<td>.599</td>
</tr>
<tr>
<td>1979a</td>
<td>175</td>
<td>.566</td>
</tr>
<tr>
<td>1980a</td>
<td>94</td>
<td>.564</td>
</tr>
<tr>
<td>1981a</td>
<td>84</td>
<td>.607</td>
</tr>
<tr>
<td>1984a</td>
<td>79</td>
<td>.304</td>
</tr>
<tr>
<td>1985a</td>
<td>111</td>
<td>.288</td>
</tr>
<tr>
<td>1983</td>
<td>1,034</td>
<td>.432</td>
</tr>
<tr>
<td>1984</td>
<td>1,158</td>
<td>.375</td>
</tr>
<tr>
<td>1986</td>
<td>1,093</td>
<td>.319</td>
</tr>
<tr>
<td>1987</td>
<td>1,136</td>
<td>.276</td>
</tr>
<tr>
<td>1988</td>
<td>1,161</td>
<td>.300</td>
</tr>
<tr>
<td>1989</td>
<td>1,166</td>
<td>.269</td>
</tr>
</tbody>
</table>


CPS = Current Population Survey; DEN = union density, $W$ = usual weekly earnings divided by usual hours worked (in 1990:4 dollars, adjusted by the quarterly Personal Consumption Expenditure component of the GNP implicit price deflator); $\ln W_u - \ln W_n$ represents the unadjusted log wage differential between union and nonunion drivers; and $\theta$ represents the regression coefficient on the union membership dummy variable UN for the corresponding period. All regressions include a constant; years of schooling, experience, and experience squared; and dummies for union, nonwhite, veteran, marital status, and region (8).
Table 2. Mean Real Wages, Union Density, and Union Wage Premium Estimates for Male Nontruck Operatives and Blue-Collar Manufacturing Workers, 1983-90

<table>
<thead>
<tr>
<th>Year</th>
<th>Nontruck Operatives</th>
<th>Blue-Collar Manufacturing Workers</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>N</td>
<td>DEN</td>
</tr>
<tr>
<td>1984</td>
<td>8,191</td>
<td>.400</td>
</tr>
<tr>
<td>1987</td>
<td>7,735</td>
<td>.348</td>
</tr>
<tr>
<td>1988</td>
<td>7,276</td>
<td>.342</td>
</tr>
</tbody>
</table>

The data are derived from the earnings microdata files from BLS containing the 12 monthly CPS surveys. N = sample size; DEN = union density, W = usual weekly earnings divided by usual hours worked (in 1990:4 dollars, adjusted by the quarterly Personal Consumption Expenditure component of the GNP implicit price deflator); ln W_u – ln W_n represents the unadjusted log wage differential; and θ represents the regression coefficient on a union membership dummy variable UN for the corresponding period. All regressions include a constant; years of schooling, experience, and experience squared; and dummies for union, nonwhite, veteran, marital status, and region (8).
Table 3.

<table>
<thead>
<tr>
<th>Year</th>
<th>θ</th>
<th>Month</th>
<th>θ</th>
<th>θ̄</th>
<th>Month</th>
<th>θ</th>
<th>θ̄</th>
</tr>
</thead>
<tbody>
<tr>
<td>1983</td>
<td>.313</td>
<td>August</td>
<td>.068</td>
<td>.70</td>
<td>December</td>
<td>.502</td>
<td>6.18</td>
</tr>
<tr>
<td>1984</td>
<td>.292</td>
<td>January</td>
<td>.045</td>
<td>.44</td>
<td>October</td>
<td>.377</td>
<td>3.70</td>
</tr>
<tr>
<td>1985</td>
<td>.262</td>
<td>May</td>
<td>.106</td>
<td>1.10</td>
<td>June</td>
<td>.525</td>
<td>5.23</td>
</tr>
<tr>
<td>1986</td>
<td>.274</td>
<td>May</td>
<td>.078</td>
<td>.74</td>
<td>December</td>
<td>.462</td>
<td>4.84</td>
</tr>
<tr>
<td>1987</td>
<td>.266</td>
<td>August</td>
<td>.103</td>
<td>1.20</td>
<td>February</td>
<td>.503</td>
<td>3.34</td>
</tr>
<tr>
<td>1988</td>
<td>.326</td>
<td>June</td>
<td>.188</td>
<td>2.00</td>
<td>November</td>
<td>.464</td>
<td>4.14</td>
</tr>
<tr>
<td>1989</td>
<td>.295</td>
<td>June</td>
<td>.074</td>
<td>.54</td>
<td>July</td>
<td>.499</td>
<td>5.53</td>
</tr>
<tr>
<td>1990</td>
<td>.269</td>
<td>March</td>
<td>-.064</td>
<td>.35</td>
<td>November</td>
<td>.412</td>
<td>4.76</td>
</tr>
</tbody>
</table>

θ is the regression coefficient on UN, and θ̄ is the unweighted annual mean of monthly θ’s. Sample sizes are approximately 100 drivers per month. All regressions include a constant; years of schooling, experience, and experience squared; and dummies for union, nonwhite, veteran, marital status, and region.
Table 4. Longitudinal Analysis of Union Wage Effects, Combined Panels for 1983-84 through 1989-90.

<table>
<thead>
<tr>
<th>Union (Variable)</th>
<th>Mean</th>
<th>ln W</th>
<th>Δ ln W</th>
<th>Δ ln W</th>
</tr>
</thead>
<tbody>
<tr>
<td>Union Membership (UN)</td>
<td>.350</td>
<td>.278</td>
<td>–</td>
<td>–</td>
</tr>
<tr>
<td>Change in UN</td>
<td>-.004</td>
<td>–</td>
<td>.049</td>
<td>–</td>
</tr>
<tr>
<td>DUN</td>
<td></td>
<td></td>
<td>(2.51)</td>
<td></td>
</tr>
<tr>
<td>Union joiner</td>
<td>.044</td>
<td>–</td>
<td>–</td>
<td>.051</td>
</tr>
<tr>
<td>UN01</td>
<td></td>
<td></td>
<td>(1.74)</td>
<td></td>
</tr>
<tr>
<td>Union leaver</td>
<td>.048</td>
<td>–</td>
<td>–</td>
<td>-.047</td>
</tr>
<tr>
<td>UN10</td>
<td></td>
<td></td>
<td>(1.66)</td>
<td></td>
</tr>
<tr>
<td>Union stayer</td>
<td>.306</td>
<td>–</td>
<td>–</td>
<td>.006</td>
</tr>
<tr>
<td>UN11</td>
<td></td>
<td></td>
<td>(.45)</td>
<td></td>
</tr>
<tr>
<td>Nonunion stayer</td>
<td>.602</td>
<td>–</td>
<td>–</td>
<td>–</td>
</tr>
<tr>
<td>UN00</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

\[ R^2 \] | .255 | .007 | .007 |
\[ N \] | 3,659 | 3,659 | 3,659 |

The first regression is a levels equation with second-year observations for each driver and ln W as the dependent variable; included variables are years of schooling, experience, and experience squared; and dummies for union, nonwhite, veteran, marital status, trucking sector, region, and year. The second and third regressions are wage change equations with Δ ln W the dependent variable (mean equals .013), corresponding to equations (3) and (4) in the text. Included variables are changes in schooling, experience squared, marital status, and sector, and dummies for region and 2 panel year periods. Nonunion stayer (UN00) is the omitted category in the third column. | is given in parentheses.