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WORKERS' COMPENSATION RECIPIENCY IN UNION AND NONUNION WORKPLACES

by

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Abstract

This study estimates union effects on workers' compensation indemnity claims during 1977-92, based on individual panel data constructed from the March Current Population Survey. Union members were substantially more likely to receive workers' compensation benefits than were similar nonunion workers, and they were more sensitive to variation in benefit levels and waiting periods. The authors suggest that differences in union, as compared to nonunion, workplaces arise because workers are provided with information from their union representatives, supervisors are more likely to inform injured workers about workers' compensation filing procedures and less likely to discourage workers from filing claims, workers are less likely to fear being penalized for filing claims, and management has less discretion and ability to monitor workers and penalize them for questionable claims. The findings suggest that communication of relevant information to workers is an important determinant of workers' compensation recipiency.

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Workers' compensation provides a form of no-fault insurance covering workplace injuries. Benefits are paid to workers injured on the job, regardless of fault, and employers receive partial protection from law suits or further liability. The U.S. workers' compensation system consists of separate state laws together with federal laws covering federal longshoring workers. Workers' compensation is compulsory in all but three states (New Jersey, South Carolina, and Texas); even in states where it is elective, most employers choose to be covered to receive limited liability. Total payments from workers' compensation are sizable, amounting to 44.1 billion dollars in 1992, of which 40.9 billion was from state and 3.2 billion from federal programs. Of that total, 41 percent (17.9 billion) was for hospital and medical payments. Payments for workers' compensation exceed those for state and federal unemployment insurance, food stamps, supplemental security income (SSI), veteran programs, or housing programs (*Statistical Abstract of the United States 1995*, Table 585).

There is an extensive literature considering the incentive effects of workers' compensation, much of it focused on estimating the responsiveness of injuries, claims, or duration to benefit levels and to differences in program parameters such as the benefit waiting period. Most studies have examined incentive effects using aggregate state or state-by-industry data, either for single or multiple time periods, or micro data from a single or selected states.

This study examines union-nonunion differences in workers' compensation recipiency. We argue that unionization should affect workers' compensation recipiency through the provision of information and because of union-nonunion differences in workplace governance. Although union status is widely believed to influence claims, the topic has not been studied thoroughly because of an absence of readily available microdata on union status and workers' compensation recipiency. We use a large database on union and nonunion workers developed from matched panels constructed from the March 1977-93 Current Population Surveys (CPS). Our study extends recent work by Krueger (1990), one of the few studies that uses a representative national data set on individuals. Krueger's analysis, however, does not consider the effects of unionization. By using data on individuals and union status, we are able to identify the direct union effect on workers' compensation recipiency, independent of the correlation of union status with industry, job hazards, and other determinants of workers' compensation.

Determinants of Workers' Compensation Indemnity Claims

The Provisions of State Workers' Compensation Laws

Workers are eligible for workers' compensation when they are disabled by injury or illness arising out of or in the course of employment. Employers are liable regardless of fault, but may dispute the severity of an injury or illness and challenge whether it is work-related. Workers' compensation claims are initially classified as "temporary total" cases, meaning that the worker's disability precludes working but is temporary. Temporary total claims constitute roughly 70 percent of all indemnity claims (Krueger, 1990, p. 76), although a far lower percentage of total costs. Disabilities that preclude working and persist beyond some time period are reclassified as "permanent total." Some states also have "temporary partial" benefits for workers who can continue work but at reduced earnings.

The typical formula for temporary total (and permanent total) benefits is two- thirds of a worker's pre-disability before-tax average wage (or in some states a higher percentage of the after-tax wage). All states have maximum weekly payments. These amounts vary considerably across states, and many states set the maximums as a percentage of the state average weekly wage (most percentages are in the range 67 to 150 percent of the average wage). Most states also have minimum payments, typically at 20 percent or more of the state average weekly wage (some states do not allow payments greater than the pre-disability weekly wage).

An important policy feature of workers' compensation is the "waiting period" necessary to establish eligibility for benefits. States typically have waiting periods of three or seven days (the range in our data set across all states and years is 2-14 days) before workers receive benefits. In the event that an injury extends beyond a defined "retroactive period" (typically one to several weeks), benefits are paid retroactively to the date of injury. Medical expenses are covered immediately following injury (i.e., no waiting period), although in some instances actual time to reimbursement may be lengthy. Workers with disabilities that are permanent but only partially disabling (e.g., loss of a limb) in most states receive pre-designated *scheduled* benefits corresponding to the disability. Disabilities not on the schedule are handled on an ad hoc basis.

In our study, potential benefits are calculated based on the formula for temporary total benefits and a workers' pre-disability weekly earnings. This is necessary given our inability to differentiate among the types of workers' compensation cases. Three considerations suggest that it should be a good measure of *expected* benefits: most cases are temporary total, permanent total benefits are usually calculated similarly, and there is a high correlation in benefit generosity between states' temporary total and other benefit levels.

The workers' compensation insurance system varies considerably across jurisdictions. Six states require that employers insure through a state-operated insurance system. Twenty states operate a state system but permit insurance through private insurance companies or self-insurance. In most states, the typical small employer purchases private insurance and large employers self-insure. Insurance companies set rates based on a combination of manual rates, which vary on the basis of rather detailed industry/occupation breakdowns, and experience rating. Small and newer firms will have costs based primarily on manual rates, while large established firms will have full or close to full experience rating (Chelius and Smith, 1983, provide an analysis of the effects of experience rating on injuries).

Workers' compensation insurance affects the incentives of employers and employees in different ways. Higher costs to employers for a given injury experience record should reduce both injuries and claims. Firms can increase the safety of the workplace through expenditures on plant design, equipment, production organization, training, and monitoring. Employers should expend resources on safety to the point where the marginal cost of reducing injuries (or reports of injuries) equals the expected marginal benefits from future workers' compensation insurance premium or self-insurance cost decreases and lower equilibrium wages owing to a safer workplace.

Employee response to higher benefits may increase workers' compensation claims. The number of successful claims is determined by the frequency and severity of injuries and the probability of a claim given an injury. Moral hazard problems arise in several dimensions. The availability of partial earnings replacement and reimbursement for medical expenses may increase the risk of injury by reducing care used by workers. For injuries that would permit the continuation of work, employees are more likely to file an indemnity claim when benefits replace a higher proportion of earnings. Even for an injury that precludes working, claims may not be made if benefits are low. The availability of benefits may also produce fraudulent claims – claims for injuries or disabilities not related to employment, or claims that overstate the severity of injuries.[1]

The literature generally indicates a positive relationship between workers' compensation indemnity claims and increases in benefits, suggesting that changes in behavior owing to worker incentives exceed those resulting from employer incentives.[2] Benefit elasticities, for example, have been found to be larger when claims rather than injuries are used as the outcome variable (e.g., Butler, 1994).[3] Further evidence supporting this conclusion includes the finding of a positive relationship between workers' compensation benefits and non-fatal injuries, but a negative relationship between benefits and fatalities (Ruser, 1993; Moore and Viscusi, 1989, 1990). In short, in response to higher benefit levels, workers are most likely to adjust reporting behavior, while employers respond by making the workplace safer.

The empirical literature finds a negative relationship between workers' compensation indemnity claims and the length of the waiting period (e.g., Krueger, 1990; Butler, 1994). The effect of the waiting period on claims is due to truncation on the basis of severity, with less severe disabilities not receiving compensation, and a decreased value of benefits received. The length of the retroactive period appears to have no significant effect on the number of claims (Krueger, 1990).

Worker and Job Characteristics

Of particular importance in our statistical analysis is the control for job risk through inclusion of an industry injury rate variable, industry dummies, establishment size, and measures of occupational working conditions. Such an approach permits us to more accurately measure the causal effect of individual union status on the probability of a workers' compensation claim.

Worker characteristics are likely to affect indemnity claim rates to the extent that they proxy

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[1] Reports of injuries tend to be higher on Monday mornings then at other times of the week and the ratio of "hidden" to "non-hidden" (e.g., sprains and back injuries versus lacerations) injuries is highest on Mondays (Smith, 1989). Since workers' compensation covers the medical costs of injuries, Card and McCall (1996) argue that the Monday effect should be more pronounced among workers without health insurance. Although they do observe more injuries on Monday, they find no difference among workers more or less likely to be covered by health insurance (they predict rather than directly observe insurance coverage).

[2] Surveys of the literature on workers' compensation are included in Butler (1994), Krueger (1990), Moore and Viscusi (1990), Ehrenberg (1988), and Chelius (1983). We do not examine benefit *duration* in this study. For a recent analysis, see Meyer, Viscusi, and Durbin (1995). In a series of recent papers, Fishback and Kantor examine the evolution of state workers' compensation laws; see, for example, Fishback and Kantor (1994).

differences in job risks or in costs and benefits facing workers. Variables measuring worker earnings, schooling, age, gender, marital status, part-time status, and occupational skills control for the opportunity cost of foregone earnings in the event of a job absence, as well as for worker experience, skills, and attitudes toward risk-taking that affect behavior. Skill-related variables should be negatively related to workers' compensation recipiency for several reasons. Because more highly skilled workers have higher earnings and are more likely to have their benefits capped, their opportunity cost of missing work is high. Skilled workers also tend to take some of their higher potential compensation in the form of a safer work environment, hence lowering injury rates. Such workers are also likely to take greater safety precautions and use better judgement, both because they have better information and tend to have lower discount rates (and hence more schooling).

The relationship between worker age and workers' compensation costs is not straight-forward. On one hand, we expect age to be negatively correlated with accidents to the extent that worker skills increase with experience, risk-taking behavior declines with age, and older workers are less likely than young workers to be matched to dangerous jobs. On the other hand, deterioration of physical skills with age may lead to more injuries, and older workers may be less able to withstand a given level of physical demands without injury. Although in our work we do not measure the severity or duration of workers' compensation indemnity claims, it is likely that older workers have a longer duration of injury, conditional on a claim, and have a lower probability of returning to work owing to a decline in the discounted present value of future earnings as one ages.

Differences in workers' compensation claims correlated with gender, race, marital status, and part-time status reflect a variety of factors, some of which we discuss in the context of the empirical evidence presented below.

The Role of Union Status on the Recipiency of Workers' Compensation

There are a number of reasons for expecting a positive association between union status and receipt of workers' compensation. Union workers are more likely than nonunion workers to be in jobs with dangerous or

[3] Kniesner and Leeth (1989) and Butler (1994) examine differences between the use of claims and injuries as outcome variables, as well as discuss relevant literature.

unpleasant working conditions (Duncan and Stafford, 1980; Leigh, 1982; Worrall and Butler, 1983). We can control for that difference, however, by including the industry injury rate, occupational working conditions, and industry of employment in the estimating equation. An important advantage of individual data over industry data is that with the use of the former, one can provide detailed industry and occupation controls and examine whether there are union-nonunion differences in claim rates *within* sectors. By contrast, the use of industry-level data makes it difficult to distinguish industry effects from union effects.

Potentially important is the independent role of union voice and union-nonunion differences in workplace governance (see Freeman and Medoff, 1984).[4] Union workplaces tend to be more highly structured, with contract terms and the relations of exchange formalized explicitly and implicitly. Compensation for workplace injuries is likely to be understood by both workers and management as a right. In the event of an injury, either workers already are aware of the availability of workers' compensation benefits or they are quickly made aware by co- workers, shop stewards, or supervisors. Managers are not likely to discourage legitimate claims for workers' compensation, since such actions would be known to the union and could constitute a grievance. Moral hazard may be a more serious problem in unionized establishments if management has less discretion and ability to monitor and penalize workers for questionable claims. Some possible penalties – for example, differential treatment with respect to future promotions and raises, or reassignment of responsibilities following the return to work.[5] – are less likely in union than in nonunion workplaces, since in the former there exist grievance procedures and the terms of compensation are typically determined contractually and based on job position and seniority, rather than being set on an individual basis. In short, individual union workers are unlikely to be penalized or to believe they will be penalized in the event they file a workers' compensation claim.

In contrast to workers in a unionized establishment, those in nonunion workplaces are less likely to know or be informed of the availability of and procedures for obtaining workers' compensation in the event of an injury. Management or supervisors may be less likely to facilitate a workers' compensation claim and, in

[4] Viscusi (1980), among others, has emphasized the importance of information about injury risks in achieving an optimal level of workers' compensation. Here we emphasize the role of unions in providing information to workers and union constraints on supervisors' behavior following occurrence of an injury.
[5] Butler, Johnson, and Baldwin (1995) examine the return to work following disability using Canadian data. They note

some cases, may directly discourage it. Nonunion workers are more likely to be penalized, or to believe they may be penalized, for filing a claim. Managers are not likely to be penalized if they discourage a claim, and individual workers may have little recourse if they persuaded not to file a claim or if they do file it and are subsequently penalized.[6] We are not suggesting that moral hazard is eliminated in nonunion settings. If it were, we would expect a negative claims-benefits relationship. Such a relationship is not observed. But we do expect to observe lower levels of moral hazard and the discouragement of more legitimate claims in nonunion than in union establishments.

Working to reduce workers' compensation claims in unionized workplaces are several factors. First of all, the same union voice that can facilitate a claim in the event of injury may also facilitate through bargaining a safer work environment and training on risk-avoidance. Although evidence indicates that union workplaces are more dangerous than nonunion workplaces, we know of no evidence indicating that the causal effect of unions on workplace safety is negative (e.g., as a partial offset to union wage premiums).[7] In addition, turnover is lower in union workplaces and safety may well increase with employee experience.

Although there is only limited evidence on the effect of union status on workers' compensation claims, that which exists suggests that indemnity claims are higher in union than in nonunion establishments. What is not clear is whether the union effect is causal or, instead, union claims are higher because union jobs are more dangerous. Butler and Worrall analyzed claims data provided by the National Council on Compensation Insurance (NCCI) for 196 state-by-year observations (a subset of 35 states from 1972-78). They found claims positively related to state union density, but stated that "whether this is because unions are formed in more risky places, or because unions facilitate the filing of claims given any level of risk is unclear" (Butler and Worrall, 1983, 587). Elsewhere, examining microdata on about 2,608 individuals from the 1978 Social Security Survey of Disability and Work, Worrall and Butler (1983) confirmed that blue-collar union

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that unions provide protection to workers insuring their ability to return to the same job.

[6] Budd and McCall (1994) use the NLSY and find that among workers eligible for unemployment insurance (UI), union workers in blue-collar jobs are more likely to receive benefits than are nonunion workers. Similar to the arguments above, they emphasize the role of unions in informing eligible workers about the availability of benefits and in facilitating the filing of claims. Unions' informational role affects receipt of UI benefits *following* the loss of work. In the case of workers' compensation, the union role is likely to be even greater since unions affect management and worker behavior that determines whether an injury will be reported and lead to a loss of work, as well as receipt of benefits given the loss of work.

members reported greater exposure to job hazards (see, also, Leigh, 1982; Duncan and Stafford, 1980) and had a greater likelihood of a health condition caused by a job hazard than did similar nonunion workers. Union members were more likely to be employed in risky industries, measured by injury rates or lost workdays. Finally, Krueger and Burton (1990) found that states with higher union density also had higher workers' compensation costs, consistent with a higher claims rate among union workers.[8]

In a study potentially relevant for our analysis, Gleason and Roberts (1993) interviewed 174 Michigan workers' compensation *recipients* in order to explore whether union members perceive the process to be more procedurally fair than do nonunion workers. They found few important differences between union and nonunion recipients. In particular, the two groups had similar perceptions of the attitude of their supervisors and of threats to their job security. Unfortunately, in order to test directly the theses proposed in our paper, one would also need to record the perceptions of workers who do *not* make claims. Most interesting from our standpoint is that union members comprised 58 percent of Gleason and Robert's (very small) sample of recipients.

Taken together, the limited evidence available supports the proposition that workers in union jobs are more likely than those in nonunion jobs to make workers' compensation claims. The difference in claims appears to result partly from greater risk in union jobs. We are not aware of any previous study that simultaneously provides detailed industry controls and good measures of union status and working conditions, thus permitting inferences about the causal role of unions in workers' compensation claims.

Data and Methodology

In this study, we investigate the determinants of workers' compensation indemnity claims, with particular emphasis on the role of union status and state laws. Specifically, we examine the factors associated

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[7] Fairris (1995) explicitly tests this hypothesis, but finds no support.

[8] There has been research on unions and other aspects of workers' compensation. Moore and Viscusi (1990, Ch. 8) and Dorsey and Walzer (1983) conclude that there is a smaller tradeoff between workers' compensation benefits and wages for union than nonunion workers. Moore and Viscusi also address the issue of union endogeneity in the wage equation. Hamermesh and Wolfe (1990) conclude that premiums for injury risk are higher among union than nonunion workers (see, also, Moore and Viscusi, 1990, Ch. 8). Hamermesh and Wolfe (1990, p. S187) interpret this result as support for the notion that unions provide information to workers about what are otherwise poorly perceived risks in the workplace. Butler and Appel (1990, Appendix) consider the effects of state union density on changes in state workers' compensation minimum and maximum benefit payments. They find that union density is positively associated with increases in the minimum, but not with increases in the maximum.

with receipt of workers' compensation income by individual workers in a given year *t*, provided that workers' compensation income was not received in year *t-1*.

The principal data source for our analysis is the March CPS. We follow an approach used previously by Krueger (1990). The March CPS provides information on respondents' income, earnings, and principal occupation and industry in the previous year. Included also is a question on types of income, including a separate category identifying receipt of workers' compensation.[9] The CPS sample design is such that households are included for four consecutive months in a year, followed by eight months out of the survey, followed by four months in. Ignoring for the moment sample attrition, half of the sample in each March survey may be included in the previous year's March survey, while the other half will be included in the next year's March survey. Although the CPS is not designed as a panel data set, the inclusion of household identifiers and individual characteristics allows reliable matching of individuals across most adjacent years (Krueger, 1990; Peracchi and Welch, 1995).

We construct a sample comprising 14 matched panels providing information for the *calendar years* 1976/77- 1980/81, 1982/83-1983/84, and 1985/86-1991/92. The database is constructed from the March CPS surveys for 1977/78 through 1981/82, 1983/84, 1984/85, and 1986/87 through 1992/93 (the March surveys report income from the *previous* year). A data appendix provides a detailed description of how the sample was constructed and the matching of the union status variable to the March CPS.[10]

The choice of years was dictated by data availability. Union status information was not available on a regular basis prior to May 1973. Because of inconsistent household identifiers in the March 1976 and March 1977 surveys (which precluded a match of individuals across years), we chose to begin our panel with calendar years 1976/77, based on the matched March 1977 and 1978 surveys. The 1981/82 panel was excluded due to the absence of information on union status in 1982. The Census ran a special test sample of

[9] Beginning in 1989, information is provided not only on the receipt but also the amount of workers' compensation income. Prior to 1989, such income is not broken out separately.

[10] Union status is reported for the full sample in the May CPS for 1973-80, for a quarter sample (the outgoing rotation groups) in May 1981, and in each month for the outgoing rotation groups beginning in January 1983. There was no union question in 1982. For the years 1977-80, half of the respondents in the March survey potentially are matched to their union status response in May (rotation groups 3, 4, 7, and 8); in 1981 a quarter sample potentially is matched between March and May (rotation groups 4 and 8); and beginning in 1983 the full March sample potentially can be matched with union status reported in their outgoing month (either March, April, May, or June).

households from July 1984 through September 1985, but did not reinterview these households a year later, making it impossible to construct a panel for calendar years 1984/85. Finally, scrambled household identifiers in the initial release of the March 1994 CPS public use file prevented the creation of a 1992/93 calendar year panel.

Our complete panel includes observations for adjacent years on 109,913 individuals for 14 of the 16 periods between 1976/77 and 1991/92. Included in our panel sample are private sector wage and salary workers ages 16 and over with positive earnings in year *t*-1. Excluded are workers for whom a matched record was not made in year *t*, workers with imputed earnings or earnings less than the minimum wage in year *t*-1, workers receiving workers' compensation in year *t*-1, workers who were likely to be exempt from a state's workers' compensation program, and matched records in which individuals had more than one employer in year *t* (multiple employers in year *t* would preclude the matching of reliable union status information for year *t*-1 based on subsequent March- June outgoing rotation group records in year *t*).[11] Panel attrition (i.e., a failure to match records across years) results when households move (which means they are not reinterviewed by the Census), when individuals leave a household, if there is a failure to reinterview for other reasons (this is fairly rare), or if a unique match cannot be made (due, for example, to the presence of two individuals in the same household with identical gender and age). Peracchi and Welch (1995) analyzed attrition rates among matched March CPS files and concluded that age is the most important determinant of a successful match. Less important factors decreasing match probabilities were poor health, low schooling, and identification of the respondent as neither a household head nor spouse, while sex and race were not important match predictors following control for other factors.

Individuals from our CPS sample are matched to information on the workers' compensation law in

11 Workers receiving workers' compensation benefits in both years *t-1* and *t* are excluded from the analysis. Most individuals with benefits in consecutive years are not likely to represent new claims but, rather, are workers injured late in a calendar year or those with an unusually long duration of benefits. Our sample includes a representative number of such workers -- individuals not receiving benefits in year *t-1* who begin receiving benefits in year *t*, and continue receiving benefits in year *t+1*. We will miss a small number of *new* indemnity claims, however, from workers filing separate claims in consecutive years. Although the exclusion of this latter group may bias our estimates, the effect is likely to be small. A total of 293 individuals, 46.4 percent of whom are union members, received benefits in years *t-1* and *t* and are excluded from the analysis. We have no way of knowing how many of these workers received benefits for separate indemnity claims. In contrast to the 293 recipients excluded, the 1,594 recipients included in our sample had a union density of 37.4 percent. If anything, union-nonunion differences in claim rates may be understated by our

their state of residence. We created a database containing provisions of state workers' compensation laws by year based on information in *Analysis of Workers' Compensation Laws*, published annually by the U.S. Chamber of Commerce. Although this publication contains information on many aspects of state laws, the primary information we use in this study is on benefit rules (e.g., replacement rates and maximum benefits), waiting periods, retroactive periods, and coverage exemptions. Based on information on individual earnings in the CPS, we are then able to construct a *worker-specific* measure of *expected* benefits in the event of an injury. Information on detailed industry, occupation, and class of worker allows us to exclude workers whose employers are exempt from their states' workers' compensation laws (e.g., agricultural workers, longshoremen, and railroad workers).[12]

Industry differences in working conditions are controlled for by the inclusion of an industry injury rate measure, days lost per 100 workers, and relatively detailed industry dummies.[13] Average establishment size in workers' industry of employment is included to account for both injury risk and reporting differences due, say, to differences in experience rating between small and large employers. That control variable is potentially important in this study because employer size is positively correlated with unionization, and its omission might therefore entangle the effects of unions and establishment size.[14] Finally, we add to the data set measures of occupational skills and working conditions constructed from the *Dictionary of Occupational Titles* (DOT), with occupational characteristics mapped to the CPS based on Census occupational codes

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(necessary) sample selection criterion.

[12] Workers are covered by laws based on employer location, whereas our matching is based on household location. A few states express benefits as a percentage of take-home rather than gross pay. For these states, we assume a combined state and federal (including OASDI-H) *average* tax rate of 22 percent in calculating potential benefits. Krueger (1990) finds that workers' compensation recipiency is not related to differences in tax rates.

[13] We include industry injury rates as fixed effects, averaged over the years 1984, 1986, and 1988. These data, provided for SIC industries, were mapped by hand to Census industry categories in the CPS. We find similar relationships between workers' compensation claims and industry injury risk when the latter is measured by injuries per 100 workers rather than by days lost per 100 workers (U.S. Department of Labor, biennial).

[14] Establishment size is measured using a method similar to that in Brown and Medoff (1989). CPS supplements for May 1979, May 1983, May 1988, and April 1993 provide establishment size as a *categorical* variable. In order to create a *continuous* establishment size variable that could be matched to workers in our March CPS sample, we proceeded as follows. Mean establishment size by 1-digit industry by size class category (corresponding to the CPS categories) was obtained from appropriate years of *County Business Patterns*. These values were assigned to each worker in the CPS supplements containing establishment size information. Using the CPS supplements, mean establishment size was computed by 3-digit industry by region (empty cells were assigned the national industry value). Sample weights were used throughout the calculation. Values of mean establishment size were assigned to workers in the March CPS based on 3-digit industry and region. Mapping between 1970 and 1980 Census 3-digit industry codes was based on an algorithm provided by the Census (some randomization is involved).

(Miller et al., 1980; England and Kilbourne, 1988).[15] Relatively detailed industry and broad occupation dummies are also included.

Our principal approach is to estimate probit equations examining how union status affects the probability of a worker having a successful indemnity claim for workers' compensation during a year. In order to isolate the union effect, we include measures of state workers' compensation laws, worker characteristics, and job characteristics.[16] The probit equation, estimated by maximum likelihood, takes the general form:

(1) $p_{i,t} = \Phi(LAW_{i,t}\Gamma + X_{i,t,t\text{-}1}\beta + Z_{i,t\text{-}1}\Omega + UN_{i,t\text{-}1}\Theta + \in_{i,t})$

Here, Φ is the normal cumulative distribution function and $p_{i,t}$ is the probability that individual i receives workers' compensation income in year t. *LAW* includes a vector of workers' compensation law provisions that vary by state and year. State-by-year benefit rules, however, lead to expected log benefit levels that vary with worker i in year t (since individual earnings differ). The waiting period and the retroactive period (both measured in days) vary by state and year but not for individuals within states. Worker/job characteristics in X, measured at year t-1, are earnings and part-time status, while those measured in year t are years of schooling completed, age, marital status by gender, race, Hispanic status, and residence in a large metropolitan area. Union status (*UN*), a worker/job characteristic defined at the individual level, is broken out separately, since it is the principal focus of the paper. *UN* is recorded in March-June of year t, but is intended to measure union status in year t-1.[17]

Vector Z includes occupational and industry characteristics. At the industry level, we include detailed industry dummies, the industry injury rate, and average establishment size.[18] At the occupational level, we

[15] The so-called "Treiman" DOT values are matched to the March CPS for 1977-1982 using 1970 Census codes (Miller et al., 1980). England and Kilbourne (1988) provide a mapping of DOT values to 1980 Census codes. Changes between the 1980 and 1990 Census occupation codes are minor. The CPS began using the 1980 Census codes in 1983 and the 1990 Census codes in 1992. We make these codes time consistent and match the England and Kilbourne DOT values to the March CPS for 1983-93. As widely recognized, matching grouped variables to individual observations leads to a downward bias in standard errors on the grouped variables, but no necessary coefficient bias (Moulton, 1990). Because of this, we are cautious in ascribing statistical significance to variables measured at the industry and occupation levels.

[16] Our analysis differs from Krueger (1990) in that we include union status, include 14 rather than two panels, and include measures of job skill requirements, working conditions, establishment size, and industry injury rate.

[17] We exclude workers with more than one employer during year *t*. If we also exclude those changing jobs during year *t*-1 our sample is about 10 percent smaller, while union coefficients are slightly larger in absolute value, consistent with there being reduced measurement error.

[18] In work not shown we find no relationship between 1-digit industry-by-region annual unemployment rates (calculated by us from the CPS) and workers' compensation claims (note that year dummies are also included). High

include broad occupation dummies, plus skill and working condition variables from the DOT. In our probit model, we include only two DOT variables -- *DOT-SVP* represents specific vocational preparation and is measured by months of training required for proficiency in an occupation; *DOT-Strength* is a 1 to 5 index measuring required strength in the occupation, ranging from "sedentary" to "very heavy." Other DOT variables are included in our dataset. Because of substantial collinearity among the DOT occupational variables, the occupational dummies, and individual characteristics, these DOT variables have coefficients that are close to zero. These variables are described subsequently in Table 2, where means are presented for non-recipients and recipients of workers' compensation.

Descriptive Evidence

Before analyzing the determinants of workers' compensation claims, we provide descriptive evidence. Table 1 presents means by year of the percentages of workers receiving new workers' compensation indemnity claims. Individuals are counted as recipients if they had earnings in year *t*-1 and received workers' compensation income in year *t*. We present calculated recipiency rates for two samples – all matched March to March individuals meeting the sample criteria stated previously, and a smaller sample for whom individual level union status information is available. The latter sample is the one used throughout the remainder of the paper. Differences in claim rates across the two samples are very small, amounting to at most about .2 percentage points (in 1981 and 1983). Over the entire period, 1.45 percent of the workers in our estimating sample in year *t*-1 were new workers' compensation claimants in year *t*. Changes in incidence over time are not unlike those economy- wide, although we are reluctant to infer trends or cycles when sample sizes of claimants in any given year are so small.[19]

The rate of new indemnity claimants we estimate is slightly lower than that found by Krueger, who calculated a 1983 and 1984 rate of 1.5 percent (versus our 1.45 percent averaged over the same two years). That difference may stem from Krueger's somewhat more restrictive sample selection criteria. Krueger

unemployment can have offsetting effects on claims. On the one hand, a slack labor market may increase the attractiveness of workers' compensation benefits as an income source. On the other hand, weak demand should decrease injuries by leading to a slower work pace and the use of fewer inexperienced workers.

[19] Workers' compensation claimants account for about a .2 percentage points lower share when we retain workers in the CPS with imputed earnings and earnings less than the minimum wage. For times series figures on U.S. workers' compensation costs as a percentage of payroll, calculated from Social Security Administration figures on social

indicated that his figure was in the range of estimates of workers' compensation cases per covered worker derived from administrative records and concluded that there were not significant recall problems in the CPS. Recent information from the National Council on Compensation Insurance (NCCI) indicates a new indemnity claimant rate of 2.25 percent of covered full-time equivalent workers in 37 states plus D.C. for varying "policy years" in effect during the 1990-92 period.[20]

There are several reasons the NCCI rate will differ from that calculated from the CPS. Our figures are based on information from all states (and D.C.), we do not convert part-time workers to full-time equivalents, and our CPS sample includes workers in very small firms that are often not covered by workers' compensation. Although these factors are likely to explain some of the differences, we suspect that CPS estimates understate actual new claimant rates. The most likely explanation is that some recipients fail to report in the March CPS the receipt of workers' compensation income during the previous year. Underreporting or recall bias may be most severe among individuals who had minor injuries or received benefits early during the previous year. An understatement (as opposed to an overstatement) of the new claim rate increases our confidence that those measured as recipients are true recipients, while recipients incorrectly classified as non-recipients are far too small in number to affect mean characteristics of the non-recipient population. To the extent that the recipient population is understated, however, partial derivatives calculated from our probit models are likely to be biased toward zero.

Table 2 provides means of selected variables for the sample of workers who did and did not receive workers' compensation during year *t*. Weekly earnings are slightly higher among the 108,319 workers who did not receive benefits, averaging $551 (in 1992 dollars) in year *t*-1, as compared to $520 in year *t*-1 among the 1,594 workers who subsequently received benefits in year *t*. Average available weekly benefits are highly similar between the two groups. Comparing the means of benefits and earnings understates the average replacement rate (since the mean of a ratio is not equal to the ratio of means), which is .53 among non-

welfare expenditures, see Burton (1994, p. I-2).

[20] The NCCI rate of temporary total cases is 1.49 percent. The 2.25 percent rate presented in the text includes temporary total, permanent partial, and permanent total cases. Not included are medical-only cases and fatalities. These figures, from the NCCI *Annual Statistical Bulletin, 1995*, p. 296, were kindly provided by David Durbin of NCCI.

recipients and .52 among recipients.[21] In short, focusing on sample means provides little evidence that benefit levels and replacement rates are determinants of workers' compensation indemnity claims.

Difference between recipients and non-recipients in state mean waiting periods and retroactive periods are small, but in the expected direction. Whereas recipients have a mean waiting period of 5.1 days, non-recipients have a mean of 5.4 days (most states have waiting periods of three or seven days). Mean retroactive periods are 14.9 and 15.4 days, respectively.

As the figures in Table 2 show, 37.4 percent of recipients of workers' compensation are union members, compared to only 17.7 percent of non-recipients. To our knowledge, differences in recipiency rates based on union status have not previously been calculated with individual worker data. We find the recipiency rate for union workers to be 3.0 percent, roughly *triple* the 1.1 rate among nonunion workers (see Table 4). This large difference confirms anecdotal and prior statistical evidence indicating that claim rates are higher in union settings. Left unanswered at this juncture is the extent to which union-nonunion differences result from differences in worker characteristics, job hazards, and other job- related characteristics.

Personal and job characteristics differ considerably between workers' compensation claimants and non- claimants. Recipients have lower skill levels, as seen by lower mean schooling (11.9 versus 12.8 years completed), and employment in less skill-intensive occupations, as seen by lower mean levels of required education (*DOT-GED*) and months of training for occupational proficiency (*DOT-SVP*). Men are overrepresented among recipients, as are full-time workers. Recipients tend to be in occupations with more hazardous conditions, greater strength requirements, greater physical demands, more exposure to non-weather related environmental conditions, and work exclusively or partly outdoors. As expected, recipients work in industries with higher rates of injury and days lost. In results not shown, workers in blue-collar occupations (e.g., precision production and craft, operatives, and laborers) are overrepresented, whereas workers in most white-collar jobs are underrepresented (e.g., professional, technical, sales, and administrative support). Differences across broad industry classifications are less pronounced than across occupations, although workers in manufacturing, mining, and construction are more likely to receive benefits. There are only small

[21] Conclusions differ somewhat when we compare the mean logs of earnings and benefits. Recipients have earnings and expected benefit distributions less skewed to the right than among non-recipients.

recipient/non-recipient differences in race, marital status, age, industry establishment size, and (based on microdata beginning in 1989) firm size.

Union Status, Workers' Compensation Indemnity Claims, and the Benefit Elasticity

In this section we examine the role of union status on workers' compensation indemnity claims. The differences between recipients and non-recipients in worker and job characteristics seen in Table 2 reinforce the need to use multivariate analysis in order to infer the causal role of union membership. We focus first on the sensitivity of claims-benefit elasticity estimates to inclusion of union status, and then consider the independent effect of union status on claims.

Much previous research has focused on the response of claims to differences in benefit levels and other program parameters. Table 3 presents the probit equation results for alternative specifications of the workers' compensation claims model. Included for each variable are the probit coefficient, the asymptotic t-ratio, and the partial derivative of p with respect to each X (evaluated at the means of the X's). The dependent variable is equal to one if an individual received workers' compensation benefits in year t. It is expressed as a function of potential benefits, the waiting period, and individual and job characteristics.

We provide estimates from several specifications. In column 1 of Table 3, we provide results from a sparse model including only policy variables measuring expected benefits from workers' compensation (which vary by state, year, and individual), state waiting and retroactive periods (these vary by state and year), individual weekly earnings, and year dummies. In specification 2, the union status variable is added to this sparse specification. We then provide results from a dense specification (column 3) that excludes union status but includes variables measuring individual characteristics, job characteristics, and dummies for occupation, industry, and year. In specification 4, we add union status to the relatively dense specification.

Finally, specification 5 adds a set of 50 state dummies (D.C. is the reference group) to capture any state-specific differences in workers' compensation laws, administration, or behavior that are fixed over time and not measured by included variables. State fixed effects also may account for the endogeneity of workers' compensation laws (due, for example, to state union density levels), which are determined at the state level (Butler and Appel, 1990). We are able to include both state fixed effects and dummies for state waiting and retroactive periods, since there have been modest changes in these over our sample period. The coefficients on the waiting and retroactive periods in specification 5 thus measure the effects on claims from state-legislated *changes* in these program requirements.

The responsiveness of workers' compensation claims to benefit levels is theoretically indeterminate, given opposing incentive effects on employers and employees. We find the claim-benefit elasticity to be positive but modest in magnitude, smaller than estimates found by Krueger (1990) and in many studies using aggregate state and state-by-industry data. Our point estimates indicate a benefit elasticity of .394 based on the sparse specification (1), versus .209 following inclusion of detailed personal and job characteristics other than union status (specification 3). The benefit elasticity is found to drop *substantially* once we control for union status, falling from .394 to .215 when UN is added to the sparse specification (column 2) and from .209 to only .129 when it is added to the dense specification (column 4). We find a slightly higher benefit elasticity following the inclusion of state fixed effects, increasing from .129 in specification (4) to .178 in (5).[22]

Unionization is not considered in most studies of workers' compensation. The finding of substantially lower benefit elasticities following control for union status, therefore, is of some importance, suggesting that benefit elasticity estimates may be overstated in many previous studies. Unionized workers tend to have higher benefit levels. The difference in log benefits between union and nonunion members is .244; mean weekly benefits (in 1992 dollars) for union members are $291 versus $240 among nonunion workers. Because union membership and benefit levels are positively correlated, exclusion of union status biases upward estimates of the benefit elasticity (we examine union-nonunion differences in benefit elasticities).

Evidence on the direct effect of union status on receipt of workers' compensation is unambiguous.

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[22] As shown by Krueger (1990, p. 89n), the benefit elasticity is calculated by $h=\beta\varphi(z)/p$, where β is the probit coefficient on log benefits, $\varphi(\bullet)$ is the normal density function, $\Phi(\bullet)$ is the cumulative normal distribution function, $z=\Phi^{-1}(p)$, and p is the sample mean workers' compensation participation rate (.0145 for our full sample union and nonunion sample). The benefit elasticity estimates in Krueger (1990) are larger than in much of the literature that he cites. They are also larger than we obtain using a similar methodology and specifications as does Krueger. For example, he obtains an elasticity of .45 in a model similar to (3) (as compared to our estimate of .21), and an elasticity of .74 following control for state fixed effects (as compared to .23 in our state model without union, not shown). When we restrict our analysis to the same two years (1983/84 and 1984/85) used by Krueger, we obtain results highly similar to his. The CPS sample used by Krueger (with two panel periods versus the 14 used here) turns out to produce results nonrepresentative of those obtained from CPS samples from either earlier or later periods. It is not surprising that a relatively small sample could produce unrepresentative estimates (Krueger had 290 recipients in his sample of 19,082 persons). Apart from the benefit elasticities, Krueger's other results for 1982/83-1983/84 (based on the 1983-85 March CPS files) are similar to what we obtain with our larger 1976/77-1991/92 sample. Krueger does not consider the effects of union status.

When union status is included in a sparse specification controlling only for benefits, earnings, the waiting period, and year, its partial derivative is .0140, implying (given the .0145 mean claim rate) that union workers are roughly twice as likely (96.6 percent) as the *mean* worker to receive workers' compensation during a year.[23] Of course, much of the union-nonunion gap may be accounted for not only by differences in benefit levels, but also by differences in worker and job characteristics. The union coefficient falls by roughly fifty percent as we move to specification (4). The partial derivative of .0067 implies that union workers are 46.2 percent more likely then the average worker, and 60.4 percent more likely than a nonunion worker, to receive workers' compensation. Union coefficients change little when state fixed effects are included (model 5), the partial decreasing from .0067 to .0062. In work not shown, the union coefficient was scarcely affected when we added nearly 200 industry dummies to the model (variables measured at the industry level were omitted).

In short, union status appears to have a large direct effect on the recipiency of workers' compensation, even following a careful accounting for union-nonunion differences in working conditions, worker characteristics, and program attributes (i.e., benefit levels and waiting periods). These results provide strong support for the hypothesis that union representation increases workers' compensation recipiency by providing information to workers that facilitates claims, and through contractual protection (both explicit and implicit) against employer monitoring and penalties for worker claims.

To explore further the question of how unionism affects workers' compensation indemnity claims, we provide separate estimates of union and nonunion equations, allowing coefficients to differ between the union and nonunion sectors. In Table 4, we present coefficients on the workers’ compensation policy variables (benefits, waiting period, and retroactive period) from the three specifications shown previously – the sparse specification and dense specifications with and without state dummies. The results indicate that union workers are more responsive to differences in benefit levels than are nonunion workers. In the dense specification without state dummies, the partial of claims with respect to log benefits is 3.6 times larger in the union than in the nonunion equations, and 5.1 times larger in the specification including state dummies. That

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[23] The recipiency rate is 3.0 percent among the union sample, versus 1.1 percent of the nonunion sample. If one evaluates the partial relative to the nonunion mean, the estimate from the sparse specification is that union workers are more than twice as likely (126 percent) as nonunion workers to receive workers' compensation.

being said, standard errors are large, with no statistically significant benefit coefficients. Differences in benefit *elasticities* (shown toward the bottom of the table) are more modest, the union elasticity being roughly double the nonunion elasticity in the dense specification without state dummies, and nearly triple when state dummies are added.[24]

Although worker responses to differences in benefit levels appear weak, results in Table 3 indicate a strong relationship between receipt of workers' compensation benefits and the length of the waiting period before one is eligible for benefits. Longer waiting periods (say, seven rather than three days) tend to eliminate a relatively large number of minor injuries that might otherwise be compensated. The coefficient on waiting period is highly stable across specifications (1) through (4). It is effectively zero following inclusion of state dummies (specification 5), which is not surprising since there exists little intertemporal variation in state-mandated waiting periods. The estimates from specification (4) imply that an increase in the waiting period from three to seven days would decrease workers' compensation claims by 22.1 percent.[25] Our results support the conclusion reached previously in the literature (e.g, Krueger, 1990; Butler, 1994) that increasing the waiting period can lead to a substantial reduction in workers' compensation claims without compromising protection of workers against losses from serious workplace disabilities.

Table 4 provides separate waiting period estimates for union and nonunion workers. Focusing on the dense specification without state dummies (model 2), since waiting periods change little over time, the effect of a one-day difference in the waiting period on the probability of a claim is twice as high for a union worker as for a nonunion worker (-.00133 versus -.00067). The larger absolute response for union than nonunion workers further supports the premise that unionism affects claims through the provision of information to its members and through an increased sensitivity to changes in program parameters.

Although the absolute response is larger for union workers, this result implies a somewhat smaller

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[24] Following Krueger (1990, p. 89n) we emphasize benefit elasticities computed using the total sample recipiency rate. Elasticities computed using separate union and nonunion recipiency rates are also shown in Table 4. Partial derivatives for each of the variables are evaluated at the full sample means in order to determine the effects of an average worker. Changes in the partials are very small when we substitute separate union and nonunion means.
[25] This is calculated by (-.00080*4)/.0145=.226, where -.00080 is the partial of claims with respect to days, 4 is the change between a three and seven day waiting period, and .0145 is the mean of the dependent variable.

percentage effect, since union recipiency rates are roughly triple those among nonunion workers. As found by Krueger (1990), state differences in the retroactive period do not affect the probability of a workers' compensation claim. The greater sensitivity of union than nonunion workers to differences in benefit levels and waiting periods reinforces our previous conclusion that union representation increases workers' compensation claims, at least in part by informing workers about the availability of benefits in the event of a disabling injury.

Workers' Compensation Indemnity Claims and Worker and Job Characteristics

In this section, we provide a brief overview of results from our model other than the effects of policy variables and union status. As seen in Table 3, workers' compensation recipiency is not statistically related to the level of weekly earnings. Absent controls for worker and job skills (specifications 1 and 2), claims are inversely related to weekly earnings, reflecting the higher opportunity cost of work absence for higher-wage workers. Once we control for skill-related wage determinants of earnings such as schooling, age, and SVP, all of which are inversely related to claims, there is no remaining claims-earnings relationship. This is not surprising, because in the dense specification, which controls for several measurable determinants of earnings, the earnings variable primarily captures residual earnings emanating from unmeasured skills and working conditions or from rents.

In our summary of the relationship between workers' compensation claims and personal and job characteristics, we focus attention on the results from model (4), the dense specification without state dummies. As discussed above, skill-related variables are negatively related to the probability of a workers' compensation claim. This is most evident from the schooling variable, each additional year of schooling lowering the probability of receipt of workers' compensation by 4.3 percent (i.e., .00063/.0145). We also obtain negative coefficients on years of training required for occupational proficiency (*DOT-SVP*), and lower claims rates among workers in white-collar occupations (these latter results are not shown). Schooling and other skill-related variables may proxy higher opportunity cost earnings, safer working conditions not measured fully by other variables, greater awareness of risks and injury prevention, lower discount rates, and less risk-taking behavior. On the other hand, more educated workers are likely to have greater knowledge of, and ability to deal with, the workers' compensation system.

Workers' compensation claim rates vary with age, being low for (the relatively small sample of) teenagers, highest for workers ages 20-49, and lower among older workers.[26] The low rate among older workers may reflect greater job experience and skills or less risk-taking behavior. Black workers are about 19.3 percent less likely than the mean worker in the sample, other things equal, to receive workers' compensation income (the standard error on the estimate is relatively high). This difference is interesting, given that *unadjusted* claim rates are similar. Together, those two results are consistent with black workers being employed in more hazardous jobs, but not more likely to receive workers' compensation. A possible explanation is that black workers have or are provided with less information about the availability of workers' compensation benefits, or injuries by black workers are less likely to be certified as compensable than are injuries by white workers. Hispanics exhibit a probability of recipiency 13.1 percent lower than among non-Hispanics, although this difference is not significant at standard levels. There are no significant differences in recipiency between whites and nonwhite non-blacks (primarily Asians).

We find interesting and rather robust results with respect to gender and marital status. Although unadjusted workers' compensation claim rates for women are lower than for men, women have a higher probability of a claim than men, following control for individual and job characteristics. Although not all differences are significant, among both women and men, workers never married have the lowest probability of a claim, followed by those married with spouse present, with those separated, divorced, or widowed having the highest claim probability. For example, as compared to never-married men, women with spouse present have a 18.9 percent higher probability of recipiency, while women separated, widowed, or divorced have a 39.9 percent higher probability. Although patterns with respect to marital status are the same for men and women, differences among men are small and insignificant, whereas differences are more substantial among women. Our results are consistent with Krueger's "surprising" finding of higher probabilities of claims for women and for married workers.

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[26] A finer breakdown by five-year age interval indicated little difference in coefficients among the age 20-34 and age 35-49 groups. There was variation among the groups of older (age 50 and over) workers, but sample sizes of older recipients broken down finely by age category are rather small. As mentioned previously, we cannot examine differences in the duration or cost of injuries among young and old workers.

We are reluctant to speculate on explanations for the gender and marital status results. It is possible that single workers lack good knowledge that benefits are available in the event of an injury, feel less need to obtain replacement income, or value time spent off the job less highly. A higher valuation placed on non-market time may explain higher claims probability for women and for married as opposed to never-married workers; it is not clear that it can account for the high rate among those once married but without spouse present. While the pattern found is consistent with higher claim rates among those with larger income "needs," one might expect those with the greatest needs to take fewer risks and have greater incentive not to forego earnings in the event of minor injuries. An even more speculative suggestion is that the higher claim rates among those previously married may result from unmeasured personal characteristics that make both a separation or divorce and a workplace injury claim more likely.

Another possibility is that differences in annual hours worked explain differences in claim rates. Hours differences cannot explain the higher claim rate among women than men, however, since women work fewer hours. We examined gender-specific differences in annual hours among workers in our sample by marital status. Women, for whom claim rates vary substantially with respect to marital status, have mean annual hours worked in year *t*-1 of 1,807, 1,739, and 1,855 among never married, married spouse present, and previously married women, respectively (the corresponding figures for men are 1,869, 2,183, and 2,098 hours). Such small hours differences cannot account for the large estimated differences in claim rates by marital status among women. At a minimum, our rather robust results with respect to gender and marital status warrant attention and deserve less speculative explanations.[27]

A factor not considered by Krueger is whether workers' part-time status affects workers' compensation indemnity claims. Following control for age, gender and other worker and job characteristics, part-time workers are substantially less likely than full-time workers to receive workers' compensation benefits, about 23.0 percent below the mean receipt rate (i.e., .00333/.0145).

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[27] Jim Chelius suggests that married workers have a higher probability of a claim because the effective tax rate is lower, and real benefit level higher, for married workers. Given the magnitude of benefit elasticities, we doubt that this is a sufficient explanation. Nor does it explain higher rates for those separated, divorced, or widowed than for married workers with spouse present.

There are obvious differences between part- and full-time workers, the former being more likely to be young, never married, and female. But each of these characteristics are controlled for in our model. The principal explanation for lower part-time claims is likely to be that such individuals work fewer hours. In our sample, part-time workers' average hours are approximately half those of full-time workers (21.7 versus 42.7 usual hours per week), a larger spread than the 23 percent claims differential. Our evidence suggests that workers' compensation claims rise less than proportionately with respect to hours worked. This lends some credence to a "warm-up" view of injuries, in which the appropriate margin for accidents is a day of work rather than hours of work, although this thesis cannot be tested directly absent information on days worked.[28] In addition to fewer hours (and days) among part-time workers, lower incidence rates may reflect lower injury risks in part-time jobs, workers who are more poorly informed about the availability of workers' compensation benefits (both self- information and information from their employer), and less need for replacement income in the event of an injury.

We find a negative relationship between workers' compensation claims and average establishment size in the workers' industry (for previous literature, see, among others, Ruser, 1991, 1993). The magnitude of the effect is small, an increase in average establishment size of 100 employees being associated with only a 1.3 percent decrease in claims (i.e., .00189 times .10 divided by .0145). Use of an aggregate industry size variable masks what may be a low claim rate among very small establishments, due in part to lack of coverage and lack of information or experience in using the system. Most interesting is that the coefficient on establishment size becomes substantially more negative following control for union status (see the change between models 3 and 4), owing to the positive correlation between size and union status. Thus, prior studies not controlling for union status are likely to obtain coefficients on establishment size that are biased upward. In data beginning in 1989, we find little relationship between workers' compensation and *firm* size, measured at the individual level (see Table 2 for means by recipiency status).

Occupational strength requirements have a substantial effect on workers' compensation indemnity

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[28] We thank a referee for suggestions in interpreting our part-time results. If accidents were to increase with length of shift owing to fatigue, then one would expect claims to rise more rather than less than proportionately with respect to hours worked.

claims. Using point estimates from equation (4), a one grade difference in occupational strength is associated with a .0045 increase in the probability of recipiency, an increase of 30.8 percent relative to mean recipiency. The magnitude of the *DOT-Strength* coefficient may reflect the relative importance of back injuries among workers' compensation claims, and the correlation between strength requirements and other physical demands.

The positive relationship between workers' compensation claims and industry days lost from injuries requires little direct comment. We would emphasize the value of this variable as a control in our analysis of union effects. After accounting for the riskiness of a workers' industry of employment (as well as job characteristics of their occupation), we find that union workers are substantially more likely than similar nonunion workers to receive workers' compensation benefits. One might also interpret the industry days lost and other selected variables as controlling for the inherent riskiness of a job (e.g., Butler, 1994). By this interpretation, coefficients on such variables as union status, potential benefits, and the waiting period reflect these variables' impact on claims owing to moral hazard, incentives, and information.

Conclusion

We have developed a large micro database from the CPS in order to examine the determinants of workers' compensation recipiency during 1977-92. Using individual-data over time, we have been able to examine the effect of important individual and job characteristics on the likelihood of recipiency, and to analyze differences in outcomes resulting from variation in benefit levels across time, states, and individuals within states.

An important finding from our study is that unionization has a substantial effect on indemnity claims. Union workers are far more likely than nonunion workers, other things equal, to receive benefits from workers' compensation, and the likelihood of a claim is more responsive to differences in benefit levels among union than among nonunion workers. Moreover, the inability to control for union membership in previous studies seems to impart an upward bias to benefit elasticity estimates, leading studies to overstate the responsiveness of indemnity claims to changes in benefit levels. Although much of the three to one spread in recipiency rates between union and nonunion workers is the result of differences in benefit levels, job structure and working conditions, we estimate that union workers are about 60 percent more likely than nonunion workers to receive workers' compensation payments, following control for a rather extensive assortment of personal, labor market, occupation, and industry characteristics.

We have argued that the evidence on union-nonunion differences in workers' compensation indemnity claims supports a view that the likelihood of claims is sensitive to information available to workers and the expected response of employers to worker claims. We speculate that unions provide workers with information about the availability of benefits following an injury or illness, prevent supervisors from withholding information or discouraging workers' compensation claims, and protect workers from penalty in the event that a claim is filed. An alternative (although not mutually exclusive) interpretation is that union members are "abusing" the system, reporting and receiving compensation for marginal injuries, and at the same time receiving contract protection from management attempts to monitor and reduce overreporting. The evidence may be more consistent with the latter interpretation. A more generous workers' compensation system provides employers with incentives to provide a safer work environment. Clear-cut evidence of a positive overall claims-benefit elasticity and of a larger elasticity for union than for nonunion workers suggests that there exists a large scope for moral hazard.

While we are unable to distinguish unambiguously between these explanations of union-nonunion differences in claims for workers' compensation, either one implies unequal access between union and nonunion workers. The routes through which unions affect claims has implications for public policy. If moral hazard and an "abuse" of union power is the principal explanation, then state workers' compensation agencies should more closely monitor claims coming from union establishments.[29] If the other explanation is correct, it implies that workers, most of whom are now nonunion, are not being provided with as thorough coverage for injuries as policy makers might have thought. That is, the low level of union representation in the United States may imply that workers receive from employers inadequate information about the availability of compensation for workplace injuries, or make too few claims because of fear that they will be penalized. If this is so, it lends support to calls for the development of expanded employee voice in the workplace, be it through

[29] The scope of this "problem" is of course decreasing over time as union density declines in the private sector. The percentage of private sector (agricultural and nonagricultural) wage and salary workers who are union members decreased from 21.7 percent in 1977 to 10.3 percent in 1994 (Hirsch and Macpherson, 1996, p. 10).

traditional union representation or alternative mechanisms of worker participation.[30] Policies that improve information to nonunion workers about the availability of benefits or would discourage employer monitoring and penalties against workers making claims could well entail nontrivial increases in total costs. Such policies would face vigorous opposition from employers.

[30] Such issues have been at the forefront of the dialogue surrounding the report of the (Dunlop) Commission on the Future of Worker-Management Relations (1994). For a recent analysis of the literature on worker participation, see Levine (1995).

Appendix: Construction of Longitudinal Samples from the March CPS

The longitudinal March/March matched CPS file was created in the following manner. 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 entire sample is asked questions about work activity, industry, occupation, and sources of income (including workers' compensation) the previous calendar year. A quarter sample (the outgoing rotation groups) in each monthly CPS are asked questions from an earnings supplement, including current union status, along with current weekly earnings and hours worked.

The CPS contains household identification numbers (ID) and record line numbers, but not individual identifiers. Individuals potentially can be identified for the same month in consecutive years; that is, individuals in rotations 1-4 in year 1 can be matched to individuals in rotations 5-8, respectively, in year 2. Separate data files were created for males and females, and for pairs of years. 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, an individual record from rotations 5-8 in year 2 had to be matched with an individual record from rotations 1-4 (5 with 1, 6 with 2, etc.), 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.

The matching took place as follows. For each second year rotation individual, the search was made through all first year rotation individuals with the same ID to make sure there was only 1 possible match; the file was resorted in reverse order and each selected first year rotation individual was checked to insure a unique second year rotation 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 principal reasons that matches cannot be made are if a household moves (and thus is not reinterviewed in its new location), if an individual moves out of a household, or if the Census is unable to reinterview a household and/or receive information on the individual. Peracchi and Welch (1995) analyze attrition rates among matched March CPS files and conclude that age is the most important determinant of a successful match. Other factors that lessen match probabilities are poor health, low schooling, and not a household head, while sex and race are unimportant match predictors following control for other factors.

In this paper, we construct a sample comprising 14 matched panels based on the March CPS surveys for 1977/78 through 1981/82, 1983/84, 1984/85, and 1986/87 through 1992/93, corresponding to the calendar years 1976/77-1980/81, 1982/3-1983/84, and 1985/86-1991/92. Years were excluded for reasons discussed in the text.

No direct information is available from the March CPS on the union status of respondents' longest job the previous year. But during all months for years since 1983, the outgoing rotation groups were asked union status on their *current* job. All rotation groups were asked a union membership question in the May surveys for 1973-80, and a quarter sample in May 1981 (there were no union questions in 1982). For calendar years since 1983, we retrieve all rotation groups' union status on their current job, as reported in March, April, May, or June of year t, and assign this value to a variable measuring *individual* union status on the job in year t-1 if and only if workers report that they had only one employer during the year. For individuals in 1977-1980, we retrieve union status reported in May for the *half* March sample still in the CPS in May (March rotation groups 1, 2, 5, and 6). For 1981, we are able to match a quarter sample of the March respondents to their union status question in May. Again, we retain individuals if and only if they report only one employer during the year.

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Table 1: Means of Workers' Compensation Utilization, by Year

	Full Sample		Estimation Sample	
Year	Sample Size	Workers' Comp. Receipt (percent)	Sample Size	Workers' Comp. Receipt (percent)
1977	11,040	1.32	3,882	1.31
1978	10,743	1.52	3,941	1.55
1979	10,551	1.88	3,773	1.99
1980	13,219	1.51	4,686	1.49
1981	12,675	1.57	2,386	1.80
1982	12,768	1.36	–	–
1983	12,757	1.37	9,184	1.58
1984	12,344	1.20	9,011	1.32
1985	–	–	–	–
1986	12,769	1.25	9,145	1.27
1987	13,030	1.42	9,163	1.46
1988	12,562	1.47	8,806	1.60
1989	15,017	1.32	11,017	1.23
1990	16,195	1.46	11,864	1.53
1991	16,085	1.36	11,753	1.36
1992	15,455	1.46	11,302	1.43
1977-92	197,210	1.42	109,913	1.45

Source: 1977/78 to 1992/93 matched March/March CPS Panels for calendar years 1976/77 to 1991/92. Workers' compensation receipt rate represents the percentage of workers in year *t*-1 receiving income from workers' compensation in year *t*. The estimation sample is the subset of the full sample for whom union status information could be matched. The union sample is used in all subsequent analysis. See the text and data appendix for details on construction of the data set.

Table 2: Means of Selected Variables Among
Workers' Compensation Recipients and Non-Recipients

Variable	Recipients Mean	Recipients Std. Dev.	Non-Recipients Mean	Non-Recipients Std. Dev.
Benefits (1992 $)	252.707	106.046	248.984	117.328
Log Benefits	5.441	0.441	5.399	0.511
Weekly Wage (1992 $)	519.851	289.904	550.748	468.113
Log Weekly Wage	6.120	0.525	6.094	0.677
Benefits/Wage	0.521	0.125	0.531	0.273
Waiting Period (days)	5.109	1.922	5.373	1.894
Retroactive Period (days)	14.892	7.035	15.356	7.315
Union Member	0.374	0.484	0.177	0.382
Female	0.349	0.477	0.443	0.497
Years Schooling	11.862	2.434	12.838	2.583
Age	39.274	12.144	39.677	12.359
Black	0.074	0.262	0.075	0.263
Other Race	0.026	0.160	0.025	0.157
Hispanic	0.058	0.233	0.048	0.215
Married, Spouse Present	0.725	0.447	0.714	0.452
Ever Married, Other	0.082	0.274	0.073	0.260
Part-time	0.087	0.282	0.136	0.343
Large Metropolitan Area	0.288	0.453	0.304	0.460
DOT-SVP (Months Training)	19.549	16.331	25.131	19.581
DOT-GED (Gen Educ Dev)	3.361	0.757	3.744	0.822
DOT-Strength	2.623	0.691	2.199	0.742
DOT-Physical Demands	2.104	0.789	1.715	0.825
DOT-Environmental Cond.	0.544	0.548	0.329	0.449
DOT-Hazards	0.212	0.280	0.135	0.235
DOT-Indoors & Outdoors	0.231	0.268	0.164	0.225
DOT-Outdoors	0.057	0.128	0.040	0.109
IND-Days Lost per 100	88.422	55.649	66.264	46.791
IND-Injuries per 100	10.036	4.707	7.963	4.725
IND-Establishment Size	357.159	388.084	348.924	379.643
IND/REG-Unemployment	0.068	0.031	0.062	0.029
Firm Size <25	0.163	0.370	0.206	0.405
Firm Size 25-99	0.161	0.368	0.154	0.361
Firm Size 100-499	0.165	0.372	0.169	0.375
Firm Size 500-999	0.080	0.271	0.065	0.246
Firm Size 1000+	0.431	0.496	0.406	0.491
Sample Size	1,594		108,319	

Source: Micro variables - March CPS Panels, 1976/77 to 1991/92; *DOT* variables - *Dictionary of Occupational Titles*; IND variables, see text. Firm size variables are based on a post-1989 CPS sample of 34,919 individuals, 503 and 34,416 recipients and non-recipients, respectively. Occupational skill variables are DOT-GED, a 1-6 index of general educational development measuring necessary reasoning, writing, and mathematical skills; and DOT-SVP, specific vocational preparation measured by months training required for occupational proficiency. Occupational working condition variables are DOT-Environment, the sum of the proportions in jobs with four extreme environmental disamenities; DOT-Hazards, the proportion in hazardous jobs; DOT-Strength a 1 to 5 index ranging from "sedentary to "very heavy"; DOT-Physical, the sum of the proportions in jobs with four physical demands (e.g., climbing, stooping, reaching, and seeing); and DOT-Indoors & Outdoors and DOT-Outdoors, measuring the proportion of workers whose work is 25-75 percent and 75 plus percent outdoors, respectively.

Table 3: Determinants of Workers' Compensation Recipiency: Probit Results, 1977-92

Variable	Coeff. $\lvert t \rvert$ $\partial P/\partial X$ *(1)*	Coeff. $\lvert t \rvert$ $\partial P/\partial X$ *(2)*	Coeff. $\lvert t \rvert$ $\partial P/\partial X$ *(3)*	Coeff. $\lvert t \rvert$ $\partial P/\partial X$ *(4)*	Coeff. $\lvert t \rvert$ $\partial P/\partial X$ *(5)*
Log Benefits	.15535 (3.717) [.00563]	.08458 (1.966) [.00287]	.08236 (1.762) [.00240]	.05064 (1.073) [.00145]	.06999 (1.193) [.00193]
Waiting Period	-.02537 (4.742) [-.00092]	-.02639 (4.887) [-.00090]	-.02795 (4.986) [-.00082]	-.02778 (4.946) [-.00080]	-.00506 (0.116) [-.00014]
Retroactive Period	-.00077 (0.526) [-.00003]	.00020 (0.132) [.00001]	-.00109 (0.705) [-.00003]	-.00037 (0.238) [-.00001]	.00137 (0.225) [.00004]
Log Weekly Wage	-.08309 (2.623) [-.00301]	-.08931 (2.710) [-.00303]	.01419 (0.350) [.00041]	-.00834 (0.203) [-.00024]	-.03109 (0.662) [-.00086]
Union Membership (individual)	–	.41301 (18.484) [.01403]	–	.23187 (9.166) [.00666]	.22558 (8.712) [.00621]
Years of Schooling	–	–	-.02272 (4.582) [-.00066]	-.02180 (4.357) [-.00063]	-.02220 (4.398) [-.00061]
Age 20-34	–	–	.27067 (2.910) [.00789]	.26963 (2.892) [.00774]	.27103 (2.885) [.00746]
Age 35-49	–	–	.16919 (1.783) [.00493]	.16000 (1.682) [.00460]	.16017 (1.671) [.00441]
Age 50 & over	–	–	.16491 (1.729) [.00481]	.15149 (1.584) [.00435]	.14674 (1.523) [.00404]
Black	–	–	-.08410 (2.117) [-.00245]	-.09750 (2.446) [-.00280]	-.06329 (1.525) [-.00174]
Other Race	–	–	.01388 (0.218) [.00040]	.01404 (0.220) [.00040]	-.07639 (1.095) [-.00210]
Hispanic	–	–	-.06190 (1.316) [-.00181]	-.06598 (1.399) [-.00190]	-.13919 (2.780) [-.00383]
Part-time	–	–	-.09543 (2.336) [-.00278]	-.11610 (2.819) [-.00333]	-.12813 (3.076) [-.00353]
Large Metropolitan Area	–	–	.04867 (2.034) [.00142]	.03568 (1.483) [.00102]	-.04362 (1.508) [-.00120]

Table 3 (continued): Determinants of Workers' Compensation Recipiency, Probit Results, 1977-92

Variable	Coeff. |*t*| $\partial P/\partial X$ *(1)*	Coeff. |*t*| $\partial P/\partial X$ *(2)*	Coeff. |*t*| $\partial P/\partial X$ *(3)*	Coeff. |*t*| $\partial P/\partial X$ *(4)*	Coeff. |*t*| $\partial P/\partial X$ *(5)*
Male: Married, Spouse Present	–	–	.01945 (0.544) [.00057]	.01235 (0.345) [.00035]	.02586 (0.716) [.00071]
Male: Separated, Divorced, Widowed	–	–	.02568 (0.361) [.00075]	.01797 (0.252) [.00052]	.02193 (0.305) [.00060]
Female: Never Married	–	–	.05481 (1.119) [.00160]	0.4649 (0.947) [.00134]	.04626 (0.936) [.00127]
Female: Married, Spouse Present	–	–	.10119 (2.453) [.00295]	.09545 (2.311) [.00274]	.10207 (2.452) [.00281]
Female: Separated, Divorced, Widowed	–	–	.20665 (3.587) [.00602]	.20112 (3.484) [.00578]	.20912 (3.602) [.00576]
IND-Days Lost per 100 Workers	–	–	.00174 (6.769) [.00005]	.00165 (6.340) [.00005]	.00172 (6.538) [.00005]
IND-Establishment Size/1,000	–	–	-.03647 (1.095) [-.00106]	-.06567 (1.950) [-.00189]	-.05478 (1.603) [-.00151]
DOT-Months Training Required (SVP)	–	–	-.00302 (2.837) [-.00009]	-.00244 (2.289) [-.00007]	-.00244 (2.269) [-.00007]
DOT-Strength	–	–	.17410 (7.138) [.00508]	.15543 (6.328) [.00446]	.15873 (6.421) [.00437]
Year (13)	yes	yes	yes	yes	yes
Industry (38)	no	no	yes	yes	yes
Occupation (10)	no	no	yes	yes	yes
States & D.C. (50)	no	no	no	no	yes
Benefit elasticity	.3941	.2146	.2089	.1285	.1775
Log Likelihood	-8297.5	-8136.1	-7870.3	-7829.1	-7760.5
Sample Size	109,913	109,913	109,913	109,913	109,913
Mean Dep. Variable	.0145	.0145	.0145	.0145	.0145

Source: March/March CPS Panels, 1976/77 to 1991/92; Dictionary of Occupational Titles; U.S. Chamber of Commerce, Analysis of Workers' Compensation Laws; and selected CPS supplements. The dependent variable is recipiency of workers' compensation income in year *t*.

Table 4: Union versus Nonunion Determination of Workers' Compensation Recipiency:
Policy Variable Probit Results, 1982-92.

	Union			Nonunion		
Variable	Coeff. $\lvert t \rvert$ $\partial P/\partial X$ *(1)*	Coeff. $\lvert t \rvert$ $\partial P/\partial X$ *(2)*	Coeff. $\lvert t \rvert$ $\partial P/\partial X$ *(3)*	Coeff. $\lvert t \rvert$ $\partial P/\partial X$ *(1)*	Coeff. $\lvert t \rvert$ $\partial P/\partial X$ *(2)*	Coeff. $\lvert t \rvert$ $\partial P/\partial X$ *(3)*
Log Benefits	.12830 (1.492) [.00849]	.08194 (0.918) [.00374]	.15518 (1.315) [.00653]	.06917 (1.385) [.00198]	.04066 (0.718) [.00104]	.05280 (0.752) [.00128]
Waiting Period	-.03024 (3.208) [-.00200]	-.02917 (2.967) [-.00133]	-.12374 (1.077) [-.00520]	-.02509 (3.795) [-.00072]	-.02604 (3.769) [-.00067]	.03102 (0.662) [.00075]
Retroactive Period	-.00179 (0.602) [-.00012]	-.00091 (0.290) [-.00004]	.00482 (0.397) [.00020]	.00075 (0.436) [.00002]	-.00048 (0.266) [-.00001]	.00037 (0.052) [.00001]
Year (14)	yes	yes	yes	yes	yes	yes
Industry (38)	no	yes	yes	no	yes	yes
Occupation (10)	no	yes	yes	no	yes	yes
State (49)	no	no	yes	no	no	yes
Benefit elasticity (overall means)	.3254	.2263	.3936	.1755	.1031	.1339
Benefit elasticity (union/nonunion means)	.2908	.1857	.3517	.1819	.1069	.1388
Log Likelihood	-2660.9	-2587.8	-2555.7	-5468.7	-5206.8	-5139.9
Sample Size	19,792	19,792	19,792	90,121	90,121	90,121
Mean Dep. Variable	.0301	.0301	.0301	.0111	.0111	.0111

Source: March/March CPS Panels, 1976/77 to 1991/92; Dictionary of Occupational Titles; U.S. Chamber of Commerce, Analysis of Workers' Compensation Laws; and selected CPS supplements. The dependent variable is recipiency of workers' compensation income in year *t*. There are a total of 49 rather than 50 state/DC dummies owing to an empty union recipient cell.