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ABSTRACT

Steady increases in the costs of medical care, coupled with a rise in the fraction of workers who lack medical care insurance, have led to a growing concern that the Workers' Compensation system is paying for off-the-job injuries. Many analysts have interpreted the high rate of Monday injuries -- especially for hard-to-monitor injuries like back sprains -- as evidence of this phenomenon. In this paper, we propose a test of the hypothesis that higher Monday injury rates are due to fraudulent claims. Specifically, we compare the daily injury patterns for workers who are more and less likely to have medical insurance coverage, and the corresponding differences in the fraction of injury claims that are disputed by employers. Contrary to expectations, we find that workers without medical coverage are no more likely to report a Monday injury than other workers. Similarly, employers are no more likely to challenge a Monday injury claim -- even for workers who lack medical insurance.

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Is Workers' Compensation Covering Uninsured Medical Costs?
Evidence from the 'Monday Effect'

Any targeted social program is vulnerable to abuse or even outright fraud in the determination of benefit eligibility. It is widely believed, for example, that a sizeable fraction of Disability Insurance recipients are able to work -- and are therefore technically ineligible for benefits -- but claim a disability in order to receive benefits.[1] Similar concerns are expressed about other targeted programs, including Aid to Families with Dependent Children, Unemployment Insurance, and Workers' Compensation.[2] In the case of Workers' Compensation, secular increases in the costs of medical care, coupled with recent rises in the fraction of workers who lack medical insurance (Olson (1994)), have heightened concerns that the program is paying for off-the-job illnesses and injuries. Difficulties in policing the boundary between on-the-job and off-the-job injuries have even led some analysts to propose "24-hour" medical coverage that incorporates Workers' Compensation into a universal health care program and eliminates the special status of work-related injuries.[3]

Possibly the most striking evidence of fraudulent claim activity in the Workers' Compensation (WC) program arises from the unusual pattern of Monday accident claims. In a seminal paper, Smith (1989) showed that WC claims for strains and sprains are more likely to arise on a Monday, whereas harder-to-conceal injuries like cuts and lacerations are about equally

[1]See e.g. Parsons (1980). Bound (1989) presents a dissenting view.

[2]See Wolf and Greenberg (1986) for an analysis of fraud in the AFDC program, and Burgess (1992) for an analysis of compliance with unemployment insurance job search requirements.

[3]See Burton (1992) and Baker and Krueger (1994) for discussions of "24 hour coverage" proposals.

as likely on a Monday as on other weekdays.[4] Figure 1 presents corroborating evidence of this pattern drawn from a large sample of injury claims for the state of Minnesota.[5] As in Smith's analysis, the incidence of burns, cuts, and fractures is roughly constant across workdays. By comparison, strains and back injuries are about 25 percent more likely to occur on Mondays than other weekdays. Although circumstantial, this evidence is consistent with the view that some workers have "postdated" weekend back injuries and strains in order to obtain medical coverage and indemnity benefits through WC.

In this paper we present a more direct test of the hypothesis that the "Monday effect" in WC claims arises because of higher rates of fraudulent claims. A simple equilibrium model of claim-filing behavior by injured workers and claim-monitoring activity by employers suggests employees who lack medical insurance coverage for off-the-job injuries will file more fraudulent Monday claims, and employers will monitor these claims more carefully. We test these predictions using administrative data on WC claims for the state of Minnesota combined with data on medical insurance coverage patterns from the March Current Population Survey (CPS). A two-sample estimation technique enables us to pool the two data sources and study the effect of insurance coverage on the timing of injury claims and the likelihood that employers challenge their liability for a WC claim.

Our empirical findings suggest that higher Monday injury rates are unrelated to the presence or absence of medical insurance. Although medical coverage rates vary markedly

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[4]The observation that accidents are more likely on Mondays is an old one: see Vernon (1921, chapter 10).

[5]The data set is described more fully in the next section of the paper.

across the population, the fraction of excess Monday injuries is constant across groups. We also find that employers are no more likely to deny liability for Monday accident claims -- even among workers with the lowest rates of off-the-job medical coverage. Both findings call into question the hypothesis that higher Monday injury rates arise from higher fraudulent claim rates on that day.

I. The 'Monday Effect' in Injury Rates -- Theoretical Issues

To set the stage for our empirical analysis it is useful to begin by laying out a simple theoretical model of injury reporting and claims monitoring that incorporates the possibility of fraudulent claims. We model the decisions of workers who are injured off-the-job about whether to initiate a WC claim, and the decisions of employers (or insurers) to monitor potentially fraudulent claims. We focus on two questions: how will the fraction of off-the-job injuries reported on a given workday vary with the insurance status of workers and the rate of off-the-job injuries? and how will the detection rate of fraudulent claims vary with the same factors?

Consider an employer with a stable workforce who expects $\mathbf{v}$ valid injury claims per workday. Let $\mathbf{u}$ represent the number of off-the-job injuries that occur prior to the workday ($\mathbf{u}$ will vary by weekday, with a higher value of $\mathbf{u}$ on Monday). Finally, let $f \leq \mathbf{u}$ represent the number of fraudulent claims arising from off-the-job injuries that are reported to the employer. The fraction of fraudulent claims is $\mathbf{r} = f/(v + f)$. Suppose that the employer (or the insurer) spends an amount $\$\mathbf{e}$ per claim investigating its validity, and assume that this monitoring activity results in a detection probability of $p(e)$ for fraudulent claims. Thus, a total of $p(e)\mathbf{f}$ invalid claims are rejected by the employer. If the cost is \$C for each of the remaining claims, then the employer's total costs are

$$C\,\{\ (1 - p(e))f\ +\ v\ \}\ +\ (f + v)e\ .$$

Assuming that p(e) is strictly concave, and that p'(0) is sufficiently large relative to **C**, the first-order condition $p'(e) = 1/C{\cdot}r$ is necessary and sufficient to describe the firm's optimal choice of **e**, given the rate of fraudulent claims **f** and the other parameters of the model. This implies that

$$(1)\quad e\ =\ \eta(f)\ ,$$

where $\eta'(f) = -\,v\ /\ [\ C\ f^2\ p''(e)\ ]\ >\ 0$. The firm will increase its monitoring effort, the greater the expected number of invalid injury claims.

Assume that an employee's medical insurance covers a fraction α of off-the-job injury costs, where α can range from 0 for an uninsured worker to 1 for an employee with full medical insurance and no deductible. If the worker is injured off-the-job, he or she can either obtain treatment directly, at a net cost of $C(1-\alpha)$, or report to work and file a fraudulent claim. Ignoring any penalty for filing a claim that is ultimately rejected, the net expected cost of this strategy is $\delta + p(e)\{\ C(1-\alpha)\ \}$, where δ represents the subjective cost (i.e. pain and suffering) associated with delaying treatment until sometime after the start of work. An employee who is injured off-the-job will choose to delay treatment and file a false WC claim if

$$\delta\ <\ (1 - p(e))\ C\ (1-\alpha)\ .$$

We assume that δ varies across injuries, depending on the time of the off-the-job injury and other factors. For example, it may be more painful to delay the treatment of an injury that occurs many hours before the start of work. If $H(\delta)$ represents the distribution function of delay costs (for injuries of a given cost level **C**), then

$$(2)\quad f = \phi(e) = u \cdot H(\,(1 - p(e))\, C\, (1-\alpha)\,)\ ,$$

is the expected number of fraudulent claim filings, conditional on the employer's monitoring effort **e** and the other parameters of the model. Note that $\phi'(e) \leq 0$, implying that workers are less likely to file a fraudulent claim the higher the rate of monitoring activity.

Equations (1) and (2) together determine the equilibrium behavior of workers and employers.[6] Three types of solutions are possible: a solution with **f = u** (every worker who is injured off the job files a claim); a solution with f = 0 (no fraud); and an interior solution with $0 < f < u$. For simplicity, we shall concentrate on the latter case. Nevertheless, we suspect that the "no fraud" solution is relevant for many, if not most, off-the-job injuries.[7] Furthermore, for many types of injuries, it is presumably difficult to conceal the fact that the injury occurred off-the-job. Thus our analysis is implicitly focused on injuries like muscle strains and back injuries that are <u>not</u> immediately life-threatening, and that are not necessarily the result of a verifiable accident.

Starting from an equilibrium in which some but not all off-the-job injuries are reported at work, it is easy to show that an increase in the rate of off-the-job injuries (**u**) leads to an increase in the equilibrium number of off-the-job injury filings (f^*) and an increase in the

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[6]We are ignoring the possibility that employees "strategically delay" the filing of an accident claim for more than one day. Technically, such a possibility may arise if the cost of delaying treatment is relatively low, and if the level of claims monitoring activity is so much higher on Monday that employees believe it is worthwhile to delay filing a fraudulent claim until Tuesday.

[7]Examination of equations (1) and (2) shows that the "no fraud" solution is more likely, the higher are the costs of delaying treatment relative to the costs of treatment.

equilibrium level of the employer's monitoring effort (e^*) (i.e., $df^*/du > 0$ and $de^*/du > 0$). The equilibrium rate of "rejected claims" (the fraction of claims that are determined to be fraudulent) is $p(e^*) \cdot f^*/(f^* + v)$. Since $p(e)$ is increasing in e, and both f^* and e^* are increasing with u, an increase in the number of off-the-job injuries will lead to a higher rate of rejected claims.

To interpret these results, consider the comparison between injury claims filed on Monday and those filed on Tuesday (or any other weekday). Assume that the number of on-the-job injuries is the same on different workdays, whereas the number of off-the-job-injuries occurring prior to work is higher on Monday.[8] According to the model, we would then expect to see a greater number of total accident claims on Monday, a higher employer monitoring rate for Monday claims (manifested, for example, by a higher probability that the employer contests the validity of Monday injuries), and a higher fraction of Monday claims that are ultimately rejected. Of course, these predictions depend on the maintained assumption that underlying on-the-job injury rates are the same on Monday as on other work days.

Even if on-the-job accident rates are different on different workdays, it is possible to test for the presence of fraudulent claims by comparing the relative fraction of claims on Mondays for workers with different levels of off-the-job insurance coverage (i.e., different levels of α). In particular, suppose that off-the-job and on-the-job injury rates are similar for all workers,

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[8]Assuming that the off-the-job injury rate is approximately constant per hour, a typical worker with an 8-to-5 Monday-to-Friday work schedule has a 420% higher probability of an off-the-job injury before the start of work on Monday morning than before the start of work on Tuesday morning. The relative rate of weekend injuries may be even larger if weekend activities (sports, home repair) are more likely to result in an injury than activities during a normal weekday evening.

regardless of their medical coverage, and that more off-the-job injuries occur over the weekend. Comparing the magnitudes of the derivatives df^*/du and de^*/du with respect to the level of off-the-job insurance coverage, it is possible to show that

$$\frac{d^2f^*}{d\alpha du} < 0$$

and

$$\frac{d^2e^*}{d\alpha du} < 0 .$$

The first of these inequalities implies that workers with lower off-the-job insurance coverage will have a bigger 'Monday effect' in their injury claims, while the second implies that employers of workers with lower medical coverage rates will expend relatively more resources monitoring their Monday claims. Taking these two predictions together, we would therefore expect to see a greater relative rate of disputed claims on Mondays for workers with lower off-the-job medical coverage, and a greater relative fraction of Monday claims that are ultimately rejected. In the empirical analysis below we test these predictions by comparing the relative probability of a Monday injury among workers with different likelihoods of off-the-job medical insurance, and the relative probability that the employer denies liability for Monday injuries filed by employees with different probabilities of medical insurance.

II. Initial Data Description

Our empirical analysis of Workers' Compensation claims is based on a 10 percent random sample of the "first reports" of injury filed with the Minnesota Department of Labor and Industry between 1985 and 1989. A first report is normally posted for any serious injury, and is legally required for all injuries that result in more than three days of lost work time.[9] The data set thus excludes minor injuries that only required medical treatment and/or up to three days of lost work time. Some 50,000 first reports were filed annually in the mid-1980s in Minnesota, resulting in a total of 25,563 injuries in our sample.

The first column of Table 1 presents descriptive statistics for the overall sample of claims, including a 10 percent subsample for which no wage data are available. The other columns of the table show the characteristics of the subsample of injuries with a valid pre-injury wage, classified by the day of the week on which the injury occurred. The level of wages is a key predictor of the likelihood of medical coverage (see below). Hence, for most of our analysis we concentrate on injuries records with valid wage information.

The first report forms classify injury claims by type of injury (e.g. burn or fracture), body part (e.g. upper back), and cause (e.g. struck by falling object). The most likely injury is a back strain caused by a slip or fall. Interestingly, back injuries, strains, and injuries caused by a slip or fall are all more prevalent on Mondays than on other weekdays.

The average employer and employee characteristics in our injuries sample differ somewhat from the average characteristics of the Minnesota workforce, reflecting the non-random incidence

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[9]Because of a waiting period for disability benefits, injuries that result in no more than 3 days of lost work time (including the day of the injury) do not generate an indemnity claim and do not require a first report of injury.

of injuries across workers and jobs. Construction and manufacturing jobs, for example, are over-represented in the claims sample relative to their shares of total employment in Minnesota, whereas trade and services are under-represented.[10] By the same token, female and white collar workers, who account for 48 percent and 54 percent of the Minnesota labor force, respectively, account for much lower shares of WC claims. The average weekly wage of injured workers ($358) is slightly below the average for all Minnesota workers ($382 per week in March 1987). Virtually all of the wage differential between injured workers and a random sample of workers in the state, however, is explicable by a small set of demographic, industry, and occupation controls (see below).

Rows 14-17 of Table 1 show the percent of injury claims with positive indemnity payments (including temporary total and temporary partial benefits paid to workers during their recovery period, permanent partial benefits paid as lump sums or continuing benefits post-recovery, and other lump sum payments), the mean payment conditional on positive payments, the percent of claims with temporary total benefits, and the mean duration of temporary total disability. The subsample of injuries with a valid wage observation includes a higher fraction of cases with temporary total benefits (71.4 versus 65.9 percent overall). This differential reflects the fact that the temporary total benefit rate is a direct function of the pre-injury wage: the administrative files are therefore more likely to include the injured worker's wage rate in cases where temporary total benefits were paid.

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[10]In the March 1987 CPS, construction and manufacturing account for 5.2% and 20.5% of Minnesota employment, whereas trade and services account for 21.7 and 33.7% of Minnesota employment.

Mean indemnity payments and the duration of benefits are very similar for injuries that occur on Mondays or other weekdays.[11] Weekend injuries, by comparison, have significantly lower mean indemnity payments and significantly shorter benefit periods.[12] In part these differences reflect the higher concentration of weekend injuries among retail trade and service workers, and the lower average severity of injuries in these industries. Even controlling for industry, however, weekend injuries are more likely to involve female, white collar, and lower-wage workers who tend to have lower cost claims. In view of the distinctive character of weekend workers and weekend injuries, we focus exclusively on weekday (i.e. Monday-Friday) injuries in the remainder of this paper.

Across all types of weekday injuries, 22.95% occur on a Monday. If work hours were equally distributed across the weekdays (see below) then one would expect exactly 20 percent of weekday injuries to arise on Mondays. On this assumption, the "excess fraction" of Monday injuries is 2.95% and is significantly different from zero at any conventional significance level.[13] As shown in Figure 1, however, the magnitude of the Monday effect varies from essentially 0 for fractures to approximately 5 percent for strains and back injuries. Appendix Table 1 presents the excess fraction of Monday injuries for 14 categories of injuries, along with t-statistics for the test that one-fifth of weekday injuries occur on Monday. The pattern of the

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[11]A t-test for a difference in the mean indemnity payment between Monday and Tuesday-Friday injuries has a value of 0.66. A t-test for a difference in the corresponding durations of temporary total benefits has a value of 0.39.

[12]A t-test for a difference in the mean indemnity payment between weekend and Tuesday-Friday injuries has a value of 3.78. A t-test for a difference in the corresponding durations of temporary total benefits has a value of 2.86.

[13]The t-statistic for the hypothesis that 20% of all weekday injuries occur on Mondays is 10.77.

data is suggestive: the excess fraction of Monday injuries is higher for non-life-threatening and easy-to-conceal injuries, and lower for directly visible and/or extremely painful injuries like burns, lacerations, and fractures. Work-related occupational injuries (such as carpal tunnel syndrome) are also significantly more likely to occur on Mondays. We hypothesize that this pattern is driven by the arbitrary nature of the injury date for an occupational disease and a tendency to begin a spell of lost work time on Monday.

III. Medical Coverage and the Monday Effect

As illustrated by the simple model in section I, one explanation for the Monday effect in injury rates is that workers post-date their weekend injuries in order to recover their medical costs through the workers' compensation system. A critical check on this interpretation is that Monday injury claims are more likely among workers who lack medical insurance coverage. Unfortunately, our WC claims data set contains no direct information on the medical insurance status of injured workers. We proceed by using a two-sample estimation technique that combines information on medical insurance coverage from the March Current Population Survey with data on the timing of WC injury claims from our administrative data files.[14]

Consider a sample of weekday injury claims, and let $y_i=1$ if the ith injury claim is reported on a Monday, and 0 otherwise. Assume that π_i, the probability that $y_i=1$, is a function of a set of characteristics of the worker involved in the injury (x_i), and an indicator for whether the individual has off-the-job medical coverage (m_i):

[14]Two-sample estimation methods are analyzed by Murphy and Topel (1985), Angrist and Krueger (1992) and Arellano and Meghir (1988).

(3) $\pi_i = x_i'\beta + m_i\gamma$.

Our theoretical model suggests that if the Monday effect is attributable to the fraudulent filing of WC claims for weekend injuries, then $\gamma < 0$, since uninsured workers have a higher incentive to file a false Monday claim. Actual medical coverage is unobserved in our sample of injury claims. Suppose that a secondary sample is available, however, that includes medical coverage information as well as data on a vector of instruments z_i (some of which may be included in x_i) that are correlated with medical insurance coverage status. Let

(4) $P(m_i=1 \mid z_i) = z_i'\theta$.

The coefficients of equation (3) can then be estimated consistently by a simple two-step procedure. The first step is to estimate equation (4) on the secondary sample. In the second step, equation (3) is estimated by ordinary least squares, replacing unobserved medical coverage with its imputed value $(z_i'\hat{\theta})$. This procedure is similar to conventional two-stage least squares, with two important differences: (1) the "first-stage" equation is estimated on the secondary sample, rather than the main sample; and (2) the full set of "exogenous determinants" of π_i (the full set of x's) is not necessarily included in the vector of predictors z_i. Nevertheless, it is easy to show that this two-sample two-stage estimation method is consistent, and to derive appropriate standard errors for the estimated coefficients of equation (3). Details are relegated to the statistical appendix.[15]

[15]This procedure is a special case of the two-step estimation procedure discussed by Murphy and Topel (1985). Our standard error formulas account not only for the estimation of the first-stage equation in the secondary sample, but also for the fact that both (3) and (4) are linear probability models, and are therefore conditionally heteroskedastic.

Our secondary source of medical insurance information is the March 1987 Current Population Survey (CPS). Supplementary questions in this survey enable us to determine whether or not a given individual has any form of medical insurance coverage (through their own job, a government program, or another family member). We fit equation (4) to the CPS subsample of employed individuals in the 12 Midwestern states, using as predictors of medical coverage a quadratic function of age, a set of 3 gender/marital status interaction dummies, 6 occupation dummies, 8 industry dummies, and interactions of the log weekly wage with marital status, gender, and industry. This equation is reasonably successful in predicting medical insurance coverage, with an R-squared coefficient of 0.13.[16] The most important predictors of insurance coverage are the marital status/gender interactions and the wage interaction terms. For reference, the estimated coefficients of the prediction equation are reported in Appendix Table 2.

A maintained assumption in the two-sample procedure is that medical coverage status has the same relationship with the predictor variables in the CPS sample as in the WC claims sample.[17] In order to assess the plausibility of this assumption, we used a similar two-stage procedure to first estimate a weekly <u>wage</u> equation for the CPS sample, and then predict a weekly wage for each individual in the WC claims file.[18] The estimated coefficients from the

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[16]Although the R-squared coefficient may seem low, we emphasize that the prediction equation is a linear probability model in which the mean of the dependent variable is 0.89. In such a model, the R-squared is theoretically bounded far below 1 -- see Morrison (1972).

[17]In other words, if we could estimate equation (4) using observations from the WC claims sample, we would get the same coefficient estimates as we obtained in the CPS sample, apart from sampling error.

[18]We used only age, age-squared, marital status/gender dummies, occupation dummies, and industry dummies to predict the wage.

CPS sample provide a remarkably accurate wage forecast for injured workers. The mean forecast error is less than 0.3 percent; and the correlation of the predicted and actual wages for individuals in the claims file is 0.57. These findings suggest that the two samples are quite similar (conditional on observable worker and job characteristics), and that the assumptions needed to justify the two sample procedure are reasonable.

Table 2 illustrates the variation in medical insurance coverage rates across various employee groups and the corresponding variation in the size of the Monday effect in injury rates. Column 1 gives the percentage of individuals with medical insurance coverage in each group, estimated from the March 1987 CPS sample. Columns 2 and 3 show the percent of all weekday injuries and the percent of all weekday back injuries that occur on Monday for each group. As shown in column 1, medical insurance coverage rates are substantially lower for younger and single workers, and for workers with lower weekly wages. Perhaps surprisingly, however, the fraction of Monday injuries is virtually constant across demographic groups and wage quartiles. These simple tabulations provide little support for the hypothesis that the Monday effect in injury rates is attributable to the post-dating of weekend injuries by uninsured workers.

A potentially stronger test of the link between medical insurance coverage and the Monday effect is obtained by stratifying workers into groups based on their predicted probability of insurance coverage, and then comparing the fraction of Monday injuries across groups. Rows 4a to 4d present medical insurance coverage rates and percentages of Monday injuries for workers grouped into quartiles by the imputed probability of insurance coverage. Again there is no evidence that workers with lower coverage rates have a higher fraction of Monday injuries. This point is illustrated graphically in the upper panel of Figure 2. Here we show the fraction of weekday back injuries on different days of the week for the four coverage-probability quartiles. The distributions of injuries by weekday are quite similar for all four groups, with no indication that workers with a lower probability of insurance are more likely to report a Monday claim.[19]

The Distribution of Work Hours over the Week

An important assumption underlying the comparison of injury rates by day of the week is that the distribution of work hours is constant across weekdays. If the probability of working on Mondays varies with the same characteristics that determine the probability of medical insurance, then the simple comparisons in Table 2 and Figure 2 may be invalid. To assess this possibility we used information on weekly work schedules from the May 1985 CPS to construct a sample of individuals who usually work at least one weekday per week.[20] (People who work only on the weekends are excluded, since these individuals would never report a weekday injury). We then computed the probabilities of working on different weekdays for this sample of weekday workers.

The upper panel of Figure 3 shows the fraction of weekday workers at work by day of the week for employees in each of the 4 quartiles of predicted medical coverage. For the three

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[19]Exactly the same pattern emerges when we include all injuries, and not just back injuries. Since the Monday effect is most pronounced for back injuries, however, we decided to illustrate the patterns for back injury claims.

[20]The "Work Schedule and Dual Job Supplement" of the May 1985 CPS asks each individual which days of the week they normally work on their main job. Our analysis is based on non-self-employed workers who report an hourly or weekly wage for their main job, and report that they usually work at least one regular workday per week.

upper quartiles the probability of being at work on a given weekday is roughly constant. For the lowest quartile, however, the fraction at work rises over the week. Further investigation revealed that this pattern is driven by work schedules in the retail trade industry. Low-wage workers in retail trade have a relatively low rate of medical insurance coverage, and are also less likely to work on Mondays than later in the week. Within the retail trade sector, then, the expected fraction of Monday injuries for workers with a low probability of medical coverage is less than 20 percent. As a result, a comparison of excess Monday injuries by quartile of predicted insurance coverage may fail to show a higher Monday effect for the lowest quartile group, even if these workers are more likely to post-date weekend injuries. An obvious correction for the differential probability of Monday work is to exclude retail trade employees from the analysis. As shown in the lower panel of Figure 3, this exclusion effectively equalizes the probability of working on different weekdays for the first quartile group.

We also conducted a more formal analysis of the relationship between medical insurance coverage rates and the relative probability of working on Monday. Specifically, we fit a series of linear probability models for the event of working on different weekdays (among the sample of people who usually work at least one weekday), including as an explanatory variable the estimated probability of medical coverage ($z_i'\hat{\theta}$) formed from the coefficient estimates of equation (4).[21] As suggested by the pattern in the upper panel of Figure 3, the results show that workers with a higher probability of medical insurance are more likely to work on any weekday. Moreover, the effect of the estimated medical coverage variable is larger on Mondays. This

[21]The estimates are reported in an earlier draft of this paper (Card and McCall (1994)).

pattern persists in models that include demographic and industry controls in addition to the predicted coverage variable. When we exclude retail trade workers from the sample, however, the estimated effect of the medical coverage variable is virtually constant across 5 weekdays. Based on these findings, we conclude that the assumption of an equal distribution of work hours across weekdays is valid, providing that retail trade employees are excluded from the sample.

The lower panel of Figure 2 shows the distribution of weekday injury claims by quartile of predicted medical coverage for the sample of back injuries that excludes retail trade workers. As would be expected from the patterns in Figure 3, the exclusion of retail trade workers raises the fraction of Monday injuries for the lowest quartile group and lowers the fraction of Friday injuries. Nevertheless, there is still no indication that workers with lower medical coverage rates have more Monday injuries. Indeed, workers with the lowest rates of medical insurance coverage have the lowest relative fraction of Monday injury claims.

Although the May CPS data suggest that workers with different rates of medical coverage have similar relative probabilities of working on Mondays, it should be emphasized that these data pertain to scheduled rather than actual work hours. If absentee rates are higher on Mondays, and the differential is correlated with the determinants of medical coverage, then our analysis may understate the effect of medical insurance coverage on Monday injury rates. Holidays are one source of differential absenteeism rates across weekdays. A holiday weekend not only reduces the expected number of Monday injury claims, but may also lead to an increase in the number of Tuesday claims (Smith (1989)). In the analysis below we test for the effect of holidays by comparing specifications that exclude major holidays (Memorial Day, Labor Day, Fourth of July, Thanksgiving), the entire week between Christmas and New Years, and all post-holiday workdays.

A second possibility is that non-holiday-related absences are higher on Mondays than other weekdays. We are aware of only one recent study that reports absenteeism rates by day of the week. This study (Barmby, Orme, and Treble (1991)) concludes that absenteeism rates are about the same or slightly lower on Mondays than on Tuesday-Thursday, and actually peak on Fridays.[22] Given this finding, we present some specifications below that exclude Friday injuries. We have been unable to find any studies or data sources that break down absenteeism patterns by day of the week and demographic characteristics. Thus we cannot directly test whether workers with lower medical coverage rates have higher Monday absenteeism. This limitation must be kept in mind in interpreting our results.

IV. Models for the Relative Probability of a Monday Injury

Table 3 presents estimates of the effect of medical insurance coverage on the relative probability of a Monday injury. Specifically, the table reports estimates of the coefficient of imputed medical coverage from linear probability models for the event of a Monday claim, estimated on various samples of weekday injury claims. Column 1 gives the estimated coverage coefficients from models with no other control variables. The estimates in column 2 are obtained from models that include a set of demographic and industry control variables, while the models in column 3 include 25 additional variables describing the nature and cause of the injury. The

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[22]Barmby, Orme and Treble (1991) analyze data for a single British firm.

estimates in the upper panel of the table are based on samples that include retail trade workers, whereas the estimates in the lower panel are based on samples that exclude these workers.

We emphasize that the coefficients in these conditional probability models measure the effect of insurance coverage on the relative fraction of weekday injuries that occur on Monday: they provide no information on the relation between medical insurance coverage and overall injury rates. In fact, tabulations of the March 1987 CPS suggest that workers without medical insurance coverage have slightly lower overall probabilities of a WC injury claim.[23] However, our interest here is in the effect of medical insurance on the timing of weekday injuries, rather than the overall number of such injuries.

Most of the estimated coefficients in Table 3 are positive -- the opposite of the sign suggested by the hypothesis that workers without medical insurance are more likely to report a Monday injury. Consistent with the fact that retail trade employees with lower coverage rates are less likely to work on Mondays, the exclusion of retail trade workers leads to some reduction in the estimated coverage coefficients. Even when retail trade workers are excluded from the sample, however, the coefficients tend to be close to zero -- even for subsamples of back injuries and sprains.

Two other conclusions emerge from Table 3. First, the exclusion of claims filed on holidays or on the day after a holiday has little effect on the estimation results (compare the estimates in rows 2 with those in row 1). Likewise, redefining the pool of weekday injuries to exclude Friday injuries has no little or no effect on the results (compare row 3 to row 1 or 2).

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[23]In the entire CPS sample of adult workers with earnings in the previous year, 1.76% report receiving WC payments. This fraction is 1.57% for workers without medical insurance coverage and 1.79% for workers with medical coverage.

Second, although we expected to see a bigger effect of insurance coverage on the weekly pattern of back injuries and strains, the data do not confirm this prediction. If anything, our findings suggest that the presence of medical insurance coverage increases the relative fraction of back injury claims on Monday (see rows 5 and 8).

We have estimated a variety of alternative specifications to probe the robustness of these conclusions. In particular, we investigated the effects of adding two additional control variables to our analysis: the pre-injury wage; and a set of dummy variables representing the worker's benefit-replacement rate while on temporary disability. Our analysis of the replacement rate is motivated by the observation that employees with higher replacement rates who are injured off the job have a stronger incentive to file a fraudulent claim and receive temporary disability payments, rather than work through the recovery period. It is therefore interesting to check whether our inferences about the effect of medical coverage on the magnitude of the Monday effect are robust to the inclusion of measures of the replacement rate.

In Minnesota, the WC benefit rate is fixed at two-thirds of the pre-injury wage, subject to a maximum and minimum linked to the state average weekly wage.[24] The combination of minimum and maximum rates implies that the replacement rate falls into 5 ranges: greater than 1 (for the small percentage of workers who earn less than 20 percent of the state average weekly wage); exactly 1 (for the 10 percent of workers whose wage is between 20 and 50 percent of the state average weekly wage); between 2/3 and 1 (for the 20 percent of workers who earn between

[24]Minnesota laws during our sample period set a subminimum benefit ($75.20 per week in October 1987) as a lower bound on all benefits, and a primary minimum such that claimants whose benefits would be below the primary minimum under the two-thirds formula receive the lower of the primary minimum benefit amount and their weekly wage.

50 and 75 percent of the state average wage); exactly 2/3 (for roughly 50 percent of workers who earn between 75 and 150 percent of the state average wage); and less than 2/3 (for the 20 percent of workers who earn more than 150 percent of the state average wage).

Our findings from these extended specifications are presented in Table 4. For brevity we report only the results obtained on samples that exclude workers in the retail trade industry. (Results for the overall sample are similar). In general, neither the level of wages nor the range of the benefit replacement rate exerts an independent effect on the probability of a Monday claim, and the addition of these variables has no effect on our conclusion that medical coverage is unrelated to the Monday effect in WC claims.

V. Denial of Liability

Just as employees who are injured off the job have an incentive to file fraudulent WC claims, employers and insurers have an incentive to screen out these claims. In Minnesota, employers who intend to dispute the validity of a claim begin the process by filing a "Notice of Denial of Liability".[25] The pattern of denial rates by day of the week and probability of medical coverage provides further evidence on the hypothesis that the Monday effect in injury rates is attributable to the post-dating of weekend injuries by uninsured workers. As emphasized by our theoretical model, if the Monday effect is attributable to fraud, we would expect employers to monitor Monday claims of uninsured workers more carefully, and to be more likely to deny liability for their injuries.

[25]See chapter 4 of Minnesota House of Representatives Research Department (1988) for an analysis of litigation involving issues of primary liability in the WC system of the state.

Employers filed a notice of denial of liability for about 10 percent of the injury claims in our sample. Figure 4 shows the rates of denial of liability for back injuries by day of the injury and quartile of predicted medical coverage. Perhaps surprisingly, the overall denial rate is no higher for Monday injuries than for injuries on other weekdays, even for injured workers with the lowest probabilities of medical coverage. These patterns do not suggest that employers and/or insurers are more likely to question the legitimacy of Monday claims by groups of employees with low insurance rates (or indeed by any group of employees).

We have also conducted a more formal analysis of the determinants of the probability of denying liability, based on the following model:

(5) $P(\text{deny liability}) = x_i'\mathbf{a} + m_i\mathbf{b} + \text{Monday}_i\mathbf{c} + m_i{*}\text{Monday}_i\mathbf{d}$,

where x_i is a vector of characteristics of the ith injury claim, m_i is an indicator for whether or not the worker who filed the claim has medical insurance coverage, and Monday_i is an indicator for a Monday injury. The coefficient **d** measures the relative effect of medical coverage on the probability that the employer denies liability for a Monday injury. Our theoretical model suggests that **d** is negative, since employers will monitor the Monday injury claims of uninsured workers more closely if these workers are more likely to claim fraudulent Monday injuries. As in our analysis of Monday injury rates, we can estimate equation (5) by replacing m_i with a consistent estimate of the probability of medical coverage $(z_i'\hat{\theta})$. The results of this exercise are reported in Table 5.

Columns 1-3 of Table 5 present estimation results for the overall injury sample. The model in column 1 excludes any additional control variables, while the model in column 2 adds controls for the characteristics of the injury and the worker, as well as dummy variables for injuries reported on holidays, post-holiday workdays, and Fridays. Finally, the model in column 3 adds the injured worker's weekly wage, and indicators for the range of the worker's benefit-replacement rate (RR). Parallel sets of models are reported in columns 4-6 for the subsample of claims that excludes injuries on holidays, post-holiday workdays, and Fridays, and in columns 7-9 for the subsample of back injuries.

With respect to the presence of a Monday effect in denial rates, the results in Table 5 are clear-cut. There is no indication of higher denial rates for Monday injuries, nor of a differential Monday effect in the denial rate for uninsured workers. Contrary to our expectations, employers do not seem to scrutinize Monday injuries more carefully than injuries on other weekdays.

On the other hand, the results suggest that employers are more likely to deny liability for the injuries of uninsured workers, and particularly workers with higher replacement rates, regardless of the day of their injury. The models in columns 2, 5, and 8 show a highly significant reduction in denial rates for insured workers. Once controls for the wage and replacement rate are introduced (columns 3, 6, and 9), the insurance coverage effect falls in magnitude and is no longer statistically significant. In these specifications, however, the replacement rate variables are highly significant, and show a consistent pattern of higher denial rates for workers with higher replacement rates.[26]

One explanation for these findings is that workers with higher replacement rates are more likely to file questionable or fraudulent injury claims (on any day of the week) in hope of beginning a spell of WC benefits. We would then expect to see a higher probability that

[26]We also estimated specifications that included interactions of the Monday indicator with indicators for the different ranges of the replacement rate. These models show no indication of a differential Monday effect in denial rates for workers with different replacement rates.

employers contest the injury claims of workers with higher replacement rates.[27] Even if workers with different replacement rates have the same probability of filing a fraudulent injury claim, however, employers may be more likely to dispute the claims of workers with higher replacement rates, since the employer's net cost of an injury spell (the WC benefit minus the savings in wages) is higher for these workers. Based on the evidence in Table 5 it is difficult to distinguish between these alternative explanations.

In summary, the patterns of denial for liability of WC injury claims show virtually no evidence of a Monday effect, nor of bigger Monday effect for workers who lack medical coverage for their off-the-job injuries. It is hard to imagine why employers do not dispute a higher fraction of Monday injury claims, if a higher fraction of these claims are truly fraudulent. Thus, the absence of a Monday effect in denial rates is consistent with our findings on the relative rate of Monday injuries for workers with higher and lower probabilities of medical insurance. In neither case do the results support the view that the higher overall rate of Monday injuries is driven by a higher rate of fraudulent claims.

V. Summary and Conclusions

This paper is motivated by a simple observation: certain types of injuries are more likely to arise on Mondays than on other weekdays. This 'Monday effect' has been interpreted as evidence that some employees who are injured off-the-job during the weekend report their

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[27]Chelius (1982) analyses the effect of replacement rates on the frequency of WC injury claims, and argues that a higher rate of injuries for workers with higher replacement rates may reflect lower safety incentives for these workers. The higher injury rate may also reflect a higher rate of fraudulent claims among workers with higher replacement rates.

injuries as having occurred at work (Smith (1989)). Workers without medical insurance have a particularly strong incentive to "postdate" weekend injuries and file an injury claim on Monday. To evaluate the effect of this incentive we use a two-sample estimation strategy to combine injury data by day of the week from the Minnesota Workers' Compensation system with medical insurance coverage data from the March Current Population Survey. Contrary to our expectations, we find that employees with low rates of medical insurance coverage are no more likely to file a Monday injury claim than other workers.

One explanation for this finding is that employees with low probabilities of medical insurance coverage are less likely to work on Mondays. Indeed, low-wage workers in retail trade have below-average medical coverage rates and are less likely to work earlier in the week. When we exclude retail trade employees from our analysis, however, we continue to find that medical insurance coverage rates are unrelated to the relative fraction of Monday injuries. We also check for the effect of holiday weekends by excluding injuries filed on major holidays and post-holiday work days. Again, we find no indication that workers who lack medical insurance file more Monday injury claims.

Just as employees have an incentive to report off-the-job injuries as having occured at work, employers and insurers have an incentive to screen out fraudulent claims. In fact, employers deny liability for about 10 percent of the injury claims in our sample. If a higher fraction of Monday injuries are fraudulent, we would expect to see higher denial rates for these injuries, especially for injured workers with the lowest probabilities of off-the-job medical insurance. Consistent with our conclusions based on the Monday effect in injury rates, however, we find that employers are no more likely to deny liability for Monday injury claims -- even for workers with low probabilities of medical coverage.

These findings point to two conclusions. First, the interpretation of the 'Monday effect' in injury rates as evidence of fraudulent claim behavior may be inappropriate. A higher fraction of back sprains, strains, and similar injuries occurs on Monday than other weekdays. However, these injuries are very evenly distributed across the workforce, and are not associated with a higher level of employer screening activity. An alternative explanation for the 'Monday effect' is that a higher fraction of sprains and strains truly arise on Mondays: perhaps as a consequence of the return to work after a weekend hiatus. Recent research suggests that a similar Monday effect arises in the weekly pattern of heart attacks among the working population -- an effect that is surely unrelated to fraud (Willich et al (1994)). We believe that the evidence on the Monday effect in injury claims is more consistent with a physiologically-based explanation than with alternatives based on fraudulent claim filing.

Second, concern that the Workers' Compensation system is covering the costs of off-the-job injuries for workers without medical insurance has led to growing interest in "24 hour" coverage plans and other alternatives to the current WC system. Our findings suggest that more evidence is needed to firmly establish the rate of fraudulent claim activity, and to evaluate the benefits of any reform in the WC insurance system.

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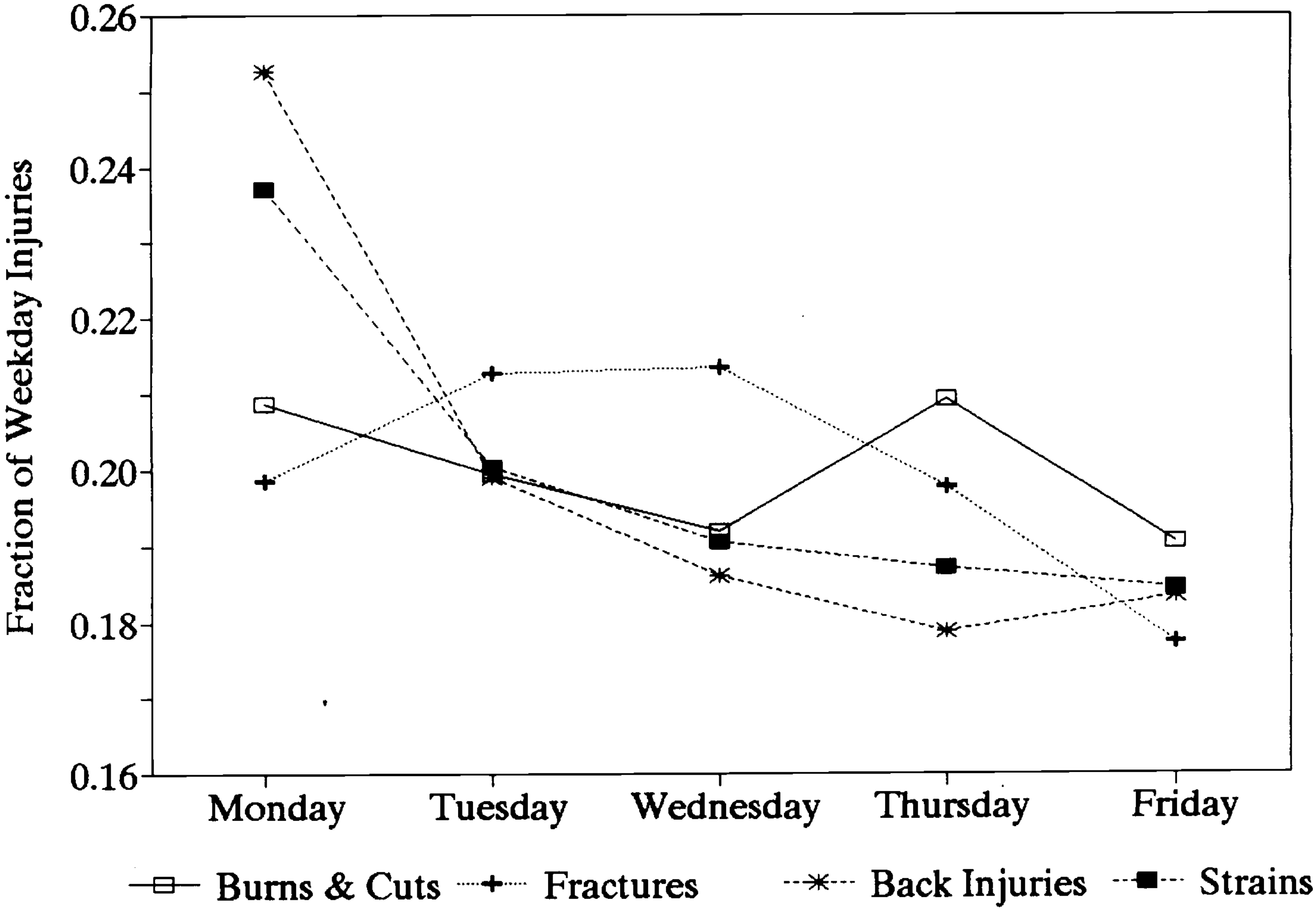


Figure 1: Injury Claim Distribution

weekday wk1

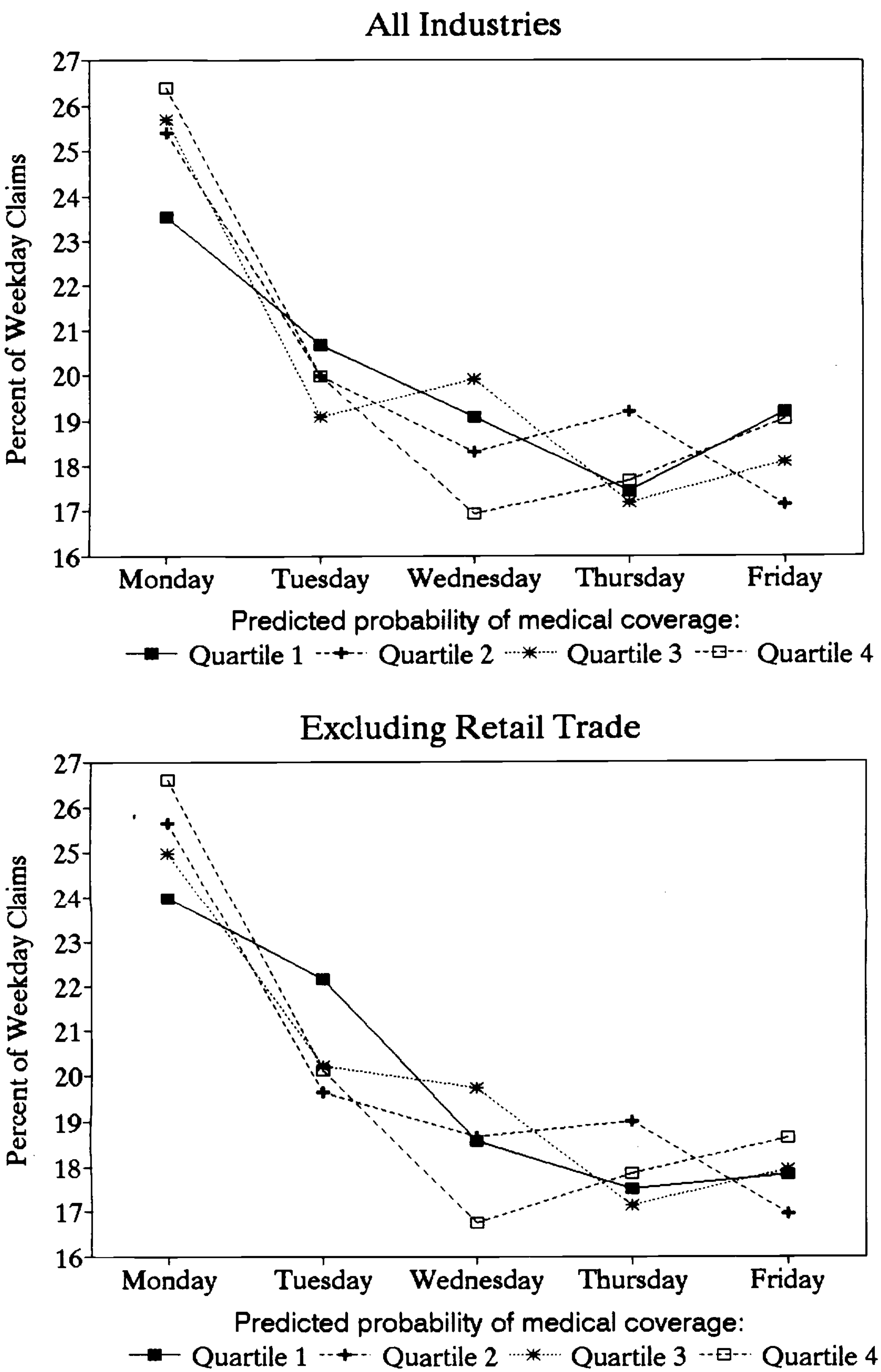
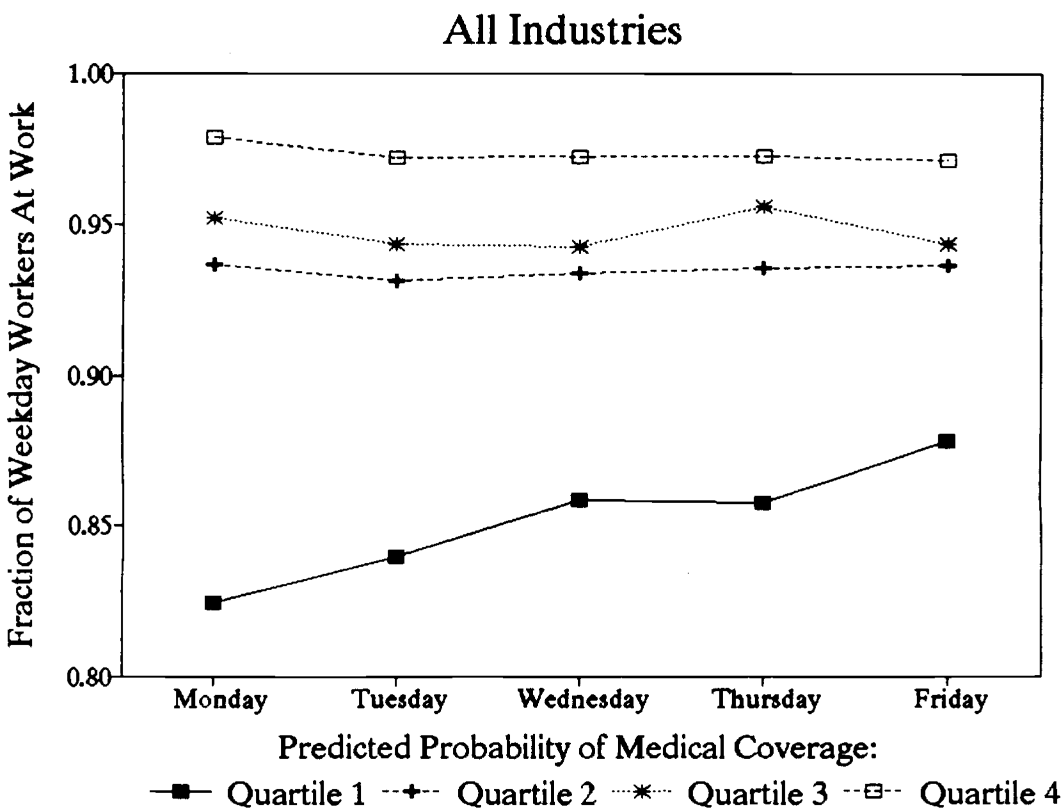


Figure 2: Distribution of Weekday Back Injuries, By Quartile of Predicted Medical Coverage



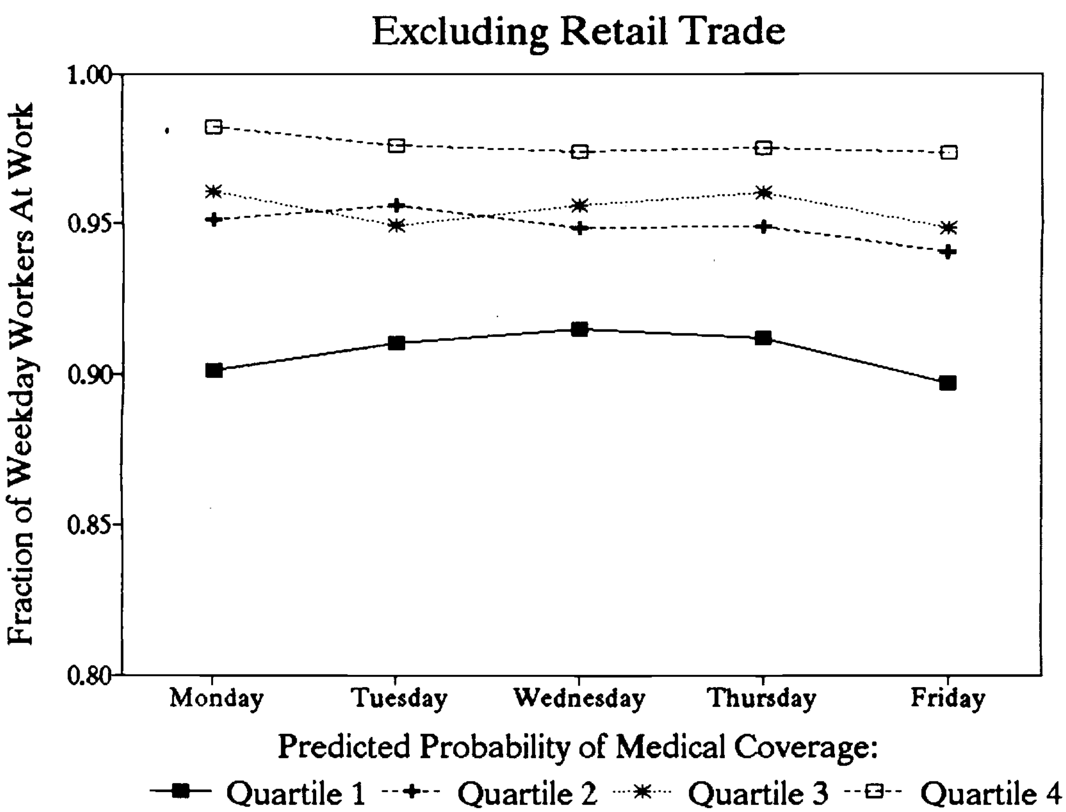


Figure 3: Fraction of Weekday Workers At Work, By Quartile of Predicted Medical Coverage

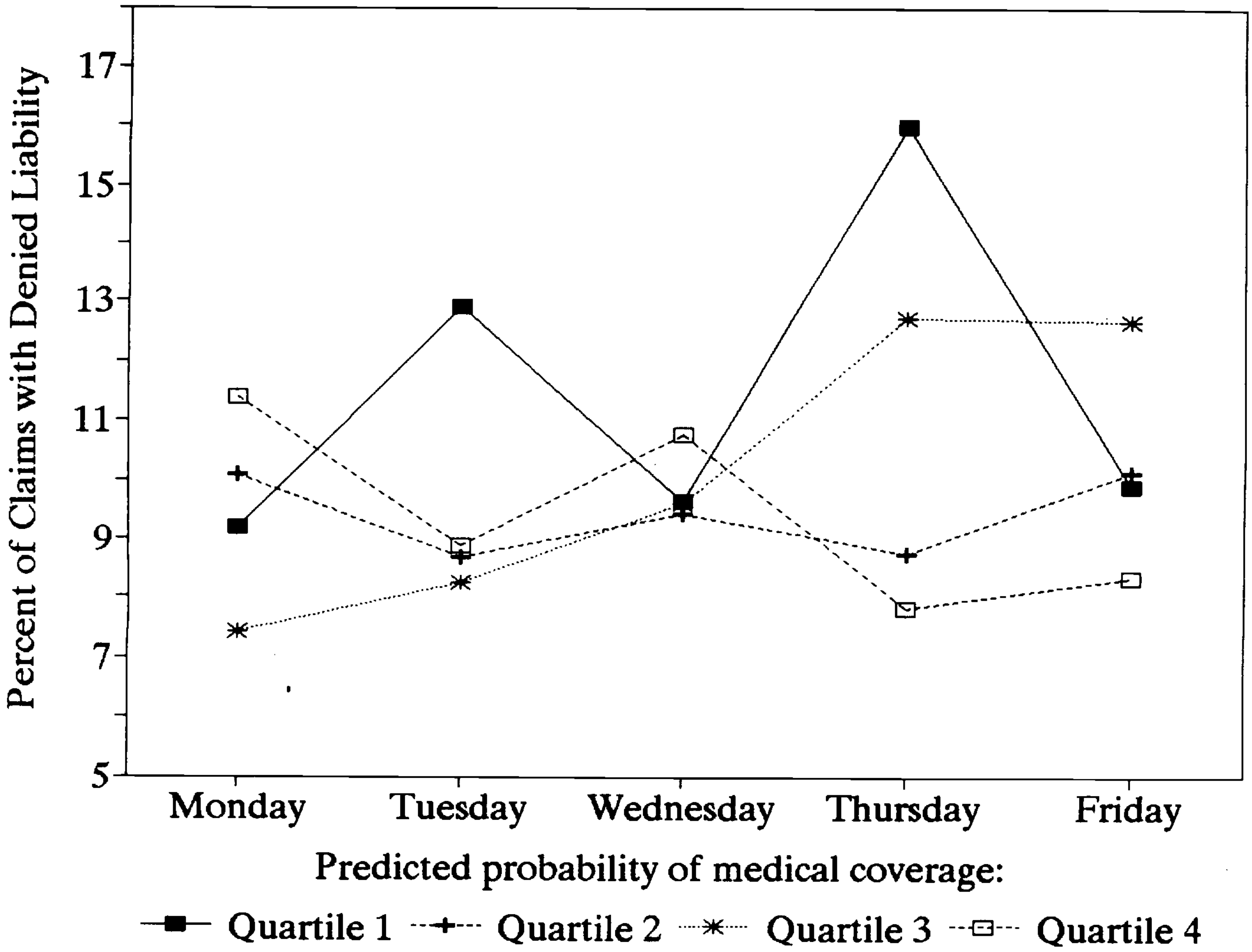


Figure 4: Rate of Denial of Liability for Back Injuries, By Quartile of Predicted Medical Coverage

Table 1: Characteristics of Injuries by Availability of Wage Data and Day of Injury

	All Injuries	Injuries With Valid Wage Data: All	Monday	Tuesday-Friday	Weekend
Injury Characteristics:					
1. Percent Back Injuries	29.8	31.3	34.3	30.2	33.1
2. Percent Burns & Cuts	13.4	13.2	12.0	13.6	13.4
3. Percent Fractures	5.6	5.9	5.2	6.2	5.6
4. Percent Strains	43.0	45.0	46.4	44.4	47.0
Employer Characteristics:					
5. Percent Self-insured	20.3	20.6	21.1	20.3	21.4
6. Percent Construction	11.3	11.9	12.6	12.6	5.0
7. Percent Manufacturing	29.7	31.5	33.6	32.7	19.1
8. Percent Trade	18.9	19.2	18.3	18.3	27.3
9. Percent Services	22.9	22.8	21.7	21.5	33.8
Employee Characteristics:					
10. Percent Female	31.1	31.0	29.3	29.9	42.2
11. Average Age	35.1	35.2	35.4	35.3	33.6
12. Percent White Collar	35.4	35.4	32.0	33.1	58.7
13. Average Weekly Wage	$358	358	367	362	312
Claim Characteristics:					
14. Percent With Indemnity	71.2	75.8	76.8	75.6	76.1
15. Mean Indemnity Amount (for positive claims)	$6488	6336	6667	6429	4998
16. Percent With Temporary Total (TT) benefits	65.9	71.4	72.9	71.2	71.3
17. Mean Duration of TT Benefits (weeks)	10.8	10.7	10.7	10.9	9.2
18. Sample Size	26,563	23,747	4,892	16,422	2,360

Notes: Sample consists of 10% sample of injuries reported to Minnesota Department of Labor and Industry between 1985 and 1989.

Table 2: Probability of Medical Coverage and Relative Probability of Monday Injury, by Worker's Characteristics

	Probability of Medical Coverage (1)	Percent of Weekday Injuries On Monday (2)	Percent of Weekday Back Injuries On Monday (3)
1. All Workers	89.4	23.0	25.3
2. By Age/Marital Status/Sex:[a]			
a. Younger Single Men	74.5	23.1	25.9
b. Older Single Men	85.6	22.9	25.2
c. Younger Married Men	89.9	23.0	26.5
d. Older Married Men	96.2	23.3	25.8
e. Younger Single Women	77.2	22.6	24.6
f. Older Single Women	85.7	22.5	21.8
g. Younger Married Women	91.9	23.7	25.7
h. Older Married Women	95.2	22.3	25.0
3. By Quartile of Weekly Wage:			
a. Quartile 1	76.7	23.0	23.2
b. Quartile 2	87.0	22.2	25.8
c. Quartile 3	95.7	22.6	25.9
d. Quartile 4	97.5	24.0	26.1
4. By Quartile of Predicted Probability of Medical Coverage:[b]			
a. Quartile 1	69.1	22.2	23.5
b. Quartile 2	89.1	23.5	25.4
c. Quartile 3	95.1	22.6	25.7
d. Quartile 4	98.0	23.5	26.4

Notes: Entries in column 1 are for Midwestern workers in March 1987 Current Population Survey who report earnings and weeks of work for the previous year. Entries in columns 2-3 are for injuries in Minnesota during 1985-89.

[a] Younger workers are those with age under 30. Older workers are those age 30 or higher.

[b] Probability of medical coverage is imputed using data on age, gender, marital status, average weekly wage, industry, and occupation. Individuals are then sorted into quartiles based on their predicted probability of medical coverage.

Table 3: Estimated Effect of Medical Coverage on the Conditional Probability of a Monday Injury

		Models with Additional Controls for:	
	No Controls (1)	Worker Demographics & Industry (2)	Injury Type/Cause, Worker Demographics & Industry (3)
I. All Industries:			
1. All Injuries	0.04 (0.03)	0.06 (0.04)	0.05 (0.05)
2. Exclude Major Holidays and Post-Holidays	0.04 (0.03)	0.07 (0.04)	0.06 (0.05)
3. Exclude Fridays, Major Holidays and Post-Holidays	0.05 (0.03)	0.08 (0.05)	0.06 (0.06)
4. Back Injuries Only (All dates)	0.09 (0.06)	0.16 (0.09)	0.16 (0.09)
5. Sprains Only (All dates)	0.03 (0.05)	0.05 (0.08)	0.05 (0.08)
II. Excluding Retail Trade:			
6. All Injuries	0.02 (0.03)	0.01 (0.05)	-0.02 (0.06)
7. Exclude Fridays, Major Holidays and Post-Holidays	0.03 (0.04)	0.03 (0.07)	-0.01 (0.07)
8. Back Injuries Only (All dates)	0.07 (0.06)	0.12 (0.10)	0.11 (0.10)
9. Sprains Only (All dates)	0.02 (0.05)	-0.02 (0.09)	-0.02 (0.09)

Notes: Standard errors, corrected for heteroskedasticity and two-step estimation method (see text) are in parentheses. Table entries are estimated coefficients of imputed medical coverage from linear probability models for the event of a Monday injury, estimated on the sample of weekday injuries. Models in column 2 include control variables for gender, age, age-squared, marital status (interacted with gender), industry (8 categories), and occupation (6 categories). Models in column 3 include 25 additional controls for the nature and cause of the injury.

Table 4: Estimated Effect of Medical Coverage on the Conditional Probability of a Monday Injury -- Further Results

	All Injuries			Back Injuries		
	(1)	(2)	(3)	(4)	(5)	(6)
1. Medical Coverage (Imputed)	0.047 (0.065)	0.048 (0.059)	0.046 (0.065)	-0.035 (0.200)	0.065 (0.164)	-0.030 (0.199)
2. Log Weekly Wage	-0.007 (0.011)	--	0.001 (0.018)	0.023 (0.027)	--	0.030 (0.038)
3. <u>Replacement Rate:</u>[a]						
a. RR > 1	--	0.074 (0.032)	0.075 (0.043)	--	0.010 (0.067)	0.044 (0.081)
b. RR = 1	--	0.010 (0.013)	0.010 (0.020)	--	-0.006 (0.028)	0.013 (0.038)
c. RR Between 0.67 and 1	--	0.010 (0.009)	0.010 (0.012)	--	-0.011 (0.018)	-0.001 (0.022)
d. RR < .66	--	0.017 (0.010)	0.016 (0.012)	--	0.012 (0.020)	0.004 (0.022)

Notes: Standard errors, corrected for heteroskedasticity and two-step estimation method, are in parentheses (see text). Table entries are estimated coefficients from linear probability models for the event of a Monday injury, estimated on the sample of weekday injury claims, <u>excluding claims in retail trade</u>. All models include controls for gender, age, age-squared, marital status (interacted with gender), industry (8 categories), and occupation (6 categories).

[a] Replacement rate (RR) is the ratio of the injured worker's weekly benefit amount to his or her pre-injury wage. The replacement rate is statuatorily determined as a function of the pre-injury wage. An indicator for individuals with a replacement rate equal to two thirds is excluded.

Table 5: Estimated Effect of Medical Coverage on the Probability that the Employer Denies Liability for the Injury

	All Injuries			All Injuries Excluding Fridays Holidays, etc[a]			Back Injuries		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
1. Monday Injury (1=Yes)	0.01 (0.04)	0.02 (0.04)	0.01 (0.03)	0.01 (0.04)	0.01 (0.04)	0.01 (0.04)	-0.08 (0.03)	-0.10 (0.08)	-0.07 (0.09)
2. Medical Coverage (imputed)	0.06 (0.02)	-0.14 (0.04)	-0.07 (0.07)	0.05 (0.02)	-0.15 (0.04)	-0.08 (0.07)	-0.08 (0.05)	-0.17 (0.07)	-0.07 (0.14)
3. Monday Injury * Medical Coverage	-0.02 (0.05)	-0.02 (0.04)	-0.02 (0.04)	-0.01 (0.05)	-0.02 (0.05)	-0.02 (0.04)	0.09 (0.09)	0.11 (0.09)	0.07 (0.10)
4. Log Weekly Wage	--	--	0.03 (0.01)	--	--	0.03 (0.01)	--	--	0.04 (0.02)
5. Replacement Rate:[b]									
a. RR > 1	--	--	0.09 (0.02)	--	--	0.09 (0.01)	--	--	0.16 (0.05)
b. RR = 1	--	--	0.06 (0.01)	--	--	0.06 (0.01)	--	--	0.07 (0.02)
c. RR Between 0.67 and	--	--	0.02 (0.01)	--	--	0.02 (0.01)	--	--	0.03 (0.01)
d. RR < .66	--	--	-0.02 (0.01)	--	--	-0.02 (0.01)	--	--	-0.02 (0.01)
6. Controls for Personal and Injury Characteristics[c]	no	yes	yes	no	yes	yes	no	yes	yes
7. Controls for Fridays, Holidays, and Post-Holidays[d]	no	yes	yes	no	yes	yes	no	yes	yes

Notes: Estimated standard errors corrected for heteroskedasticity and two-step estimation method (see text) are in parentheses. Models are linear probability models for the event that the employer files a Denial of Liability form, disclaiming responsibility for the injury.

Notes continue...

Notes to Table 5, continued

[a]Sample excludes all injury claims filed on Fridays, major holidays, or the weekday immediately following
a major holiday.

[b]Replacement rate (RR): see note to Table 4.

[c]Controls for gender, age, marital status, industry, occupation, and nature and cause of the injury.

[d]Controls for injury claims filed on Fridays, major holidays, or the weekday immediately following
a major holiday.

Appendix Table 1: Excess Fraction of Weekday Injuries on Monday, By Nature of the Injury

	Number of Injuries	Percent on Monday	T-test for Excess Mondays[a]
1. Burns	443	19.19	0.43
2. Lacerations	2,375	21.18	1.44
3. Fractures	1,274	19.86	0.12
4. Strains	9,560	23.73	9.12
5. Contusions	1,453	23.33	3.17
6. Dislocations	602	24.75	2.91
7. Hernias	416	20.91	0.46
8. Inflammations	642	23.21	2.03
9. Heart/Vascular	145	19.31	0.21
10. Occupational Injuries	860	26.51	4.77
11. Insect Bites, Ulcers Seizures	153	22.88	0.89
12. Other Injuries: Amputations, Electric Shock, Concussions, etc	935	21.93	1.48
13. Unknown	2,312	23.17	3.81
14. Not Coded	143	18.18	0.54
15. All Injuries	21,314	22.95	10.77

Notes: Based on sample of weekday injuries with valid weekly wage data. See Table 1 for source.

[a]T-test for hypothesis that 20% of injuries occur on Monday.

Appendix Table 2: Estimated Linear Probability Model for Medical Coverage

Explanatory Variable	Coefficient	Explanatory Variable	Coefficient
Constant	0.431 (0.141)	Log Wage*Agriculture & Mining	0.083 (0.024)
Female (1=yes)	-0.105 (0.084)	Log Wage*Construction	0.233 (0.024)
Age (coefficient x10)	-0.061 (0.015)	Log Wage*Manufacturing	0.127 (0.013)
Age Squared (coefficient x 1000)	0.082 (0.017)	Log Wage*Utilities & Transportation	0.168 (0.021)
Married (1=yes)	-0.516 (0.075)	Log Wage*Wholesale Trade	0.134 (0.021)
Sales and Clerical Occupation	0.004 (0.007)	Log Wage*Retail Trade	0.120 (0.013)
Craft Occupation	-0.031 (0.009)	Log Wage*Finance, Insurance, Real Estate	0.120 (0.019)
Operative Occupation	-0.024 (0.010)	Log Wage*Public Adminstration	0.138 (0.012)
Transportation Operative Occupation	-0.067 (0.014)	Log Wage*Service Industry	0.105 (0.020)
Laborer Occupation	-0.037 (0.014)	Log Wage*Married	-0.071 (0.012)
Service Occupation	-0.053 (0.007)	Log Wage*Married* Female	-0.045 (0.017)

Note: Table continues.

Appendix Table 2, continued

Explanatory Variable	Coefficient	Explanatory Variable	Coefficient
Construction Industry	-0.924 (0.190)	Log Wage*Female* Agriculture & Mining	-0.004 (0.011)
Manufacturing Industry	-0.186 (0.143)	Log Wage*Female* Construction	0.031 (0.008)
Utilities & Trans- portation Industry	-0.456 (0.177)	Log Wage*Female* Manufacturing	0.004 (0.003)
Wholesale Trade Industry	-0.230 (0.178)	Log Wage*Female* Utilities & Transport	0.009 (0.004)
Retail Trade Industry	-0.186 (0.145)	Log Wage*Female* Wholesale Trade	0.005 (0.005)
Finance, Insurance & Real Estate Industry	-0.150 (0.169)	Log Wage*Female* Retail Trade	--
Public Administration Industry	-0.287 (0.140)	Log Wage*Female* Finance, Insurance	0.008 (0.004)
Service Industry	-0.015 (0.172)	Log Wage*Female* Public Administration	0.007 (0.003)
Female*Married	0.234 (0.010)	Log Wage*Female* Service Industry	-0.004 (0.004)
Log Wage*Female	0.022 (0.015)	Minnesota Resident	0.007 (0.009)

Notes: Standard errors corrected for heteroskedasticity are in parentheses. Dummies for managerial and professional occupations and agriculture and mining industry are excluded. Estimated on 15,701 Midwest workers in March 1987 CPS. Mean probability of any medical coverage is 0.894.

Statistical Appendix

This appendix outlines the two-step procedure used in the estimation of equation (3) in the text. To recapitulate the model, let y_i denote an indicator variable for the event that the ith weekday injury occurs on a Monday and assume that

$$P(y_i = 1 \mid x_i, z_i, m_i) = x_i'\beta + m_i\gamma$$

where x_i is a vector of observable characteristics of the injury, m_i is an indicator which equals 1 if the injured worker has medical insurance, and z_i is a vector of other characteristics (possibly including some or all of the elements of x_i) that are correlated with medical coverage but uncorrelated with the relative probability of a Monday injury, conditional on x_i and m_i. This equation implies that

$$\text{(A1)} \quad y_i = u_i'\alpha + \varepsilon_i ,$$

where $u_i' = (x_i', m_i)$, $\alpha' = (\beta', \gamma)$, and ε_i is a conditionally heteroskedastic error with $E(x_i \varepsilon_i) = E(z_i \varepsilon_i) = 0$. The variables y_i and x_i are observed in the primary sample (i.e., the injury claims sample). Medical coverage is unobserved in the primary sample, but is observed in the secondary sample (i.e., the CPS sample). Assume that

$$P(m_i = 1 \mid x_i, z_i) = z_i'\theta ,$$

where z_i includes at least one element that is excluded from x_i. This expression implies that

$$\text{(A2)} \quad m_i = z_i'\theta + \mu_i .$$

where μ_i is a conditionally heteroskedastic error with $E(z_i \mu_i) = E(x_i \mu_i) = 0$. We assume that the data vector (y_i, x_i, z_i, m_i) is independently and identically distributed in the primary and secondary samples with finite fourth moments. The joint distribution of (y_i, x_i, z_i, m_i) is not necessarily the same in the two samples, but we assume that (A2) is valid in both samples (for the same value of θ).

Let $\hat{\theta}$ denote the (consistent) estimate of the coefficient vector θ when equation (A2) is estimated by OLS in the secondary sample, and let V_θ denote the asymptotic variance-covariance matrix of $\hat{\theta}$:

$$\sqrt{n_2}\ (\theta - \hat{\theta}) \xrightarrow{d} N(0,\ V_\theta)\ ,$$

where n_2 is the sample size of the secondary sample. Finally, let

$$\hat{u}_i' = (\ x_i',\ z_i'\hat{\theta}\ )$$

represent the vector of observed covariates for observations in the primary sample, when unobserved medical insurance status is replaced by its imputed value using the coefficient vector $\hat{\theta}$.

Equation (A1) implies that

$$y_i = \hat{u}_i'\alpha + (m_i - \hat{m}_i)\gamma + \varepsilon_i\ .$$

The two-stage two-sample estimator of α is

$$\begin{aligned}\hat{\alpha} &= (1/n_1\ \Sigma_i\ \hat{u}_i\ \hat{u}_i'\ )^{-1}\ 1/n_1\ \Sigma_i\ \hat{u}_i\ y_i \\ &= \alpha + (1/n_1\ \Sigma_i\ \hat{u}_i\ \hat{u}_i'\ )^{-1}\ 1/n_1\ \Sigma_i\ \hat{u}_i\ (\ (m_i - \hat{m}_i)\gamma + \varepsilon_i\ ) \\ &= \alpha + (1/n_1\ \Sigma_i\ \hat{u}_i\ \hat{u}_i'\ )^{-1}\ 1/n_1\ \Sigma_i\ \hat{u}_i\ (\ z_i'(\theta - \hat{\theta})\gamma + \mu_i\gamma + \varepsilon_i\ )\ .\end{aligned}$$

Hence:

$$\text{(A3)}\quad \hat{\alpha} - \alpha = (1/n_1\ \Sigma_i\ \hat{u}_i\ \hat{u}_i'\ )^{-1}\ \{\ 1/n_1\ \Sigma_i\ \hat{u}_i\ z_i' \cdot (\theta - \hat{\theta})\gamma + 1/n_1\ \Sigma_i\ \hat{u}_i\ (\ \mu_i\gamma + \varepsilon_i\ )\ \}\ .$$

Let h denote the ratio of the sample sizes n_1/n_2. We assume that the samples are drawn in such a way as to insure that as n_1 tends to infinity, h tends to a finite constant h^o. Under this assumption we can consider asymptotic arguments as $n_1 \to \infty$ and $n_1/n_2 \to h^o$. Let $\nu_i' = (x_i', z_i'\theta)$, and assume that

$$1/n_1 \Sigma_i \nu_i \nu_i'$$

and

$$1/n_1 \Sigma_i \nu_i z_i'$$

both tend to finite matrices of full row rank as $n_1 \to \infty$, where the sums are taken over observations in the primary sample. Denote these limiting matrices by M_1 and M_2, respectively. Note that $\hat{u}_i = \nu_i + \xi_i$, where $\xi_i = (0, z_i'(\hat{\theta} - \theta))$. It follows from the assumptions made so far that as $n_1 \to \infty$ with $n_1/n_2 \to h^o$,

$$1/n_1 \Sigma_i \hat{u}_i \hat{u}_i' \to M_1,$$

$$1/n_1 \Sigma_i \hat{u}_i z_i' \to M_2,$$

and

$$\text{plim} \{ 1/n_1 \Sigma_i \hat{u}_i (\mu_i\gamma + \varepsilon_i) \} = 0.$$

These facts, together with the consistency of $\hat{\theta}$, establish the consistency of $\hat{\alpha}$.

To derive the asymptotic distribution of $\hat{\alpha}$, note that (A3) implies:

$$\sqrt{n_1} (\hat{\alpha} - \alpha) = \{1/n_1 \Sigma_i \hat{u}_i \hat{u}_i' \}^{-1} \{ \sqrt{h} \cdot 1/n_1 \Sigma_i \hat{u}_i z_i' \cdot \sqrt{n_2}(\theta - \hat{\theta})\gamma + 1/\sqrt{n_1} \Sigma_i \hat{u}_i (v_i\gamma + \varepsilon_i) \} .$$

Finally, let

$$C = E \{ \nu_i \nu_i' (v_i\gamma + \varepsilon_i)^2 \} ,$$

where the expectation is taken with respect to the joint distribution of

(y_i, x_i, z_i, m_i) in the primary sample. Then $\sqrt{n_1}(\alpha - \hat{\alpha}) \overset{d}{\to} N(0, D)$, where

$$D = M_1^{-1} (h^o M_2 V_\theta M_2' \gamma^2 + C) M_1^{-1} .$$

Under the preceding assumptions M_1 and M_2 can be consistently estimated by the corresponding sample outer-products with ν_i replaced by $\hat{u}_i$, V_θ can be estimated by the conventional heteroskedasticity-consistent covariance matrix for the OLS estimator $\hat{\theta}$, γ can be estimated by its estimate from the two-stage procedure, h^o can be estimated by n_1/n_2, and C can be consistently estimated by

$$1/n_1 \Sigma_i \hat{u}_i \hat{u}_i' \hat{r}_i^2 ,$$

where $\hat{r}_i$ is the estimated residual for the ith observation from the second-stage estimation equation.