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THE INCIDENCE OF MANDATED EMPLOYER-PROVIDED INSURANCE: LESSONS FROM WORKERS' COMPENSATION INSURANCE

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EXECUTIVE SUMMARY

Workers' compensation insurance provides cash payments and medical benefits to workers who incur a work-related injury or illness. Many features of the workers' compensation program parallel features of proposed mandated employer-paid health insurance plans. This paper empirically examines the incidence of the workers' compensation program

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to infer the likely consequences of mandated health insurance proposals. In certain industries, such as trucking and carpentry, workers' compensation insurance costs are quite large, and vary tremendously within states over time, and across states at a moment in time. This variation is used to identify the incidence of the program. Empirical analysis of two data sets suggests that changes in employers' costs of workers' compensation insurance are largely shifted to employees in the form of lower wages. In addition, higher insurance costs are found to have a negative but statistically insignificant effect on employment. The implied elasticity of labor demand from our results is about −.50.

There are two primary methods that a government can use to provide universal access to a good or service: it can provide the good or service directly, as in the case of public education and national parks, or it can mandate that employers arrange for provision of the good or service for their workers and dependents, as in the case of workers' compensation insurance and certain pension safeguards. These two approaches may have different implications for the efficiency and equity of a public program. Interest in understanding the economic impact of employer mandates has risen in recent years as several mandated employer-provided health insurance proposals have gained support.

An estimated 15.2% of Americans under age 65 (some 33 million people) lack health insurance coverage.[1] Bills that would require employers to provide the uninsured with a minimum level of medical insurance are currently pending before the Congress and many state legislatures, and such legislation has already been enacted in Hawaii and Massachusetts. These bills are markedly different from national health insurance plans in most European countries, which are funded by general revenues and administered directly by the state.

Several factors explain the current political popularity of mandated health insurance proposals. More than two-thirds of the uninsured are full-time, full-year workers or are in families that are headed by a full-time, full-year worker. Moreover, nearly 80% of those with medical insurance already obtain their insurance from an employer-sponsored plan.[2] Consequently, employer mandates have the potential to extend health insurance coverage to a large number of uninsured individuals without radically restructuring the insurance industry. Furthermore, mandating that employers provide health insurance for their workers is

[1] This estimate was calculated by the authors from the March 1989 Current Population Survey, and pertains to calendar year 1988.

[2] These statistics are from Chollet (1987).

a way for the government to expand health insurance coverage without raising taxes. This feature of mandated benefits takes on added significance in an era of tight budget constraints.

Although the government may be able to shed responsibility for the cost of providing health insurance in a mandated program, there may nevertheless be substantial societal costs to such a program. In particular, requiring employers to pay health insurance premiums may increase the cost of hiring workers and result in lower employment. Indeed, Dr. Lewis Sullivan, Secretary of Health and Human Services, recently based his opposition to mandated health benefits on the presumption that: "By adding overly burdensome mandates on business, we could retard economic growth and constrict employment opportunities."[3]

But will mandated health insurance dramatically reduce employment? Traditional payroll tax incidence models suggest that some portion of the rise in employer costs may be shifted to wages, mitigating the fall in employment. And perhaps more importantly, if employees value health insurance, the theory of compensating wage differentials suggests that wages will fall even further than in the case of a pure payroll tax, further reducing the employment decline. In equilibrium, if the value employees place on health insurance is equivalent to the employers' costs of providing insurance, wages will be reduced by the full cost of the benefit and employment will be unchanged.

Although economic models provide a clear framework for analyzing the impact of compulsory insurance, the available empirical evidence on the trade-off between fringe benefits and wages provides little evidence of wage offsets. In fact, an impressive number of published papers report the "wrong" signed coefficient in analyses of the wage–fringe benefit relationship.[4] More generally, the labor economics literature has been largely unsuccessful in documenting consistent evidence of compensating wage differentials for a variety of nonpecuniary factors, such as the risk of unemployment and workplace hazards.[5] The findings of this literature challenge the view that the costs of mandated employer-provided health insurance will be shifted to employees in the form of lower pay.

Nevertheless, there are several reasons why the empirical evidence may fail to find evidence of a trade-off between fringe benefits and wages even if such a trade-off exists. First, available data sets do not permit researchers to control for all aspects of worker productivity. This

[3] Quoted from *The New York Times*, July 24, 1990, p. A10.

[4] A good example of this literature is Smith and Ehrenberg (1983). Additional references are given in Section II.

[5] See Brown (1980) and Abowd and Ashenfelter (1981) for examples.

is a problem because more productive employees may prefer to take their compensation in the form of both higher wages and higher fringe benefits. Second, most of the literature has examined fringe benefits that are voluntarily chosen by employers and employees. With voluntary fringes, it is difficult to identify exogenous variation in benefits that should induce wage offsets. Finally, the studies that have examined the effect on wages of mandated benefits (e.g., unemployment insurance) are identified by relatively small differences in program costs across states and over time. As a result, any trade-off between fringes and wages could easily be swamped by greater variation in omitted, state-level factors, such as union power.

This paper presents new evidence on the incidence of mandated employer-provided insurance by examining the experience of the workers' compensation insurance program. Workers' compensation laws require employers to secure insurance to provide a minimum level of cash payments and medical benefits in the event of work-related injuries and illnesses. The laws are exclusively administered by the states, which leads to wide variation in the generosity of the program across the states at a point in time, and within states over time. For example, the insurance rate in the trucking industry in 1987 ranged from 3% of payroll in Indiana to 25% of payroll in Montana. We use two data sets to estimate the effect of increases in workers' compensation costs on wages and employment. Throughout much of the analysis we focus on five high-risk industries (truck drivers, carpenters, plumbers, gasoline station employees, and nonprofessional hospital employees) that have great cost variability.

The structure of the workers' compensation program enables us to overcome at least some of the limitations of the past literature. First, workers' compensation benefits are not voluntary fringe benefits. Second, we test for insurance shifting in narrowly defined industries and occupations, so bias due to omitted worker and job characteristics poses less of a problem. Third, most of the analysis uses the wide variation in the growth of insurance rates in states over a 10-year period to identify the extent of insurance rate shifting. We are aware of no other payroll "tax" that varies as much across states or over time as workers' compensation insurance rates.

Finally, we note that the workers' compensation insurance program has much in common with proposed mandated health insurance schemes. In both programs, employers are required to secure a minimum level of insurance for their workers, and employers directly pay the insurance premiums. Furthermore, a substantial component (over 35% of costs) of the workers' compensation program is medical insurance for workplace injuries and illnesses. Consequently, inferences drawn from the workers' compensation program may be relevant for mandated health insurance.

I. CONCEPTUAL FRAMEWORK AND INSTITUTIONAL ANALYSIS

Summers (1989) argues that, because of wage offsets, the incidence and welfare costs of employer mandated benefits are different from that of a pure payroll tax, which is used to finance public provision of benefits.[6] The basic point of his argument is illustrated in Figure 1.[7] The imposition of an employer mandate will shift the labor demand curve downward (from *DD* to *D′D′*) and, with a fixed labor supply schedule, employment would fall from E_0 to E_1 and wages from W_0 to W_1. However, if workers value the benefit that they are receiving, labor supply will shift outward, and employment will fall by a lesser amount (to E_2 in Figure 1), whereas wages will fall by an even greater amount (to W_2). Thus, wage offsets in response to mandated benefits have the potential to reduce the labor cost increases created by mandated benefits.[8]

Formally, suppose that labor demand (L_d) is given by

$$L_d = f_d(W + C) \tag{1}$$

and further suppose that labor supply (L_s) is given by

$$L_s = f_s(W + \alpha C) \tag{2}$$

where C is the cost of mandated health insurance, αC is the monetary value that employees place on health insurance, and W is the wage rate. Using this notation it can easily be shown that

$$\frac{dW}{dC} = -\frac{\eta^d - \alpha\eta^s}{\eta^d - \eta^s} \tag{3}$$

[6] Danzon (1989) provides a related analysis, which also models the impact of heterogeneous workers and considers the general equilibrium implications.

[7] We have drawn the labor supply curve assuming the substitution effect dominates the income effect in the relevant range.

[8] As Summers notes, this analysis must be construed differently for health care benefits, which are fixed with respect to hours of (full-time) work, and for other types of benefits. This does not affect the analysis as long as one assumes that employment, rather than hours, are represented on the horizontal axis.

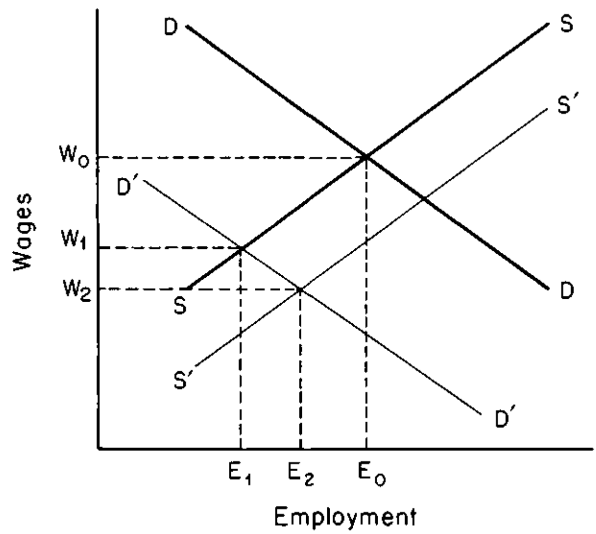

FIGURE 1. ***The Labor Market Impact of Mandated Benefits***

where η^d and η^s are the elasticities of demand for and supply of labor, respectively. From equation (3) it is clear that if α equals one, wages would fall by the full cost of the mandated benefit. In this situation, the employer's cost of providing the mandated benefit would be fully shifted to employees. On the other hand, if employees place no value on mandated health insurance, which is the case if $\alpha = 0$, the incidence of mandated health insurance is exactly analogous to that of a payroll tax.

As explained below, we rely on interstate differences in workers' compensation costs across industrial groups over a 10-year time period in our empirical work. If, as seems likely, workers are mobile across states and/or industries over this time horizon, η^s will be quite large. In the limit, as η^s approaches infinity the proportion of costs that are shifted to wages tends to α.

The effect of mandating benefits on employment is

$$\frac{dL}{L} = \frac{W_0 - W_2 - C}{W_0}\eta^d \tag{4}$$

which indicates that the amount of employment sacrificed because of mandated health insurance is inversely related to the wage offset caused by the provision of health insurance. Thus, we can summarize the labor market effects of mandated benefits in terms of the elasticities of supply and demand for labor, the cost of the benefit, and the fraction of that cost that is valued by employees.

A. Previous Evidence

What is the evidence for the shifting of employer benefit costs? A large empirical literature addresses this question by estimating hedonic labor market relationships (see Rosen, 1987 for a survey). However, past efforts to find the expected negative coefficient on fringe benefits in wage equations have been largely unsuccessful, as is discussed by Triplett (1983), Smith and Ehrenberg (1983), Leibowitz (1983), and Monheit, Hagan, Berk, and Farley (1985). Perhaps the failure to find a trade-off between fringe benefits and wages should not be surprising in view of the difficulty establishing compensating wage differentials for a variety of other nonwage aspects of work, including the risk of layoff. In fact, with the notable exception of work fatalities, the labor economics literature has not found consistent evidence of compensating wage differentials for work disamenities (Brown, 1980, Smith, 1979).

One of the difficulties in interpreting this literature, however, is that most studies of compensating differentials compare wages and working conditions among workers in different occupations and industries. There is nothing in the theory of compensating differentials to suggest that lawyers should receive less generous fringe benefits than manual laborers because of their higher wage. Indeed, if fringe benefits are a normal good, one would expect higher-paid workers to take some of their compensation in the form of better working conditions and fringe benefits. Nevertheless, the past literature on compensating differentials should challenge researchers to search for additional evidence before assuming that employee valuation of fringe benefits will be reflected in the incidence of those benefits.

Finally, we note that a related empirical literature has developed within public finance attempting to measure the incidence of payroll taxes. Unlike research on compensating differentials, most of this research is based on time-series or cross-country data. In an early cross-country study, Brittain (1972) reports evidence that he interprets as showing payroll taxes are fully offset by lower wages.[9] In an analysis of an unusual natural social experiment, Holmlund (1983) uses time-series data on payroll taxes in Sweden to examine wage growth in a period when the payroll tax increased from 14 to 40%. He estimates that

[9] More specifically, Brittain finds that capital's share does not decline with the imposition of a payroll tax. See Feldstein (1972) for a detailed critique of Brittain's work, in which he points out that Brittain's test can be viewed simply as evidence that the gross wage equals the marginal product of labor.

roughly 50% of the employer payroll tax is shifted to wages in the short run. Finally, Hamermesh (1979) uses the variation in payroll tax rates due to the social security payroll tax limit to estimate wage offsets; his estimates indicate that from 0 to 35% of the social security tax is shifted to wages.[10]

Although it would be desirable to have estimates of all of the parameters in equation (3) to forecast the incidence of mandated insurance benefits, in view of the past literature we have a more modest goal: We simply would like to examine the empirical plausibility of wage offsets in response to changes in the employers' cost of providing mandated benefits in the workers' compensation insurance program. If the evidence suggests that the employers' cost of providing workers' compensation insurance are offset by lower wages, then the expectation that mandated health insurance costs would be shifted to wages will be strengthened.

B. Description of Health Insurance Proposals

The potential importance of cost shifting of mandated health benefits is highlighted in Table 1, which summarizes several recent Federal and state proposals to compel employers to provide health care benefits. The Federal proposal (The Basic Health Benefits for All Americans Act) would mandate that all full-time workers have health insurance, with firms paying at least 80% of the premiums.[11] Similarly, the state proposals include some employer costs, either through a strict provision mandate, through a payroll tax, or through a "pay-or-play" plan, in which firms that do not provide a certain minimum level of health care must pay the difference in the form of a payroll tax. For example, in Massachusetts, any firm that does not spend an average of 12% of the Medical Security Wage Base on employee health benefits must pay the difference into a fund from which the state will purchase insurance for the uninsured.[12]

Table 1 indicates that employers will nominally bear the cost of providing health insurance for their employees according to the proposed legislation in each state. In addition, many of the proposed laws would use tax revenue generated by employer mandates to fund a larger universal insurance system, which would extend coverage to uninsured individu-

[10] Hamermesh's range of tax rates is from 0 to 5.85%. By contrast, the range in workers' compensation insurance rates across states in the industries we examine is from 3 to 26% of payroll.

[11] This bill passed the Senate Labor and Human Resources Committee, and is currently being discussed, along with alternative proposals, by a bipartisan Senate working group.

[12] This base is currently $14,000, and is indexed for medical inflation.

als who do not work. If the insurance provided to nonemployed individuals is commensurate with that provided through employment, the labor supply curve depicted in Figure 1 would not shift. More generally, the extent the labor supply curve shifts is inversely related to the quality of health insurance provided to nonworkers.[13]

The ultimate incidence of mandated employer-provided health insurance plans is difficult to predict because of the modest experience with such plans in the United States. However, the structure of workers' compensation insurance, which in many respects is similar to proposed health insurance plans, provides an opportunity to estimate the actual incidence of mandated employer-provided insurance.

II. WORKERS' COMPENSATION INSURANCE

Workers' compensation insurance is the oldest and largest mandated employer benefit program in the United States. In 1987 (the latest year with available data) workers' compensation benefits totaled $27.4 billion, and employer insurance costs equalled $38 billion.[14] The structure of the program allows for a test of the incidence of mandated benefits. Employers are required to purchase insurance or self-insure to provide a minimum level of cash benefits and medical care for workers who suffer a work-related injury or illness. Employees are covered by workers' compensation immediately on being hired.

The level of benefits that employers are required to provide varies tremendously across states, and has increased substantially in many states since the mid-1970s.[15] Table 2 documents the variability in workers' compensation benefits for a sample of 10 states. In Indiana, for example, an employee whose arm is permanently disabled in a work-related accident is entitled to a maximum indemnity benefit of $27,450. The same injury in Illinois qualifies for a maximum benefit that is more than five times as generous ($142,112). Similarly, benefits for the other types of injuries exhibit tremendous variability across states.

In addition, although it is difficult to quantify, it is widely believed

[13] Notice also that as the quality of health insurance provided to nonworkers improves, the relative efficiency gains of mandated benefits vis-à-vis public provision, which Summers (1989) emphasizes, declines.

[14] These statistics are from Nelson (1990), Tables 1 and 7.

[15] In large part, several states increased their workers' compensation benefits in the 1970s to conform with the recommendations of the National Commission on Workmen's Compensation Laws.

TABLE 1
Selected State and Federal Health Insurance Proposals

State	Coverage	Funding	Other features
U.S. Sen. Bill 768	Full time workers; may be extended to all workers above 185% of poverty line	Firms pay 80% of premiums, but pay 100% if worker earns less than 125% of minimum wage; comprehensive coverage	Federal subsidies for small firms; new small firms may offer less protection
CA Ass. Bill 3032 (1)	All employees, except new hires and those covered by workers' comp	Employers pay 75% of individual premiums, and 50% of family premiums; part-time workers are prorated	Tax credit of 25% of cost for small firms; insurers must accept small groups and must set "comparable" rates across groups
(2) "Cal-Care"	Low profit firms; low income/uncovered workers	Redirected employer tax credits and other state funds—reduced benefits	Firms with low profits are subsidized; low-wage workers are subsidized
CT Sen. Bill 342	Small firms	Employers pay 100% for small firms; waive premium tax	Limited loss ratios

HI Prepd. Health Care Act	Employees >20 hours/week and above 0.867 × minimum wage	Employers pay 50% of premiums, but a limit on worker expense of 1.5% of wages	
MA Health Security Act	Universal	12% payroll tax capped at Medical Social Security wage base, minus firm's current medical benefits	Tax credit for small firms (under 50 employees)
MI Sen. Bill 97	Minimum wage employees and low-wage workers who work more than 17.5 hours/week	3% tax on workers not covered and firms not covering (if more than 10 employees)—state contributes 3% also	
NJ Ass. 3382	All full time workers at firms with at least 6 employees	12% payroll tax up to Medical Social Security wage base (currently $14,000)	Study measures to help small firms
OR Health Partnership Act	Firms with less than 25 workers	Firm pays 75% of worker premium and 50% of dependent premium	Tax credit of 50% of cost for small firms could expand coverage to all workers

TABLE 2
Maximum Indemnity Benefits Paid to Selected Types of Work Injuries: 12 States in 1990

	Type of permanent impairment					Temporary injury
State	Arm	Hand	Finger[1]	Leg	Foot	(10 weeks)
California	$58,975	$43,450	$3,360	$64,575	$33,740	$2,660
Georgia	39,375	28,000	7,000	39,375	23,625	1,750
Illinois	142,112	114,899	24,189	120,946	93,733	6,047
Indiana	27,450	21,960	4,392	24,705	19,215	2,741
Hawaii	119,496	93,452	17,618	110,304	78,515	3,830
Massachusetts	20,402	16,132	NA	18,504	13,760	4,745
Michigan	114,863	91,805	16,226	91,805	69,174	4,270
Mississippi	42,516	31,887	7,440	37,202	26,573	2,126
Missouri	40,333	30,424	7,823	35,987	26,947	2,898
New Jersey	89,539	54,390	4,933	85,469	45,386	3,700
New York	93,600	73,200	13,800	86,400	61,500	3,000
Texas	47,600	35,700	10,710	47,600	29,750	2,380

[1] Benefit is for loss of use of index finger.

Source: Derived from *1990 Analysis of Workers' Compensation Laws.* Washington, DC: U.S. Chamber of Commerce, 1990, Chart 6 and 7.

that some states are more restrictive than others in permitting certain kinds of claims. This is especially likely to be the case for back injuries, which are difficult to diagnose objectively, and costly to treat.[16] Finally, some states require more generous medical benefits than others. For example, states vary in the extent of choice they allow employees over their treating physician.[17]

In most states workers' compensation insurance rates are established by a rating bureau. In principle, these rates are based on the actuarial cost of workers' compensation insurance, which is the expected benefit paid to each injury times the probability that a worker will incur each type of injury. A different rate is established for several hundred detailed industrial/occupational groups, known as activities. In 36 states, a national rate-making organization known as the National Council on Compensation Insurance (NCCI) pools information on risks and proposes rates for each activity. Several other states follow procedures that

[16] We also note that there is variability across states in the waiting period required before benefit payments commence, and in the retroactive period on which benefits are paid retroactively to compensate for the waiting period.

[17] See Boden and Fleishman (1989) for an analysis of interstate differences in medical costs in workers' compensation insurance.

are similar to the NCCI. The rates adopted by the states are known as "manual" rates.

Manual rates are the initial rate posted to firms, but they are not the actual, bottom line cost of insurance in many cases. For many employers, manual rates are adjusted in response to the specific firm's loss experience (e.g., experience rating), dividends are paid by insurance companies, and large firms may be retrospectively rated.[18] Because manual rates and the bottom line insurance costs are highly correlated, manual rates provide a reasonable approximation of the actuarial cost of benefits for each activity.[19] Furthermore, unlike in unemployment insurance, there is no cross-industry subsidization in workers' compensation rates.

Table 3 summarizes the interstate variability in workers' compensation rates for 47 states with comparable rating systems in 1987.[20] The table focuses on five activities that are the focus of our empirical analysis on premium shifting (carpenters, gasoline station workers, nonprofessional hospital employees, plumbers, and truck drivers).[21] These activities were selected because they comprise a large sample of workers in most states, and because they concord well with the occupational and industry definitions used by the Census Bureau. Furthermore, we selected these activities because work-injury rates are high in these jobs. For example, Table 3 shows that 1 in 20 truck drivers receives workers' compensation in a year, as opposed to less than 1 in 50 workers in all jobs. Because work injuries are a prominent feature of employment in these jobs, it is more likely that workers will be aware of their benefits under workers' compensation insurance and that the cost of providing these benefits will be shifted onto wages.

Table 4 further illustrates the variability in insurance costs by listing the manual rates for truck drivers in each state in 1978 and 1987. The spread in insurance rates for truck drivers across the states is enormous. For example, in 1987 the rate for truck drivers in Minnesota was 21% of

[18] The rate-making process is described in detail in Burton, Hunt, and Krueger (1985). Although on net experience rating hardly changes the average cost of workers' compensation insurance in a state, the other rating devices tend to reduce employers' costs, on average.

[19] For example, the correlation between the average manual rate in a state and the average manual rates adjusted to reflect dividends, experience rating, premium discounts, and rate deviations is 0.85. (*Source:* authors' estimates based on Table 4 and Table 22 of Burton, Hunt, and Krueger, 1985.)

[20] The insurance rates are described in more detail in the Data Appendix.

[21] The rates for carpenters are a weighted average of two carpentry classes (NCCI class 5403 and 5645), where the weights are the payroll attributed to each class nationally.

TABLE 3
Characteristics of State Workers' Compensation Insurance Rates[1]:
Selected Activities, 1987

Activity	(1) Mean[2] rate	(2) SD[2]	(3) Minimum	(4) Maximum	(5) Incidence rate[3]
1. Carpenters	11.32	4.86	3.35	26.25	3.28
2. Gasoline stations	5.18	1.85	1.73	11.09	2.16
3. Plumbers	6.02	2.45	1.74	15.05	3.34
4. Truck drivers	10.76	4.69	3.01	25.40	5.64
5. Nonprofessional hospital workers	4.43	2.32	1.42	13.00	2.93
6. All activities[4]	2.12	NA	NA	NA	1.72

[1] The insurance rates described in the first four columns refer to the "manual" rate for each state. No adjustments have been made for premium discounts, experience rating, dividends, and other competitive devices.

[2] Means and standard deviations (SD) are weighted by each industry's employment in the state. The sample consists of 46 states and the District of Columbia.

[3] The incidence rate measures the percent of private, non-self-employed workers who received income from workers' compensation insurance in calender years 1986 and 1987. These estimates were derived from the March 1987 and 1988 Current Population Survey tapes by the authors.

[4] The average manual rate for all activities is a payroll-weighted average of 44 activities, which account for 61% of payroll covered by workers' compensation insurance. This estimate is taken from Burton and Schmidle (1989, Table 1).

TABLE 4
Workers' Compensation Rates as a Percent of Payroll in the Trucking Industry (Class 7219)

State	1978	1987	Change 1987–1978
Alabama	4.49	10.07	5.58
Alaska	10.55	17.41	6.86
Arkansas	15.94	10.86	−5.08
Arizona	11.68	11.22	−0.46
California	10.04	17.26	7.22
Colorado	5.88	11.91	6.03
Connecticut	6.78	12.91	6.13
Delaware	10.45	9.79	−0.66
D.C.	15.04	16.04	1.00
Florida	17.71	15.12	−2.59
Georgia	4.70	7.73	3.03
Hawaii	9.71	20.29	10.58
Idaho	6.39	15.50	9.11
Illinois	6.01	11.45	5.44
Indiana	2.39	3.01	0.62

Iowa	5.89	8.77	2.88
Kansas	4.59	6.85	2.26
Kentucky	7.04	8.05	1.01
Louisiana	10.66	10.65	−0.01
Maine	7.05	9.16	2.11
Maryland	5.85	11.09	5.24
Massachusetts	5.50	8.48	2.98
Michigan	9.24	15.05	5.81
Minnesota	11.5	20.93	9.43
Mississippi	6.27	7.98	1.71
Missouri	NA	5.16	NA
Montana	8.27	25.40	17.13
Nebraska	5.04	6.47	1.43
New Hampshire	4.16	12.55	8.39
New Jersey	7.36	7.89	0.53
New Mexico	8.6	12.23	3.63
New York	9.62	5.97	−3.65
North Carolina	2.42	5.16	2.74
Ohio	5.32	12.20	6.88
Oklahoma	7.81	11.55	3.74
Oregon	14.68	23.46	8.78
Pennsylvania	NA	15.97	NA
Rhode Island	5.15	7.27	2.12
South Carolina	3.68	8.12	4.44
South Dakota	5.87	8.22	2.35
Tennessee	2.88	4.37	1.49
Texas	6.83	9.98	3.15
Utah	4.92	9.23	4.31
Vermont	3.11	6.53	3.42
Virginia	4.28	6.51	2.23
West Virginia	NA	5.67	NA
Wisconsin	3.41	8.86	5.45

payroll, while the rate for the same category of workers in Indiana was only 3% of payroll.[22] Furthermore, there is great diversity among the states in the growth in these insurance rates over the last 10 years. For example, rates soared from 9.7 to 20.3% of payroll in Hawaii between 1978 and 1987, while they fell by nearly 4 points in New York over the same time period. An important source of variation for within-state changes in workers' compensation costs over time is changes in benefit generosity (see Krueger and Burton, 1990; Butler and Worrall, 1990).

We are aware of only one study that uses interstate variation in work-

[22] Insurance rates for truck drivers are especially high because they have a high injury rate, and because their injuries tend to be relatively costly. Interestingly, data for Minnesota indicate that only 14% of truck drivers' injuries are from highway accidents, while almost half are from falls or strains and sprains (see Lewis, Meyers, and Senese, 1988).

ers' compensation costs to estimate the extent of shifting of mandated benefits.[23] Dorsey and Walzer (1983) use the May 1978 Current Population Survey to estimate the trade-off between employer liability for injuries as measured by workers' compensation costs and earnings. They find a large, negative effect of workers' compensation costs on wages in the nonunion sector, and a large, positive effect of workers' compensation costs in the union sector. Our analysis has several different features than Dorsey and Walzer's. Most importantly, we estimate the extent of cost shifting for narrowly defined activities that have wide interstate variability in costs, and we incorporate permanent state effects in the analysis. In addition, we provide estimates of the effect of workers' compensation insurance costs on employment.

III. ESTIMATES OF THE INCIDENCE OF WORKERS' COMPENSATION INSURANCE

We base our initial analysis of insurance cost shifting on microdata from the merged outgoing rotation group (OGRG) samples of the Current Population Survey (CPS). The CPS is a monthly survey of a rotating sample of over 55,000 households containing approximately 110,000 individuals. The Outgoing Rotation Group files consist of all individuals who are in their last survey month. The extracts of the CPS that we use are described in detail in the Data Appendix. Briefly, our sample contains privately employed carpenters, truck drivers, nonprofessional hospital employees, gasoline station employees, and plumbers. Self-employed workers are excluded from the sample. The data are taken from the 1979, 1980, 1981, 1987, and 1988 CPS files. For each worker in the sample we merge on the corresponding workers' compensation manual rate in the worker's state and industry/occupation group (activity).[24] A worker is observed in only one year, but we can control for permanent *state* effects because each year we have a sample of individuals from a constant set of states.

The CPS data have several advantages over other microdata sets for

[23] Moore and Viscusi (1990, Chapter 2) review several studies that estimate the relationship between wages and workers' compensation benefit levels. Most of these studies find a negative relationship.

[24] To be precise, we assign the 1978 workers' compensation rates to individuals in the 1979–1981 CPSs, and the 1987 rates to individuals in the 1987–1988 CPS. Data on workers' compensation rates for most states were provided to us by John F. Burton, Jr., from his ongoing research on measuring workers' compensation costs. Tracking down rates for all the intervening years would be a difficult task, and we suspect that using rates for 1978 and 1987 does not greatly affect the results.

this analysis. Most importantly, the large samples provided by the CPS enable us to focus on a narrow set of activities. Furthermore, the CPS contains information on several employee characteristics, including education, potential experience, marital status, and apprenticeship status.[25]

Using the pooled time-series/cross section of CPS data, we estimate a wage equation of the following form separately for each industry/occupation group:

$$\ln W_{ijt} = \beta_0 + \beta_1 C_{jt} + \beta_2 X_{ijt} + \alpha_t + \mu_j + \epsilon_{ijt} \tag{5}$$

where ln W_{ijt} is the natural log of the usual hourly wage rate, C_{jt} is the appropriate workers' compensation rate, X_{ijt} is a vector of covariates, α_t represents a set of year dummy variables (t=*1980,1981,1987,1988*), μ_j represents a set of state dummy variables, and ϵ_{ijt} is a stochastic error term. The subscript i indicates individuals, j indicates states, and t indicates years. Coefficient estimates are denoted by βs.

In contrast to past studies of compensating differentials that pool individuals from different occupations and industries and then control for average injury rates at the three-digit occupation or industry level, we estimate the wage–cost trade-off relationship within each three-digit industry/occupation group. Furthermore, since state and year dummy variables are included in equation (5), the incidence of workers' compensation costs is identified by varying patterns in workers' compensation rates within states over time. An estimate -1 for β_1 would indicate that increases in workers' compensation costs are fully offset by lower wages, while a coefficient of 0 would imply no wage offsets.

Workers' compensation benefits are not subject to federal income tax. Therefore, one might expect that the appropriate specification of equation (5) would utilize the after-tax wage as the dependent variable. However, if the tax rate for individuals in these activities is proportional to wages, then because the dependent variable is specified in log form the tax rate will be absorbed by the intercept term. Furthermore, the assumption of proportional taxes is likely to be approximately correct in this sample because most workers in these activities are below the earnings ceiling for the OASDHI payroll tax, and because workers in these activities are likely to be in a common tax bracket for the federal income tax.

Estimates of equation (5) are reported for each activity in the first five columns of Table 5. In four of the five activities we find a negative relationship between workers' compensation insurance rates and wages, but the relationship is statistically significant only for truck drivers. Also, the

[25] One shortcoming of the OGRG CPS files is that union status is not available before 1983.

TABLE 5

OLS Regression Estimates of Wage Equations[1]

Variable		*Occupation/industry group* Carpenters (1)	Gasoline stations (2)	Hospital workers (3)	Plumbers (4)	Truck drivers (5)	*All Activities* (6)	(7)
1.	Workers Compensation	−.517	−.651	−.119	1.729	−.966	−.187	−.865
	insurance rate	(.327)	(.839)	(.577)	(1.050)	(.331)	(.136)	(.184)
2.	Male	.021	.096	.029	.191	.078	.049	.050
	(1=yes)	(.065)	(.020)	(.016)	(.355)	(.059)	(.014)	(.014)
3.	Black	−.108	−.098	−.005	−.197	−.082	−.067	−.078
	(1=yes)	(.027)	(.027)	(.014)	(.041)	(.022)	(.011)	(.011)
4.	Other nonwhite	−.022	−.049	−.045	−.149	−.108	−.064	−.056
	(1=yes)	(.037)	(.039)	(.032)	(.076)	(.064)	(.021)	(.021)
5.	Years of	.036	.040	.021	.044	.031	.035	.033
	education	(.002)	(.003)	(.002)	(.005)	(.003)	(.001)	(.001)
6.	Potential	.026	.021	.008	.032	.022	.022	.021
	experience	(.002)	(.002)	(.001)	(.003)	(.002)	(.001)	(.001)
7.	Potential experience	−.400	−.382	−.075	−.476	−.329	−.327	−.323
	squared/1000	(.030)	(.039)	(.029)	(.063)	(.043)	(.018)	(.018)

8.	Apprentice (1=yes)	−.075 (.039)	—	—	−.203 (.040)	—	−.165 (.027)	−.188 (.027)
9.	Part-time worker (1=yes)	−.199 (.021)	−.126 (.015)	−.092 (.015)	−.240 (.056)	−.081 (.029)	−.128 (.010)	−.122 (.010)
10.	Metropolitan dummy (1=yes)	.113 (.012)	.087 (.014)	.068 (.014)	.123 (.022)	.131 (.014)	.109 (.007)	.111 (.007)
11.	Married (1=yes)	−.040 (.108)	.033 (.026)	−.004 (.013)	.091 (.435)	−.015 (.083)	−.017 (.015)	−.015 (.015)
12.	Married × Male	.160 (.109)	.055 (.029)	.120 (.023)	.035 (.435)	.095 (.084)	.139 (.017)	.126 (.017)
13.	44 state dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes
14.	44 state dummies × 4 activity dummies	—	—	—	—	—	No	Yes
14.	R^2	.395	.294	.332	.433	.213	.514	.532
15.	Sample size	4,784	2,708	1,928	1,556	4,268	15,244	15,244

[1] Dependent variable: log usual hourly wage. Each equation also includes an intercept and 4 year dummies. Columns 6–7 also include five dummy variables for each activity. Data are from the 1988, 1987, 1981, 1980, and 1979 full-year CPS files; see the Data Appendix for additional details. Standard errors are in parentheses.

wage–costs trade-off is largest for truck drivers, with an estimated coefficient of $-.97$ (t-ratio $= -2.9$). The other results suggest that approximately half of workers' compensation costs are shifted to wages for carpenters and gasoline station workers, but these estimates are fairly imprecise. On the other hand, the estimated trade-off between wages and insurance costs for plumbers is positive (1.73), but the coefficient has a large standard error.

The other variables in the wage equation generally have their expected signs. In all of these activities more education is associated with higher earnings; there is a quadratic relationship between earnings and potential experience; part-time workers earn less per hour than full-time workers, and apprentices earn less than fully trained workers. In addition, the equations include but do not report year dummy variables. The estimated coefficients for the year dummies indicate that real wages fell by roughly 20% in these activities between 1979 and 1988.

Columns (6) and (7) of Table 5 contain estimates for the pooled sample of the five activities. These equations also include four activity dummy variables. In column (6) we include 44 state dummy variables. In column (7) we interact the state dummy variables with the activity dummy variables. The estimated trade-off between insurance costs and wages depends critically on the specification of the state effects. If each activity is constrained to have the same effect (column 6), about 20% of workers' compensation costs are estimated to be offset by lower wages. However, if the activity effect is freed-up by state (column 7), fully 86.5% of workers' compensation costs are shifted to employees. Furthermore, an F-test indicates that state-by-activity effects explain a significant fraction of wage variability over the constrained state effects.

Evidently, the fixed state wage effects are also related to the level of workers' compensation costs. This relationship would occur, for example, if in some states unions have traditionally been influential in raising wages in some industry/occupation groups but not in others, and if influential unions are able to lobby for relatively more generous benefits for the types of injuries experienced by their members (e.g., back sprains vs. carpal tunnel syndrome). In this scenario, the less-restrictive model in column (7) would provide a more appropriate estimate of the trade-off between wages and the employers' cost of workers' compensation insurance.

A. The Importance of State Effects

The discussion above emphasizes the importance of the specification of the state effects in the pooled sample. It should further be noted that if the wage equations in Table 5 are estimated with the state effects omitted from the equation, the estimated wage–cost relationship is positive rather than negative in each of the five activities, as well as in the pooled sample of activities. For example, the coefficient on costs in the wage equation for truck drivers is positive .34 with a standard error of .16 when state effects are omitted.

There are several reasons to include state effects in the wage equations. First, state effects will pick up permanent cost of living differences between areas. Second, and perhaps more important, omitted state factors that traditionally have determined the generosity of workers' compensation benefits are likely to be correlated with the wage level of a state. As mentioned previously, benefits are likely to be greater in states that have a more powerful union movement, and strong union's are also likely to raise wages for nonunion as well as union members in the state. Another potential state-by-activity fixed effect is the risk of injury: Insurance rates will be especially high for an activity in states where that activity has an unusually high injury rate. Moreover, if there are compensating wage differentials for injury risk, we would also expect to find a positive association between insurance costs and wages.[26]

To the extent that injury risk is permanently greater in some states than in others (e.g., roads are always icier in Minnesota than in Florida, which affects truck drivers but not plumbers), including state-by-activity effects will control for within-industry state-level injury rate differences. Furthermore, the recent growth in workers' compensation costs is not likely to be due to union power, since union's influence has waned considerably in the 1980s. Instead, recent growth in insurance costs is likely to be due exogenous changes in medical costs, and to changes in benefit levels in response to the National Commission on Workmen's Compensation Laws' recommendations. In view of these considerations, the within-state variation in costs is probably a more appropriate source of identifying information.

B. Analysis of BLS Industry-Level Wage and Employment Data

In addition to the CPS, we analyze state-level, employer-reported data collected by the Bureau of Labor Statistics. This data set is based on employers' quarterly ES-202 reports, which are filed by all establishments covered by unemployment insurance laws. In 1988, these data contained

[26] We tried to assess directly the importance of this explanation by including the car accident rate in the wage equations for the truck drivers, as a measure of injury risk. Without including state effects, the car accident rate has a positive and statistically significant effect, but its inclusion only slightly reduces the positive wage–cost relationship. Because the vast majority of truck drivers' injuries are unrelated to driving accidents, the car accident rate may be a poor measure of injury risk for these workers.

information on 87.4 million private sector workers in 5.6 million establishments nationwide, or 99 percent of the private sector workforce.[27] The Bureau of Labor Statistics (BLS) compiles state-level averages of employment and wages by four-digit Standard Industrial Classification (SIC) from the ES-202 reports. Thus, we can use the BLS data to examine *both* wage and employment effects of workers' compensation insurance.

In comparison to the CPS, the BLS data have the advantage of being based on a virtual census of employers in the state. A disadvantage of the BLS data, however, is that the NCCI activity definitions do not match as well with the BLS data as they do with the CPS data because the BLS data are classified by industry alone. Therefore, to some extent the BLS data blur the occupational distinctions in the NCCI activity classifications.

We matched the state-level data for 1979 and 1988 to workers' compensation rates in the preceding year for four of the five activities in the sample. Unfortunately, it was necessary to exclude nonprofessional hospital employees because the SIC classification system does not distinguish between professional and nonprofessional employees. However, another six high-cost industries that can be matched between the NCCI and SIC classifications were added to the sample. The additional industries are agricultural machinery, excavation, gas and oil dealerships, lumber yards, masonry, and road and street construction.[28]

Figure 2 presents a scatter plot of the change in log average wage versus the change in insurance rates from 1978 to 1987 based on the BLS state-level data. The size of each observation in the plot is proportional to the number of truck drivers employed in the state. The graph also displays the fitted line from an employment-weighted regression of change in log wage on change in costs. A strong negative relationship is apparent. The slope of the regression line is -1, which is very close to the estimated extent of shifting in the trucking industry based on CPS data. Furthermore, most of the states (especially the large ones) tend to cluster fairly close around the regression line.

Table 6 reports estimates of regressions of the 10-year change (1979–1988) in the log of the average wage on the change in insurance rates for the pooled sample of activities. In levels, this specification is analogous to the within-state-by-activity specification shown in column 7 of Table

[27] The ES-202 data and the publicly available state-level data are described in BLS (1988). We are grateful to Michael Buso of the BLS for creating an extract of these data for us.

[28] These activities were excluded from the microanalysis either because they do not match well with the three-digit census classifications and/or because the CPS sample sizes are too small to give precise estimates.

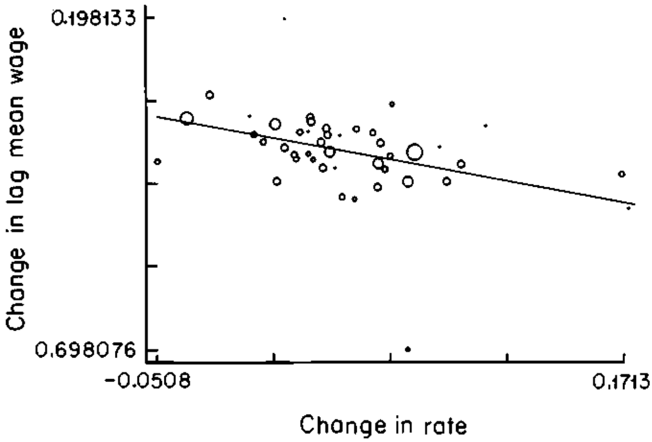

FIGURE 2. ***Trucking industry, 1979–1988.***

5, only we do not have controls for individual characteristics in the state-level analysis. We estimate a similar equation with the change in log employment as the dependent variable. In the first two columns we limited the sample to the original sample of activities (excluding nonprofessional hospital workers), and in the last two columns we use the expanded set of 10 activities. The regressions are weighted by the number of reporting establishments in 1988.

TABLE 6
Aggregate Wage/Employment Data Coefficients with Standard Errors in Parentheses[1]

	Sample and Dependent Variable			
	4 Industries		*10 Industries*	
Variable	(1) Wage change	(2) Employment change	(3) Wage change	(4) Employment change
Intercept	−0.065	0.182	−0.062	0.163
	(0.011)	(0.025)	(0.008)	(0.019)
Change in rate	−0.863	−0.112	−0.561	−0.248
R^2	(0.289)	(0.664)	(0.209)	(0.503)
	0.05	0.0002	0.018	0.001
Sample size	176	176	388	388

[1] Equations are estimated by weighted least squares, where the weights are the number of reporting units in 1988.

For the original sample, the wage regression yields virtually the same coefficient on the workers' compensation cost variable as was found in the CPS microdata (compare column 7 of Table 5 with column 1 of Table 6). The results indicate that 86% of increases in workers' compensation costs are shifted to workers. The results for the expanded set of activities indicate a smaller trade-off between wages and insurance costs (−.56), but the estimate is still statistically significant and sizable. The similarity between the aggregate results based on BLS data and the microlevel wage regressions based on CPS data is reassuring, especially in light of their different unit of observation and industry definitions.

Columns 2 and 4 report results of regressing the change in log employment on the change in insurance rates. Because insurance costs appear to be only partially offset by lower wages, one would expect a negative coefficient on insurance rates in these regressions. The results provide some evidence of a negative effect of higher workers' compensation rates on employment growth, but the estimated coefficient is less than half its standard error in each sample.[29] If one assumes that 85% of workers' compensation costs are shifted to wages and that the remaining 15% are borne by employers, then the coefficient estimate in column 2 suggests that the elasticity of labor demand is −.75. On the other hand, if we assume that employers bear 44% of workers' compensation costs (1−.56), then the estimate in column 4 implies an elasticity of labor demand of −.56. Although these estimates are extremely imprecise, they are within the range of typical estimates of the elasticity of labor demand (Card, 1990; Clark and Freeman, 1980).

The following calculation gives an indication of the magnitude of the effect on employment of recent increases in workers' compensation insurance costs. Between 1972 and 1987, nationwide workers' compensation costs increased by about 1% of payroll, from .722 to 1.785% (Burton and Schmidle, 1989). If 85% of this increase was borne by labor and .15% by employers, and if we assume an elasticity of labor demand of −.75, then this increase in program costs would have reduced employment by .11%. Since approximately 90 million workers are covered by workers' compensation insurance, this means a loss of a little over 100,000 covered jobs.[30]

[29] We note, however, that the estimated wage trade-off and employment trade-off for the sample of 10 activities is sensitive to the weight used. The estimated wage–cost coefficient ranges from −.50 to −.89 depending on whether the equation is weighted by number of reporting units in 1979, employment in 1980, or average employment in 1979 and 1988. The coefficient on costs in the employment equation ranges from −.03 to −.50, depending on the weight used.

[30] According to Nelson (1990), in 1987 approximately 88.4 million workers were covered by workers' compensation insurance.

IV. APPLICATION: HEALTH INSURANCE

What can we infer about the likely consequences of mandatory employer-provided health insurance from the estimated incidence of the workers' compensation insurance program? Because employers who already provide health insurance are unlikely to be affected by such a mandate it is useful to consider the sectors of the workforce that have a low rate of coverage by healthy insurance. Table 7 reports the percent of various subgroups of the population that was covered by some form of health insurance in 1988. The relatively low rates of coverage for workers in the construction and agricultural industries, in small firms, and in low-wage jobs have been documented previously (e.g., Chollet, 1987).

The low rate of health insurance coverage in low-wage jobs is potentially important for our purposes because the floor established by the minimum wage may impede wage offsets that would otherwise result from mandating health insurance. The jobs we have examined in the workers' compensation program typically have wage rates well above the minimum wage. If an uninsured worker's wage is at or slightly above the minimum wage, the minimum wage may prevent wage offsets from occurring in response to benefit mandates. The existence of this kind of an institutional impediment would cause a greater employment reduction than suggested from our analysis of workers' compensation insurance.[31] Several authors have noted that the minimum wage may inhibit wage offsets in response to mandated health insurance (see Reinhardt, 1987; Monheit, Hagan, Berk, and Farley, 1985; Chollet, 1987). An important question is: How large is the share of uninsured workers that is potentially constrained by the minimum wage?

In an influential study based on the 1986 March CPS, Chollet (1987) claims that as many as 50% of uninsured workers earn less than 125% of the minimum wage, and that 35% of uninsured workers earn the minimum wage or less. Although we will present evidence suggesting these figures exaggerate the impact of the minimum wage, if a great many of the uninsured earn the minimum wage it is hard to imagine that important wage offsets would occur in response to mandated benefits.

To explore the impact of the minimum wage on health insurance shifting, we use the March 1989 CPS to estimate the earnings distribution of uninsured workers. We restrict the sample to non-self-employed individuals between age 16 and 65 who worked for pay in 1988. For each

[31] We note, of course, that insurance costs could be shifted by means of nonwage elements of the compensation package for minimum wage workers. Holzer, Katz, and Krueger's (1990) analysis of application rates suggests incomplete rent dissipation in minimum wage jobs, however.

TABLE 7
Percent of Various Groups Covered by Health Insurance

	Percent covered by			
Characteristics	Health insurance	Private health insurance	Insurance through own or dependent's job	Insurance through own job
Sex				
Male	84	83	76	65
Female	87	84	75	50
Age				
16–24	78	75	59	31
25–35	85	82	77	64
35–45	90	88	83	67
45–55	90	90	83	70
55–65	92	91	80	68
Race				
White	87	86	77	58
Black	78	72	64	53
Other	82	79	69	54
Hours				
>35 hours, 48 weeks	90	90	85	77
>18 hours, 26 weeks	87	86	80	67
Wages				
W < \$3.35/hour	71	62	45	11
\$3.35 ≤ W < \$5/hour	73	68	55	26
\$5 ≤ W < \$10/hour	86	85	76	58
W ≥ \$10/hour	95	95	91	81
Education				
Less than HS	73	68	60	36
Finished HS	85	83	75	58
Some college	89	88	76	58
Post-college or more	94	94	87	75
Employer size				
<25 employees	74	71	56	30
25–100 employees	83	80	72	54
100 + employees	91	89	83	70
Industry				
Agriculture, mining	69	65	52	37
Construction	74	73	65	50
Manufacturing	90	89	85	77
Transportation	91	90	86	77
Trade	81	78	66	39
Services	87	85	75	55

Source: Authors' calculations from the March 1989 Current Population Survey.

TABLE 8
Percentage of Uninsured Workers Falling into Various Wage Intervals[1]

	Wage ≤ 3.35	Wage between $3.00 and $3.70	Wage between $2.75 and $3.95
	A. All Workers		
Uncovered	19.4	10.5	18.6
Uncovered by employer plan	19.5	6.4	17.1
	B. Full Time, Full Year Workers		
Uncovered	16.8	9.0	16.9
Uncovered by employer plan	15.2	8.3	14.8

[1] Sample excludes self-employed workers, and those who earn less than $1.00 per hour or more than $100 per hour. Sample size is 8506 for Part A and 6482 for Part B.

Source: authors' calculations from the March 1989 CPS.

worker in the sample, an average hourly wage rate is calculated by dividing annual earnings by the product of annual weeks worked and usual weekly hours. To trim outliers, we exclude observations on individuals whose derived hourly earnings are below $1 per hour, or greater than $100 per hour.[32] This sample is used to calculate the share of uninsured workers that is likely to be affected by the minimum wage.

The first row of Table 8 gives the fraction of *all* uninsured workers who fall into various wage intervals.[33] The second row gives the fraction of workers who are not covered by an employer-sponsored insurance plan who fall into each wage interval. Part B contains the same set of estimates for full-time, full-year workers (i.e., those who worked at least 26 weeks in the year and usually worked 18 hours or more per week). The results indicate that 19.4% of all uninsured workers in 1988 earned the minimum wage or less. The estimate is 16.8% when we limit the sample to full-time, full-year workers, which is the group most likely to be covered by a mandated health insurance law.

For several reasons, we prefer estimates that place a window around the minimum wage, rather than the open-interval estimates in column 1. First, the annual average hourly wage rate is likely to be measured with

[32] Eliminating outliers in this fashion does not qualitatively affect the results once the self-employed are excluded from the sample.

[33] The various sources of health insurance coverage are Medicare, Medicaid, CHAMPUS/VA care, employment-based insurance, and private health insurance. Coverage may be in one's own name or as a dependent.

considerable noise because it is based on self-reported annual hours of work, which is notoriously poorly measured (Duncan and Hill, 1985). Second, some workers who earn more than the minimum wage may be constrained by the minimum wage because their earnings are prevented from falling below the wage floor in response to mandated health insurance. Finally, we wish to put a lower bound below the minimum wage because there is a tremendous amount of noncompliance with the minimum wage.[34] If some individuals are paid less than the minimum—either legally or illegally—there is no reason to suspect that their wages will not be reduced if their employers are required to provide them with health insurance.

Column (2) places a 35¢ window around the minimum wage, and column (3) places a 60¢ window around the minimum wage. Since the estimated hourly cost of meeting the Federal health insurance mandate is 55¢ for a full-time, full-year worker (U.S. Congress, 1989), the wider interval in column (3) is probably a slight overstatement of the percent of uninsured workers who are constrained by the minimum wage. No matter how the wage intervals are defined, however, the results in Table 8 are quite different from Chollett's estimate that 50% of uninsured workers will be constrained by the minimum wage. According to our estimates, less than 20% of uninsured workers earn within 60¢ of the minimum wage. Although this is a nontrivial share of the uninsured, it suggests that for the majority of uninsured workers there is scope for wage offsets engendered by mandated employer-provided health insurance.

Why are our estimates considerably lower than Chollett's, even though both are based on March CPS data? The discrepancy is mainly due to the fact that we exclude self-employed workers. For example, including the self-employed raises the estimate of the number of uninsured workers below the minimum wage by nearly 10 percentage points. However, we believe that three reasons justify excluding the self-employed from these tabulations. First, the proposals outlined in Table 1 do not apply to self-employed workers. Second, self-employed workers are typically exempted from State and Federal minimum wage laws. Third, the earnings of the self-employed reflect returns and losses to capital investments.

Finally, we note that Table 8 may give an overestimate of the impact of the minimum wage for mandated benefits because many of the proposed mandated health insurance laws provide subsidies to small employers. Since relatively many low-wage workers are employed by small

[34] The extent of noncompliance with minimum wage was first documented by Ashenfelter and Smith (1979).

establishments, these subsidies would reduce the employers' costs of providing health care to low-wage workers. Tabulations from the March 1989 CPS indicate that 42% of uninsured individuals who earn less than the minimum wage are in firms with 25 employees or fewer.

On the other hand, the real minimum wage has increased since March 1989, and thus may be more of a constraint. Nevertheless, our results suggest that the minimum wage may not be as much of an impediment to wage adjustments in response to mandated benefits as previously believed.

V. CONCLUSION

This paper has analyzed the impact of cost shifting in response to increases in mandated workers' compensation insurance costs. The results suggests that a substantial portion of the cost to employers of providing workers' compensation benefits is shifted to employees in the form of lower wages. Given the similarity between workers' compensation insurance and many proposed employer-mandated health insurance plans, our findings suggest that a large share of the employers' cost of meeting health insurance mandates may be borne by employees. Furthermore, our tabulations of the share of uninsured workers whose earnings are near the minimum wage suggest that in 1988 less than one-fifth of uninsured workers were likely to be constrained by the minimum wage. Although the nominal burden of mandated employment-based health insurance will be borne by firms, if the experience of health insurance is similar to that of workers' compensation insurance, our estimates suggest that employees will ultimately bear a large fraction of the burden of financing mandated health insurance through lower wages.

In spite of our main conclusion that a sizable portion of the cost of mandated benefits is likely to be shifted to employees, we should also stress that the shifting of workers' compensation costs is incomplete. Employers bear at least some additional cost because of mandated work-injury insurance. As a consequence, we find that increases in workers' compensation costs are associated with reduced employment growth. Although extremely imprecise, our estimates suggest that every one percentage point increase in workers' compensation rates is associated with an employment decline of .11%. The adverse employment effects of mandated health insurance may well be larger than those in workers' compensation insurance because the minimum wage is likely to be more of a constraint for uninsured workers, especially in view of recent increases in the real minimum.

VI. DATA APPENDIX

A. Workers' Compensation Insurance Rates

The workers' compensation insurance rates used in Table 3–6 and Figure 2 are state manual rates, collected from each state's rate pages. Manual rates are expressed as a proportion of payroll. Manual rate data were generously provided to us by John F. Burton, Jr. In some cases, if manual rates were not available for a state we used the rate for the assigned risk pool, less 10%. The rate for the Minnesota trucking industry in 1987 is from Lewis, Myers, and Senese (1988), and is based on insurance industry data. Data are available for a maximum of 44 states and the District of Columbia in both 1978 and 1987. The states in the sample use the NCCI rating classification system, or a comparable system. The NCCI codes for the five activities used in the analysis are truck drivers (NCCI 7219), plumbers (NCCI 5183), gasoline service station employees (NCCI 8387), nonprofessional hospital employees (NCCI 9040), and carpenters (weighted average of NCCI 5403 and 5645).

B. CPS Data

The CPS data used in Table 5 are from the outgoing rotation group files for 1979, 1980, 1981, 1987, and 1988. The sample was limited to individuals between age 18 and 65 who were privately employed. The sample only includes individuals from the 45 jurisdictions that have manual rate data in 1978 and 1987. The 1978 manual rate was merged to observations in the 1979, 1980, and 1981 CPS samples; the 1987 manual rate was merged to observations in the 1987 and 1988 CPS samples. The "usual hourly wage" is the ratio of usual weekly earnings to usual weekly hours. Individuals with allocated weekly earnings were deleted from the sample, as were those who earned less than $1.67 per hour or more than $150.00 per hour in 1988 dollars.

The "married" dummy variable equals one if the worker is married with his or her spouse present. The "part-time" dummy variable equals one if the worker usually works less than 35 hours per week. "Potential experience" is age minus education minus 6. The "apprentice" dummy variable equals one if the worker's occupation code is an apprentice.

The following 1970 Census industry (CIC) and occupation (COC) codes were used to define the activities: truck drivers (CIC 417; COC 715), plumbers (CIC 67–78, COC 522 or 523), gasoline station employees (CIC 648), nonprofessional hospital employees (CIC 838; COC 630, 690, 694, 762, 901–903, 912–916, or 950), and carpenters (CIC 67–78, COC 415 or 416). The following 1980 Census industry (CIC) and occupation

(COC) codes were used to define the activities: truck drivers (CIC 410; COC 804 or 805), plumbers (CIC 60, COC 585 or 587), gasoline station employees (CIC 621), nonprofessional hospital employees (CIC 831; COC 435–437, 439, 443–444, 449, 453, 748, 777, 883), and carpenters (CIC 60, COC 567 or 569).

C. BLS Data

The BLS data are annual totals derived from quarterly ES-202 reports. The BLS data consist of state-level averages for 3- and 4-digit Standard Industrial Classification (SIC) groups. The dependent variable in columns (1) and (3) of Table 6 is the change in the log of the arithmetic average annual wage in each industry. The annual wage is in 1988 dollars. The dependent variable in columns (2) and (4) is the change in log employment. The SIC codes for the four activities in the basic sample in Table 6 are truckers (SIC 421), plumbers (1711), gas stations (SIC 5541), and carpenters (SIC 1751).

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