The Labor Market Effects of Rising Health Insurance Premiums

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We estimate the effect of rising health insurance premiums on wages, employment, and the distribution of part-time and full-time work using variation in medical malpractice payments driven by the recent "medical malpractice crisis." We estimate that a 10% increase in health insurance premiums reduces the aggregate probability of being employed by 1.2 percentage points, reduces hours worked by 2.4%, and increases the likelihood that a worker is employed only part time by 1.9 percentage points. For workers covered by employer provided health insurance, this increase in premiums results in an offsetting decrease in wages of 2.3%.

I. Introduction

In the United States, two-thirds of the nonelderly population is covered by employer-provided health insurance (EHI), either directly or as a de-

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pendent through a family member's coverage.[1] According to a national survey conducted by the Kaiser Family Foundation, the cost of EHI has increased by over 59% since 2000, with no accompanying increase in the scale or scope of benefits; between 2003 and 2004 the price of premiums increased 11.2%, 9 percentage points more than the 2.3% increase in workers' hourly earnings.[2] Increases in health insurance premiums may have significant effects on labor markets, including changes in the number of jobs, hours worked per employee, wages, and compensation packages. Indeed, it is possible that a significant portion of the increase in the uninsured population may be a consequence of employers shedding this benefit as health insurance premiums rise (Porter 2004). Simple correlations are consistent with this story: despite strong economic growth in the 1990s, the number of nonelderly uninsured grew by 3 percentage points to 15.7% of the population, while the price of health insurance premiums grew by 34%.

Understanding the relationship between health insurance costs and labor markets is of growing policy importance. Proposals to cover the uninsured often rely on "employer mandates" that would require employers to cover eligible workers.[3] For example, California's Senate Bill 2 (also known as Proposition 72, narrowly defeated in November 2004) would have required all employers with more than 20 employees to provide health insurance to their workers (who work more than 100 hours per month). Other policy proposals include the provision of tax credits for the purchase of nongroup health insurance, which would differentially change eligible employees' valuation of benefits provided by their employer versus wages.

The magnitude of the effects of increases in benefit costs on employment, wages, and health insurance coverage will be driven by the elasticities of labor supply and demand, institutional constraints on wages and compensation packages, and how much workers value the increase in health insurance costs. Since employers currently provide such coverage voluntarily, if workers fully value these benefits then they will bear the cost of the increase in reduced wages, with no accompanying change in

[1] These tabulations are from the annual demographic files of the Current Population Survey (CPS) for 1988–2003. We define the nonelderly population as those under the age of 65.

[2] These figures are obtained from Kaiser Family Foundation, Kaiser Commission on Medicaid and the Uninsured (2003). In fig. 1, we use these data to illustrate the growth of premiums since 1996 for family and single-person policies.

[3] Yelowitz (2004) provides a thorough discussion of this legislation and estimates its economic impact.

employment, employment costs, or employee utility.[4] Thus, in a world where workers value benefits at their cost and are able to sort between firms based on their preferences—and without other institutional constraints—increases in the costs of benefits should be fully offset by decreases in wages.

There are many reasons, however, to believe that firms have limited ability to offset increases in the price of health insurance premiums through lower compensation. Institutional constraints, such as the minimum wage, union rules, and other provisions of labor law that prohibit different demographic groups from being paid differently, limit a firm's ability to reduce compensation.[5] Nondiscrimination provisions increase the incentives to move workers with different preferences for benefits than their coworkers, that is, "anomalous workers," into part-time positions. Carrington, McCue, and Pierce (2002) provide evidence of such behavior. They note that anomalous workers (i.e., low-pay employees in high-pay firms) are more likely to be part-time employees than they are in firms where they resemble the typical worker. They demonstrate that Internal Revenue Service nondiscrimination rules limit the extent of within-firm inequality in benefits but increase (within-firm) wage inequality. Their study demonstrates that firms skirt these nondiscrimination rules by moving workers with unusual benefits into part-time positions. In the context of our research, such pressures would imply that employers move married women (who typically decline EHI) working in firms where the typical worker is a married man (who elects coverage) into part-time jobs. For these reasons, increases in the cost of providing health insurance may affect employment and the structure of work.

Identifying the magnitude of these effects empirically, however, is difficult. Data on premiums and wages are usually not jointly available at the individual level. Additionally, most micro–data sets (such as the CPS) do not allow the researcher to control adequately for worker characteristics, such as ability, that might simultaneously influence both health insurance costs and the outcome under study. In this article we uncover the causal effect of increases in the cost of benefits on labor market out-

[4] This view is explicitly studied in the literature estimating the wage-fringe tradeoff (e.g., Gruber 1994). A \$1 increase in the value of fringes may be offset by a \$1 reduction in fringe benefits—or, in the case of most tax-favored benefits, a \$1/$(1 - t)$ reduction, where t is the tax rate.

[5] We used the February CPS supplement to examine the degree to which low-wage workers have EHI and are therefore at risk of losing such coverage. In these data (weighted using CPS weights), we focused on two samples: (*a*) workers who earn within \$3 of the minimum wage in their state and (*b*) workers who earn less than \$10 per hour (which corresponds with the 25th percentile of wages). For these samples, the fractions that are eligible for EHI are 52.6% and 60.5%, respectively, and the fractions covered by EHI are 31.7% and 38.5%, respectively.

comes by exploiting an exogenous source of variation in the cost of providing health insurance: the recent "medical malpractice crisis," where malpractice costs for physicians grew dramatically in some states but not in others. As we discuss in more detail below, the growth of malpractice payments affects both malpractice insurance premiums and the cost of health insurance: if the demand for health care is relatively inelastic (because of private or public health insurance), the increased cost of malpractice will be borne by consumers in the form of higher health insurance premiums, rather than primarily by physicians in the form of lower compensation (see Baicker and Chandra 2005).

In Section II we outline a conceptual framework for our analysis. Section III examines econometric challenges to estimating the hypothesized effects and provides a justification for our use of malpractice payments as an instrumental variable for health premiums. In Section IV we describe the data that we use. In Section V we present empirical results, including specification checks that provide validation for our use of malpractice payments as a plausible instrument for health insurance premiums. We find that the cost of increases in health insurance premiums is borne by workers through decreased wages for those with EHI and through decreased hours for those moved from full-time jobs with benefits to part-time jobs without. As we discuss in Section VI, these results have strong implications for the effectiveness and distributional impact of many different health-care reform proposals.

II. Conceptual Framework

The effect of increased benefits costs on employment outcomes depends on several factors. Summers (1989) examines the effects of mandated benefits (versus taxes) on wages and employment, highlighting the importance of the employees' valuation of the benefit. The provision of a benefit that is fully valued by workers should not change employment—but should decrease wages by the cost of the benefit. Gruber and Krueger (1991) discuss this model more formally. Following their framework, let $L_s = f(W + \alpha C)$ and $L_d = (W + C)$ be the labor supply and demand curves, respectively. The variable W represents wages, C represents the premium for the health insurance provided by the employer (thus incorporating both the quantity of insurance and the price for that policy), and αC represents employees' monetary valuation of that health insurance. It is straightforward to demonstrate that

$$dW/dC = -(\eta^d - \alpha\eta^s)/(\eta^d - \eta^s), \tag{1}$$

where η^d and η^s are the price elasticities of labor demand and supply. If $\alpha = 1$, then wages fall by the full cost of the mandated benefit, and if $\alpha = 0$, then the results are identical to those obtained for the incidence of a payroll tax. Additionally, the proportional change in employment will be given by

$$dL/L = \eta^{d}(W_0 - W_1 - dC)/W_0, \qquad (2)$$

where W_0 and W_1 represent the initial and final levels of wages.

Equation (2) demonstrates that the effect of rising health insurance premiums on employment is inversely proportional to the wage offset caused by the employer provision of health insurance (which means that it is also a negative function of α) and is proportional to the elasticity of labor demand. Note that as the premium (C) is the product of the quantity of health insurance provided and the price of that bundle, changes in C may reflect changes in the price of EHI or changes in quantity (directly or as an indirect effect of changes in price), although in practice employers may be limited in their ability to change the health insurance bundle.

An increase in the price of EHI results in an inward shift of the labor demand curve, unambiguously lowering employment and wages ceteris paribus. However, there may also be a corresponding shift outward of the labor supply curve: if the cost of both EHI and health insurance in the nongroup market increase together (and α is positive), workers will value the increase in employer expenditures on benefits and be willing to accept lower wages in exchange for higher spending on health insurance. It is important to note that this outward shift of the labor supply curve will not occur if the price of EHI increases while the price of health insurance in the nongroup market is unchanged, because workers could then purchase the health insurance on their own without using extra resources (making their outside option more attractive relative to working and being compensated in part with EHI) and would thus be unwilling to accept an offsetting reduction in wages.

The effect of an increase in the cost of EHI on the labor market equilibrium, again assuming α is positive, is thus a fall in wages but an ambiguous change in employment. Employment falls less or rises more, the greater is the worker valuation of health insurance, α. There are a few groups for whom we might, however, particularly expect to see employment losses: employers may not be able to reduce wages for workers near the minimum wage. There are also other antidiscrimination provisions of labor law that prohibit different demographic groups from being paid differently. Union rules and workplace norms about differential pay for different demographic groups (who may have heterogeneous demands for EHI) make it difficult for an employer to perfectly fine-tune the wage response to rising EHI premiums. For such groups, we expect to see a decrease in employment.

In the Summers (1989) model, labor supply is treated as a discrete choice; there is no distinction between employment and hours worked. If this divisibility is introduced, the effect on hours worked and employment are ambiguous. This is because increases in the price of EHI raise the fixed costs of employment, which suggests that employers would want more hours from fewer workers. In this case, hours for workers with EHI would increase, while their wages decrease; this mechanism is emphasized by Cutler and Madrian (1998). By contrast, since part-time workers are typically not covered by EHI, we might also expect employers to replace full-time workers eligible for EHI with part-time workers without EHI. In this situation, total employment may increase. In both cases, we do not expect to see an increase in wages or a decrease in hours for workers with EHI. Nor do we expect to see an increase in full-time employment, although the total effect on employment is ambiguous.

In light of these ambiguous analytical predictions of the effect of health insurance costs on hours, employment, and the fraction of full-time and part-time jobs, assessing the labor market effects of increases in health insurance premiums is fundamentally an empirical question.

III. Empirical Strategy

Empirically evaluating the effect of rising health premiums on employment, wages, hours worked, and the composition of employment (the share of jobs that are full time or part time) is an exercise with numerous challenges. Data sets such as the census and the CPS do not contain information on the employer costs of health insurance or the generosity of plans. Additionally, even if this information were available, such data sets do not allow the researcher to control adequately for worker characteristics that might also influence the outcome under study.[6] In principle, a large-scale social experiment that does not suffer from attrition or agents attempting to compensate for their treatment regime may solve selection problems of this nature, and an instrumental variable estimation strategy can reproduce the experimental estimate if the underlying assumptions behind instrumental variables (IV) estimation are satisfied.

To motivate our estimation strategy, consider the following structural equation for a worker i in state j and in year t:

$$\text{Outcome}_{ijt} = \beta_0 + \beta_1 HI_i + X_i\Pi + S_j + T_t + \varepsilon_{ijt}. \tag{3}$$

Here, Outcome_i is the labor market outcome of interest (hours worked, wages, wage income, unemployment, part-time/full-time status, or receipt of health insurance). The expression X_i measures person-level covariates, including controls for family structure, marital status, and industry. Expressions S_j and T_t are state and year fixed effects, respectively, and ε_{ijt} is

[6] These limitations are identical to those that have plagued the literature on identifying the wage-fringe trade-off. Currie and Madrian (2000) provide a comprehensive overview of this literature.

a person-specific idiosyncratic term.[7] Expression HI_i measures the premium associated with providing individual i with EHI.[8] This equation can be modified to include interaction effects and indicator variables for certain demographic groups that may be of particular interest (e.g., hourly workers, married women, workers near the minimum wage, or workers with EHI).

The first problem inherent in an ordinary least squares (OLS) estimation of equation (3) is that Cov(Cost of $HI_i, \varepsilon i) \neq 0$. For example, workers with high ability may work at firms that offer generous health benefits and, therefore, high premiums. Second, data sets such as the Survey of Income and Program Participation (SIPP) or CPS do not report the value or generosity of the health insurance plan received by a worker.[9] Empirical researchers have responded to this limitation by imputing health insurance premiums to each respondent based on industry (Cutler and Madrian 1998) and based on industry, firm size, and family/single status (Yelowitz 2004). These imputations solve the missing data problem and can in principle reduce the potential endogeneity problem. We first discuss the iden-

[7] Cutler and Madrian (1998) do not include state fixed effects. Yelowitz (2004) examines data from California; his analysis is therefore comparable to one where state fixed effects are included.

[8] Technically, HI_i should measure the difference between EHI premiums and premiums for policies purchased in the nongroup market. The nongroup market is small relative to the size of the group market, and it is also highly individually rated. Using data from the 2005 March CPS, we found that 16.9 million individuals under the age of 65 (6.6% of this group) were covered by directly purchased health insurance. In contrast, 161 million individuals (63.2%) were covered by EHI. This 10-fold difference suggests that the nongroup market is small. In contrast to the group market, prices in the nongroup market are affected by smoking, age, alcohol consumption, weight, and preexisting conditions such as diabetes, hepatitis, AIDS, asthma, high cholesterol, and liver conditions. We made this determination by pricing individual premiums on http://www.eHealthInsurance.com. For the purpose of estimation, the above discussion makes it clear that we must include state fixed effects. These fixed effects will control for state-level differences in the price of EHI and insurance purchased in the nongroup market. However, to the extent that the price of insurance in the nongroup market is highly idiosyncratic to the individual, our IV strategy will remove any correlation between premiums for EHI and premiums for nongroup health insurance.

[9] The SIPP offers several advantages over the CPS, and we explored the possibility of using the SIPP in the early stages of our project. The ability to use a fixed-effects estimator makes its use extremely appealing, as it would control for time-invariant heterogeneity. To use the SIPP we would need persons to switch jobs, industry, or family structure, so that we can observe changes in the price of EHI overtime; if there were no such changes, the imputed premium for EHI would be perfectly collinear with the fixed effect, and these observations would not contribute to the estimation of the coefficient on EHI. However, if the factors that induce job changes or family transitions (such as illness or economic downturns) are also correlated with changes in labor market outcomes, then we would need an IV strategy.

tification strategy implicit in this approach and then contrast it with our alternative strategy.

Imputed premium data may be thought of as representing premiums that have been obtained using the match characteristics (such as industry and family structure) as instruments. That is, if data on HI_i were available, we could, in principle, estimate:

$$HI_{dfs} = \gamma_0 + \text{Industry}_d + \text{Firm Size}_f + \text{Family Structure}_s + v_{dfs}. \quad (4)$$

The *dfs* subscripts make explicit the notion that equation (4) is estimated at the level of industry *d*, firm size *f*, and family structure *s* and not at the level of a person *i*. (Equation [4] could also be estimated at the state level.) One could use the fitted values from equation (4), $\overline{HI}_{dfs}$, as the key regressor in the estimation of equation (3). Implicitly, these fitted values may be thought of as characterizing the relationship

$$HI_i \equiv \overline{HI}_{dfs} + m_i. \quad (5)$$

Here, m_i represents the portion of health insurance premiums that is idiosyncratic to person *i*. It is probably the case that $\text{Cov}(\varepsilon_{ijt}, m_i) \neq 0$ because more productive workers who can command large compensation packages will choose packages with more health insurance. This is the principal reason for why the OLS estimation of equation (3) will not produce consistent estimates. If $\text{Cov}(\varepsilon_{ijt}, \overline{HI}_{dfs}) = 0$, as would be the case if industry, firm size, and family structure were valid instruments, then we may estimate:

$$\text{Outcome}_{ijt} = \beta_0 + \beta_1 \overline{HI}_{dfs} + X_i \Pi + S_j + T_t + \varepsilon_{ijt}. \quad (6)$$

The central problem with estimating equation (6) in lieu of equation (3) is the possibility that $\text{Cov}(\overline{HI}_{dfs}, \varepsilon_i) \neq 0$. This would be true if the "instruments" (industry, firm size, and family structure) are correlated with ε_i, the unobservable characteristics of the worker. If workers in a certain sector of the economy or those who are married are systematically more likely to have different levels of unobservable characteristics that affect health insurance premiums, then such a correlation is possible.[10] This problem is identical to the standard endogeneity problem in program evaluation, where receipt of the treatment is correlated with unobservable characteristics of the person receiving treatment.

[10] We examined whether these instruments failed a test of overidentifying restrictions. To perform this test we regressed the two-stage least squares residuals on all the exogenous variables and examined whether the R^2 exceeded the critical values of a χ^2 distribution with degrees of freedom equal to the number of instruments. For each of our key dependent variables (wages and salary, hours, part-time status, and employment), the test statistic using industry and firm size as instruments exceeded a value of 1,000, several times higher than the critical values of the χ^2 distribution required to reject these variables as instruments.

A solution to this problem is to instrument for imputed premiums using variables that are uncorrelated with ε_i and m_i but are correlated with imputed health insurance premiums. In our analysis we use state-level, per-capita medical malpractice payments as an instrument for imputed premiums. For malpractice payments to provide a valid instrumental variable for imputed premiums, it must be the case that the instruments affect health premiums. Second, it must also be the case that malpractice payments are not correlated with the unobservable characteristics of workers. In the next subsection, we explore the prima facia validity of these assumptions.

The Medical Malpractice Crisis

The "medical malpractice crisis" that began at the turn of the twenty-first century refers to the dramatic increase in physician premiums for malpractice insurance. Chandra, Nundy, and Seabury (2005) and Mello, Studdert, and Brennan (2003) provide an overview of this crisis and its underlying causes and consequences. Both the American Medical Association and the Physician Insurers Association of America attribute the dramatic increase in physician malpractice insurance premiums to the growth in malpractice payments (see Smarr 2003; American Medical Association 2004a, 2004b). Whereas other factors such as declines in insurers' investment income—including the presence of an underwriting cycle, a less competitive insurance market, and climbing reinsurance rates—are acknowledged to have contributed to this medical malpractice crisis, insurer losses from increases in malpractice payments are believed to be the primary contributor to the growth of malpractice premiums. Indeed, a General Accounting Office study of seven states concluded that the growth of insurer's losses from payments is the primary driver of the growth of premiums (see U.S. General Accounting Office 2003a, 2003b).

If the demand for health services is inelastic, then the effect of increasing malpractice payments on malpractice premiums will have little effect on net physician compensation. Indeed, Baicker and Chandra (2005) argue that because of the nature of health insurance (which insulates the patient from the marginal costs of seeking care, and which is subsidized by the tax code), the demand for medical services is relatively inelastic. The demand for health services by Medicare beneficiaries is likely to be even less elastic, as they are further insulated from even a wage-fringe benefit trade-off. Consumers of health care are therefore likely to bear the brunt of the cost through increases in the price of health care (and, consequently, health insurance premiums).[11]

[11] In theory we could also use malpractice premiums as an instrument for health insurance premiums. However, there is less systematic data on malpractice insurance premiums. There is an annual survey conducted by the publication *Med-*

With this preliminary validation, we use increases in malpractice payments as an instrument for health insurance premiums to estimate the following first-stage equation:

$$\overline{HI}_{ijt} = \gamma_0 + \gamma_1 \text{ Malpractice Payments}_{jt} + X_i\Pi + S_j + T_t + v_i, \qquad (7)$$

where, as discussed below, malpractice payments are broken down by the size and number of payments for different specialties. Instrumenting for imputed premiums removes the bias from any residual correlation between ε_i and $\overline{HI}_{ijt}$. This is because the instrument only picks up that part of the (within-state) variation in imputed premiums that is attributable to (within-state) changes in the malpractice climate. It may be tempting to reason that the correlation of premiums with the instrument, malpractice payments, is potentially spurious because states with high malpractice payments may have workers who are systematically more or less abled. This is not the case, however, as all of our specifications include state fixed effects.

It is particularly important in the context of this source of variation to understand the way that workers will value benefits. Our use of this instrument does not rely on the fact that workers get more or better health care as their premiums rise. Rather, as malpractice costs rise, the price of purchasing health care through any source—employer insurance, nongroup insurance, or out of pocket—will increase. Workers may be willing to accept lower wages in exchange for costlier health insurance because they would have to pay more on the open market for it, whether or not the increase in premiums is associated with higher value health care.

IV. Data

A. Health Insurance Premiums

We use annual state-level data on health insurance premiums by type of policy (family or single) and employer size from the Medical Expenditure Panel Survey for 1996 to 2002. We assign premiums to workers based on their state of residence and year and on their family structure (with single respondents given the single premium). We obtain very similar results when we do not match on family structure or when we further match workers based on firm size (with employees of small firms given the small firm premium and unemployed respondents given the average premium).

In figure 1 we illustrate the steady growth in premiums for family premiums and single premiums over the time period of our study. All

ical Liability Monitor, but the survey does not rely on administrative data, does not cover all states or medical specialties, and varies year to year in the number of insurers that are surveyed.

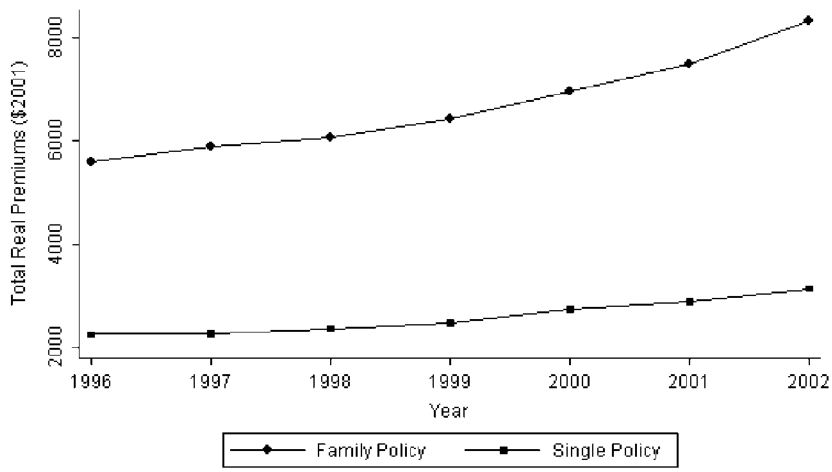

FIG. 1.—Growth in health insurance premiums. The data are from the Medical Expenditure Panel Survey. The premiums are expressed in year 2001 dollars.

dollar figures are expressed in year 2001 dollars. Family premiums grew from an average of \$5,000 in 1996 to well over \$8,000 in 2002. Premiums for single policies also grew substantially—from an average of \$2,000 in 1996 to over \$3,000 in 2002. In figure 2 we illustrate the details of family and single policies for the 10 states with the largest population in 2000—panel A reports the level of premiums in 1996, and panel B reports the level of premiums in 2002. We see that family premiums grew between 40% and 60% over this time period in these states. The growth in single premiums was smaller but still considerable: in states such as Florida, Georgia, Michigan, and Ohio, premiums for single people grew by over 40%. Both panels also show the share of total premiums that was paid for by employee and employer contributions—even though premiums increased substantially, the share paid by employees remained relatively stable.

B. Labor Market Data

The CPS is a monthly survey of about 50,000 households conducted by the Bureau of the Census for the Bureau of Labor Statistics. The survey has been conducted for more than 50 years and is the primary source of information on the labor-force characteristics of the U.S. civilian noninstitutional population. The March (annual demographic survey) files of the CPS contain information on hours worked, wage and salary income, unemployment, and health insurance coverage in the past year.

We use data from the 1996–2002 March CPS. We use information on demographics (such as age, gender, race, marital status, family size, and

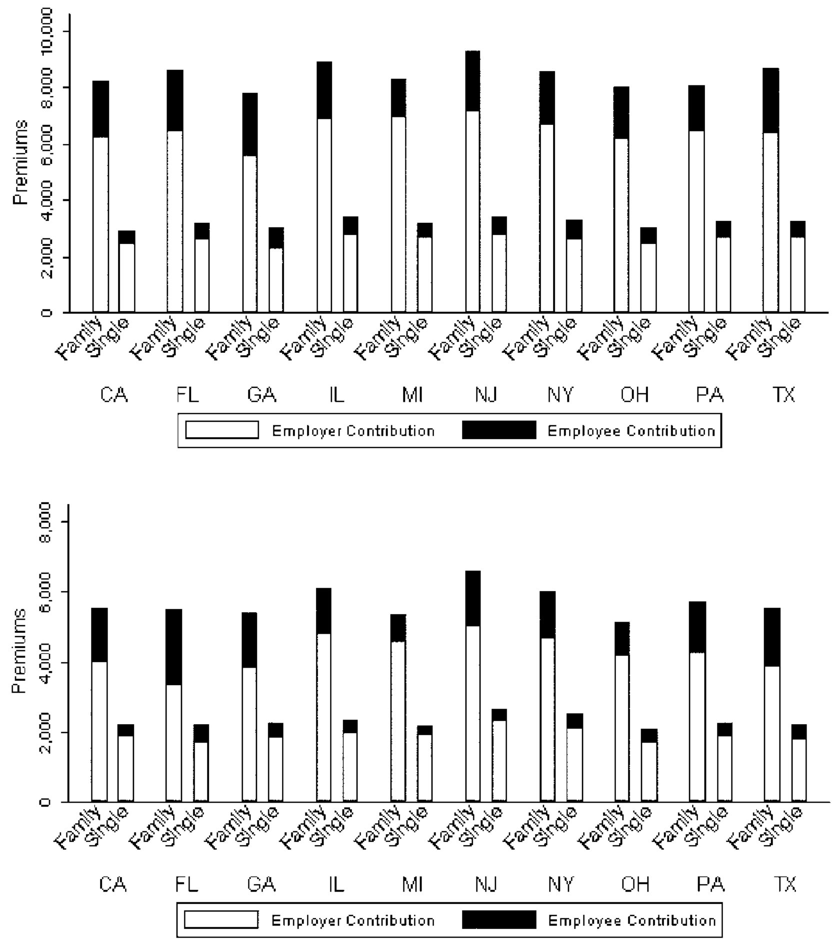

FIG. 2.—Health insurance premiums in 1996 (panel A), and health insurance premiums in 2002 (panel B). The data are from the Medical Expenditure Panel Survey. The premiums are expressed in year 2001 dollars. The 10 largest states (by population) are shown.

education), labor market variables (such as wage and salary, employment status, firm size, and hours worked), and health insurance coverage (such as source of coverage). Because we expect premiums from last year to affect current labor market outcomes, we measure hours worked, full-time/part-time status, and unemployment during the reference week of the survey (typically the second week of March). We include all respondents between the ages of 22 and 64, although we further limit the sample in some of our analyses. Our data are summarized in table 1.

Table 1
Summary Statistics

	All		1996–99		2000–2002	
	Mean	SD	Mean	SD	Mean	SD
HI variables:						
Premiums	5,259	1,968	4,773	1,675	5,938	2,139
HI from employer	.62	.49	.61	.49	.62	.49
Any HI	.84	.37	.83	.37	.84	.36
Labor market outcomes:						
Hours	37.2	16.7	35.9	17.8	39.1	14.5
Wage and salary income (real)	33,902	37,152	32,896	36,818	35,310	37,570
Part time (<30 hours per week)	.14	.35	.15	.36	.14	.35
Employed	.88	.33	.88	.33	.88	.33
Malpractice payments (real per capita $):						
Total	3.27	1.94	3.13	1.79	3.47	2.11
Internal medicine	1.74	1.31	1.59	1.18	1.96	1.44
Obstetrics-gynecology	7.36	4.44	6.91	4.00	8.00	4.93
Surgery	.25	.21	.26	.23	.23	.18
Malpractice payments (no. per 100,000 population):						
Total	1.6	.7	1.6	.7	1.6	.7
Internal medicine	.4	.2	.4	.2	.4	.3
Obstetrics-gynecology	3.1	1.4	3.1	1.4	3.1	1.5
Surgery	.2	.2	.2	.3	.1	.1

NOTE.—HI = health insurance. Individual-level observations are from the 1996–2002 Current Population Survey (CPS). The sample is limited to those ages 22–64. The HI premiums are from the Medical Expenditure Panel Survey (state-year data on premiums by policy type and employer size). The malpractice payments are from the National Practitioner Data Bank. Labor market outcomes and employer health insurance information are from the CPS (March and February). Sample sizes were 194,739 for the 1996–99 period and 151,785 for the 2000–2002 period.

C. Medical Malpractice Payments

We obtain data on malpractice payments from the National Practitioner Data Bank (NPDB) for the 1996–2002 period.[12] We calculate the size and number of payments resulting from medical treatments (including diagnosis, medication, and other medical treatment), surgical treatments (including surgery and anesthesia), obstetrical treatments, and other treatments (including monitoring, equipment, intravenous and blood, and all

[12] Noncompliance is subject to civil penalties codified in 42 U.S.C. 11131–52. We examine payments that resulted from either a court judgment against the provider or a settlement made outside of the courts. We exclude payments that were linked to dentists, pharmacists, social workers, or nurses. In a small fraction of payments, there are multiple physician defendants (and thus multiple reports), but only the total payment by all defendants is reported. In these cases we average the payment by the number of physicians involved. In the NPDB, 5% of payments are made by state funds in addition to other payments made by the primary insurer for the same incident. We match such payments based on an algorithm that uses unique physician identifiers, state of work, state of licensure, area of malpractice, type of payment (judgment or settlement), and year of occurrence.

others). Table 1 shows the growth of per capita malpractice payments at the state level between 1996–99 and 2000–2002. The variability of payments (over time within states) is the source of our identification. For example, over the 2001–3 period, per capita payments were highest in the states of New York, Pennsylvania, New Jersey, Connecticut, West Virginia, and Delaware. In these states the burden of malpractice liability was almost twice the U.S. average of $13.50 per person. See Chandra et al. (2005) for more details on the growth of malpractice payments as measured by the NPDB.[13]

V. Results

We begin with an examination of the effect of increases in health insurance premiums on employment, wages, and hours worked. The odd columns of table 2 show the results of the OLS estimation of equation (3). All regressions include state and year fixed effects and the individual-level controls outlined above, and they are weighted using the March CPS final weights. Standard errors are clustered at the state level. Premiums, income, and hours are all measured in logs. The OLS effect of increases in health insurance premiums on labor market outcomes suggests that a 10% increase in premiums leads to a 1.3% decrease in wage and salary income and a 0.1% decrease in hours worked. As discussed above, however, the OLS results are likely to be biased by omitted individual characteristics (such as ability) and economic conditions. We use medical malpractice payments (including real per capita dollars and the number of payments per capita, by specialty, current and lagged) to instrument for health insurance premiums. Table 3 reports several variants of the first stage regression shown in equation (5). For the two-stage least squares estimates that follow, we use the most flexible form of the instruments (which provides the greatest power in the first stage, with the instruments jointly significant at $p < .0001$ and with a first-stage F-statistic of 4.45), but here we show more constrained forms to aid in interpretation (since payments by different specialties are highly correlated). The results suggest that when per capita malpractice payments double, health insurance pre-

[13] One of the major criticisms of the NPDB is the "corporate shield" (U.S. General Accounting Office 2000). This is a loophole that makes payments made on behalf of a hospital or other corporation exempt from inclusion in the NPDB, as long as any individual practitioner is dropped as part of a settlement agreement. Chandra et al. (2005) provide evidence that this is not a first-order source of bias in the data bank. Even if it were these omissions would only serve to weaken the estimated first stage. A more problematic source of bias occurs if there is state-level variation in the magnitude of the corporate shield (a hypothesis on which there is no formal or anecdotal evidence). We include state-fixed effects in our analysis to help ameliorate this potential problem. We are grateful to Aaron Yelowitz for recommending this discussion.

Table 2
Effect of Premiums on Labor Market Outcomes

	Ln (Wage and Salary Income)		Ln (Hours) (If Employed)		Part Time (If Employed)		Employed		Have HI through Employer		Employee Share of HI Premium	
Sample: All (with Positive Hours Last Year)	OLS (1)	IV (2)	OLS (3)	IV (4)	OLS (5)	IV (6)	OLS (7)	IV (8)	OLS (9)	IV (10)	OLS (11)	IV (12)
Ln (HI premium)	−.129	−.088	−.009	−.236	.012	.190	−.022	−.120	−.173	−.063	−.129	−.031
	(.14)	(.133)	(.023)	(.103)	(.026)	(.072)	(.014)	(.065)	(.130)	(.077)	(.032)	(.088)
N	346,524	346,524	304,744	304,744	304,744	304,744	346,524	346,524	346,524	346,524	346,524	346,524
Covariates and fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

NOTE.—OLS = ordinary least squares; IV = instrumental variables; HI = health insurance. The individual-level observations are from 1996–2002 Current Population Survey (CPS). The sample is limited to those ages 22–64. The HI premiums are from the Medical Expenditure Panel Survey (state-year data on premiums by policy type and employer size). The malpractice payments are from the National Practitioner Data Bank. The labor market outcomes and employer health insurance information are from the CPS (March). Part-time workers work less than 30 hours per week. Covariates include race, age, age^2, age^3, marital status, education, gender, and health status. Instruments include real dollar amount and number of medical malpractice payments per capita for different specialties (surgery, obstetrics-gynecology, internal medicine, and other) for current year and previous year. Premiums assigned based on state, year, and family structure (and employer size for employed). Regressions are weighted by March CPS weights, and standard errors are clustered at state level.

Table 3
First Stage Regressions

	Effect of Growth in Malpractice Payments on Health Insurance Premiums (Log-Log Specification)		
	(1)	(2)	(3)
All payments	.011	.021	
	(.009)	(.008)	
Surgical payments			.014
			(.006)
Obstetrics-gynecology payments			−.003
			(.003)
Internal medicine payments			.005
			(.007)
Other payments			−.001
			(.001)
No. of surgical paymets			−.022
			(.008)
No. of obstetrics-gynecology payments			.005
			(.005)
No. of internal medicine payments			.008
			(.012)
No. of other payments			−.004
			(.003)
F-statistic on instruments	2.820	4.100	4.450
F-test significance	.185	.054	.007
Covariates	No	Yes	Yes
State and year effects	Yes	Yes	Yes

Note.—The individual-level observations are from the 1996–2002 Current Population Survey. The sample is limited to those ages 22–64. The health insurance premiums are from the Medical Expenditure Panel Survey (state-year data on premiums by policy type and employer size). The malpractice payments are from the National Practitioner Data Bank. Covariates include race, age, age^2, age^3, marital status, education, gender, and health status. Standard errors are clustered at state level. The growth of payments is measured as the change in real payments per capita or the number of payments per capita.

miums increase by 1%–2%. This is consistent with previous estimates that malpractice payments comprise around 1% of total health expenditures (Kessler and McClellan 1996). We show the results of a similar regression at the state-year (rather than individual) level graphically in figure 3.

Results from two-stage least squares estimation of equation (6) are shown in the even columns of table 2. Here, a 10% increase in premiums reduces wages and salary by 0.9%. There is a large effect of premiums on hours worked—coming partly from increases in the probability of unemployment but also through increases in the probability of part-time work. This is consistent with our expectation that as the cost of providing health insurance benefits increases, firms will substitute part-time workers with limited benefits for full-time workers. In fact, in our data only 22% of part-time workers have employer health insurance, while 64% of full-time workers do. Consistent with the reduction in full-time jobs, there is also an overall decline in employment rates and a (statistically insig-

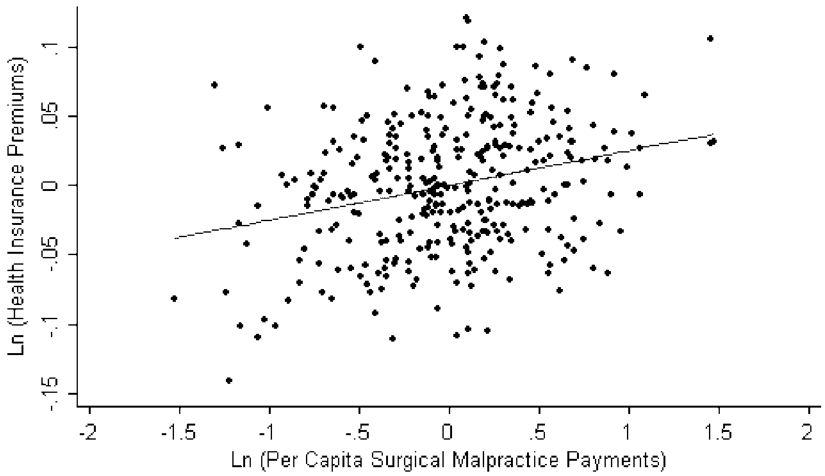

Fig. 3.—The effect of malpractice payments on health insurance premiums. The regression line represents an adjusted coefficient of 0.03 (clustered SE 0.01). The sample includes annual observations of 48 continental U.S. states from 1996 to 2002. The controls include categorical measures of educational attainment, race, age, marital status, health status, and gender mix (at the state-year level), as well as state and year fixed effects. Standard errors are clustered at the state level. The covariates are aggregated to state-year level from the annual March CPS.

nificant) decline in the probability of being covered by EHI.[14] In contrast to the OLS results, all the IV results are consistent with the predictions of a model where workers partially value health benefits or where firms are constrained in their ability to adjust wages. For each IV specification we evaluated the potential endogeneity of the imputed health insurance premium variable by performing a Durbin-Wu-Hausman test. In each of these tests, we rejected the null hypothesis that the OLS estimates were

[14] Ideally, we would also examine the extent to which employers offer EHI. To shed more light on this issue we examined the results of the Kaiser/Health Research Educational Trust survey. These data show that between 1996 and 2002 (the period of our study), the percentage of firms offering EHI increased from 59% to 66%. There was no change for larger firms (defined as those who employ over 200 workers), 99% of whom offer coverage. These firms account for approximately 60% of total employment. Unfortunately, we are unable to explore the relationship between an employer's decision to offer EHI and increases in the price of EHI in more detail. Questions about whether or not an employee is offered EHI are only asked in the 1997, 1999, and 2001 February supplements. In these years, we can match almost three-quarters of the March respondents to the February survey, but this represents less than 30% of our original sample. In this smaller sample, however, we do not have the power to detect changes in EHI being offered to employees: our results were simply too noisy to allow us to conclude that we were learning anything systematic.

not systematically different from the IV estimates (each test statistic had a *p*-value of less than 0.0001).

As health insurance costs have risen, popular concern has grown over increases in required employee contributions to health insurance premiums. In fact, between 1996 and 2002, the employee share of health insurance premiums remained relatively stable at just under 20%. This fraction does not seem to respond to increases in health insurance premiums—using our IV specification in table 2, column 12, the fraction of premiums paid by employees (and consequently employers) does not respond to increases in premiums.

We were concerned that the first-stage *F*-statistic of 4.45 could be indicative of a problem of weak instruments (Staiger and Stock 1997).[15] Following Stock, Wright, and Yogo (2002), we reestimated the models in table 2 with limited information maximum likelihood (LIML) methods as operationalized by Moreira (2003) and Moreira and Poi (2003). This treatment constructs tests of coefficients based on the conditional distributions of nonpivotal statistics. The LIML point estimates were virtually identical to the IV results reported in table 2 (SEs in parentheses): wages −0.080 (0.143), hours −0.218 (0.072), part time 0.162 (0.058), any health insurance 0.067 (0.054), and EHI −0.040 (0.073).

We also examined whether our results were robust to the inclusion of industry and occupation controls and state-year unemployment rates, and we found very similar results with the inclusion of these additional variables. Specifically, the IV coefficients (SEs) on the probability of being part time with these additional controls entered simultaneously were 0.17 (0.07), on hours they were −0.199 (0.106), on employment they were −0.10 (0.065), and on the probability of receiving EHI they were −0.04 (0.07). These estimates are very similar to those reported in table 2.[16] We report these results as a specification check rather than our primary specification because the additional covariates are all potentially endogenous variables that may change in reaction to increases in health insurance premiums.

As previous models and empirical research have suggested, we might expect certain groups to be more sensitive in changes in the cost of health

[15] It should be noted that without clustering the standard errors, the first stage *F*-statistics exceed 500. We rely on the smaller, more conservative *F*-statistic with clustered standard errors, but we report this statistic as well because it is not clear what the appropriate *F*-stage is with clustered data. We are grateful to Douglas Staiger for recommending the use of LIML to examine the potential for weak instruments.

[16] Results with the inclusion of region-specific time trends were also quite similar (SEs in parentheses): wages and salary −0.095 (.177), hours −.18 (.158), part time 0.196 (0.096), employed .160 (.094), any HI −0.065 (0.105), EHI −0.104 (.096), and employee share 0.054 (.125).

insurance. First, workers with health insurance should see a much bigger offset in their wages than workers without, who should see none. The first column of table 4 tests this hypothesis by including the interaction of health insurance premiums with an indicator for coverage by EHI. We see that, in fact, all of the reductions in income are borne by employees with health insurance. The magnitude of the elasticity of −0.20 is consistent with dollar-for-dollar offset (since premiums, paid with pretax dollars, are about 20% of wage and salary income at the mean)—implying that covered workers bear the full incidence of increases in health insurance premiums. The results for part-time workers in table 4, columns 3–4, are consistent with part-time workers seeing an increase in wages and a decline in coverage when health insurance premiums increase (which would be the case if workers were moving from full-time jobs with benefits to part-time jobs with higher wages instead of benefits), but these results are not statistically significant.

If our predictions are correct, we should also see declines in wage and salary income (and health insurance coverage from an employer) for workers in sectors where the demand for labor is particularly elastic. Because manufacturing goods are nationally traded and the labor demand for manufacturing workers is a derived demand, we would expect the local demand for such workers to be particularly sensitive to the price of health insurance. Table 4, columns 5–7, shows that in addition to the larger predicted wage declines for this group, we also observe a decrease in the probability of employment and an increase in the probability of being shifted to part-time work.[17]

Which workers would be most likely to give up EHI (in exchange for higher wages) as premiums increase? Married women are likely to have a lower value of EHI, as they may have access to insurance though their husbands (Berger et al. 2003). As shown in the last panel of table 4, these women are indeed more likely to lose EHI when premiums go up. They are also more likely to be employed and more likely to be in part-time jobs (recall that our model predicted that total employment might increase if there was a shift to part-time jobs). In contrast to the results for married women, we find that there is a large wage offset for married men (coefficient on interaction term of −0.04, not shown in table) but that there is no effect on the probability of being part time (coefficient on interaction term of 0.009, insignificant) or on the probability of being employed for this group (coefficient on interaction term of 0.021, insignificant). That the results for married men and women are different in a manner that accords with eco-

[17] Employment in the manufacturing sector is measured as major industry of employment last year, while the dependent variable is measured as employment last week. For this reason, we cannot include interactions with hours or EHI, also measured currently as opposed to last year, using this variable.

Table 4
Differential Effects of Premium Increases (Health Insurance Premiums Instrumented with Malpractice Payments)

	Interaction Term and Outcome										
	Workers with EHI		Part-Time Workers		Manufacturing Workers			Married Women			
	Ln(Wage and Salary) (1)	Pr(Employed) (2)	Ln(Wage and Salary) (3)	Pr(EHI) (4)	Ln(Wage and Salary) (5)	Pr(Employed) (6)	Pr(Part Time) (7)	Ln(Wage and Salary) (8)	Pr(EHI) (9)	Pr(Employed) (10)	Pr(Part Time) (11)
Ln (HI premium)	.05	−.19	−.12	−.02	−.12	−.11	.17	−.08	.03	−.19	.15
	(.14)	(.08)	(.18)	(.10)	(.13)	(.06)	(.07)	(.15)	(.09)	(.07)	(.07)
Main effect of interaction term	2.32	.34	−.99	.01	1.41	.48	−.42	−.11	.96	−.57	−.19
	(.37)	(.18)	(.46)	(.23)	(.37)	(.14)	(.14)	(.37)	(.32)	(.19)	(.18)
Ln (HI premium) × interaction term	−.20	−.01	.03	−.04	−.14	−.05	.04	−.04	−.14	.04	.04
	(.04)	(.02)	(.06)	(.03)	(.04)	(.02)	(.02)	(.04)	(.03)	(.02)	(.02)
R^2	.39	.23	.33	.12	.26	.03	.05	.26	.09	.17	.05

NOTE.—EHI = employer-provided health insurance; HI = health insurance. The individual-level observations are from the 1996–2002 Current Population Survey (CPS). The sample is limited to those ages 22–64. Health insurance premiums are from Medical Expenditure Panel Survey (state-year data on premiums by policy type and employer size). Malpractice payments are from National Practitioner Data Bank. Labor market outcomes and employer health insurance information are from the CPS (March). Part-time work is defined as less than 30 hours per week. Instruments include real dollar amount and the number of medical malpractice payments per capita for different specialties (surgery, obstetrics-gynecology, internal medicine, and other) for current year and previous year. Premiums assigned based on state, year, and family structure (and employer size for employed). Regressions are weighted by March CPS weights, and standard errors are clustered at state level.

nomic theory is reassuring. It provides further evidence that our instrument is not picking up spurious underlying trends in labor market outcomes.

In table 5 we further explore the robustness of our identification strategy. We study the relationship between predicted health insurance premiums and variables, which should not be affected by our instruments. (The predicted premium on the right-hand side captures the variation in our instruments that is used in the IV estimation.) Columns 1–5 of table 5 demonstrate that the instruments are unable to predict variation in percentage black, educational attainment, gender, marital status, or health. Compositional changes in the levels of these variables could potentially affect the labor market outcomes that we study but should not be affected by the increase in malpractice payments—and they are not. To further test our identification strategy, we also include as a dependent variable the probability that an employee is included in an employer pension plan, shown in column 6 of table 5. This could be viewed as a falsification test—health insurance premiums might not be expected to affect pension benefits—but it is possible that when health plan costs go up, all other forms of compensation (wages and other benefits) are reduced to absorb the cost. This does not seem to be the case: the probability of an employee having a pension benefit does not respond to increases in health insurance premiums in the IV specification, with an insignificant coefficient estimate. Finally, in the last three columns on the right, we note the lack of relationship between predicted premiums and health outcomes (measured at the state-year level).[18] This finding rules out a class of explanations wherein the population of states with relatively higher malpractice payments is relatively sicker—and as sickness levels increase, health premiums rise. Population illness levels are not the driving factor here. Table 5 also notes that predicted premiums are not associated with higher cesarean-section rates (a procedure that is widely believed to be affected by the use of "defensive medicine").[19]

We can use our estimates to study the economy-wide impact of the growth of health insurance premiums. Using the estimates in tables 2 and 4, we can calculate the effect of rising health insurance premiums on the probability of being employed, employed as a full-time worker, average hours worked, and annual income. These estimates are summarized in table 6. A 20% increase in health insurance premiums (smaller than the

[18] Data on aggregate mortality come from the Area Resource File (National Center for Health Workforce Analysis 2003; reported at the county-year level, aggregated to the state-year level by the authors)

[19] Baicker and Chandra (2005) demonstrate that increases in medical malpractice liability are not associated with changes in physician flows or the greater use of surgical procedures. This finding rules out a situation where increases in malpractice payments affect both the price and quantity of health care received by workers; changes in the malpractice climate appear only to affect the price of health care as measured by health insurance premiums.

Table 5
Specification Checks

Sample: All	Black (1)	College Education (2)	Female (3)	Married (4)	Good Health (5)	Employer Pension (6)	Mortality Rates: Overall (7)	Mortality Rates: Cancer (8)	Cesarean-Section Rates (9)
Ln (HI Premium)	−.090	.048	.002	−.090	−.104	.180	.001	.0003	−.072
	(.061)	(.094)	(.062)	(.078)	(.099)	(.137)	(.002)	(.0004)	(.144)
R^2	.076	.068	.011	.002	.072	.038	.990	.091	.984
N	346,524	346,524	346,524	346,524	346,524	346,524	240	206	240
State and Year FEs	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Sample	CPS Micro	CPS Micro	CPS Micro	CPS Micro	CPS Micro	CPS Micro	State-year	State-year	State-year

NOTE.—The individual-level observations are from the 1996–2002 Current Population Survey (CPS). The sample is limited to those ages 22–64. The health insurance premiums are from the Medical Expenditure Panel Survey (state-year data on premiums by policy type and employer size). The malpractice payments are from the National Practitioner Data Bank. The labor market outcomes and employer health insurance information are from the CPS (March). The instruments include the real dollar amount and the number of medical malpractice payments per capita for different specialties (surgery, obstetrics-gynecology, internal medicine, and other) for the current year and the previous year. The premiums assigned are based on state, year, family, and structure (and employer size for employed). The regressions are weighted by March CPS weights, and standard errors are clustered at state level. Covariates include race, age, age^2, age^3, marital status, education, gender, and health status, excluding the dependent variable.

Table 6
Effects of 20% Increase in Premiums

	Mean	Coefficient	Effect
Probability of being employed (percentage point)	73%	−.12	−2.4%
Probability of working full time, conditional on working (percentage point)	84%	−.189	−3.8%
Average hours per week, conditional on working	41	−.237	−1.9
Average annual income (insignificant)	33,750	−.088	−594
Average annual income, conditional on working and having employer health insurance	41,442	−.205	−1,699

increase seen in many areas in the past 3 years) would reduce the probability of being employed by 2.4 percentage points—the equivalent of approximately 3.5 million workers. A similar number of workers would move from full-time jobs to part time, reducing the average number of hours worked per week by a little over an hour. Annual (wage and salary) income would be reduced by $1,700 for those who are employed and have EHI.[20] Together, these estimates demonstrate that the labor market effects of rising health insurance are far from neutral.

VI. Conclusion

Rising health insurance premiums, unemployment, and uninsurance have led to increased scrutiny of the labor market consequences of rising benefits costs. These relationships are, however, difficult to disentangle without a source of exogenous variation. We use variation in medical malpractice payments to derive the causal effect of rising health insurance premiums on wages, employment, and health insurance coverage.

We find that the cost of increasing health insurance premiums is borne primarily by workers in the form of decreased wages for workers with EHI—so that they bear the full cost of the premium increase. Our analysis implies that workers do at least partially value health insurance benefits but that there are impediments to full adjustment through wages, particularly for certain groups. Institutional constraints that prevent firms from discontinuing coverage only for those workers who value it least mean that as the costs of benefits rise, firms and workers have an incentive to move from full-time jobs with benefits to part-time jobs without. We see this increase in part-time work. Workers who value coverage the least will have the greatest incentive to move into jobs that do not offer coverage

[20] Ideally, we would have also been able to determine whether wage changes in "protected groups," such as union workers and those near the minimum wage, are smaller, with correspondingly higher negative employment effects. We performed a subgroup analysis to explore this hypothesis and noted that our results were too noisy to conclude anything systematic.

as premiums rise. We find that groups like married women, who are likely to have lower value of health insurance coverage through their employer, are more likely to lose coverage as premiums rise.

Our results on wage shifting are consistent with those in Gruber (1994): for workers with EHI, we observe full shifting of the increased price of health insurance onto wages. In addition, our results provide further evidence that the effects of increasing costs are borne disproportionately by particular groups.[21] In contrast to Gruber's study and to the results in Gruber and Krueger (1991), we find effects on both hours and employment. These results may appear to be contradictory, but they are not: in Gruber's study workers receive new maternity benefits, and in Gruber and Krueger they receive more generous workers compensation; both benefits are probably valued by workers, and the empirical finding of the full shifting of increased costs to wages with no effect on overall employment is consistent with the insights of Summers (1989). In our article, however, the increase in the price of health insurance premiums is driven by the medical malpractice crisis, a change that may not enhance the value of health benefits.[22] It is therefore unsurprising that workers do not value this increase in costs as highly and that the labor market responds with decreased wages and labor utilization.

These results have strong implications for policies designed to cover the uninsured. For example, if employer health insurance mandates raise the cost of employing workers, we should expect most workers to bear the cost through reduced wages. If some classes of workers are exempt from the mandate (such as part-time workers or those at particularly small firms), employers are likely to substitute uncovered jobs for covered ones, undermining the net effect of the mandate on insurance rates. More generally, rising health insurance premiums will place an increasing burden on workers and increase the ranks of both the uninsured and the unemployed.

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[21] Gruber (1994) finds that the cost of the maternity benefits are fully borne by married women. Gruber (1997) finds the payroll decline that accompanied Chilean privatization of social security was fully manifested in wages. Sheiner (1995) finds that demographic groups with higher ex ante insurance costs (such as older workers) experience full wage shifting when the price of health insurance increases.

[22] Other work by Kessler and McClellan (1996) did not find an effect of the increased utilization of medical services stemming from tort reforms on mortality from heart attacks and ischemic heart disease. There is also a related literature in medicine that notes that increases in marginal medical spending are not associated with improvements in patient satisfaction (Fisher et al. 2003a, 2003b; Baicker and Chandra 2004).

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