The Effect of an Employer Health Insurance Mandate on Health Insurance Coverage and the Demand for Labor: Evidence from Hawaii

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ABSTRACT

Over the past few decades, policy makers have considered employer mandates as a strategy for stemming the tide of declining health insurance coverage. In this paper we examine the long term effects of the only employer health insurance mandate that has ever been enforced, Hawaii’s Prepaid Health Care Act, using a standard supply-demand framework and Current Population Survey data covering the years 1979 to 2005. The coverage gap between Hawaii and other states increased over the sample frame, as did real health insurance costs, implying a rising burden of the mandate on Hawaii’s employers. We use permutation (placebo) tests across all states to examine the magnitude and statistical properties of these growing coverage differences and their impacts on labor market outcomes, conditional on an extensive set of covariates. As expected, the coverage gap is larger for less educated workers, who tend to have low rates of coverage in the voluntary market. Because of the high variability of wages over time, we are not able to identify an effect of Hawaii’s mandate on wages. A parallel analysis of workers employed fewer than 20 hours per week suggests that the law increased employers’ reliance on such workers in order to reduce the burden of the mandate. We find no evidence suggesting that the law reduced employment probabilities.
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I. INTRODUCTION

Over the past few decades, policy makers have considered employer health insurance mandates as a strategy for stemming the tide of declining health insurance coverage. Roughly twenty years before the Clinton Administration’s failed healthcare reform plan in the early 1990s, the Nixon administration proposed a health care initiative, the Comprehensive Health Insurance Plan, which included an employer mandate. At the state level, Massachusetts (1988), Oregon (1989), and Washington (1993) also passed employer-sponsored insurance (ESI) mandates, though each of these state laws was overturned in subsequent votes or voided due to conflicts with the federal Employee Retirement Insurance Security Act (ERISA) (Oliver 2004). More recently, The California Health Insurance Act of 2003 passed the state legislature and would have required firms with at least 20 employees to provide state-approved health coverage or pay a fee to a state-sponsored purchasing pool. This law was struck down by a relatively narrow margin (1.8 percentage points) in a direct referendum in November 2004.¹

Since these laws have not been adopted, direct evidence regarding the effects of an ESI mandate is scarce. As a result, analyses of the potential effects of an employer mandate tend to be based on simulations that assume that these effects would be similar to those of other policies, such as an increase in the minimum wage (Yelowitz 2003; Baicker and Levy 2007; Burkhauser and Simon 2007), or to other market changes that affect health insurance premiums, such as

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¹ Less extensive legislation exists in other states. For example, in 2006 Maryland passed the “Fair Share Health Care Fund Act” which would have required very large employers to spend at least 8 percent of payroll on health benefits. The law, dubbed the “Wal-Mart Bill” because it was written to target that company, was later struck down for violating ERISA. Similar bills have been proposed in other states.
increases in medical malpractice costs (Meara, Rosenthal and Sinaiko 2007). The validity of these assumptions and hence the validity of the inferences drawn from these simulation models is not known.

One state law that has been enforced for over two decades, and therefore provides a potential source of information on the effects of ESI mandates on coverage and labor market outcomes, is Hawaii’s Prepaid Health Care Act (PHCA). Hawaii’s mandate requires that virtually all private sector employers provide health insurance coverage to all employees working at least 20 hours per week. Despite the significance of this law and its relevance to ongoing policy debates, research on its effects is quite limited. The main reason for this is that the PHCA legislation was passed in 1974, five years before any national survey provided state-level information on health insurance coverage. Because of this timing, it is difficult to compare outcomes before and after the passage of the law.

While this is a significant data limitation, it does not preclude a meaningful analysis of the PHCA, for several reasons. First, because of legal challenges the validity of the law was in question until 1983. We show that in the years just prior to 1983, the percentage of workers receiving ESI coverage was very similar in Hawaii and the rest of the US. Around that time, however, coverage rates in Hawaii and other states began to diverge. By the late 1980s, the percentage of workers with health insurance through their own employer was more than 10 points higher in Hawaii; the most current data show a 13-point gap. This pattern suggests that if an employer health insurance mandate has significant effects on labor market outcomes, we should see differences between Hawaii and the rest of the US beginning to emerge in the early 1980s and grow more pronounced in subsequent years.
Moreover, while the provisions of the PHCA have remained unchanged since 1983, growth in health care costs has outpaced the general rate of inflation and the growth in average wages in subsequent years, which means that the cost to employers of complying with the law has increased over time. As such, data pre-dating the original PHCA legislation is not necessary for testing the hypothesis that the growing cost associated with the mandate has led to effects on wages, hours or employment.

In this paper we use data from March Current Population Survey (CPS) for the years 1980 to 2006 (data years 1979 to 2005), combined with data from the Monthly Outgoing Rotation Group (MORG) files from the CPS, to compare trends in health insurance coverage and three labor market outcomes—wages, hours and employment—in Hawaii and the rest of the US. The effect of an employer mandate on these outcomes should be strongest for workers who are unlikely to receive health insurance benefits in a voluntary market, such as lower skilled workers and young adults with a relatively weak demand for insurance. In contrast, a mandate should have little effect on workers who have high rates of ESI coverage in the absence of a mandate. To account for this heterogeneity in treatment effects we stratify our analysis by education, contrasting trends between workers with and without a college degree.

Our main analyses rely on a variation of Fisher’s permutation test, which entails comparing the usual difference in means between the U.S. and Hawaii to “placebo” comparisons between each of the 50 states plus DC and the remainder of the U.S., conditional on an extensive set of control variables. We find that by this metric Hawaii has an unusually high fraction of individuals with ESI coverage compared with each of the remaining states, consistent with the expected effect of the mandate on coverage. Using the same framework, we also find that Hawaii is not unusual in its distribution of wages and employment, although we uncover
relatively weak evidence suggesting that employers may have attempted to reduce the burden of the mandate by slightly increasing the fraction of workers employed in part-time positions that are exempt from the law.

II. BACKGROUND AND PREVIOUS LITERATURE

II.A. Health Insurance in Hawaii

Hawaii’s employer mandate legislation (PHCA) was passed in 1974, the same year that the US Congress passed ERISA, which established Federal regulation of employer-sponsored benefit programs, including health benefits. ERISA preempts state laws relating to benefit plans and has been interpreted by the courts as prohibiting state laws mandating employers to provide health insurance (Mariner 1992). Shortly after it went into effect in January of 1975, the PHCA was challenged on these grounds. The lawsuit was brought in 1976 by Standard Oil of California, which offered an employee health benefit program that did not comply with the standards set out by the Hawaii law. In 1977, the US District Court of Northern California ruled in the company’s favor. This decision was upheld by the US Court of Appeals in 1980 and by the US Supreme Court in 1981 (The Hawaii Uninsured Project 2004). In 1983, the U.S. Congress granted an ERISA exemption to PHCA. Because that Federal legislation specified that substantive changes to PHCA would void the exemption, the law has remained essentially unchanged since then (Oliver 2004).

The Act requires private-sector employers in Hawaii to provide health insurance coverage containing a minimum level of benefits (including inpatient hospital coverage, emergency room care, maternity care, and medical and surgical services) to employees working
20 or more hours per week. Other than part-time employees, exemptions also apply to new hires (employed less than four weeks), seasonal employees, commission-only workers, and “low-wage” employees. Employers must finance at least 50% of the premium cost and can require employees to contribute an amount no greater than 1.5% of their wages. Employers that fail to follow the requirements of the Act can be prevented from doing business in the state and required to pay for any health care costs incurred by their employees during the period of noncompliance.

The fact that the legal challenge to PHCA came so soon after its enactment combined with a dearth of data on insurance coverage prior to 1979 makes it difficult to determine when the law was truly in effect. According to some accounts the initial impact of PHCA on insurance coverage was quite modest. One estimate is that after the law went into effect, private insurance enrollment increased by no more than 5,000 individuals, slightly less than one percent of the state’s working age population (Friedman 1993, p. 54). Similarly, figures reported by the Health Insurance Association of America (HIAA) in their annual *Source Book of Health Insurance Data* do not indicate a major increase in ESI coverage in Hawaii around 1975.

It is not clear how vigorously the law was enforced during the period between the first court ruling in 1977 and exhaustion of the state’s judicial appeals in 1981. On one hand,

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2 To qualify a plan must provide benefits that are comparable to plans that have the largest number of subscribers in the state (Section 393-7). The determination of whether a plan meets this standard is made by the State Department of Industrial and Labor Relations in consultation with an expert advisory panel.

3 PHCA defines low-wage workers as those whose monthly earnings are less than 86.67 times the legislated hourly minimum. These typically would be workers not covered by the minimum wage or exempted from PPHCA by the part-time provision.

4 The same source estimates that between 10,000 and 30,000 previously insured workers saw the details of their coverage change as employers modified their health insurance offerings to comply with the law.

5 In fact, according to the HIAA publications, the number of people with health insurance actually declined between 1973 and 1975. However, over the entire decade of the 1970s, the HIAA data indicate that the number of people with insurance increased by between 2 and 3 percent, which is slightly higher than the state’s population growth rate (1.7% per year).
opponents of the PHCA did not file for an injunction to prevent it from being enforced while it was under appeal (Friedman 1993, p. 66). On the other hand, after the 1977 ruling the state Department of Labor and Industrial Relations suspended employer compliance audits (Agsalud 1982, p. 14). As we show in Section IV, by 1979 the percentage of Hawaiian workers with ESI in their own name was only slightly higher than the percentage for the rest of the country. After 1983, when Federal legislation established once and for all the legality of the PHCA, the gap widens.

**II.B. Theory: The Economics of Employer Mandated Health Insurance**

Summers (1989) showed how the labor market effects of an employer benefit mandate can be analyzed using a simple supply and demand framework. A benefit mandate will cause the labor demand curve to shift back and the labor supply curve to shift out, causing wages to fall. The magnitude of the wage change, and the effect of the mandate on hours and employment, will depend on how workers’ valuation of the benefit compares with the employers’ cost of provision. Let $B$ represent the employer’s cost of providing the benefit and $\alpha B$ refer to workers’ valuation of that benefit, i.e., their willingness-to-pay. The effect of a mandate on wages and employment will depend on $\alpha$.

If $\alpha \geq 1$, and absent labor market barriers, a mandate will cause wages to fall by the full cost of the employment provision or more, thereby preventing any reduction in the quantity of labor hired. However, in the absence of a pre-existing market failure, we would expect workers whose valuation exceeds the cost to receive the fringe benefit in a voluntary market. As a result, such workers should not be directly affected by an employer mandate. Instead, the most important effects of a mandate will be on workers who would not otherwise receive the benefit. In the simple supply and demand model, the absence of voluntary coverage can occur for two
reasons. One is that workers’ willingness to pay for the benefit is less than the employer’s cost—i.e., $\alpha<1$. Another is that workers’ wage is close to the minimum wage. In both cases, employers will not be able to reduce wages enough to offset the cost of the benefit, leading to a possible reduction in the quantity of labor demanded. In addition, if certain types of employees are exempt from the mandate, employers will have an incentive to substitute exempt for covered workers. Because the PHCA’s main exemption is for employees working fewer than 20 hours per week, the law may lead to increased reliance on such part-time workers. If this type of hours adjustment is not sufficient to offset the remaining costs of the mandate, employers are likely to hire fewer workers.

Several recent studies test for the effect of mandated insurance benefits on wages and employment. Gruber and Krueger (1991) analyzed the wage incidence and employment effects of state-mandated workers’ compensation insurance. They find that higher workers’ comp premiums are almost perfectly offset by wage reductions, leaving total compensation and employment essentially unchanged. Similarly, Gruber (1994) found that state and later federal laws mandating that private health insurance policies cover maternity benefits led to wage reductions for those workers most likely to be affected by the laws: married women of child-bearing age. Because the wage reductions were comparable to the cost of the benefit mandate, the laws had no significant effect on employment.

Kaestner (1996) examined the effect of state-mandated workers’ compensation and unemployment insurance on labor market outcomes for teenagers and young adults. Because young workers are less likely to value these benefits at their full cost and are more likely to have wages near the legal minimum, their wages may not adjust sufficiently to offset the cost of the mandates. As a result, for this population there is greater reason to expect negative employment
effects. Kaestner’s empirical results are mixed. For teenagers, the pattern of his results is consistent with a simple demand and supply model in the presence of wage constraints: he finds no evidence of a wage offset and a negative effect on employment. He also finds negative employment effects for older workers, though his wage results suggest that the cost of the mandated benefits were more than offset by a reduction in wages. This combination is inconsistent with the demand-supply model.

An important consideration in analyzing the possible effects of a health insurance mandate like the PHCA is the increase over time in the cost of the mandated benefit. As such, the mandate is unlikely to cause a one-time shift in the labor demand and supply curves. Rather, the potential impact of the policy on wages, hours and employment is likely to grow over time as the cost of the mandated benefit rises.

**II.C. Past Research on Hawaii’s Mandate**

There are several significant challenges associated with analyzing the impact of Hawaii’s PHCA. The first is the ambiguity concerning when the policy truly went into effect. It is not clear whether the biggest changes in employer and worker behavior should have taken place just after 1975, when the law was initially implemented, or after 1983, when the legal issues concerning its validity were finally resolved. A second and related complication is the dearth of data on key outcomes prior to the passage of the legislation. The most commonly used source of individual-level information on health insurance is the March CPS, which did not include questions on health insurance prior to the 1980 survey. Information on labor market variables is available for earlier years, but individual states cannot be identified prior to 1977. Because of these data limitations, the small number of studies that analyze the mandate’s effect on insurance coverage and labor market outcomes are based mainly on data from after 1983.
Any analysis of the PHCA must begin by estimating the law’s effect on insurance coverage. As noted above, an employer health insurance mandate is not likely to have an effect for the majority of workers who would receive ESI in a voluntary market. For other workers, exemptions and non-compliance by employers may mute the effect of a mandate on the number of workers receiving health insurance from their employers. The effect of the mandate on insurance coverage may be further reduced if some workers who gain health insurance coverage through their employer would have otherwise received insurance through another source, such as their spouse’s employer, a private non-group policy, or public insurance.

Several prior studies compared insurance coverage in Hawaii and the rest of the US, with mixed results. Based on an analysis of CPS data from the mid-1980s, Dick (1994) concluded that the PHCA did not raise insurance rates in Hawaii relative to other states. However, the other studies, which use later CPS data, found that ESI coverage is significantly higher in Hawaii and attributed this result to the PHCA. Thurston (1997), who used data from 1990 to 1993, reported an 11-point difference between Hawaii and the rest of the United States in the percentage of workers receiving health insurance through their employer. While he did not adjust for observable differences in worker characteristics across states, he provided two additional pieces of information to support his interpretation that this difference is the result of Hawaii’s mandate. First, he presented aggregate data from 1969 showing that the percentage of Hawaiians with hospital insurance (88.3%) was roughly the same as the percentage for the US as a whole (87.2%). Second, he presented a scatter plot showing sizeable Hawaii-U.S. differences in coverage rates for industries where coverage rates are low in the rest of the United States and essentially no difference for industries where ESI coverage is most common. This pattern is
consistent with mandate effects that were concentrated mainly on firms that would not provide insurance in a voluntary market.

Kronick et al. (2004) assessed how much of the cross-state variation in the extent of ESI coverage can be explained by individual and area-level characteristics. Based on these explanatory variables, Hawaii’s predicted ESI rate is slightly below the average for all states. This indicates that while Hawaii may have a unique geography and history, it is not an outlier in terms of economic factors that influence health insurance coverage. Kronick et al. showed that Hawaii’s actual coverage rate is 11 points higher than this predicted rate. Lee et al. (2005) also found that the percentage of Hawaiian workers with ESI is substantially higher than would be expected based on individual and job characteristics.

Differences in data and research design make it difficult to reconcile these divergent results regarding the coverage effects of the mandate. Dick’s finding of little or no difference between Hawaii and other states may be due to his reliance on a sample that includes non-workers, whereas the other two studies analyzed workers and focused mainly on ESI. Alternatively, it is possible that the passage of PHCA did little to increase insurance coverage immediately, but once the law was in place it slowed the erosion of employment based coverage as health care costs increased, leading to growing differences between Hawaii and other states over time (Neubauer 2003).

To the extent that the PHCA prevented employers in Hawaii from dropping coverage in response to rising premiums, we should expect to see adjustments along other margins, such as wages, hours, or employment. The existing literature provides limited evidence on these labor market effects. Thurston (1997) investigated the wage effects of Hawaii’s mandate using industry-level data from 1970 and 1990 and found mixed evidence for wage reductions due to
the expansion of ESI. When he limited his analysis to data from Hawaii, he found that wages grew less in industries that should have been most strongly affected by the PHCA as a result of their low rates of voluntary ESI provision. However, when he used other states to control for industry-specific changes over this twenty year period, the results implied that Hawaii’s mandate had either no effect on wages or that the law contributed to significantly higher wage growth in Hawaii relative to the rest of the United States.

Thurston also tested for an effect on part-time work, again using data from the 1970 and 1990 Censuses aggregated to the industry level. His results suggested that after the passage of Hawaii’s law, the percentage of workers with low hours increased more in Hawaii than in the rest of the US. One limitation of his analysis, however, is that his measure of “low hours” does not match up exactly with the 20 hour cut-off used in Hawaii to distinguish between covered and exempt workers. Lee et al. (2005) constructed matched samples to compare the distribution of hours worked in Hawaii and other states. Pooling data for the period from 1994 to 2004, they found that the percent of individuals working less than 20 hours per week was nearly the same in Hawaii (4.07%) as in their matched sample of residents of other states (4.42%). However, the percent of workers reporting exactly 20 hours was higher in Hawaii. Given that hours data are subject to rounding, they interpreted this as suggesting a tendency for employers in Hawaii to hold worker hours just below the level determining coverage by the PHCA.

Much of the discussion and critique of employer mandates has centered on their potential negative effects on employment. Several papers argued that for workers near the minimum wage, an employer mandate will have effects similar to an increase in the minimum wage (Yelowitz 2003; Baicker and Levy 2007; Burkhauser and Simon 2007). Taking parameter estimates from studies finding a negative employment effect of the minimum wage, they forecast
that an employer mandate would reduce employment of low skill workers. However, none of the
prior studies on Hawaii’s PHCA tested whether the law had such an effect.

III. DATA AND RESEARCH STRATEGY

Our main source of data is the Current Population Survey (CPS). The analysis of health
insurance uses data from the March CPS, which provides information on health insurance
coverage during the prior calendar year, including whether or not a respondent received ESI
coverage on the longest job held during that year. For the analysis of labor market outcomes, we
use data from the Merged Outgoing Rotation Group (MORG) files; compared with the March
data, the MORG files provide larger sample sizes for point-in-time measures of wages and
employment status, which are not subject to the recall bias that may affect the retrospective data
from the March survey. For both sets of analyses, we constructed 27-year repeated cross-
sections that include all available years for which the health insurance questions were asked in
the March survey (survey years 1980-2006, which correspond to the reference years of 1979-
2005, the same as in the MORG files). In both data sets, we focus on workers ages 18-64 who
are employed in the private sector; we exclude government employees from the analyses because
the law does not apply to them.

As discussed above, our data starts after the PHCA legislation initially went into effect
but before the legal issues about its validity were resolved with the 1983 ERISA exemption.
Thus, a true “pre/post” analysis is not possible. However, even if pre-1974 data were available,
this approach or a standard “difference in differences” model would not be an ideal research
design. These approaches are most reasonable for policies that are expected to have a one-time
permanent effect. The effect of a benefit mandate will depend on the costs that the mandate
imposes on employers. In the case of a health insurance mandate, this cost will grow over time as health care costs increase in real terms. Indeed, it is conceivable that the immediate impact of Hawaii’s mandate was minimal because very few workers gained coverage through the law and because in the 1970s health insurance premiums were small relative to wages.

Figure 1 provides a sense of how the cost to employers of the PHCA has increased over time. The graph plots real ($2006) single coverage health insurance premiums in Hawaii expressed as an hourly cost per month for a full-time worker for the years 1974 to 2006. The premium data are from the Hawaii Medical Service Association (HMSA), the largest private insurer in Hawaii, and they correspond to the most popular community-rated plan that HMSA sells to employer-sponsored groups of 100 or less. In 1974, the single coverage premium for a plan meeting the standards of the PHCA was $15.96, which for an employee working 40 hours per week translated to a nominal cost of 9 cents per hour, or 38 cents in current dollars. The cost per hour was only slightly higher in 1980 (46 cents), but then increased steadily thereafter. By 2006 the monthly premium was $261, or $1.51 per hour for a full-time employee, more than three times the real cost in 1980.

Another way to gauge the cost of the mandate is to compare these figures to the minimum wage, which fell in real terms over this period. In 1974, the hourly cost of single coverage health insurance in Hawaii was roughly 5% of the state’s minimum wage of $2.00. Thus, for firms employing workers at the minimum wage, the initial cost associated with the mandate was small. For workers with hourly wages of $2.09 and more, there was no legal impediment preventing employers from reducing wages to fully offset the cost of health insurance. By 2006, however, the mandate increased the cost of employing workers at Hawaii’s minimum wage (of $6.75) by

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6 We thank Jennifer Diesman of HMSA for providing these data.
22%. The steady increase over time in the costs of complying with the mandate indicates that the effect of the PHCA on labor market outcomes is likely to be reflected in a long-run divergence between Hawaii and other states rather than a one-time shift.

As noted, the economic theory regarding mandates predicts that the effect of an employer mandate on health insurance coverage should be greatest for workers who have low rates of ESI coverage in a voluntary market, such as younger and less skilled workers. By the same logic, any effect of the mandate on hours or employment should also be concentrated on these workers. In contrast, the mandate should not affect the demand for higher skill workers who would receive ESI offers in the absence of the law. Thus, models that test for employment or hours effects for high skilled workers represent a type of falsification test.

The contrast between high and low skill workers may be less informative for the analysis of wages. Recall that the PHCA not only requires employers to offer insurance but also caps the amount that employees can be required to pay for that coverage. If this cap is binding, wages will respond more strongly to rising health insurance premiums in Hawaii than in other states even among workers who would receive coverage in the absence of the mandate. In fact, data from the Medical Expenditure Panel Survey-Income Component (MEPS-IC), a national survey of establishments, show that among establishments that offer health insurance, employee premium contributions for single coverage are substantially lower in both dollar and percentage terms in Hawaii than in the rest of the US. For example, in 2005 among establishments with 100 to 999 employees, the average employee contribution for single coverage in Hawaii was 7% of the total premium, compared to an average of 19% across all states.7

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7 Larger establishments allow for a clean comparison because essentially all firms with 100 or more employees offer health insurance. State-level data for all establishment sizes are available at
IV. INSURANCE COVERAGE EFFECTS

IV.A. Trends in ESI Coverage

We begin our analyses by comparing trends in insurance coverage in Hawaii and the rest of the United States. Figure 2 displays data on the percentage of private sector workers who received health insurance coverage through their own employer, for the years 1979 to 2005. The first thing to note is that in the early years the coverage rates in Hawaii and the rest of the country are very similar. The plots begin to diverge in 1982, the year before Congress granted Hawaii's ERISA exemption. From that year on, coverage fell slightly in Hawaii but less than in other states. By 1988, the ESI coverage rate was 11 percentage points higher in Hawaii. After 1988, there is some year-to-year variability in the gap, presumably because of sampling error in Hawaii's relatively small sample. By 2005, 70% of private sector workers in Hawaii received health insurance through their employer, compared to only 57% in the rest of the US.

To account for the fact that the employer mandate should have affected high and low skill workers differently, we stratify the analysis by education, presenting trends for workers with a high school degree or less (Figure 3A) and those with at least a college degree (Figure 3B). The results are consistent with our expectations. The general pattern for less educated workers is similar to the full sample results. Whereas in 1979, Hawaii's coverage rate was comparable to the rest of the United States, in 2005 the rate in Hawaii was 18 percentage points higher than in the rest of the nation. The graph for college educated workers is quite different. Apart from a decline in the ESI coverage rate in 1987, which may be related to changes in the CPS insurance

questions that were implemented in the 1988 CPS, there is little trend in these data for either group.

It may seem surprising that the coverage rate in Hawaii is well below 100 percent, given that there are few exemptions to the PHCA. Because the March CPS provides no information on whether a worker was offered insurance, it is not possible to assess employer compliance using these data. Other data sources, however, shed light on this issue. According to data from the MEPS-IC, 99% percent of Hawaiian establishments with between 10 and 24 employees offer health insurance. Three quarters of Hawaiian establishments with fewer than 10 employees report offering insurance, compared to 35% of comparably sized establishments in other states.8

Worker-level information on ESI offers, eligibility and take-up is available in the CPS Contingent Worker Supplement (CWS) fielded in 1995, 1999, 2001 and 2005. According to these data, over 60% of Hawaiian workers without own-name ESI coverage work for a firm that offers insurance. Roughly three-fifths of this group declines coverage, almost always because they are covered through a family member. Most workers whose employers offer ESI but themselves are not eligible cite low hours or job tenure as the reason. Roughly 30% of workers in firms not offering insurance would not be eligible anyway, mainly because they work fewer than 20 hours per week. The remaining workers, whose lack of own-name ESI coverage cannot be explained by variables in the CWS, represent 7% of all private sector employees in Hawaii. This figure probably overstates the degree of non-compliance because even in the CWS it is not possible to fully account for the PHCA’s coverage rules.

IV. B. Regression-Adjusted Differences in ESI Coverage

8 These MEPS figures are from tables available online. It is reasonable to assume that a substantial number of the very small firms that don’t offer coverage employ only workers who are exempt from the PHCA because they work short hours, are hired on a seasonal basis or have an alternative source of coverage. However, we are not able to test that assumption directly.
The data in Figures 2 and 3 are not adjusted for differences between Hawaii and the rest of the United States in terms of worker or firm characteristics. To control for these factors we can estimate regressions of the form:

$$I_i = X_i \beta^I + \delta H_i + u_i$$

(1)

where the dependent variable equals one if individual i receives own-coverage ESI and zero otherwise, $X$ is a standard set of covariates, and $H$ is an indicator variable for observations from Hawaii. Under the assumption that the model is correctly specified, the simplest inference procedures to test the null hypothesis that $\delta=0$ assume that the only reason we observe departures from the null hypothesis is due to sampling variability. Another possibility is that equation (1) is misspecified and we might obtain a non-zero estimate of $\delta$ even if the null were true and there was no sampling variability. To address this possibility, usual practice (see for example, Moulton (1990)) is to estimate a variant of (1),

$$I_i = X_i \beta^I + \delta H_i + e_s + u_i$$

(1')

where $e_s$ is a state random effect, i.e., $E(e_s|X, H) = 0$ and $\sigma_s^2 \neq 0$. Sometimes parametric assumptions about the distribution of $e$ and $u$ are made, as in the classical “random effects” model (e.g., Searle, Casella, and McCulloch 1992). A key weakness of this approach is that it is often unclear how sensitive inferences are to the specific (and largely untestable) assumptions about the structure of the unobserved components of variance; moreover, frequently choices
researchers have made about these assumptions appear to have lead to insufficiently conservative
inferences (e.g., Moulton 1990).

We therefore rely on an alternative approach that is a variant of Fisher’s permutation or randomization test (Fisher 1935).9 The classical permutation test is the “two sample problem” with random assignment to either a treatment or a control group. The null hypothesis is that the treatment has no effect and that the two samples come from the same population. An appropriate statistic (such as the difference in means) is computed for the two samples, call it $d$. Under the null hypothesis the labels—“treatment” or “control”—are arbitrary and can be reassigned to the units. If the size of the two groups is $n$ and $m$ it is possible to compute the “difference in means” associated with all possible $(n+m)/(n!m!)$ samples created by permuting the labels. The test is performed by computing the percentile that $d$ represents in the distribution of all the “placebo” differences. If the observed difference $d$ is large relative to these placebo differences, the treatment difference is judged “significant.” As a practical matter, in such two sample problems the conclusion derived from the permutation test is no different than the t-test, and Fisher’s intent in deploying it was merely to demonstrate the utility of conventional t-statistics even when the underlying data was not normal.

Identical application of Fisher’s test to our situation (merely permuting the labels “Hawaii” and not “Hawaii”) confirms Fisher’s finding that it is similar to the usual t-test. We identify the most important problem as the failure of the i.i.d assumption, specifically that because of inadequate controls, misspecification, etc., errors for individuals within the same state are more correlated than errors for individuals in different states. To remedy this problem we

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9 For a simple application to an economic question and some discussion see Johnston and DiNardo (1997) Chapter 11.2. For a careful formal discussion of the statistical assumptions underlying the test see Lehman (1959), Chapter 5.7. For some colorful *ad hominem* about those who might use such “non parametric” permutation tests, see Fisher (1935, chapter 3).
consider a variant of equation (1’) where we treat the parameters \( E[e_s] \) as sharing a common
distribution with zero mean under the null hypothesis that the Hawaii mandate had no effect, and
we rely on the “large” number of U.S. states (and the large number of observations per state) to
provide 50 (estimated) values of \( e_s \).

The intuition underlying this test is straightforward: imagine that a researcher misapplied
the state labels so that the dummy variable denoted by \( H_i \) in equation (1) actually was set to one
for some other state (not Hawaii). Under the assumptions of the research design (and given our
large samples), the coefficient in this “placebo” case should be very close to zero. The
permutation test we employ uses all possible placebo estimates as the standard to judge the
significance of what is essentially the “Hawaii effect” by comparing it to these 50 placebo
estimates—one for each other state plus the District of Columbia. If the coefficient on Hawaii is
similar to these other placebo estimates, our inference is that the effect is not “real”, even if it
were to pass a conventional test of significance using the usual inference procedures.\(^{10} \) We
discuss the details of the test in Appendix B as well report some Monte Carlo results on the
performance of the estimator.

In Table 1 we report estimates of \( \delta \) from regressions that compare Hawaii to the other 50
states. To ensure a reasonable number of observations from Hawaii, we pool data from the 2003
through 2006 March CPS files (reference years 2002-2005). Instead of reporting standard errors,
we display the 2.5 and 97.5 percentile of the placebo estimates. The range between the two can
be thought of as the 95 percent confidence interval for our state-level estimates. If the point
estimate for Hawaii lies within this interval, we interpret the results as failing to reject (at

\(^{10} \) Indeed, taking the usual expedient of using “clustered” standard errors at the state level, many of our
“placebo” estimates are significantly different from zero at conventional levels; for some outcomes we
investigate, all of the placebo estimates are significant.
conventional significance levels) the null hypothesis that the true effect is zero. If it lies outside this interval we interpret this as evidence against the null and in favor of the alternative that Hawaii is truly different from the comparison group with respect to the outcome.

Unadjusted results are presented in the top panel. For the full sample, the Hawaii coefficient lies well outside the distribution for the other states, providing strong evidence that Hawaii has significantly higher ESI coverage than does the rest of the United States. The same is true for the less educated subsample. Hawaii is not unusual, however, in terms of the ESI coverage of college-educated workers. Although the unadjusted own name coverage rate for Hawaii is higher than the mean for all other states, there are several other states with comparable rates.

The estimates presented in the middle panel of the table condition on worker demographics (see the table note for the complete list of controls). The full distributions of placebo effects for this model are presented in Figure 4. The two thin red lines show the 2.5 and 97.5 percentile of the distribution for other states and the thick blue line shows the coefficient for Hawaii. When we condition on worker demographics, the point estimates for Hawaii are smaller, but the range of estimates is smaller as well, and the qualitative story is the same. In the full sample, the coefficient on the Hawaii dummy is .095. When we run the regression on the other states, the next highest estimate is .058 and all the remaining estimates are less than .05. In the less educated sample, the adjusted gap between Hawaii and all other states is 15 percentage points. The next highest state is Nevada, with an estimated coefficient of .107; no other state has a coefficient higher than .05. Thus, for the full sample and for workers with a high school degree

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11 The results in Table 1 are coefficients from a linear probability model. We obtain essentially identical results by estimating a logit model, and calculating for each observation the difference between the predicted probability if the Hawaii indicator is set to one and the prediction if it is set to zero.
or less we can easily reject the null that the difference between Hawaii and the other states is zero. The data do not support this conclusion for college-educated workers.

When we add job characteristics (dummies for industry, occupation, and firm size), the estimated gap between Hawaii and other states is larger. In this specification, the estimate for college-educated workers meets our standard for significance, though the gap between Hawaii and the next closest state remains substantially smaller than for the less educated sample.

While we cannot use this research design to determine how precisely much of Hawaii’s higher rate of own-name ESI coverage should be attributed to the state’s employer mandate, these results are strongly consistent with the hypothesis that the law significantly raised own-name ESI coverage relative to the complete range of counterfactual outcomes for the other states.

IV.B. The Distribution of Insurance Coverage (all sources), 2002-2005

The results in Figures 2 and 3 and Table 1 are for own-name ESI coverage, the outcome that should be most directly affected by an employer mandate. However, the ultimate goal of the policy is to increase insurance coverage overall. The effect of an employer mandate on the number of people with health insurance will be muted if the policy’s main effect is to shift the source of coverage—for, example from a spouse’s employer or an individually purchased policy—rather than reduce the number of people who are uninsured. Therefore, it is important to consider not only the effect of an employer mandate on the number of workers receiving insurance through their own employers, but on the distribution of coverage from all sources. To this end, Table 2 reports the distribution of coverage in Hawaii and the rest of the US. Again, to ensure adequate sample sizes, we pool data for reference years 2002 to 2005.

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12 The main reasons for this are that the average firm size is lower in Hawaii and a smaller percentage of workers are employed in manufacturing. Apart from these differences, however, the distribution of job characteristics in Hawaii is quite similar to the distribution in the rest of the country.
Hawaiian workers are less likely to have ESI coverage through a dependent or to purchase non-group insurance, though in both cases the differences are very small. As a result, the 13 point gap in own name ESI coverage translates to an 11 point gap in overall ESI coverage and a 9.4 point gap in the percentage with any private insurance. Because the percentage of workers with public insurance is the same for Hawaii and the rest of the country, the 9 point difference in the percent uninsured is entirely explained by the higher rate of ESI coverage in Hawaii.

When we cut the data by education we see again that the PHCA has very different effects on workers of different skill levels. Among workers with a high school degree or less, the percentage of Hawaiian workers receiving health benefits from their employers is 16 percentage points (or 35%) higher than the rate for the rest of the US. Less than one percentage point of this difference can be attributed to workers shifting the source of their coverage from the employer of a family member to their own employer. Less educated workers in Hawaii are less likely to have public insurance than those in other states. The net effect of these differences is that among workers with a high school degree or less the percent uninsured is 13.7 percentage points lower in Hawaii than in the rest of the U.S. Relative to the rate for the rest of the country, this is a 40% effect.

In contrast, the figures in the last two columns suggest that Hawaii’s employer mandate has not affected the insurance coverage of higher skilled workers in Hawaii. Among college-educated workers, the slightly higher rate of own name ESI coverage in Hawaii is mostly offset by a lower rate of coverage as a dependent. This pattern is consistent with the incentives generated by the PHCA rules on premium contributions. The law constrains the amount that employees can be required to contribute toward single coverage premiums but places no limits
on contributions for family coverage, which means that the incremental cost of adding dependents is higher in Hawaii than in other states. Therefore, married couples with two sources of ESI are generally better off taking two single policies (or one single and one employee plus children) than to obtain family coverage through one spouse’s employer. Data from the CPS Contingent Worker Supplements provides corroborating evidence of this type of behavior: in those data the percentage of college-educated workers offered health insurance by their employer is essentially identical for Hawaii and the comparison group, ranging from 85 to 90 percent depending on the year of the survey.

The percentage of college educated workers with other types of coverage is also similar in Hawaii and other states. As a result, the percent uninsured is essentially the same as well.

V. LABOR MARKET OUTCOMES

The coverage results strongly suggest that after the 1983 the PHCA was a binding constraint for a significant number of Hawaiian employers, especially those employing less-skilled workers. While the Hawaii-U.S gap in coverage has increased only slightly since the early 1990s, the cost associated with the mandate has continued to grow because the growth in health insurance premiums has exceeded general inflation. The basic demand-supply model predicts three possible labor market effects. First, to the extent that employers are able to pass the cost of the mandate on to workers, wage growth should be lower as a result of the mandate. To the extent that the wage offset is incomplete, perhaps because of a binding minimum wage, the mandate may affect labor demand by raising the cost of employing less skilled workers. In this case, we may see an effect on hours if employers make greater use of part-time workers who are exempt from the mandate. Alternatively, firms may simply hire fewer workers.
A common approach for analyzing the effect of state policies is to use a difference-in-differences model in which changes for the state enacting the policy are compared to changes for a set of “control states”. A typical assumption is that the policy induces a one-time shift in the outcome of interest. As noted above, such a specification does not make sense for analyzing a policy like the PHCA as the cost of the mandate has grown over time. Rather, to test whether Hawaii’s employer mandate has led to reduced wage growth over time, a shift to part-time employment or a long-run decline in employment, we can estimate models of the form:

\[ Y_i = X_i' \beta + \gamma_{\text{Hawaii}} + \delta \text{TREND} + \theta \text{TREND} \times \text{Hawaii} + \epsilon_i + u_i. \] (2)

where the vector of control variables \( X \) contains the same set of worker demographic characteristics used for the analysis of coverage outcomes in the preceding section. The coefficient of interest is \( \theta \), which indicates whether the trend in the labor market outcome of interest was different in Hawaii than in other states. Based on the basic demand-supply model and the features of Hawaii’s law, we would expect \( \theta \) to be negative in wage and employment regressions and positive for low-hours regressions in which the dependent variable equals one for individuals working fewer than 20 hours per week.\(^{13}\)

As discussed in the previous section, the results of a simple regression may be misleading if there are other factors that vary across states other than Hawaii and affect trends in wages, hours, or employment. For this reason, we assess the “significance” of the parameter estimates

\(^{13}\) In principle, an alternative specification would be to include additional interactions to account for the period between 1981, when the PHCA was finally repealed, and 1983, when it was reinstated. However, given how short the “repeal period” was and given the ambiguity about how strongly the law was enforced between 1979 and 1981, we are doubtful that such a model would produce meaningful results. As an alternative approach, we varied the start year of the analysis period. Doing so does not materially affect the results.
by comparing them to the distribution of estimates obtained by running each regression 51 times, once for each state (plus the District of Columbia). As in the case of the ESI regressions discussed in the preceding section, if the parameter estimate for Hawaii falls outside the range of the estimates obtained for these “placebo” treatments, we conclude that the outcome in Hawaii is significantly different than in other states; given the extensive controls employed in these regressions, such differences are likely due to the ESI mandate. If on the other hand the results we obtain for Hawaii lie within the range of results obtained when other states are treated like Hawaii, statistical inferences regarding the impact of the PHCA are not warranted. As usual, such a null finding would not necessarily imply that the law had no effect; it is also consistent with the conclusion that we are not able to detect an effect in the data with our research design.

V.A. Wages

Results for the wage analysis are listed in Table 3. We report estimates for the full sample and the sample stratified by educational attainment, listing the results for two parameters: the coefficient on the state dummy and the interaction between that variable and the linear time trend. In the table, we report the 2.5 and 97.5 percentiles of the empirical distribution of parameter estimates from our placebo regressions. Figure 5 is a histogram for the full distribution of the placebo estimates of the two coefficients.

For each sample, the coefficient on the Hawaii dummy is near the midpoint of the 95 percent confidence interval obtained from the 50 placebo regressions. This implies that before the PHCA was reinstated in the early 1980s, the distribution of wages in Hawaii was not significantly different from other states.

Our primary focus is on the interaction between the state dummies (Hawaii or other states) and the time variable. In the case of Hawaii, this coefficient is interpreted as representing
the cumulative impact of the state’s ESI mandate on wages over our sample frame. Consistent with the expected negative wage effect of the mandate in a simple supply-demand framework, this coefficient is negative. The point estimate is between -.002 and -.003, depending on the estimation sample. In the full sample, the t-statistic (based on robust standard errors clustered by states) is -4.21.

While a t-statistic of this magnitude would typically support a rejection of the null hypothesis, the results from our placebo regressions suggest otherwise. When we run this regression 51 times, each time comparing one state to the others, the absolute t-statistic for the state/time interaction is greater than two in 37 cases. For all three estimation samples, the coefficient for Hawaii lie near the midpoint of the range of coefficients obtained by applying the placebo test to the other 50 states. As such, our test strongly fails to reject the null hypothesis that Hawaii exhibits the same pattern of changes in relative wages over time as any other state. It is important to note that this confidence interval includes negative wage effects that are sufficiently large as to completely offset the cost of the insurance mandate. Thus, we should not interpret this result to mean that the PHCA had no effect on wages, but rather that any effect the mandate may have had on wages is not detectable in our data.

**V.B. Part-Time Employment**

In Figure 6 we present trends in the percentage of all private sector employees working fewer than 20 hours per week, which is the threshold that determines whether a worker is covered by or exempt from the PHCA. The first thing to note is that such short hours are relatively uncommon: for the full time period, fewer than 5% of private sector workers in either Hawaii or the rest of the country report having usual weekly hours of less than 20. This suggests that the potential scope for employers to avoid the cost of the mandate by shifting certain
employees into exempt part-time positions is limited. In the rest of the U.S., there is a very slight downward trend in low-hour employment. The percentage peaks at 5.7 in 1983-1984 and then decreases gradually thereafter. In 2006, 3.9% of private sector workers in states other than Hawaii report working fewer than 20 hours per week. Because of smaller sample sizes, the data for Hawaii show more variability. In the early years, the graph for Hawaii is generally below the graph for the other states, while in later years the percent working low hours is slightly higher in Hawaii. However, in each year, the difference is quite small.

Table 4 reports results from regressions that test whether there has been a long run trend toward low hour work in Hawaii relative to other states. As with the wage regressions we report estimates of two parameters: the coefficient on the state dummy and the interaction between that variable and the linear time trend. In the full sample, the Hawaii dummy has an estimated coefficient of -.011, which is well within the 95 percent confidence interval implied by the 50 placebo regressions. This implies that before the PHCA was reinstated in the early 1980s, Hawaii was comparable to other states in terms of low-hour employment and that these other states represent a reasonable baseline period control for Hawaii.

The estimated coefficient on the interaction term is .006, indicating that over the entire period the percentage of adults in short-hour jobs grew faster in Hawaii than in all other states combined. The fact that the estimate is slightly larger for less educated workers (.007) and zero for college-educated workers is consistent with the predictions of the basic supply-demand model that some employers in Hawaii will respond to the health insurance mandate by hiring low-skill workers on a part-time basis in order to avoid the cost of providing them insurance. The model does not predict a similar effect for high-skill workers because the costs of the mandate are lower for them.
On net, the comparison of the Hawaii results to the placebo estimate is not as striking as it is for own-name ESI coverage, though it is highly suggestive of a shift toward part-time employment in response to the cost of the mandate. There is only one state, New Hampshire, where the percentage of employees working fewer than 20 hours per week grew faster than it did in Hawaii.

However, even if we take the point estimates at face value, it is not clear how economically significant this effect is. The results for the less educated subsample imply an increase in low-hour employment of less than 2 percentage points over the full 27 year period. While this might be seen as a large relative to the mean of the dependent variable, it is small relative to the number of workers at risk for being shifted into exempt employment.

**V.C. Employment**

The final outcome we consider is employment. The unadjusted data displayed in Figure 7 show that in the early 1980s the employment rate was about 3 percentage points higher in Hawaii than the rest of the U.S. After 1990, the average gap was 1 percentage point in Hawaii’s favor. Over the whole period, the employment rate for Hawaii tracks quite closely with the overall rate for the other states.

Table 4 reports results from a set of employment regressions. The specification of our employment regressions is similar to the model used for wages and part-time work, though now the sample consists of all working-age adults (18 to 64 years old). The coefficient on the Hawaii dummy indicates that in the baseline year of 1979 the employment rate among adults in Hawaii was 4.6 points higher than the average for all other states, controlling for demographic factors. Only one state, North Carolina, had a higher adjusted employment rate in that year. The coefficient on the interaction term for Hawaii is -.0005. The robust standard error is .0002,
which implies a t-statistic of -3.23. While this might suggest that Hawaii’s employment rate fell significantly relative to other states in the years after the PHCA was reinstated, the results of the placebo tests caution against this interpretation. Hawaii ranks 20th among all states in terms of the linear trend in the employment rate over the period 1979 to 2005. Based on this, we cannot reject the null hypothesis that the trend in Hawaii was no different than the trend in other states. The interaction coefficient is larger in magnitude for the less educated subsample, though again it falls well within the confidence intervals implied by the placebo tests.

V. DISCUSSION AND CONCLUSIONS

In 1979—several years after the passage of Hawaii’s Prepaid Health Care Act but before the legal issues regarding the law’s validity were resolved—the percentage of Hawaiian workers receiving health insurance through their employers was only slightly higher than the percentage in the rest of the United States. In the decade that followed, Hawaii did not experience the same decline in ESI coverage as the rest of the US. By the mid-1990s, the percentage of workers with ESI from their own employer was 11 or more points higher in Hawaii. The most current data show a gap of 13 points. The difference is substantially larger for workers with personal attributes that imply low coverage probabilities in the remainder of the United States. This higher rate of ESI coverage brought about by Hawaii’s employer mandate did not lead to universal health insurance coverage in Hawaii, though if we take the experience of the rest of the United States as a counterfactual, Hawaii’s law did significantly reduce the number of uninsured workers in the state.

Critics of employer mandates argue that even if an employer mandate can increase insurance coverage, these gains come at the cost of labor market distortions. In particular, to the
extent that it raises the cost of certain types of workers, an employer mandate may affect labor demand. Employers may respond by either hiring fewer low skill workers or by moving such workers into part-time jobs that are exempt from the mandate. Hawaii’s experience does not provide an ideal natural experiment for testing for such effects. Ambiguity concerning the enforcement of the law while its legality was being challenged combined with a lack of appropriate data for years before 1979 rule out a traditional pre/post or difference-in-differences research design. We are, however, able to analyze trends in labor market outcomes since 1979, a period during which the cost of complying with the mandate grew steadily.

The results from our permutation (placebo) tests, which treat Hawaii as a random draw from the distribution of states, suggest no discernable wage effect of the mandate: the changes in Hawaii’s distribution of wages as the costs of the mandate rose are indistinguishable from changes in the wage distribution in other states analyzed in a parallel fashion. Using the same framework, we also found no statistically detectable changes in relative employment probabilities in Hawaii.

However, when we analyzed trends in the percentage of individuals working less than 20 hours per week, the threshold distinguishing covered and exempt workers under Hawaii’s mandate, we found evidence of a shift towards such workers in Hawaii. Our estimate of this effect lies near the tail of the distribution of the placebo estimates we generated for all states. One interpretation of this finding is that employers have responded to the growing cost burden of Hawaii’s mandate by substituting exempt for covered employees. However, given that the estimated size of this shift is quite small, this interpretation of the finding also is consistent with the view that the mandate imposed relatively small costs.
REFERENCES


Figure 1.
Single Coverage Health Insurance Premiums in Hawaii on a per-hour basis ($2006)
Figure 2: ESI Coverage, 1979-2005.

Source: Author’s tabulations (weighted) from March CFS files, 1980-2006.
Figure 3A: ESI Coverage, Workers with less than a College Degree

Figure 3B: ESI Coverage, College Educated Workers

Source: Author’s tabulations (weighted) from March CPS files, 1980-2006.
Figure 4. Summary of the Placebo Tests for Own-Name Health Insurance Coverage

The Distribution of State Effects: Own Name ESI Coverage
Full Sample, Conditional on Demographics

The Distribution of State Effects: Own Name ESI Coverage
HS Degree or Less, Conditional on Demographics

The Distribution of State Effects: Own Name ESI Coverage
College-Educated, Conditional on Demographics
Figure 5. Summary of the Placebo Tests for Log Wages

The Distribution of State Dummies: Log Wages
Full Sample, Conditional on Demographics

The Distribution of State x Time Effects: Log Wages
Full Sample, Conditional on Demographics
Figure 6. Percent Working Less than 20 Hours per Week
Private Sector Workers

Year of CPS survey

Hawaii
Rest of the U.S.
Figure 7. Unadjusted Unemployment Rates
Table 1. The Gap in Own-Name ESI Coverage: Hawaii vs. Other States, 2002-2005

<table>
<thead>
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<th></th>
<th>Full Sample</th>
<th>HS Degree or Less</th>
<th>College-Educated</th>
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<tr>
<td>Hawaii less all others</td>
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<td>97.5\text{th} percentile</td>
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<td>97.5\text{th} percentile</td>
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<td><strong>Adjusted for Demographics and Job Characteristics</strong></td>
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<td>Hawaii less all others</td>
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<td>97.5\text{th} percentile</td>
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Notes: Estimated using March CPS data. The estimation sample consists of private-sector workers age 18-64. The demographic controls are education (5 categories; 2 categories in the education sub-samples), a quartic in age, gender, gender by age quartic interactions, married, married by gender interaction, race (4 categories), residence in an urban area, and veteran status. The job characteristics are industry (12 categories), occupation (13 categories), and firm size (5 categories).
Table 2. Health Insurance Coverage in Hawaii and the rest of the US: 2002-2005

<table>
<thead>
<tr>
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<th>All Private-Sector Workers</th>
<th>HS Degree or Less</th>
<th>College- Educated</th>
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<tr>
<td></td>
<td>US</td>
<td>Hawaii</td>
<td>US</td>
</tr>
<tr>
<td>ESI—Own Name</td>
<td>54.6</td>
<td>67.8</td>
<td>46.7</td>
</tr>
<tr>
<td>ESI—Dependent</td>
<td>12.3</td>
<td>10.1</td>
<td>10.8</td>
</tr>
<tr>
<td>Total ESI</td>
<td>67.0</td>
<td>77.8</td>
<td>57.5</td>
</tr>
<tr>
<td>Private Non-group</td>
<td>3.7</td>
<td>2.3</td>
<td>3.0</td>
</tr>
<tr>
<td>CHAMPUS</td>
<td>0.9</td>
<td>1.8</td>
<td>0.9</td>
</tr>
<tr>
<td>Medicaid</td>
<td>3.4</td>
<td>2.6</td>
<td>5.2</td>
</tr>
<tr>
<td>Medicare</td>
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<td>0.1</td>
<td>0.3</td>
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<tr>
<td>Total Public</td>
<td>4.5</td>
<td>4.5</td>
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<tr>
<td>Uninsured</td>
<td>24.9</td>
<td>15.3</td>
<td>33.1</td>
</tr>
</tbody>
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Notes: Estimated using March CPS data (workers age 18-64).
Table 3. Log Wage Regression Results

<table>
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<th>Full Sample</th>
<th>H.S. Degree or Less</th>
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<td><strong>State Dummy</strong></td>
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<td>97.5 Percentile</td>
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<tr>
<td>97.5 Percentile</td>
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Notes: Estimated using CPS MORG data. The estimation sample consists of private-sector workers age 18-64. The demographic controls are education (5 categories; 2 categories in the education sub-samples), a quartic in age, gender, gender by age quartic interactions, married, married by gender interaction, race (4 categories), residence in an urban area, and veteran status.
<table>
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Notes: Estimated using CPS MORG data. The estimation sample consists of private-sector workers age 18-64. The demographic controls are education (5 categories; 2 categories in the education sub-samples), a quartic in age, gender, gender by age quartic interactions, married, married by gender interaction, race (4 categories), residence in an urban area, and veteran status.
Table 5. Employment Regression Results

<table>
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Notes: Estimated using CPS MORG data. The estimation sample consists of all individuals age 18-64. The demographic controls are education (5 categories; 2 categories in the education sub-samples), a quartic in age, gender, the gender by age quartic interactions, married, married by gender interaction, race (4 categories), residence in an urban area, and veteran status.