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IZA DP No. 4152

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April 2009

Forschungsinstitut
zur Zukunft der Arbeit
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IZA Discussion Paper No. 4152
April 2009

ABSTRACT

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

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 in the United States, Hawaii's Prepaid Health Care Act, using a standard supply-demand framework and Current Population Survey data covering the years 1979 to 2005. During this period, the coverage gap between Hawaii and other states increased, as did real health insurance costs, implying a rising burden of the mandate on Hawaii's employers. We use a variant of the traditional permutation (placebo) test 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 workers who tend to have low rates of coverage in the voluntary market (primarily those with lower skills). We also find that relative wages fell in Hawaii over time, but the estimates are statistically insignificant. By contrast, a parallel analysis of workers employed fewer than 20 hours per week indicates that the law significantly 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.

JEL Classification: J32, I18, J23

Keywords: health insurance, employment, hours, wages

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* The authors thank Meryl Motika, Jaclyn Hodges, Monica Deza, Abigail Urtz, and Aisling Cleary for excellent research assistance, Jennifer Diesman of HMSA for providing data, and Gary Hamada, Tom Paul, and Jerry Russo for providing essential background on the PHCA. Thanks also to Nate Anderson, Julia Lane, Reagan Baughman, and seminar participants at Michigan State University, the University of Hawaii, Cornell University, and the University of Illinois (Chicago and Urbana-Champaign) for helpful comments and suggestions on earlier versions of the manuscript. The views expressed in this paper are those of the authors and should not be attributed to the Federal Reserve Bank of San Francisco or the Federal Reserve System.

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

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 reversed 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 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 2004; Baicker and Levy 2007; Burkhauser and Simon 2007), or to other market changes that affect health insurance premiums, such as increases in medical

[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.

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 direct 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 on a regular basis. Despite the significance of this law and its relevance to on-going policy debates, research on its effects is quite limited. A key reason for the paucity of research is timing: the PHCA legislation was passed in 1974, five years before any national survey provided information on health insurance coverage for individuals, making it difficult to compare outcomes before and after the passage of the law.

While the early passage of the PHCA creates significant research hurdles, it does not preclude a meaningful analysis of the law's impact, for several reasons. First, because of legal challenges the validity of the law was in question until 1983. As we show, in the years just prior to 1983 the percentage of workers receiving ESI coverage was not substantially larger in Hawaii than in the rest of the United States. However, the conditional and unconditional coverage gaps between Hawaii and other states widened over time, particularly for worker groups with low rates of ESI coverage, such as young and lower skilled workers. This growing divergence in coverage, combined with substantial growth in the relative price of health care over the same period, implies that the cost of complying with the mandate has grown over time. As a result, data pre-dating the original PHCA legislation is not necessary for testing the hypothesis that the mandate has had labor market effects.

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 for the same data years, to compare trends in health insurance coverage and three labor market outcomes—wages, hours and employment—in Hawaii and the rest of the United States. 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. 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 policy effects we stratify our analysis by the probability of receiving ESI (estimated using data for states other than Hawaii).

Given the likely influence of unobserved effects at the state level, we take a conservative approach to statistical inference by relying on a variant of Fisher's permutation test (Fisher 1936; see also Johnston and DiNardo 1997, Bertrand, Duflo, and Mullanainthan 2002). Our tests entail comparisons of the usual difference in means between the United States and Hawaii to parallel "placebo" comparisons between each of the other 50 states (plus DC) and the remainder of the United States, 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. The estimated coverage gap is largest for worker groups with low coverage rates and is essentially zero for workers with high coverage rates, which supports the interpretation that Hawaii's higher coverage rates are attributable to the mandate. Using the same framework, we also find that Hawaii's distribution of wages and employment has not diverged from that in other states, although we uncover a precisely estimated and rising tendency for Hawaiian employers to reduce the burden of the mandate through increased reliance on part-time positions that are exempt from the law. Like the coverage results, the effects on part-time worker are concentrated on workers with low probabilities of receiving health benefits in the absence of an employer mandate.

II. BACKGROUND AND PREVIOUS LITERATURE

II.A. Hawaii's Health Insurance Mandate

Hawaii's employer mandate legislation (PHCA; Hawaii Revised Statutes, chapter 393) was passed in 1974, the same year that the U.S. Congress passed the Employee Retirement Income Security Act (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 that employers provide health insurance benefits (Mariner 1992). Shortly after it took effect in January of 1975, the PHCA was challenged on ERISA 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 required by Hawaii's law. In 1977, the U.S. District Court of Northern California ruled in the company's favor. This decision was upheld by the U.S. Court of Appeals in 1980 and by the U.S. Supreme Court in 1981 (The Hawaii Uninsured Project 2004). In 1983, the U.S. Congress granted a permanent 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 PHCA requires private-sector employers in Hawaii to provide health insurance coverage containing a minimum level of benefits to employees working 20 or more hours per week.[2] 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.[3] Employers must finance at least 50% of the premium cost, and the employee contribution is limited to an amount that is 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 can be required to pay for any health care costs incurred by their employees during the period of noncompliance.

Available evidence suggests that the legal uncertainty surrounding PHCA's status sharply limited the law's impact between its 1975 enactment and the U.S. Congressional intervention in 1983. For example, the state Department of Labor and Industrial Relations suspended employer compliance audits between the first court ruling in 1977 and exhaustion of the state's judicial appeals in 1981 (Agsalud 1982, p. 14). Moreover, 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).[4] Similarly, figures reported by the Health Insurance Association of America (HIAA) in their annual *Source Book of Health Insurance Data* do not indicate a significant increase in ESI coverage in Hawaii

[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), as determined 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 PHCA 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.

around 1975.[5] Finally, as we show in Section IV, in 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 the sustained legality of the PHCA, the gap widened.

II.B. 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. In his analysis, 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 employers' cost of provision.

In the absence of a pre-existing market failure, we would expect workers whose valuation strictly exceeds the cost to receive the fringe benefit in a voluntary market. Therefore, 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, either because their willingness to pay for the benefit is less than its cost or because their wage is close to the minimum wage. In such cases, the wage reduction arising from the market adjustment to the mandate will not be large enough to fully offset the cost of the benefit. If certain types of employees are exempt from the mandate, employers will face incentives to substitute exempt for covered workers. If this type of adjustment is not sufficient to offset the remaining costs of the mandate, relative labor costs rise for covered workers and employers may act to reduce

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[5] In fact, according to the HIAA publications, the number of Hawaiians with health insurance actually declined between 1973 and 1975. Dick (1994) reports data from the two largest health insurers in Hawaii that also shows little change in coverage around the time the PHCA took effect.

employment of these workers.

Various empirical studies have relied on the supply and demand framework for investigating the effects 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 found 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 married women of child-bearing age, who are most likely to be affected by the mandate. Gruber's findings suggested further that because the wage reductions were comparable to employers' costs of providing the mandated benefit, the laws had no employment effects. Kaestner (1996) examined the effect of state-mandated workers' compensation and unemployment insurance on labor market outcomes. Although his findings for teenagers and young adults were consistent with a basic demand and supply model in the presence of wage constraints, his parallel findings for older workers were inconsistent with the basic model.

II.C. Past Research on Hawaii's Mandate

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 affect 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 receiving health insurance from their employers. The effect of the mandate on overall insurance coverage may be further reduced if some workers who gain 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 have used cross-sectional data to compare 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 used CPS data from later years, found that ESI coverage is significantly higher in Hawaii and attributed this result to the PHCA (Thurston 1997, Lee et al. 2005, Kronick et al. 2004). 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 studies analyzed workers and focused mainly on ESI. If coverage mandated by the PHCA simply offsets coverage from other sources, the law could increase own-name ESI coverage but have little effect on total coverage. Alternatively, it is possible that the passage of PHCA did little to increase insurance coverage immediately after the law's legal status was resolved (1983), 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 1993; U.S. GAO 1994).

To the extent that the PHCA prevented employers in Hawaii from dropping coverage in response to rising premiums, as was done in other states over the past few decades, 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 data from the 1970 and 1990 Censuses aggregated to the industry level and found mixed evidence for wage reductions due to the expansion of ESI. His results are sensitive to assumptions regarding counterfactual wage trends.

Thurston also tested for an effect on part-time work, again using aggregated industry data from the 1970 and 1990 Censuses. 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 United States. One limitation of his analysis, however, is that the measure of "low hours" work available in his data did not closely correspond to the 20 hour cut-off that distinguishes between covered and exempt workers in Hawaii. Lee et al. (2005) compared the distribution of hours worked in Hawaii and other states. They found that the percent of individuals working less than 20 hours per week was nearly the same in Hawaii as in their matched sample of residents of other states. However, the percent of workers reporting exactly 20 hours was higher in Hawaii. Given that self-reported hours data are subject to heaping, they interpreted this finding 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, neither these studies nor the prior studies on Hawaii's PHCA tested whether the law had such a disemployment effect.[6] Moreover, none of the existing studies of Hawaii's PHCA tested for the expected differential impacts of the mandate across worker groups or investigated the rising costs of compliance over time.

III. DATA AND DESCRIPTIVE EVIDENCE

III. A. Sample Construction

Our main analyses rely on data from the Current Population Survey (CPS) March and Monthly Outgoing Rotation Group (MORG) files. The March CPS 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; we use these data to analyze the patterns and determinants of insurance coverage. For the analysis of labor market outcomes, we use data from the MORG files, which include the quarter sample of the monthly CPS to whom questions are posed regarding earnings and hours in their main job at the time of the survey. Compared with the March data, the MORG files provide larger sample sizes and 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, unless otherwise indicated we focus on workers age 18-64 who are employed in the private sector (excluding the self employed); we exclude government employees because Hawaii's PHCA law does not apply to them and their inclusion could bias the cross-state comparisons. In our analyses of wages and hours, we exclude observations with imputed values

[6] A 1994 study by the U.S. GAO (1994) examined many of the same issues that we have identified regarding the impact of PHCA but relied largely on qualitative analyses.

of those variables. Additional details regarding the characteristics of our data files, variable definitions, and treatment of imputed and top-coded data are provided in Appendix A.

Table 1 serves the dual purpose of illustrating some basic data characteristics and comparing mean values of population and job variables between Hawaii and the rest of the United States. The comparisons are made at the start and end periods of our sample frame using the March CPS files (survey years 1980-83 and 2003-06), with some variables or breakdowns not available in all years (we use the March rather than the MORG files for these tabulations because the latter do not include information on firm size). In regard to demographics, Hawaii's population is somewhat more educated on average than the population in other states, with a smaller share of individuals lacking a high school degree and a higher share with at least some college experience. Hawaii is particularly notable in regard to a very high share of individuals whose race/ethnicity is Asian/Pacific Islander (and hence are included in the "Other" category; the more detailed breakdown is not available across our complete sample frame).

In regard to job characteristics, Table 1 shows that manufacturing accounts for a smaller share of employment in Hawaii than the rest of the United States, though this difference has declined over time. The gap in manufacturing is largely offset by a higher percentage of Hawaiian employment in personal services (which includes hotels and tourism). Apart from manufacturing and personal services, the distribution of employment by industry is very similar in Hawaii and the rest of the nation. Hawaii has a slightly higher proportion of individuals employed in small firms and lower proportion in large firms than does the rest of the nation (using a cutoff of 100 employees to distinguish between small and large firms).

III.B. The Costs of the PHCA

In considering the possible labor market effects of the PHCA, it is important to recognize that the burden of the mandate has increased over time as the growth in health insurance premiums has outstripped the growth in wages. Figure 1 provides a sense of how the cost of the PHCA to employers has grown. The graph plots real ($2006) single coverage health insurance premiums in Hawaii expressed as an hourly cost 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 monthly single coverage premium for a plan meeting the standards of the PHCA was $15.96, which for an employee working 40 hours per week (assuming 4.3 weeks per month) translated to a nominal cost of 9 cents per hour, or 38 cents in 2006 dollars. The cost per hour was only slightly higher in 1979, the first year in our data set, 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 (relative to the prices of a complete bundle of goods).[7]

Another way to gauge the cost of the mandate is to compare these figures with 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

[7] An analysis by the U.S. General Accounting Office (1994) concluded that from 1975 to 1993 health care spending grew by similar rates in Hawaii and the rest of the United States. Data from the Medical Expenditure Panel Survey-Insurance Component (MEPS-IC) show that since 1996 the trend in ESI premiums in Hawaii has been similar to other states.

insurance. By 2006, however, the mandate increased the cost of employing workers at Hawaii's minimum wage (of $6.75) by 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.

III.C. Trends in ESI Coverage

We begin our analyses by comparing trends in health 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. At the beginning of the sample frame the coverage rates in Hawaii and the rest of the country are very similar. The plots begin to diverge around 1983, the year Congress granted Hawaii's ERISA exemption. In the early 1980s, coverage fell slightly in Hawaii but far 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, which results from sampling error in Hawaii's relatively small sample. By 2005, 70% of private sector workers in Hawaii received health insurance through their employer, compared with only 57% in the rest of the US.[8]

III.D. Accounting For Heterogeneous Policy Effects

As noted earlier, the basic supply and demand theory 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 the absence of a mandate, such as less skilled workers, and should have

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[8] It may seem surprising that the coverage rate in Hawaii is well below 100%, given that there are few exemptions to the PHCA. We investigated this issue by conducting supplemental tabulations using establishment level data on insurance offers from the Medical Expenditure Panel Survey-Insurance Component (MEPS-IC) and worker level data on eligibility and take-up from the CPS Contingent Worker Supplements; these data indicate that most of this gap is accounted for by workers who decline coverage (mainly because they are covered through a family member) or are not eligible (because they work too few hours or have not been with the firm long enough).

little effect on workers who have a high probability of receiving ESI in a voluntary market. A simple way to test this prediction is to stratify the analysis by education. Figure 3A presents trends for workers with a high school degree or less and Figure 3B reports trends for those with at least a college degree. The results are consistent with our expectations. The general pattern for less educated workers is similar to the full sample results. Whereas in 1979-1980 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. For college educated workers, ESI coverage rates in Hawaii and the rest of the United States are quite similar on average across the entire sample frame, suggesting little impact of Hawaii's mandate.

Stratification by education is a highly imprecise means of accounting for heterogeneous policy effects across worker groups; for example, some workers with a high school degree are older and employed in industries with high rates of ESI coverage, which may imply higher ESI coverage rates for such workers than for younger workers with higher educational attainment. We can account for heterogeneous policy effects more precisely by using a full range of explanatory variables to categorize individuals according to their probability of receiving health benefits in a voluntary market. To do this we fit the following regression on the complete sample of observations excluding those in Hawaii, using a linear probability model (LPM):

$$I_i = X_i\Gamma + \varepsilon_i \qquad (1)$$

where I is an indicator variable for own name ESI coverage, X is a vector of individual and job characteristics, Γ is a set of coefficients to be estimated, and ε_i is an error term. From this equation, we obtain the fitted probability of ESI coverage for each individual in the full sample (including Hawaii) and then sort the fitted probabilities and place individuals into quantiles of this distribution. While education plays an important role in this prediction equation, other variables are important predictors as well. The control variables used are essentially the full list from Table 1, excluding nativity and firm size, which are not available over our complete sample frame (see Appendix B, which presents the results from ESI probability equations that are discussed in the next section).[9]

Figure 4 illustrates how actual ESI coverage differs between Hawaii and the remainder of the United States across quantiles of the distribution of predicted ESI; the prediction equation is estimated using data for March CPS reference years 2002-05. We divided the data into 25 quantiles of this distribution and then calculated actual ESI coverage rates in those quantiles for Hawaii and the rest of the country. Actual ESI coverage rates in Hawaii are much higher than those in the rest of the country at the lowest quantiles of the distribution of fitted probabilities, but the rates converge as we move to higher quantiles, becoming quite similar in the top quintile. Like Figures 3A and 3B, this graph suggests that the PHCA has substantially raised coverage rates for workers who would be least likely to receive health benefits in a voluntary market but has had little effect on workers with high coverage rates in voluntary markets.

Figure 5 displays a breakdown of coverage over time, plotted separately for Hawaii and the rest of the United States, for the complete data and for each of the five quintiles obtained from the distribution of fitted ESI; the underlying equation used to form the quintiles is estimated separately for each data year and the fitted probabilities are obtained using these separate annual

[9] This approach relates closely to that of Card (1996), who formed a skill index using predicted wages from an equation estimated using nonunion workers. Card used the resulting index, which is unaffected by union wage effects, to divide union and nonunion workers into skill quintiles, for purposes of examining how the union wage effect varies by skill level. In the statistics literature, such an approach to characterizing “treatment effect heterogeneity” has been called a “prognostic score” (Hansen 2008).

estimates. The coverage gap between Hawaii and the rest of the United States is very large in the low quintiles, but it shrinks as we move to higher quintiles. Except for greater annual variation in the Hawaii line (due to its relatively small underlying sample), the two lines are nearly identical in the plot for the highest (fifth) quintile (even more similar than in the corresponding plot for college-educated workers in Figure 3B). We rely on this breakdown by predicted ESI quintiles for subsequent analyses.[10]

IV. INSURANCE COVERAGE EFFECTS

IV. A. Adjusted Differences in ESI Coverage (Permutation Tests)

We now turn to conditional analyses of the impact of Hawaii's mandate on ESI coverage. A standard approach to adjusting for differences between Hawaii and other states in regard to worker and firm characteristics is to estimate a regression equation of the form:

$$I_i = X_{is}\beta + \delta H_i + e_s + u_i \qquad (2)$$

The dependent variable, *I*, and the covariates, *X*, are defined as in (1); *H* is an indicator variable for observations from Hawaii; e_s is a vector of state-specific effects and u_i is an individual disturbance.

While estimation is straightforward in this setting, statistical inference is not, and researchers have adopted a number of approaches (see Wooldridge 2003 and 2006 for useful discussions). As Moulton (1990) demonstrates, assuming that the disturbances are i.i.d leads to

[10] All of the findings in this paper are similar when we rely on the education breakdown rather than the breakdown by ESI quintiles to distinguish between worker groups that are differentially affected by the mandate; we exclude the education breakdown in subsequent discussion purely for brevity.

insufficiently conservative inference. One response is to impose specific parametric restrictions on e_s, as in the classical random effects model (e.g. Searle, Casella, and McCulloch 1992). Another approach is to forego parametric assumptions about the error components but impose additional structure. For example, it is common to assume that the terms e_s and u_i in (1) are error components of the following form: u_i is assumed to be i.i.d. across all individuals in the sample, and e_s is a state random effect that is assumed to be i.i.d. across individuals within the same state, with $E(e_s \mid X, H) = 0$ and variance $\sigma_s^2 > 0$.

In this well-known "clustering" framework, inference relies on the asymptotic approximations associated with the assumption that the number of individuals within a state and/or the number of states grows large (Wooldridge 2006). This assumption clearly does not apply in our setting, due to our focus on a single state (Hawaii). Moreover, some researchers have argued that inference is extremely problematic and perhaps impossible in our setting, because the comparison of a single state with all others collapses the degrees of freedom in the model and creates much larger sampling variance (perhaps infinity) than is captured by the conventional asymptotic framework (Donald and Lang 2007). As noted by Wooldridge (2006), however, this argument is indistinguishable from the standard question of whether an observed conditional difference in measured outcomes is entirely due to the policy change of interest.

Rather than adopt the premise *a priori* that statistical inference is impossible, our approach is based on the assumption that a set of control variables *X may be* adequate to eliminate any undue influence of unobservables in Hawaii that are correlated with the status of its health insurance mandate. Further, our approach is relatively robust to the failure of the standard assumptions underlying the "clustering" approach. This approach is a variant of Fisher's permutation or randomization test (Fisher 1935; see also Bertrand, Duflo, and

Mullanainthan 2002).[11] The classical permutation test is the "two sample problem" with random assignment to either a treatment or a control group in an experimental setting. An appropriate statistic (such as the difference in means) is computed for the two samples, call it *d*. Under the null hypothesis that the treatment has no effect and the two samples come from the same population, the labels—"treatment" or "control"—are arbitrary and can be reassigned to the units. Assuming the sizes of the treatment and control groups are n and m, the test is implemented by re-computing the statistic *d* for each of the $(n+m)!/(n!m!)$ samples created by permuting the labels. The test statistic is formed by computing the percentile that *d* for the true treatment group represents in the distribution of all the "placebo" differences. If the treatment difference *d* is large relative to these placebo differences—i.e., if it lies beyond a critical value near the tail of the placebo distribution, such as the 97.5 percentile—the treatment difference is deemed "significant." In such a setting, the inference is "exact" and does not depend on asymptotic approximations.

Fisher's intent in deploying this test was merely to demonstrate the utility of conventional t-statistics even when the underlying data was not normal; when applied in unmodified form to our setting, his approach produces inference quite similar to conventional t-tests without adjustment for clustering. By contrast, our modified approach begins with the estimation of 50 "placebo" versions of equation (1), in each case with a different state (including the District of Columbia) serving as a stand-in for Hawaii. Given the large samples we are employing and

[11] For a simple application to an economic question see Johnston and DiNardo (1997, chapter 11.2). Lehman (1959, chapter 5.7) provides a formal discussion of the statistical assumptions underlying the test. For some colorful *ad hominem* about those who might use such "non parametric" permutation tests, see Fisher (1935, chapter 3). Bertrand, Duflo, and Mullanainthan (2002) conducted Monte Carlo tests of this approach and found that among a set of solutions to common problems with difference-in-difference estimation, randomization inference dominates with respect to coverage and robustness in small samples. They did not apply this approach to a specific empirical example, and this material did not appear in the published version of their manuscript (Bertrand, Duflo, and Mullanainthan 2004).

under the assumption that the covariate adjustment is adequate, the estimates calculated using the 50 permutations of the "Hawaii" label should be very close to zero, providing an informal specification check. We use these 50 placebo estimates to construct regions of acceptance and rejection for the null hypothesis that the Hawaii ("treatment") effect is zero. Intuitively, if this coefficient is similar to the corresponding placebo estimates for other states, our inference is that the Hawaii effect is not "real," in the sense that the effect of its health insurance mandate cannot be distinguished from the effect of unobservables in other states, even if it were to pass a test of significance using more conventional procedures for statistical inference.

It is important to note that when equation (2) is the correct specification and the error terms e_s follow, for example, a normal distribution, our method generates inferences that are similar to either the classical random effects model or the use of clustered standard errors under conventional asymptotic assumptions.[12] When the standard assumptions are violated, however, our approach produces considerably more appropriate and conservative inferences. Indeed, following the usual expedient of estimating standard errors that are clustered for observations within the same state, 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. This evidence suggests that the standard assumptions do not hold in our setting; we discuss these comparisons in more detail below.

In Table 2 we list estimates of δ from regressions that compare Hawaii to the other 50 states. We first present unadjusted differences and then use an LPM to adjust for covariates.[13]

[12] We confirmed the correspondence between our approach and conventional approaches under normality using Monte Carlo simulations, the results of which are available on request.

[13] The LPM produces coefficient magnitudes that are easily interpreted as probability effects, and it produces reliable estimates in general for dependent variables whose means are well-bounded away from

To span our sample frame and ensure a reasonable number of observations from Hawaii, we pool data for the four years at the start and end of our sample period (reference years 1979-82 and 2002-05) and report the results for each 4-year sample in separate panels. For the covariate vector (corresponding to *X* in equation 2), the demographic characteristics are education (5 categories), a quartic in age, gender, gender by age quartic interactions, married, married by gender interaction, race/ethnicity (4 categories), residence in an urban area, veteran status, and year dummies (3), plus nativity (3 categories) in 2002-05. The job characteristics are industry (13) and occupation (11) categories, plus firm size (5 categories) in 2003-06. We report results for the full sample and for five quintiles formed using the predicted probability of own name ESI coverage. For each time period, we list the estimated treatment coefficients for Hawaii for various specifications. Below them, instead of reporting conventional clustered standard errors, we display the 2.5 and 97.5 percentiles of the placebo estimates; in our framework, these values represent the lower and upper critical values for rejecting the null hypothesis that the Hawaii effect is zero, with the significance level of the test set to 5 percent.[14] For comparison purposes and also completeness, Appendix B reports results for the complete set of coefficients for the full-sample regression that includes a Hawaii dummy and the complete set of demographic and firm controls, with conventional clustered standard errors listed.

The results presented in Panel A of Table 2 provide mixed evidence on whether own name coverage was more prevalent in Hawaii during the early period (1979-1982) when the legal status of Hawaii's mandate was in doubt. When we do not adjust for covariates, the Hawaii

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0 and 1, like our ESI variable. We obtain essentially identical results in terms of magnitudes and statistical significance by estimating a logit model.

[14] This method for statistical inference, which follows the suggestion of Bertrand, Duflo, and Mullanainthan (2002), also relates closely to the approach proposed and implemented in Abadie, Diamond, and Hainmueller (2007); like us, they test for the impact of a particular policy on outcomes in one state (California's smoking ban, implemented in 1988).

coefficient for the full sample is 0.053. This estimate lies within the range formed by the upper and lower critical values of the placebo estimates. However, the estimated coefficient is larger when we condition on worker demographics (0.101) and demographics plus job characteristics (0.145); these estimates are slightly larger than the values of the corresponding 97.5 percentiles, which are 0.079 and 0.142, respectively.[15] Table 2 also displays results of the permutation tests of coverage differences using the quintile breakdown. The Hawaii coefficient is larger for workers in the lower ESI quintiles, which is consistent with our expectations. However, for the first and second quintiles the regression-adjusted gap between Hawaii and other states lies within the range of the placebo estimates, indicating that we cannot reject the null that Hawaii is no different than other states in regard to ESI coverage for those worker groups.

By contrast, the results for the end of our sample frame (2002-2005) in Panel B provide strong evidence that the PHCA has increased ESI coverage in Hawaii relative to other states. In the full sample and the first two quintiles, the Hawaii coefficients are larger than the corresponding estimates for the earlier period. The magnitude of the Hawaii estimate and its position relative to the range of placebo estimates declines monotonically across ESI quintiles. For the first four quintiles the estimate lies outside the range formed by the upper and lower critical values of the placebo estimates, but it is well-bounded by the placebo distribution in the fifth quintile.

Figure 6 provides a graphic illustration of the results in Table 2 by showing Hawaii's position in the full distributions of placebo effects for the model with complete controls, for the

[15] The inadequacy of conventional clustered errors is quite apparent in this setting, given that the absolute t-statistics exceed two for most of the placebo coefficients. We therefore obtain much wider test intervals than are produced by conventional approaches. For example, in the regression with worker and job controls for the 1979-82 sample, the range of values for the Hawaii coefficient for which we would fail to reject the null hypothesis is [-0.057, 0.142]. Based on the regression with clustered standard errors from Appendix B, column 1, the corresponding range is much narrower: [-0.029, 0.029]

five quintile groups. The two dashed lines indicate the 2.5 and 97.5 percentiles of the distribution of placebo estimates for states other than Hawaii, while the thick solid line shows the coefficient for Hawaii. For the 2002-2005 results in the bottom half of the figure, the Hawaii effect is well outside the range of placebo estimates for the first two quintiles, but the gap shrinks at higher quintiles and disappears in the fifth quintile.

While we cannot use this research design to determine precisely how much of Hawaii's higher rate of own-name ESI coverage should be attributed to the state's employer mandate, these results are 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. The results also suggest that the gap between Hawaii and other states rose over time, especially for workers with low rates of ESI coverage in the voluntary market.

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

The results discussed thus far 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. This goal may not be achieved if the policy leads to shifts in the source of coverage—for, example from a spouse's employer or an individually purchased policy. 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 3 reports the distribution of coverage in Hawaii and the rest of the United States, using the March CPS data for reference years 2002 to 2005.

Hawaiian workers are slightly less likely than workers in other states to have ESI coverage through a dependent or to purchase non-group insurance. As a result, the 13 percentage 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 find 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 United States. 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 (or 40%) lower in Hawaii than in the rest of the United States.

In contrast, 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.[16] The percentage of college educated workers with other types of coverage is also similar in Hawaii and other states, causing the percent uninsured to be quite similar as well. On net, the tabulations listed in Table 3 suggest that the intended effect of Hawaii's mandate, to increase health insurance coverage, is not substantially undermined by substitution of ESI for other sources of insurance.

V. LABOR MARKET OUTCOMES

The coverage results strongly suggest that the PHCA was a binding constraint for a significant number of Hawaiian employers, especially those employing workers with low skill who have low probabilities of receiving ESI in a voluntary market. While the Hawaii-United States gap in coverage has increased only slightly since the early 1990s, the costs associated with the mandate have grown consistently over our sample frame 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. If 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 observe an effect on hours if employers make greater use of part-time workers who are exempt from the mandate. Alternatively, firms may hire fewer workers from the available labor force, thereby reducing employment probabilities for workers subject to the mandate.

[16] 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.

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. However, such a specification is inappropriate for analyzing a policy like the PHCA, for which 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 estimate models of the form:

$$Y_{is} = X_{is}\lambda_1 + Z_s\lambda_2 + \gamma H_i + \rho T + \theta(H_i \cdot T) + \varphi_s + \eta_i . \qquad (3)$$

In this equation, Y is one of the labor market outcome variables (wages, low-hours work status, or employment status), X is the vector of individual and job controls as in equation (1), and Z_s is a vector of two variables measured at the state level: annual values of the real value of the state minimum wage and the log change in state Gross Domestic Product (GDP).[17] H is a Hawaii indicator, T is a linear time trend measured at an annual frequency, φ_s and η_i are an unobserved state effect and an i.i.d. disturbance specified as in (2), and the remaining terms are coefficients to be estimated. The parameter of interest is θ, the coefficient on the interaction between the Hawaii dummy and the time trend, 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 θ to be negative in wage and

[17] We include the state-level variables to account for broad economic factors that affect labor demand and the wage structure at the state level. Controlling for the minimum wage is important because its level is relatively high in Hawaii (about 25 cents higher than the average for other states over our sample frame). Controlling for state GDP is potentially important because Hawaii's tourism-dependent economy is subject to large swings that can be independent of national economic conditions (e.g., a prolonged slump in the state caused by a sharp drop in Japanese tourism, which lasted for much of the 1990s).

employment regressions and positive for low-hours regressions in which the dependent variable equals one for individuals working fewer than 20 hours per week.[18] For each of these outcome variables, we rely on the MORG data for estimation of the regression equations.

As discussed in the previous section regarding the effects of Hawaii's mandate on ESI coverage rates, the results of a simple regression may be misleading if there are other factors that vary across states and affect trends in wages, hours, or employment. For this reason, as before we assess the statistical significance of the estimated Hawaii effect by comparing it to the distribution of estimates obtained by running each regression 50 additional times and replacing the Hawaii indicator with an indicator for each of the other states (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 5% critical values 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 and the focus on changes over time, 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. 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.

[18] 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.

V.A. Wages

Before turning to our empirical estimates of relative wage changes, it is useful to consider what constitutes a plausible wage effect of Hawaii's mandate. We assessed this by conducting a straightforward simulation, in which we calculated the H_i*T (Hawaii*time) interaction coefficient that would be obtained if the growing costs of the mandate for employers were fully offset by an equivalent reduction in wages. We set the cost of the mandate equal to the ESI coverage differential between Hawaii and the rest of the United States multiplied by the full cost of ESI coverage in Hawaii (as displayed in Figure 1). Because our earlier estimates indicated that the ESI coverage differential is small to non-existent for higher skilled workers, we conducted the simulation only for workers in the first two ESI quintiles, using the change in coverage between the first four years and final four years of our sample frame averaged across those quintiles (see Figure 5). The calculations were performed under the assumption that employers pay 85% of the full direct costs of workers' health insurance in Hawaii and 80% of the direct costs in other states.[19] The implied decline in relative wages over time is equal to less than one-tenth of one percent annually: -0.0007, to be more precise.[20] This is the coefficient that we would expect to estimate for the Hawaii*time interaction in equation (3) for log wages if employers were able to shift the full cost of the mandate to workers through wage reductions.

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[19] We based these assumptions on tabulations of employee shares and contributions for single-coverage and family health insurance plans, using the MEPS-IC data for private-sector establishments in Hawaii and the United States as a whole (http://www.meps.ahrq.gov/mepsweb/data_stats/MEPSnetIC.jsp.).

[20] Underlying this finding is a change in the average per-worker cost of the mandate from about 8 cents per hour in 1979 (real premium costs of $0.38 per hour multiplied by a conditional coverage gap of 21 percentage points) to about 43 cents per hour in 2005 (real premium costs of about $1.51 per hour times a conditional coverage gap of about 28 points), relative to real average wages for this group of about $13.12 in Hawaii in 2005 (i.e., in 2005 the mandate cost equaled about 3.3% of hourly wages for individuals in the first two ESI quintiles).

Results for the empirical analysis of wage trends are displayed in Table 4 and Figure 7. The table lists in separate panels the results for the full sample and by quintiles of the distribution of predicted ESI probabilities. Results are shown for two parameters: the coefficient on the state dummy and the interaction between that variable and the linear time trend (see Appendix B for the complete set of regression results for the full-sample specification that includes the Hawaii dummy). In the table, we report the 2.5 and 97.5 percentiles of the empirical distribution of parameter estimates from our placebo regressions. As before, these values represent the lower and upper critical values for rejecting the null hypothesis that the Hawaii effect is zero, with the significance level of the test set to 5 percent. Figure 7 displays histograms for the full distribution of the placebo estimates for the ESI quintile breakdowns.

For each sample, the coefficient on the Hawaii dummy alone is near the midpoint of the range of estimates 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 intended to capture the annual impact of the state's ESI mandate on wages over our sample frame. Depending on the estimation sample, the point estimates for the Hawaii trend generally range from about 0 to -0.005. In all but one case these estimates are substantially larger in magnitude than the annual wage reductions that, according to our simple simulation, would be necessary to offset the cost of the mandate (-0.0007). In the full sample specification from Panel A, the t-statistic for the estimated Hawaii*time effect (-0.002) is -3.95 when we use clustered standard errors (see Appendix B, column 3).

The point estimates and conventional t-statistics might lead one to conclude that the PHCA caused wages to grow more slowly in Hawaii than in other states. However, the results listed in Table 4 argue against this interpretation, for several reasons. First, the table indicates that for each estimation sample, our permutation testing approach yields Hawaii*time coefficients that lie well within the region of acceptance formed by the upper and lower critical values of the distribution of placebo estimates. In other words, the placebo test fails to reject the null hypothesis that Hawaii exhibits the same pattern of changes in relative wages over time as the typical state.[21] Second, we find greater wage reductions over time in Hawaii for the more highly skilled sub-samples (i.e., college educated workers and those in the higher ESI quintiles), which is inconsistent with the predictions of the supply-demand framework. Given the law's larger impact on ESI coverage rates for unskilled workers, it is unlikely that it would cause larger wage offsets for skilled workers than for unskilled workers. These results suggest either that Hawaii's ESI mandate had no effect on relative wages or an effect that is not detectable in our data.[22]

Our inability to reject the null of no wage effect is unsurprising, given the very small magnitude of the simulated wage offset discussed at the beginning of this sub-section: the simulated coefficient of -0.0007 for the annual effect in the first and second ESI quintiles lies well within the 95% confidence interval of the Hawaii*time effect for those quintiles.[23] Based

[21] As with the ESI regressions, our placebo estimates suggest important violations of the assumptions required for the usual clustered standard errors to be appropriate, and our approach again produces more conservative confidence intervals: in the placebo regressions, the absolute t-statistic for this coefficient is greater than two in 37 cases.

[22] To focus more narrowly on a group most likely to be affected by the mandate, we performed the same analyses with a sample restricted to the retail industry only, with no qualitative difference in results.

[23] Our point here is similar to that made by Bhattacharya and Vogt (2000) regarding assessment of estimated employment effects of state health insurance mandates. Using statistical power functions, they

solely on this analysis one might conclude from our permutation testing approach that our design is not powerful enough to detect effects of Hawaii's mandate on *any* labor market outcomes using our data. However, the results discussed in the next section indicate otherwise.

V.B. Low Hours Employment

In Figure 8 we present trends in the percentage of all private sector employees who usually work fewer than 20 hours per week, which is the threshold that determines whether a worker is covered by the PHCA. The series shows a slight downward trend for states other than Hawaii, peaking at 5.7% in 1983-1984 and then falling gradually thereafter, to 3.9% in 2006. Because of smaller sample sizes, the plot for Hawaii shows more variability, although it generally lies below the plot for the other states early in the sample frame and above it later.

Table 5 and Figure 9 display results from regressions that test whether there has been a long-run trend toward low hours work in Hawaii relative to other states, once again showing results for the full sample and the breakdown by ESI quintiles.[24] Because the mean of the dependent variable is close to zero, we used a logit specification for this set of analyses. 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 general, the estimate for the Hawaii dummy falls well within the range formed by the upper and lower critical values obtained from 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-hours employment and that these other states represent a reasonable baseline period control for Hawaii.

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argued that the inability to reject the null of no employment effects in papers such as Gruber (1994) was largely noninformative.

[24] Complete results for the full-sample specification that includes the Hawaii dummy are reported in Appendix B. See Appendix A for discussion of the hours variable.

In the full sample, the estimated coefficient on the interaction term is positive (.017) and larger than the 97.5 percentile critical value from the distribution of placebo estimates; it thereby meets our stringent standard for statistical significance, indicating that over the entire period the percentage of adults in short-hour jobs grew significantly faster in Hawaii than in other states. The subsequent columns of the top panel of Table 5 and bottom panel of Figure 9 indicate that this full-sample result is driven by the sub-sample with the lowest probability of receiving ESI: the estimate is large and well above the top end of the range of placebo estimates for the first quintile but close to zero and near the middle of the placebo range for higher quintiles. This pattern is consistent with the expected effects of the mandate, which raised ESI coverage rates and hence employment costs substantially for low-skill workers but little for high-skill workers. Our findings indicate that Hawaiian employers increased their reliance on exempt part-time positions for low-skill workers in response to this pattern of growing mandate costs. [25]

In addition to being statistically precise, the magnitude of the estimated shift towards exempt part-time positions is economically meaningful. After proper transformation of the Hawaii*time coefficient in the logit equation, the results for the first ESI quintile imply an increase in low-hours employment of about 7 percentage points over the full 27–year sample period.[26] This estimate is large compared with the mean of the dependent variable, which is

[25] We obtained similar results to those reported here when we used an alternative weekly hours variable based on hours worked at all jobs in the reference week (see Appendix A and Autor, Katz, and Kearney 2008). We also obtained similar results when we corrected for heaping in reported hours by including in the low hours group workers who report exactly 20 hours per week and when we adjusted for the "coattail effect" by coding the low hours variable as 1 only for workers who do not receive ESI (Thurston 1997). Finally, the results were similar when we limited the sample to workers in retail, a sector where low-hour jobs are more prevalent.

[26] We computed the average partial effect (APE) by calculating the change in the probability of coverage due to the Hawaii*time effect separately for each observation and then averaging the resulting values across the relevant sub-sample. An advantage of focusing on the APE as the estimand is that we avoid

about 9 percentage points in Hawaii for individuals in the first ESI quintile. It is also large compared with the increase in relative ESI coverage that is associated with the mandate. Based on the results in Table 2, the conditional ESI coverage gap between Hawaii and other states for individuals in the first ESI quintile rose by about 6 percentage points over our sample frame, from about 21 to about 27 percentage points. The 7 percentage point increase in exempt, low-hours employment for this group is large relative to these ESI coverage gaps, implying a substantial demand shift towards jobs that are exempt from the mandate.

V.C. Employment

The final outcome we consider is employment, which has been the primary focus of recent critiques of employer mandates (Yelowitz 2003; Baicker and Levy 2007; Burkhauser and Simon 2007). The unadjusted data displayed in Figure 10 show that in the early 1980s the employment rate was about 3 percentage points higher in Hawaii than the rest of the United States. After 1990, the average gap was 1 percentage point in Hawaii's favor. Over the whole period, the employment rate for Hawaii is more variable but tracks quite closely with the overall rate for the other states. The Hawaii plot shows falling employment rates during the early to mid-1990s, a period when the state economy was struggling with a sharp drop in Japanese tourism due to the protracted economic downturn in that country, followed by an increase in the late 1990s as conditions in the state's tourism industry picked up again.

Table 6 and Figure 11 display results from regressions using an indicator variable for individual employment status as the dependent variable (see Appendix B for the complete results for the specification that includes the Hawaii dummy). The specification of our employment regressions is similar to the model used for wages and part-time work, including the breakdown

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the common pitfalls of interpreting interaction terms in such models (Ai and Norton 2003). See also Wooldridge (2009) for a more general discussion of this issue.

by ESI quintiles, though now the sample consists of all adults 18 to 64 years old. Because the mean of the dependent variable is well-bounded away from zero and one, for ease of interpretation we use an LPM rather than a logit model.[27]

For the full sample results in Table 6, the coefficient on the Hawaii dummy indicates that in the baseline year of 1979 the employment rate in Hawaii was 5.7 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 -0.0005 and its robust standard error (see Appendix B) is approximately 0.0002, which implies a t-statistic of -2.43. While this might suggest that Hawaii's employment rate fell significantly relative to other states as the costs of the PHCA mandate grew, the results of the placebo tests caution against this interpretation. For each sample, the estimated Hawaii*time effect falls well within the region bounded by the critical values obtained from our placebo estimates for other states; in the full sample, Hawaii ranks 30th among all states in regard to the magnitude of the linear trend in the employment rate over the period 1979 to 2005.[28] Even for the lower quintiles of the ESI distribution, the estimated Hawaii coefficient falls well within the acceptance region implied by the placebo tests. Based on this, we do not reject the null hypothesis that the employment trend in Hawaii was no different than in other states.

Our inability to reject the null hypothesis that Hawaii's law had no employment effects is not surprising, given the limited direct costs of the mandate. As described in footnote 20, in 2005 the mandate cost equaled about 3.3% of hourly wages for individuals in the first two ESI quintiles. Making the extreme assumption of no wage offset, and using labor demand elasticities

[27] We confirmed that the results are similar to those reported here when we use a logit model rather than an LPM.
[28] With clustered standard errors, the t-statistic on the Hawaii*time coefficient is -2.71, indicating significance at the 1% level (see Appendix B, column 5).

in the range of -0.1 to -0.2 (similar to Yelowitz 2004 and Baicker and Levy 2007), the implied reduction in employment for this group is about 0.35 to 0.65 percentage points over the complete sample frame. These cumulative effects over the sample frame imply annual effects on the order of 0.0001 to 0.00025. Such effects are an order of magnitude too small to detect in our data, given the influence of other unobservables that generate the cross-state differences in employment trends summarized in Table 6.

VI. DISCUSSION AND CONCLUSIONS

Using data for the years 1979-2005, we found that ESI coverage in Hawaii is much higher than in all other states, that the coverage gap has grown over time, and that it 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 significantly reduced the number of uninsured workers in the state. The pattern of higher ESI coverage in Hawaii passes the stringent standards for statistical significance imposed by our permutation testing framework, suggesting that it is a direct result of the mandate.

The basic labor supply-demand framework predicts that to the extent that the mandate raises the cost of employing certain types of workers, it will affect labor demand, causing wages and employment to fall and reliance on exempt workers (e.g., those working fewer than 20 hours per week) to rise. While Hawaii's experience does not provide an ideal natural experiment for testing for such effects, the increasing cost of ESI and hence the rising burden of the mandate enable us to test for the effects of Hawaii's law by analyzing trends in labor market outcomes since 1979. 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 statistically indistinguishable from changes in the wage distribution in other states analyzed in a parallel fashion. Relative wages fell in Hawaii over time, but no more so than in many other states, and the decline in relative wages in Hawaii was similar for worker groups that differ sharply in regard to the estimated effect of the mandate on ESI coverage. These results, in conjunction with our simulation of the wage effects of Hawaii's mandate using a simple direct estimate of the mandate costs, indicate that any wage reductions arising from the mandate are too small to be detectable in our data.

Although we cannot reliably estimate wage offsets resulting from Hawaii's mandate, it is likely that for some groups, such as low-wage workers for whom wage reductions are constrained by the minimum wage, any wage reductions will not be large enough to fully offset the mandate costs to employers. Under these circumstances, the basic supply and demand framework points to likely labor market distortions through adjustments along other margins. We uncovered strong evidence of such adjustments, in the form of a pronounced increase in the percentage of individuals working less than 20 hours per week in Hawaii, which is the threshold distinguishing covered and exempt worker under Hawaii's mandate. The estimated shift towards low-hours schedules was concentrated among the worker groups that were most affected by the mandate, which supports the interpretation that the shift reflects employers' direct responses to the mandate. Moreover, the effect we uncovered is large relative to the baseline incidence of low-hours employment and the level and change in conditional ESI coverage that is associated with the mandate. It also meets the stringent standard for statistical significance imposed by our permutation testing framework, providing strong evidence that the rising cost burden of the mandate caused a substantial labor demand shift towards exempt employees. The economic costs of such shifts depend on the difficulty of substituting between head counts and hours worked, which are likely to be small, particularly for the categories of low-skill work for which Hawaii's mandate was most binding.

Using a similar framework, we found no statistically detectable changes in relative employment probabilities in Hawaii. As with our findings for wages, our data and empirical design do not allow us to rule out employment reductions arising from the mandate, especially for workers who face low probabilities of receiving ESI in a voluntary market. However, the costs of the mandate to employers are small enough that the implied shifts are very small relative to the variation in other unobserved factors that affect cross-state employment trends.

Our results on net suggest that Hawaii's health insurance mandate succeeded in raising ESI coverage rates for worker groups with low coverage rates in an unconstrained environment, but the costs of this coverage expansion were relatively small. While Hawaii is a somewhat unique state in terms of geography, our analyses indicate that it is not an outlier in regard to labor market characteristics, suggesting that ESI mandates may cause limited labor market distortions in other states as well. On the other hand, Hawaii's employer mandate has left the state well short of universal coverage, suggesting a need for alternative approaches to expanding coverage if that is the ultimate policy goal.

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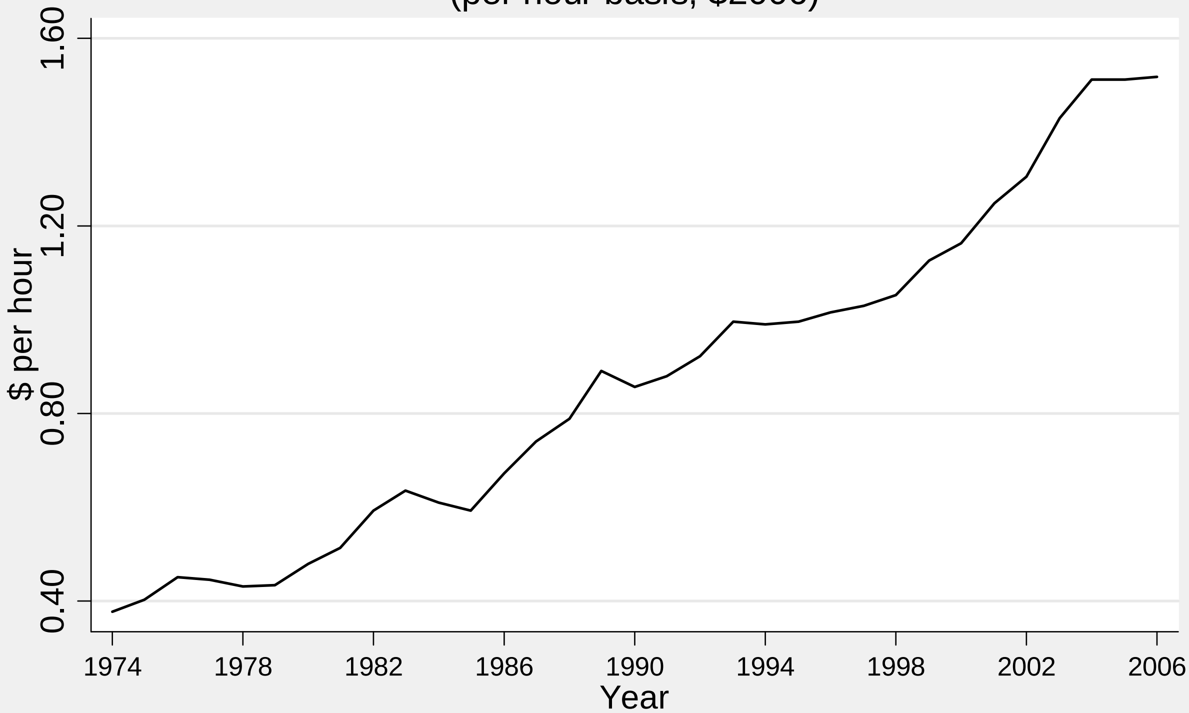
Figure 1: ESI Premiums in Hawaii

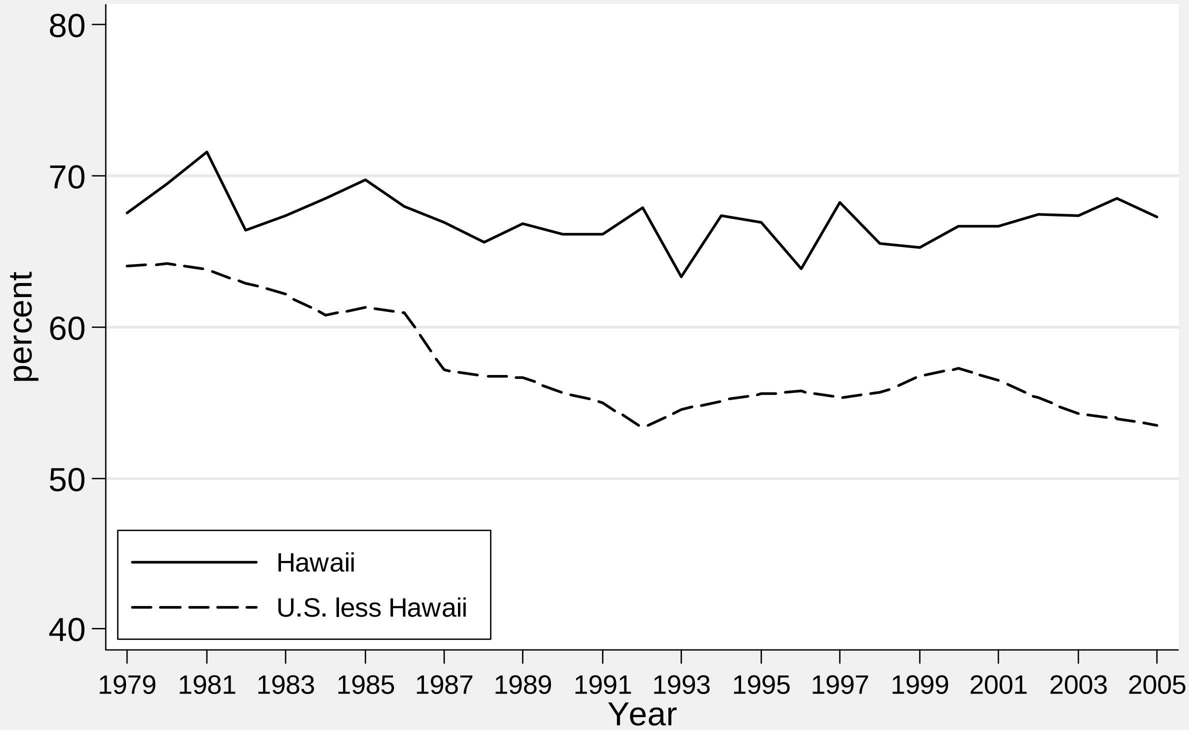
(per hour basis, $2006)



Source: HMSA, single coverage health insurance premium; deflated using the CPI for Honolulu.

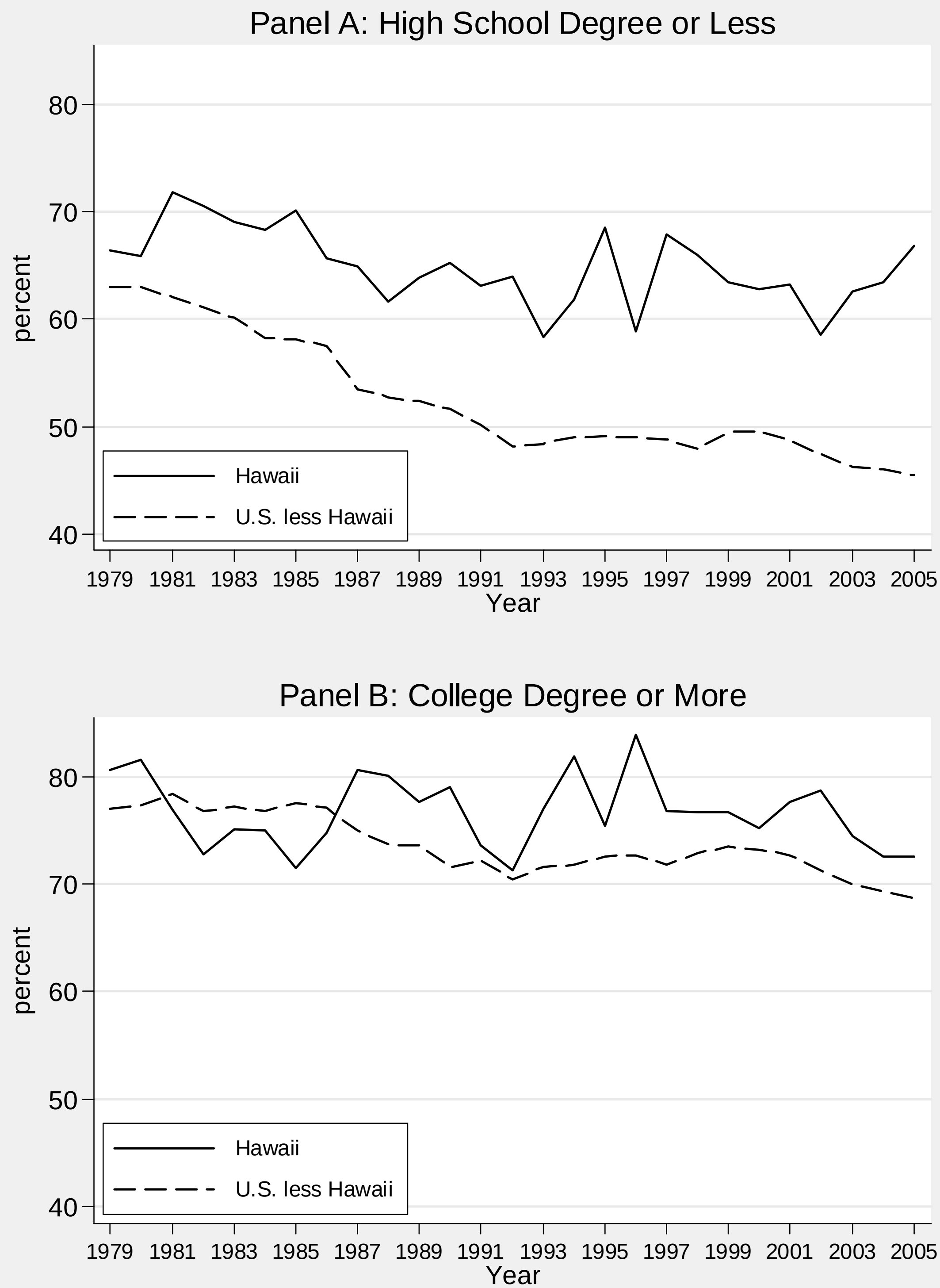
Figure 2: ESI Coverage, 1979-2005

(private sector employees)

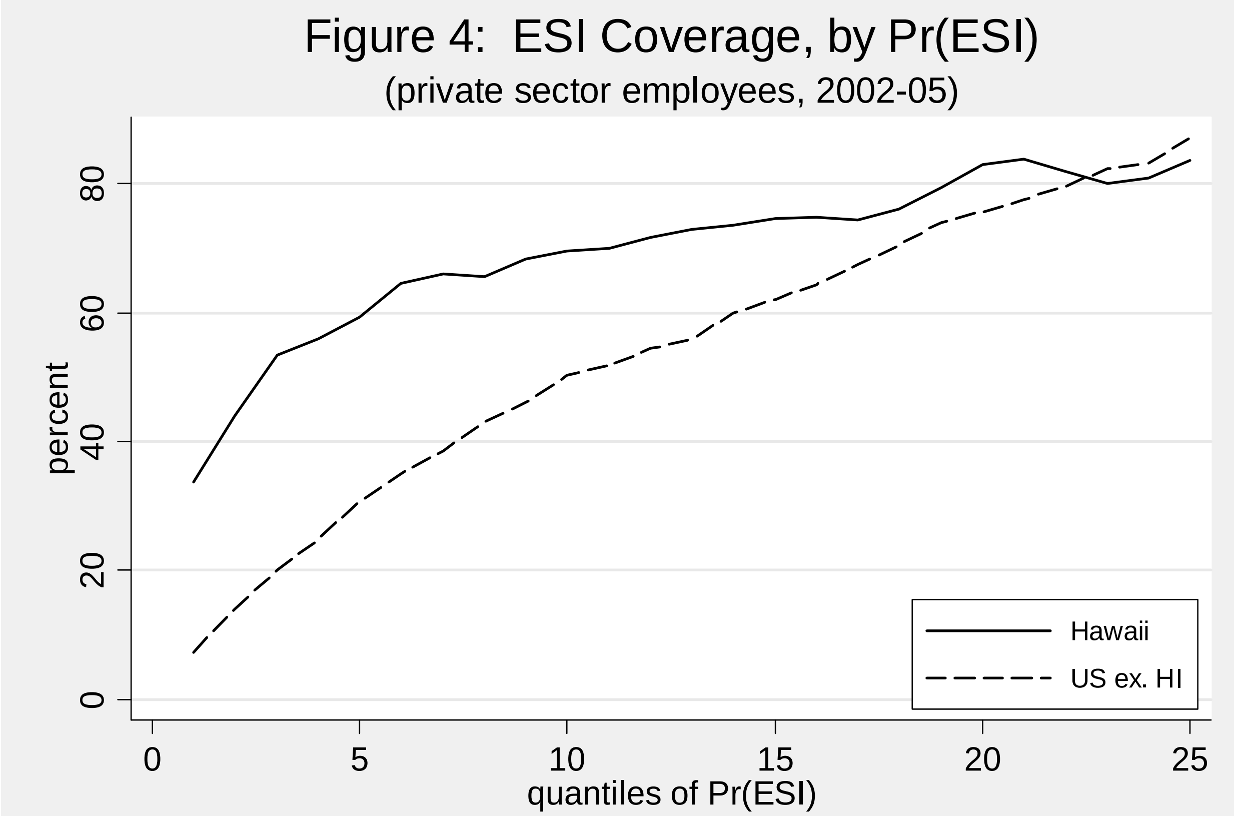


Note: Authors' tabulations (weighted) from March CPS files, 1980-2006.

Figure 3: ESI Coverage by Education, 1979-2005
(private sector employees)

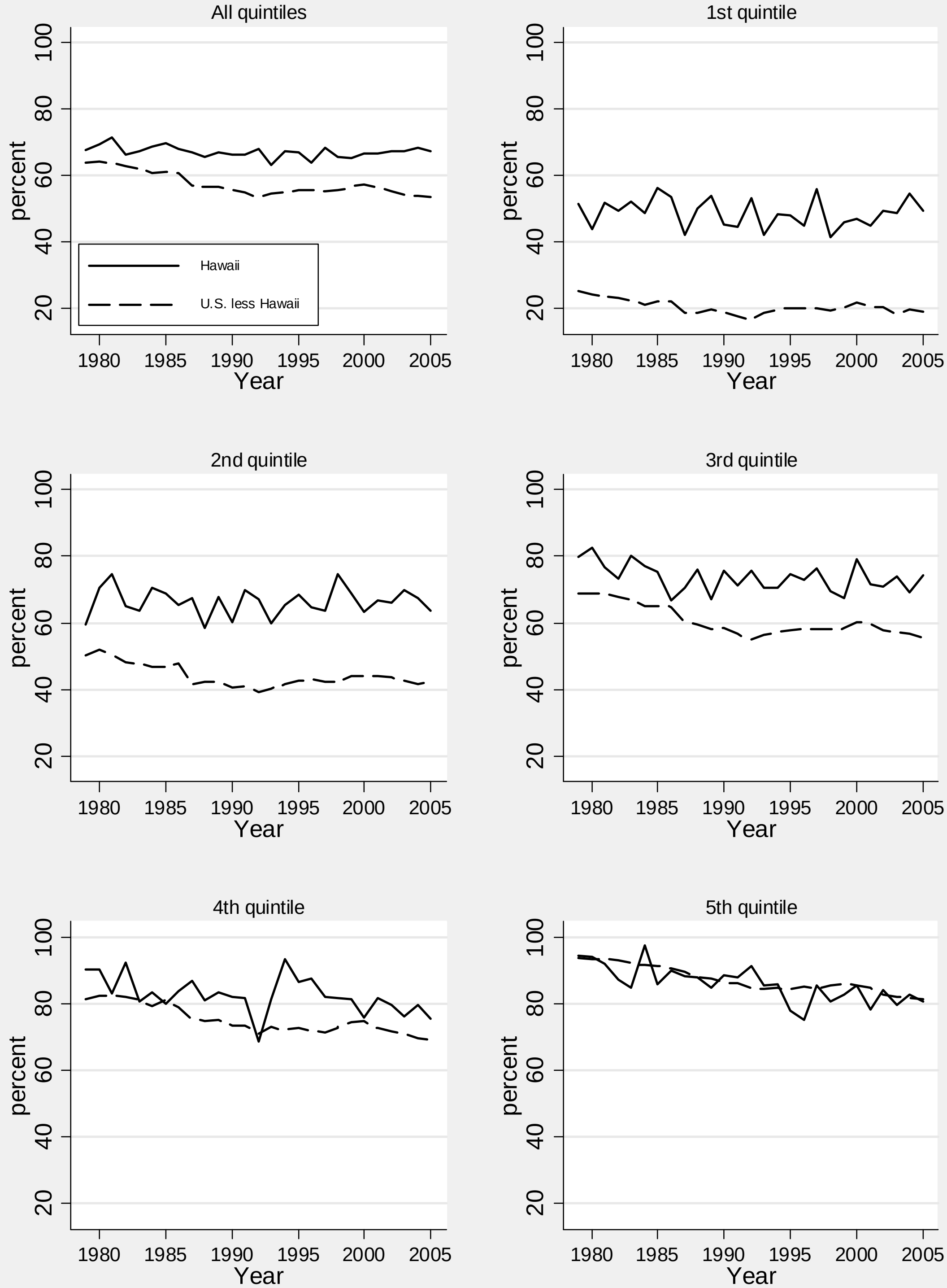


Note: Authors' tabulations (weighted) from March CPS files, 1980-2006.



Note: Authors' calculations using March CPS, 2003-2006. Quantiles formed using fitted Pr(ESI), structure estimated annually using U.S. ex. Hawaii observations (demographics, industry, and occupation controls); Hawaii line smoothed (average of 3 data points).

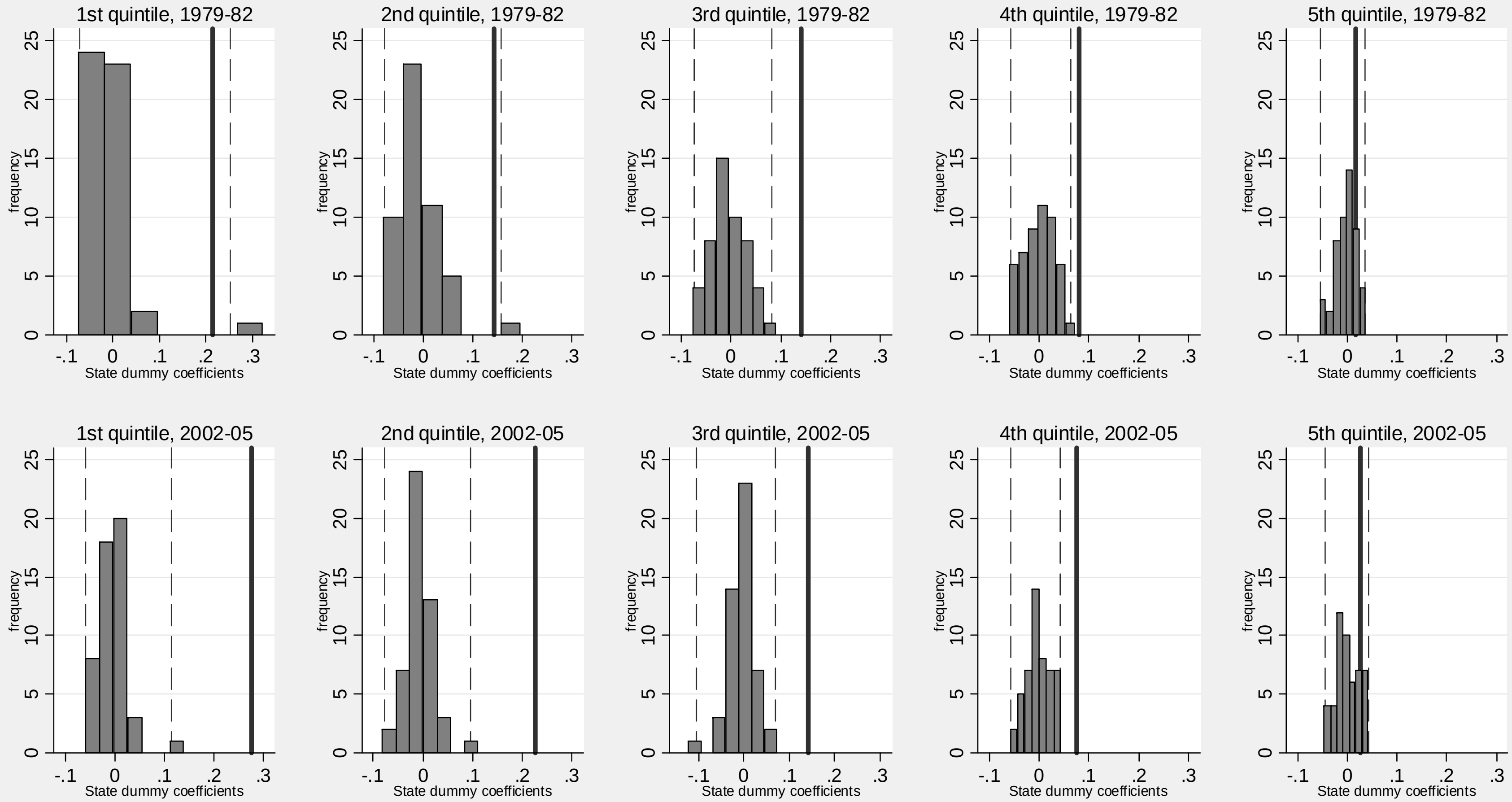
Figure 5: ESI Coverage by Pr(ESI) Quintiles, 1979-2005
(private sector employees)



Note: Authors' calculations using March CPS, 1980-2006. Quintiles formed using fitted Pr(ESI), structure estimated using U.S. excluding Hawaii observations (demographics and ind/occ).

Figure 6: Distribution of State Effects, Own ESI Coverage

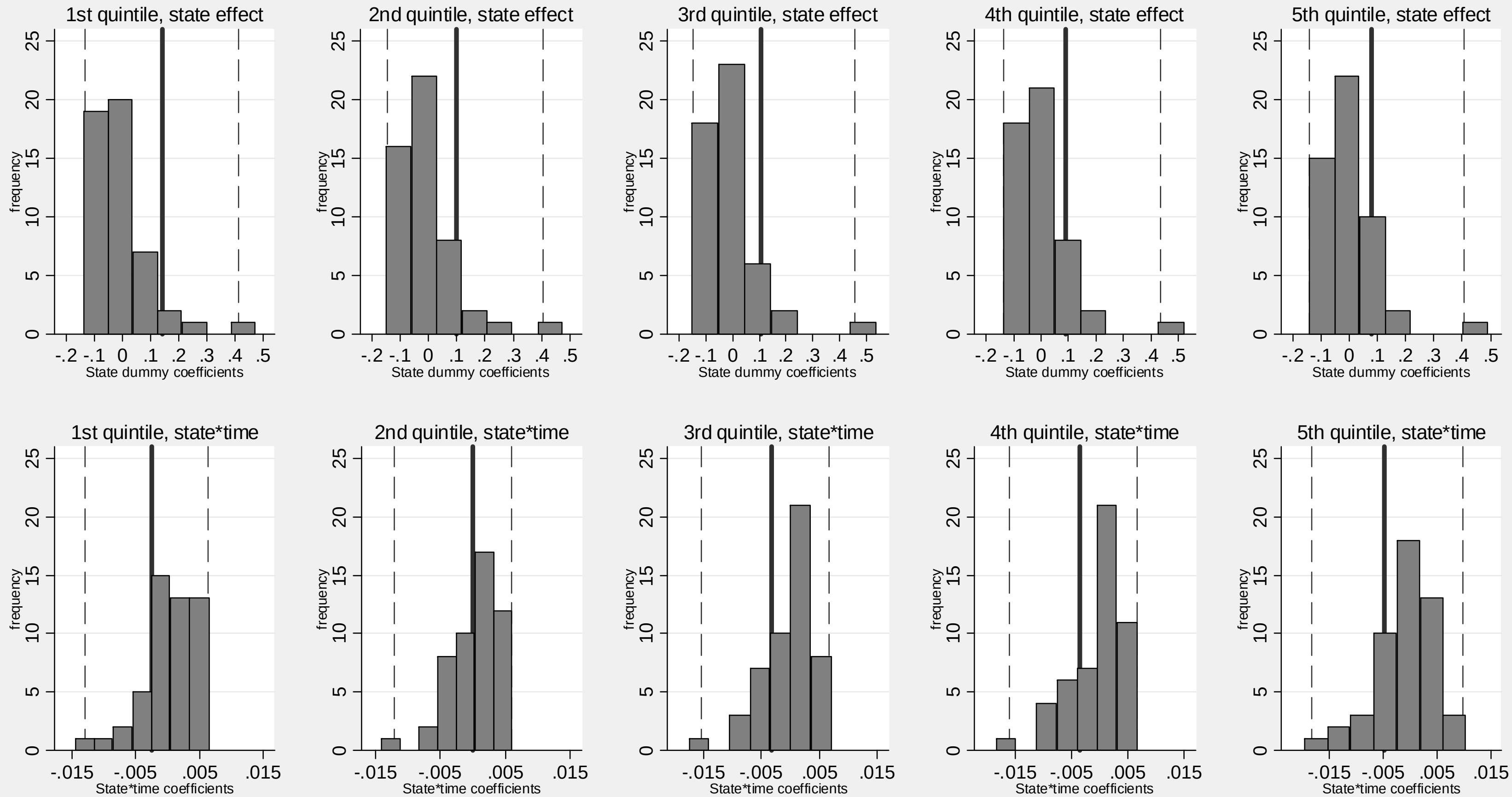
by ESI quintile, 1979-82 & 2002-05 (conditional on demographics)



Note: Estimated using March CPS data for 1979-82 and 2002-05. The thin dotted lines are the 2.5 and 97.5 percentile values (other than Hawaii), the thick solid line is the Hawaii value. Linear probability models; covariates include the complete set of demographic and job controls (see Tables 1-2). Quintiles formed as in Figure 5.

Figure 7: Distribution of State Effects, Log(Wage)

by ESI quintile, 1979-2005



Note: Estimated using CPS MORG data for 1979-2005. The thin dotted lines are the 2.5 and 97.5 percentile values (other than Hawaii), the thick solid line is the Hawaii value. Covariates include the complete set of demographic controls (excluding nativity), ind/occ, plus annual real state min. wage and GDP growth. Quintiles formed as in Figure 5.

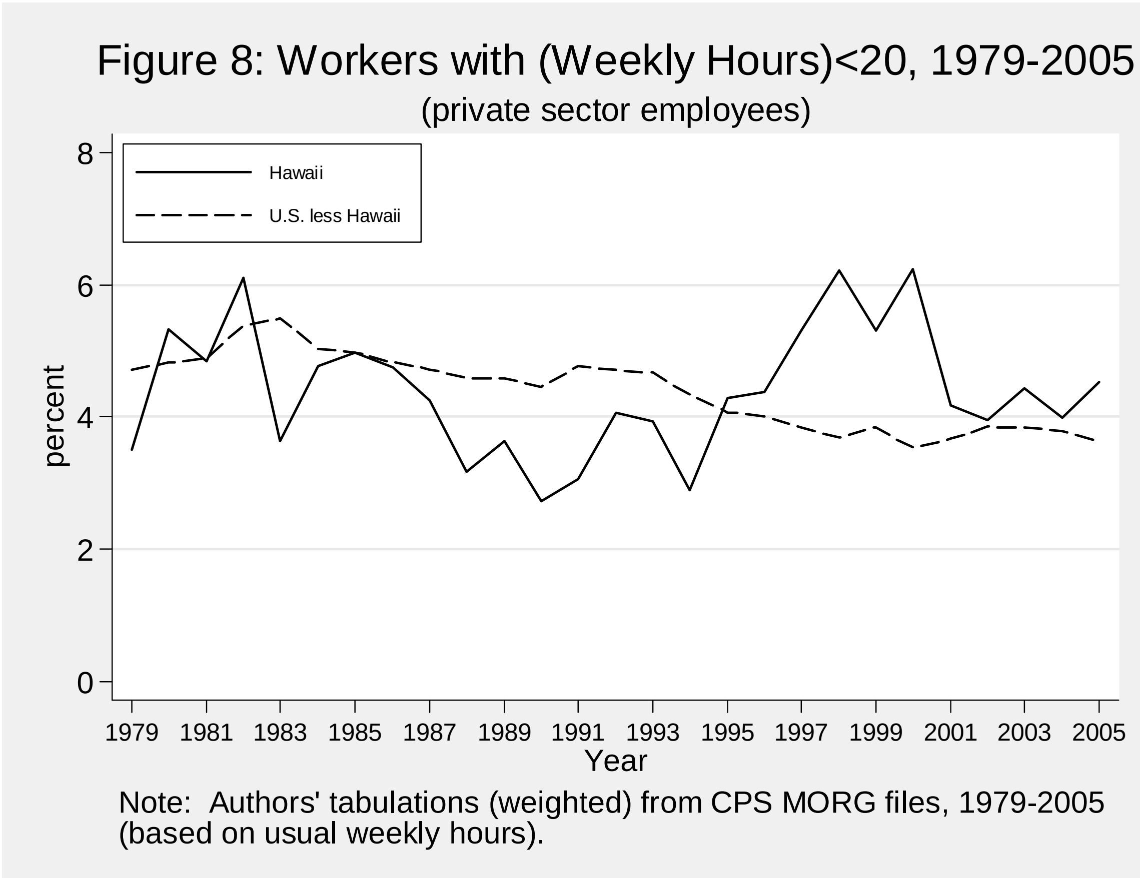
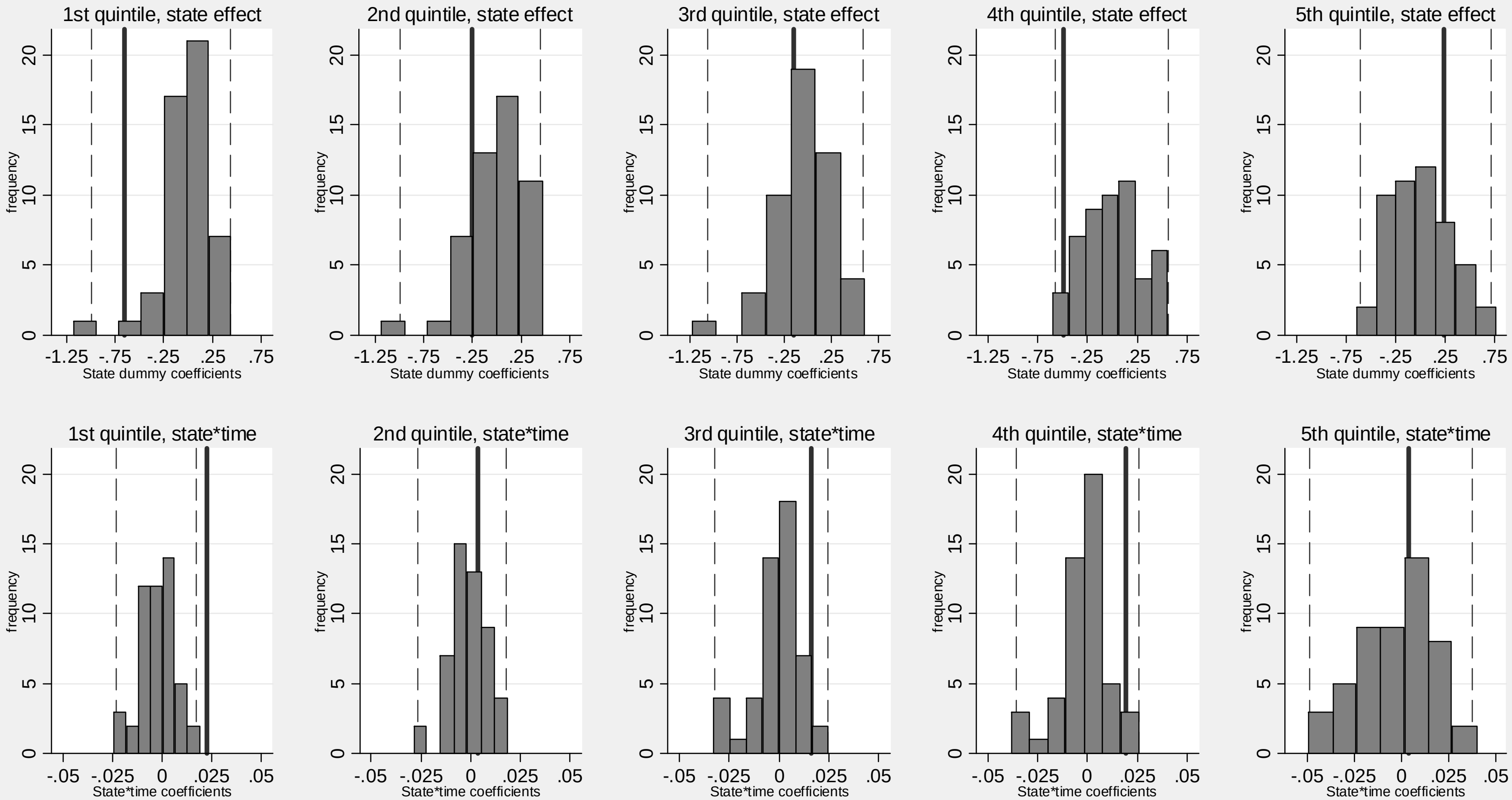


Figure 8: Workers with (Weekly Hours)<20, 1979-2005
(private sector employees)
Hawaii
U.S. less Hawaii
percent
0
2
4
6
8
1979
1981
1983
1985
1987
1989
1991
1993
1995
1997
1999
2001
2003
2005
Year
Note: Authors' tabulations (weighted) from CPS MORG files, 1979-2005
(based on usual weekly hours).

Figure 9: Distribution of State Effects, Pr(hours<20)

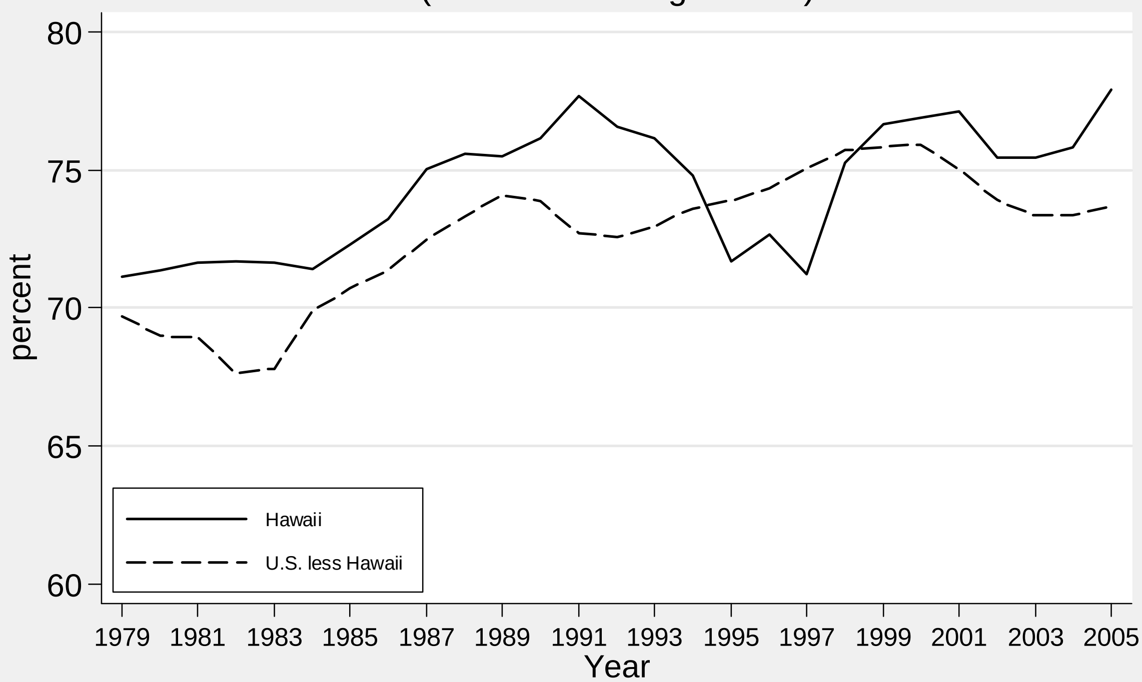
Logit regressions, by ESI quintile, 1979-2005



Note: Estimated using CPS MORG data for 1979-2005. The thin dotted lines are the 2.5 and 97.5 percentile values (other than Hawaii), the thick solid line is the Hawaii value. Logit models; covariates include the complete set of demographic controls (excluding nativity), ind/occ, plus annual real state min. wage and GDP growth. Quintiles formed as in Figure 5.

Figure 10: Percent Employed, 1979-2005

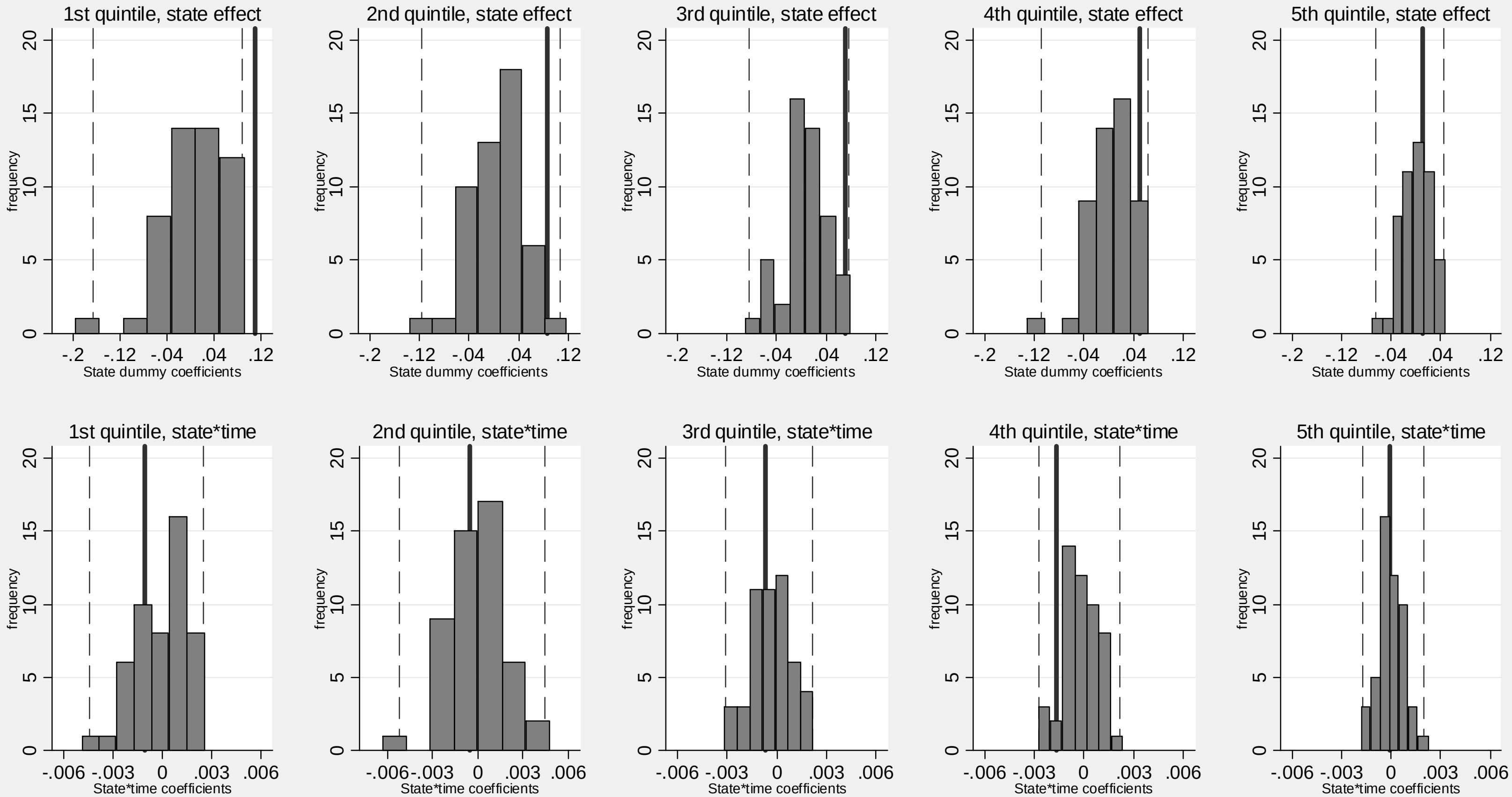
(all individuals age 18-64)



Note: Authors' tabulations (weighted) from CPS MORG files, 1979-2005 (individuals working or with a job but not at work).

Figure 11: Distribution of State Effects, Pr(Employed)

by ESI quintile, 1979-2005



Note: Estimated using CPS MORG data for 1979-2005. The thin dotted lines are the 2.5 and 97.5 percentile values (other than Hawaii), the thick solid line is the Hawaii value. Linear probability models; covariates include the complete set of demographic controls (excluding nativity), plus annual real state min. wage and GDP growth. Quintiles formed as in Figure 5.

Table 1: Population Characteristics, Hawaii and U.S. excluding Hawaii (mean values; calculated from March CPS files)

Variable	1980-1983 (survey years)		2003-2006 (survey years)	
	Hawaii	U.S. ex. Hawaii	Hawaii	U.S. ex. Hawaii
Age (years)	35.7	35.2	38.6	38.3
	(shares)			
Married	0.574	0.598	0.506	0.540
Female	0.506	0.450	0.482	0.470
MSA	0.304	0.666	0.755	0.704
Vet	0.139	0.188	0.065	0.071
Education				
<High School	0.147	0.212	0.075	0.127
High School	0.388	0.396	0.319	0.320
Some College	0.294	0.246	0.344	0.304
College Grad	0.111	0.089	0.206	0.179
>College	0.060	0.057	0.057	0.070
Race/Ethnicity				
White	0.290	0.825	0.176	0.685
Black	0.007	0.097	0.012	0.109
Hispanic	0.022	0.058	0.068	0.146
Other	0.681	0.020	0.744	0.059
Nativity				
U.S.-born	n/a	n/a	0.769	0.839
Foreign-born citizen	n/a	n/a	0.133	0.057
Non-citizen	n/a	n/a	0.098	0.104
Firm size				
<10 emps	n/a	n/a	0.164	0.155
10-24 emps	n/a	n/a	0.127	0.118
25-99 emps	n/a	n/a	0.180	0.156
100-499 emps	n/a	n/a	0.154	0.151
500+ emps	n/a	n/a	0.376	0.421

(continued)

Table 1 (con.)

Variable	1980-1983 (survey years)		2003-2006 (survey years)	
	Hawaii	U.S. ex. Hawaii	Hawaii	U.S. ex. Hawaii
	(shares)			
Industry				
Agriculture	0.045	0.021	0.017	0.011
Mining	0.001	0.013	0.001	0.005
Construction	0.077	0.063	0.093	0.077
Manufact. (non-dur)	0.060	0.115	0.024	0.057
Manufact. (dur.)	0.024	0.172	0.021	0.096
TCPU	0.086	0.070	0.080	0.070
Whlsle Trade	0.054	0.048	0.036	0.038
Retail Trade	0.244	0.194	0.150	0.142
F.I.R.E.	0.111	0.070	0.077	0.080
Business Services	0.061	0.045	0.085	0.077
Personal Services	0.085	0.038	0.187	0.103
Entertainment Serv.	0.015	0.012	0.019	0.018
Prof. Services	0.135	0.139	0.208	0.225
Occupation				
Managerial	0.097	0.090	0.124	0.128
Professional	0.105	0.098	0.114	0.135
Technical	0.015	0.018	0.020	0.029
Sales	0.103	0.082	0.131	0.129
Administrative	0.195	0.187	0.159	0.148
Services	0.196	0.134	0.225	0.163
Prod/craft/repair	0.117	0.142	0.097	0.090
Oper/fab/labor	0.059	0.135	0.041	0.086
Trans/movers	0.039	0.042	0.044	0.045
Handlers/cleaners	0.042	0.054	0.030	0.038
Farm	0.032	0.016	0.015	0.009
Sample size	2,477	230,619	4,507	297,649

Note: Authors' calculations from March CPS data, private sector workers age 18-64 (weighted).

Table 2: Estimated ESI Coverage Gaps, by ESI Quintile
(placebo tests, Hawaii vs. all other states)

Panel A: 1979-1982

	Full sample	ESI Quintiles				
		1st	**2nd**	**3rd**	**4th**	**5th**
Unadjusted						
Hawaii effect	0.053	0.238	0.162	0.098	0.065	-0.002
Placebo tests (other states)						
2.5^{th} percentile	-0.124	-0.073	-0.085	-0.069	-0.064	-0.074
97.5^{th} percentile	0.073	0.238	0.151	0.076	0.072	0.043
Adjusted for Demographics						
Hawaii effect	0.101	0.207	0.145	0.130	0.074	0.007
Placebo tests (other states)						
2.5^{th} percentile	-0.087	-0.066	-0.083	-0.077	-0.063	-0.066
97.5^{th} percentile	0.079	0.238	0.150	0.074	0.070	0.039
Adjusted for Demographics and Job Characteristics						
Hawaii effect	0.145	0.213	0.144	0.143	0.080	0.017
Placebo tests (other states)						
2.5^{th} percentile	-0.057	-0.072	-0.078	-0.075	-0.059	-0.054
97.5^{th} percentile	0.142	0.251	0.158	0.083	0.064	0.036

(continued)

Table 2 (continued)

Panel B: 2002-2005

	Full sample	ESI Quintiles				
		1st	**2nd**	**3rd**	**4th**	**5th**
Unadjusted						
Hawaii effect	0.140	0.318	0.241	0.155	0.075	-0.009
Placebo tests (other states)						
2.5^{th} percentile	-0.108	-0.058	-0.068	-0.109	-0.055	-0.064
97.5^{th} percentile	0.055	0.150	0.126	0.098	0.056	0.046
Adjusted for Demographics						
Hawaii effect	0.132	0.286	0.220	0.124	0.061	-0.014
Placebo tests (other states)						
2.5^{th} percentile	-0.080	-0.060	-0.078	-0.124	-0.061	-0.071
97.5^{th} percentile	0.072	0.133	0.130	0.092	0.060	0.049
Adjusted for Demographics and Job Characteristics						
Hawaii effect	0.168	0.276	0.228	0.142	0.076	0.025
Placebo tests (other states)						
2.5^{th} percentile	-0.061	-0.058	-0.076	-0.105	-0.057	-0.045
97.5^{th} percentile	0.073	0.115	0.095	0.069	0.043	0.042

Note: Estimated using March CPS data, private sector workers age 18-64. The demographic characteristics are education (5 categories), a quartic in age, gender, gender by age quartic interactions, married, married by gender interaction, race/ethnicity (4 categories), residence in an urban area, veteran status, and year dummies (3), plus nativity (3 categories) in 2003-06. The job characteristics are industry (13) and occupation (11) categories, plus firm size (5 categories) in 2003-06 (see Table 1). Quintiles formed using fitted Pr(ESI), structure estimated using U.S. excluding Hawaii observations (demographics and ind/occ).

Table 3: Health Insurance Coverage, All Sources, Hawaii versus Rest of U.S. (2002-2005)

	All Private-Sector Workers		HS Degree or Less		College- Educated	
	U.S. (ex. Hawaii)	Hawaii	U.S. (ex. Hawaii)	Hawaii	U.S. (ex. Hawaii)	Hawaii
ESI—Own Name	54.6	67.8	46.7	63.1	70.0	74.4
ESI—Dependent	12.3	10.1	10.8	10.0	11.6	9.0
Total ESI	67.0	77.8	57.5	73.1	81.6	83.4
Private Non-group	3.7	2.3	3.0	2.1	3.6	2.1
CHAMPUS	0.9	1.8	0.9	1.0	0.6	2.5
Medicaid	3.4	2.6	5.2	4.3	0.8	0.7
Medicare	0.2	0.1	0.3	0.3	0.1	0.0
Total Public	4.5	4.5	6.4	5.5	1.5	3.2
Uninsured	24.9	15.3	33.1	19.4	13.4	11.3

Note: Authors' calculations from March CPS data, private sector workers age 18-64 (weighted).

Table 4: Estimated Log(Wage) Effects
(placebo tests, Hawaii vs. all other states)

	Full sample	ESI Quintiles				
		1st	**2nd**	**3rd**	**4th**	**5th**
State dummy						
Hawaii effect	0.1044	0.1430	0.1014	0.1051	0.0890	0.0779
Placebo tests (other states)						
2.5^{th} percentile	-0.1202	-0.1305	-0.1451	-0.1484	-0.1387	-0.1424
97.5^{th} percentile	0.4148	0.4119	0.4068	0.4571	0.4352	0.4078
*State*time*						
Hawaii effect	-0.0020	-0.0025	0.0001	-0.0032	-0.0036	-0.0048
Placebo tests (other states)						
2.5^{th} percentile	-0.0141	-0.0130	-0.0120	-0.0156	-0.0162	-0.0184
97.5^{th} percentile	0.0064	0.0063	0.0060	0.0067	0.0068	0.0098

Note: Estimated using CPS MORG data, 1979-2005, private sector workers age 18-64. Covariates include the complete set of demographic controls (excluding nativity) and ind/occ (see Tables 1-2), plus annual values of the real minimum wage and GDP growth by state. Quintiles formed using fitted Pr(ESI), structure estimated using March CPS data, U.S. excluding Hawaii observations (demographics and ind/occ).

Table 5: Estimated Effects on Pr[(Usual Weekly Hours)<20] (Logit Regressions)
(placebo tests, Hawaii vs. all other states)

	Full sample	ESI Quintiles				
		1st	**2nd**	**3rd**	**4th**	**5th**
State dummy						
Hawaii effect	-0.4473	-0.6466	-0.2517	-0.1465	-0.4888	0.2299
Placebo tests (other states)						
2.5^{th} percentile	-0.9481	-0.9955	-0.9920	-1.0625	-0.5739	-0.6028
97.5^{th} percentile	0.4150	0.4315	0.4536	0.5859	0.5590	0.7106
*State*time*						
Hawaii effect	0.0166	0.0224	0.0037	0.0161	0.0194	0.0040
Placebo tests (other states)						
2.5^{th} percentile	-0.0195	-0.0229	-0.0267	-0.0324	-0.0360	-0.0490
97.5^{th} percentile	0.0128	0.0172	0.0182	0.0248	0.0258	0.0379

Note: Estimated using CPS MORG data, 1979-2005, private sector workers age 18-64. Logit models; covariates include the complete set of demographic controls (excluding nativity) and ind/occ (see Tables 1-2), plus annual values of the real minimum wage and GDP growth by state. Quintiles formed using fitted Pr(ESI), structure estimated using March CPS data, U.S. excluding Hawaii observations (demographics and ind/occ).

Table 6: Estimated Effects on Pr(Employed) (Linear Prob Models)
(placebo tests, Hawaii vs. all other states)

	Full sample	ESI Quintiles				
		1st	**2nd**	**3rd**	**4th**	**5th**
State dummy						
Hawaii effect	0.0572	0.1108	0.0852	0.0706	0.0506	0.0114
Placebo tests (other states)						
2.5^{th} percentile	-0.0971	-0.1661	-0.1157	-0.0838	-0.1094	-0.0658
97.5^{th} percentile	0.0602	0.0870	0.1073	0.0769	0.0628	0.0449
*State*time*						
Hawaii effect	-0.0005	-0.0011	-0.0006	-0.0007	-0.0017	-0.0001
Placebo tests (other states)						
2.5^{th} percentile	-0.0023	-0.0044	-0.0053	-0.0031	-0.0027	-0.0017
97.5^{th} percentile	0.0016	0.0025	0.0045	0.0022	0.0022	0.0020

Note: Estimated using CPS MORG data, 1979-2005, all workers age 18-64. Linear probability models; covariates include the complete set of demographic controls (excluding nativity; see Tables 1-2), plus annual values of the real minimum wage and GDP growth by state. Quintiles formed using fitted Pr(ESI), structure estimated using March CPS data, U.S. excluding Hawaii observations (demographics only).

Appendix A: CPS March and MORG Data

Our data are from the Current Population Survey's Annual Demographic Supplement (conducted each year in March, hence referred to as the March CPS) and the Monthly Outgoing Rotation Group (MORG) files. We used the processed March files from Unicon Research Corporation (www.unicon.com) and the MORG files from the National Bureau of Economic Research (www.nber.org/data/morg.html). We describe the characteristics of each source and our specific data choices in detail here. For each data set, the analyses are limited to individuals between the ages of 18 and 64 who are employed by a private sector firm (with the exception of the employment analyses, for which the sample is *all* individuals age 18-64).

We use the March CPS data for our analyses of employer-sponsored health insurance coverage (ESI). Since the 1980 survey, respondents have been asked about health insurance coverage on their longest job held in the preceding year, but the specifics of the question have changed over time. Before the 1988 survey, the ESI question was posed to all working individuals. After 1988, the question was posed to all individuals who had health insurance coverage in their own name. This change in definition creates a series break for tabulations of health insurance coverage, and the Census has chosen not to release such tabulations for calendar years prior to 1987. However, Census Bureau correspondence available in the Unicon data documentation indicates that the impact of the 1988 questionnaire change was more significant for the measurement of sources of health insurance other than ESI. Beginning in 1995, respondents were asked a series of verification questions if their initial response indicated that they were not covered by ESI; additional verification questions were added in 2000. The verification questions increase estimated ESI coverage by an amount that is small but nontrivial from the perspective of national coverage totals. To attain the highest consistency over time, we used an ESI variable that is defined without reliance on the verification questions, although we ascertained in preliminary analyses that using the verification questions does not materially affect our results.

We use the MORG data for our analyses of wage, hours, and employment outcomes. As noted in the text, compared with the March data, the MORG provides larger sample sizes for more precise point-in-time measures of these variables. (We conducted parallel outcome analyses using the March data and found results that were similar but less precise than those obtained using the MORG data, which is as expected given the smaller sample sizes in the March data.) Our hourly wage variable is defined as reported hourly earnings for those paid by the hour and usual weekly earnings divided by usual weekly hours for salaried employees. We used the CPI for all urban consumers to deflate hourly wages and other dollar-denominated figures, using the CPI for the Honolulu metro area to deflate the Hawaii figures and the all-U.S. series for other states. We limited the wage analyses to individuals whose hourly wage is greater than $1 (in 2005 dollars). We multiplied the value of top-coded earnings observations by 1.4; this largely follows Lemieux 2006, with the exception that for the sake of consistency across analyses that exclude or include imputed earnings observations (see below), we did not rely on the higher topcode enabled by the use of unedited earnings values for the years 1986-1988.

We also relied on the usual weekly hours variable to identify individuals who work fewer than 20 hours per week. Autor, Katz, and Kearney (2008) noted that this hours variable is not consistently available over time, causing missing observations for the 7-9 percent of the sample after 1994 who report that their hours of work vary. We excluded these observations from the sample rather than follow Autor et al's approach of relying on a variable that measures hours worked at all jobs during the reference week, because we need precise information on weekly hours at sample members' primary job in order to analyze the incidence of low-hours employment. As noted in the text, we obtain similar results for the hours analysis when we use the alternative variable based on hours worked in all jobs (the same is true for the wage analysis).

As discussed by Lemieux (2006) and Autor, Katz, and Kearney (2008), the incidence of imputed values of earnings and hours is substantial in both CPS data sources and has grown over time. Because the Census "hot deck" matching procedure used for imputation does not restrict donor matches to individuals in the same state, our results, which focus on one small state, could be substantially biased if we included imputed observations. We therefore dropped observations with imputed values of earnings from the analysis of earnings and observations with imputed values of hours worked from the analysis of low-hours work. We followed the procedures outlined in Lemieux (2006) for identifying imputed earnings observations (including the comparison of unedited and edited earnings values during the years 1989-1993, when the earnings imputation flags are incorrect). Supplementary analyses conducted for various specifications indicated that the results are similar when imputed values are included.

Most of our analyses include industry and occupation controls (see Table 1 for the categories). These definitions have changed over time, with especially large changes in the 2003 data year, due to the switch to the new NAICS codes. We formed consistent broad categories over time based on a comprehensive examination of the codes and consequent reclassification of some industry categories. These recoding schemes are available on request.

Appendix B: Complete Regression Results (selected equations)

VARIABLES	(1) ESI Coverage, 1979-82 (Table 2, Panel A)	(2) ESI Coverage, 2002-05 (Table 2, Panel B)	(3) Log(wage), 1979-2005 (Table 5)	(4) Low Hours (logit), 1979-2005 (Table 6)	(5) Employed (LPM), 1979-2005 (Table 7)
Hawaii dummy	0.145**	0.168**	0.104**	-0.447**	0.0572**
	(0.0145)	(0.00712)	(0.0196)	(0.0431)	(0.00878)
Time trend (annual)			0.00326**	-0.0144**	0.000978**
			(0.000646)	(0.00150)	(0.000211)
Time*Hawaii			-0.00195**	0.0166**	-0.000462**
			(0.000494)	(0.00115)	(0.000170)
Ln(state min. wage)			0.413**	-0.207	-0.00939
			(0.0659)	(0.107)	(0.0160)
Change in ln(state GDP)			-0.0329	-1.288**	0.146**
			(0.0896)	(0.335)	(0.0343)
Age	0.450**	0.342**	0.0704**	-2.449**	0.350**
	(0.0158)	(0.0164)	(0.00817)	(0.0865)	(0.0107)
$Age^2/10$	-0.162**	-0.113**	-0.0136**	0.886**	-0.137**
	(0.00610)	(0.00650)	(0.00300)	(0.0350)	(0.00387)
$Age^3/1000$	0.251**	0.164**	0.0141**	-1.424**	0.233**
	(0.0101)	(0.0110)	(0.00484)	(0.0597)	(0.00599)
$Age^4/100,000$	-0.144**	-0.0882**	-0.00838**	0.861**	-0.147**
	(0.00608)	(0.00677)	(0.00288)	(0.0364)	(0.00338)
HS degree	0.0708**	0.0836**	0.131**	-0.408**	0.150**
	(0.00416)	(0.00310)	(0.00549)	(0.0163)	(0.00423)
Some college	0.0254**	0.0855**	0.198**	0.205**	0.171**
	(0.00734)	(0.00397)	(0.00836)	(0.0235)	(0.00626)
College degree	0.0737**	0.127**	0.364**	0.0134	0.226**
	(0.00716)	(0.00488)	(0.00909)	(0.0253)	(0.00725)
Post-college degree	0.0555**	0.151**	0.475**	0.159**	0.245**
	(0.00735)	(0.00584)	(0.0108)	(0.0298)	(0.00767)
Female	-1.205**	-0.863**	-0.805**	3.808**	1.510**
	(0.149)	(0.184)	(0.0563)	(0.787)	(0.0700)
Female*Age	0.156**	0.102**	0.0743**	-0.576**	-0.177**
	(0.0178)	(0.0209)	(0.00623)	(0.0967)	(0.00706)
Female*($Age^2/10$)	-0.0690**	-0.0426**	-0.0254**	0.288**	0.0702**
	(0.00743)	(0.00834)	(0.00249)	(0.0414)	(0.00261)
Female*($Age^3/1000$)	0.122**	0.0733**	0.0324**	-0.538**	-0.116**
	(0.0130)	(0.0141)	(0.00426)	(0.0732)	(0.00429)
Female*($Age^4/100,000$)	-0.0750**	-0.0441**	-0.0130**	0.331**	0.0685**
	(0.00804)	(0.00865)	(0.00264)	(0.0456)	(0.00265)

(continued)

Appendix B (continued)

VARIABLES	(1) ESI Coverage, 1979-82 (Table 2, Panel A)	(2) ESI Coverage, 2002-05 (Table 2, Panel B)	(3) Log(wage), 1979-2005 (Table 5)	(4) Low Hours (logit), 1979-2005 (Table 6)	(5) Employed (LPM), 1979-2005 (Table 7)
Married	0.0813**	0.0149**	0.146**	-0.825**	0.115**
	(0.00389)	(0.00440)	(0.00370)	(0.0271)	(0.00307)
Married*female	-0.229**	-0.171**	-0.152**	1.434**	-0.212**
	(0.00924)	(0.00728)	(0.00325)	(0.0382)	(0.00438)
Black	-0.0166**	-0.0325**	-0.102**	-0.445**	-0.0783**
	(0.00537)	(0.00439)	(0.0119)	(0.0515)	(0.00738)
Non-white/black	-0.0393**	-0.0218**	-0.0646**	-0.201**	-0.0847**
	(0.0131)	(0.00718)	(0.0110)	(0.0409)	(0.0126)
Hispanic	-0.0295**	-0.0380**	-0.129**	-0.579**	-0.0328**
	(0.00920)	(0.00546)	(0.0171)	(0.0460)	(0.00949)
MSA residence	0.0287**	0.0164**	0.139**	-0.0835*	0.00456
	(0.00849)	(0.00399)	(0.0128)	(0.0369)	(0.00504)
Naturalized citizen		-0.00217			
		(0.00493)			
Non-citizen		-0.0766**			
		(0.00664)			
Veteran	0.000496	-0.0159**	0.0159**	0.0195	-0.00654**
	(0.00412)	(0.00468)	(0.00344)	(0.0171)	(0.00167)
Industries: Mining	0.336**	0.176**	0.467**	-1.182**	
	(0.0133)	(0.0184)	(0.0290)	(0.0978)	
Construction	0.0865**	0.0135	0.282**	-0.179**	
	(0.0140)	(0.0115)	(0.0169)	(0.0556)	
Manufacturing (non-dur)	0.318**	0.136**	0.226**	-0.760**	
	(0.0165)	(0.0155)	(0.0182)	(0.0775)	
Manufacturing (durable)	0.354**	0.144**	0.286**	-1.481**	
	(0.0129)	(0.0125)	(0.0161)	(0.0572)	
Trans/comm/public utilities	0.319**	0.0904**	0.339**	-0.498**	
	(0.0137)	(0.0138)	(0.0155)	(0.0548)	
Wholesale trade	0.268**	0.134**	0.196**	-0.701**	
	(0.0132)	(0.0124)	(0.0127)	(0.0626)	
Retail trade	0.0996**	-0.000651	-0.0221	0.133**	
	(0.0126)	(0.0126)	(0.0125)	(0.0453)	
Finance/ins/real estate	0.271**	0.0869**	0.240**	-0.653**	
	(0.0151)	(0.0138)	(0.0166)	(0.0581)	
Business services	0.0900**	-0.00537	0.146**	0.0672	
	(0.0146)	(0.0131)	(0.0144)	(0.0486)	
Personal services	0.0284	-0.0585**	-0.00954	0.520**	
	(0.0400)	(0.0147)	(0.0239)	(0.0692)	
Entertainment services	0.0460	0.00825	0.0474**	0.684**	
	(0.0370)	(0.0198)	(0.0172)	(0.0716)	

(continued)

Appendix B (continued)

VARIABLES	(1) ESI Coverage, 1979-82 (Table 2, Panel A)	(2) ESI Coverage, 2002-05 (Table 2, Panel B)	(3) Log(wage), 1979-2005 (Table 5)	(4) Low Hours (logit), 1979-2005 (Table 6)	(5) Employed (LPM), 1979-2005 (Table 7)
Professional services	0.177**	0.0574**	0.123**	0.223**	
	(0.0129)	(0.0132)	(0.0144)	(0.0554)	
Occupations: Managerial	0.242**	0.176**	0.459**	-1.599**	
	(0.0295)	(0.0169)	(0.00951)	(0.0842)	
Professional services	0.206**	0.126**	0.437**	-0.344**	
	(0.0262)	(0.0166)	(0.00872)	(0.0768)	
Technical	0.230**	0.149**	0.405**	-0.642**	
	(0.0283)	(0.0177)	(0.00808)	(0.0886)	
Sales	0.0866**	0.0589**	0.248**	-0.141	
	(0.0299)	(0.0155)	(0.00771)	(0.0783)	
Administrative	0.198**	0.116**	0.194**	-0.369**	
	(0.0267)	(0.0161)	(0.00812)	(0.0797)	
Service	0.0102	0.0131	0.0366**	0.132	
	(0.0284)	(0.0176)	(0.00924)	(0.0756)	
Prod/craft/repair	0.172**	0.113**	0.281**	-1.111**	
	(0.0270)	(0.0164)	(0.00835)	(0.0740)	
Operators	0.172**	0.0898**	0.126**	-1.035**	
	(0.0310)	(0.0145)	(0.0123)	(0.0929)	
Transportation/moving	0.102**	0.0438*	0.130**	-0.154	
	(0.0282)	(0.0169)	(0.00853)	(0.0822)	
Handlers/cleaners	0.0961**	0.00682	0.101**	-0.128	
	(0.0245)	(0.0170)	(0.00792)	(0.0743)	
Firm size 10-24		0.101**			
		(0.00377)			
Firm size 25-99		0.200**			
		(0.00388)			
Firm size 100-499		0.264**			
		(0.00434)			
Firm size 500+		0.309**			
		(0.00545)			
Year 1980	-0.000107				
	(0.00260)				
Year 1981	-0.00231				
	(0.00351)				
Year 1982	-0.00610				
	(0.00405)				
Year 2003		-0.00780**			
		(0.00258)			
Year 2004		-0.00982**			
		(0.00235)			

(continued)

Appendix B (continued)

VARIABLES	(1) ESI Coverage, 1979-82	(2) ESI Coverage, 2002-05	(3) Log(wage), 1979-2005	(4) Low Hours (logit), 1979-2005	(5) Employed (LPM), 1979-2005
Year 2005		-0.0158**			
		(0.00238)			
Constant	-4.288**	-3.622**	-0.0555	22.46**	-2.626**
	(0.149)	(0.142)	(0.146)	(0.766)	(0.117)
Observations	233096	302156	2796134	3495264	6984926
R-squared	0.261	0.245	0.470	.	0.138

** p<0.01, * p<0.05

Standard errors (clustered by state) in parentheses.