Effects of the Affordable Care Act on Part-Time Employment: Early Evidence

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ABSTRACT

The Affordable Care Act (ACA) requires employers with at least 50 full-time-equivalent employees to offer "affordable" health insurance to employees working 30 or more hours per week. Employers who do not comply may face substantial penalties, but they can circumvent the mandate by reducing employees' weekly hours below the 30-hour threshold. We examine ACA's effects on short-hours part-time employment using difference-in-differences models. We find that the ACA increased low-hours, involuntary part-time employment by 500,000–700,000 workers in retail, accommodations, and food services, the industries in which employers are most likely to reduce hours if they choose to circumvent the mandate.

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I. Introduction

The Patient Protection and Affordable Care Act (ACA) requires that employers with at least 50 full-time-equivalent (FTE) workers offer "affordable" health insurance that meets minimum standards of coverage to employees who average at least 30 hours per week at their job. Large employers who do not offer any health insurance to full-time employees, whose policies do not cover minimum benefits, or who do not cover a sufficient share of the policy's premium may face substantial financial penalties. Although the employer mandate in the ACA is intended to increase employer-sponsored health insurance and thereby improve employees' compensation and the quality of jobs, concerns emerged that it could induce some employers to hire more part-time labor that would not be subject to the mandate. If employers increase their use of part-time work to circumvent the mandate, the ACA instead could worsen employment conditions for some workers by making full-time work more difficult to find.

Reflecting such concerns, a July 2013 letter from three of the largest labor unions to congressional leaders argued that the ACA's employer mandate would "shatter not only our hard-earned health benefits, but destroy the foundation of the 40-hour work week that is the backbone of the American middle class."[1] The union leaders argued that employers would be incentivized to keep or cut workers' hours below 30 hours per week to avoid the obligation to provide insurance. Some anecdotal evidence supports these concerns. For example, as of September 2014, several months before the mandate took effect for employers with at least 100 FTE employees, *Investor's Business Daily* had compiled a list of 450 public and private employers for which it claimed there was "strong proof" (that is, official documents or accounts) that these employers had cut work hours from full- to part-time or reduced hours of new hires to be less than 30 in order to circumvent the health insurance mandate.[2] The earliest documented instances in these accounts were in 2011.

In this work, we seek to shed light on the impact of the ACA's employer mandate on part-time employment. A major challenge in studying the effect of the ACA's employer mandate involves constructing an appropriate control group, since this mandate applies nationwide. To overcome this challenge, we exploit the fact that long before the ACA was passed Hawaii implemented an employer mandate that is considerably more stringent than the ACA's mandate, which means that, unlike all other states, Hawaii should not be affected by the ACA's employer mandate. We take advantage of Hawaii's preexisting mandate by implementing a difference-in-differences strategy that uses Hawaii as an already-treated control group for the rest of the nation to estimate the effect of the ACA's employer mandate on part-time employment and on involuntary part-time employment.

A second challenge in studying the employer mandate is that, prior to the ACA, a relatively small portion of the workforce—no more than 5 percent—was in a job that did not come with an offer of health insurance and that would be covered by the employer mandate (Dillender, Heinrich, and Houseman 2015; Glied and Solis-Roman 2014).

1. http://www.forbes.com/sites/theapothecary/2013/07/15/labor-leaders-obamacare-will-shatter-their-health-benefits-cause-nightmare-scenarios/ (accessed December 27, 2021).

2. http://news.investors.com/politics-obamacare/020314-669013-obamacare-employer-mandate-a-list-of-cuts-to-work-hours-jobs.htm (accessed December 27, 2021).

Workers in these newly covered jobs are economically vulnerable, the intended target of the employer mandate, and the focus of much policy attention. However, their small number means that, for example, if a study cannot detect as statistically significant a one percentage point increase in part-time work, the study cannot detect that at least 20 percent of the vulnerable population has been shifted to part-time work because of the employer mandate, which is a sizable share of these workers. Because studies of the effect of the mandate on all workers may be underpowered to identify meaningful effects, we focus the analysis largely on the retail, hotel, and restaurant industries. These industries have characteristics and business models that are conducive to using part-time work and had high shares of workers who were not offered employer-sponsored health insurance prior to the ACA's passage and who would be covered by the employer mandate. Therefore, firms in these industries may be more likely to increase part-time staffing to avoid higher health insurance costs because of the employer mandate.

To implement the difference-in-differences research design, we draw on basic monthly Current Population Survey (CPS) data, which has information on weekly hours worked and on reasons for working part-time. Estimates from our preferred specification indicate that the ACA has increased low-hours, involuntary part-time employment in the retail and accommodation and food services (RAF) industries by 500,000–700,000 workers (or about 17–24 percent of those vulnerable to hours reductions in these industries). Although the results do not suggest that the employer mandate has led to the drastic increase in low-hours part-time work as many had feared, the documented increases are sizable in the industries that employ a high share of the at-risk population, a particularly economically vulnerable (low-wage) segment of the labor force.

The remainder of the paper is organized as follows. We begin by providing background on relevant institutions and on previous research about the effects of employer mandates for health insurance on part-time employment in Section II. In Section III, we provide an overview of the theory of employer-mandated benefits as it applies to the ACA and explain why the employer mandate is likely to have larger part-time employment effects on the RAF industries than on other industries. In Section IV, we discuss the CPS data and trends in part-time work. We detail our methodology and present results from descriptive and regression analyses in Section V. In Section VI, we consider a variety of potential threats to our empirical approach and present robustness tests. In the concluding section, Section VII, we review our findings, place them in the context of related studies, and discuss implications for future research.

II. Background on Employer Mandates for Health Insurance and Prior Research

A. The ACA's Employer Mandate

The ACA, which was signed into law in March 2010, features a mandate that large employers provide full-time workers with affordable coverage, along with a health insurance mandate for individuals (repealed in 2017), an expansion of Medicaid, and subsidies for those purchasing coverage in the individual market. The ACA employer mandate requires that companies with 50 or more FTEs offer affordable coverage to their full-time employees or face a penalty. The act defines full-time employment as employment that averages 30 hours or more per week over a 12-month "look-back" period and affordable coverage as an insurance plan that pays for at least 60 percent of covered health care expenses and for which employees pay no more than 9.66 percent of family income for the coverage.

Although originally scheduled to take effect January 1, 2014, the employer mandate was delayed in various ways. In July 2013, the effective date of the employer mandate was postponed until January 1, 2015, and in February 2014, the mandate for employers with 50–99 FTEs was further delayed until 2016 (Kennedy 2014). As of January 1, 2015, employers with 100 or more FTEs were required to offer at least 70 percent of their full-time employees and their dependents health insurance. As of January 1, 2016, that share rose to 95 percent, and the employer mandate became effective for all employers with 50 or more FTEs.

Employers who fail to comply with the mandate potentially face stiff penalties. If an employer subject to the mandate fails to offer health insurance and if at least one employee purchases insurance through an exchange and receives a subsidy or tax credit, the employer is subject to a penalty equal to $2,160 multiplied by the number of full-time employees less 30 (less 80 in 2015). If an employer offers employees health insurance, but that insurance fails to pay for 60 percent of covered expenses or the premium exceeds 9.66 percent of family income, the employer may also be subject to a penalty. Specifically, if, under these circumstances, employees receive subsidized insurance in the exchanges, then the employer is subject to a penalty of $3,240 for each employee receiving a tax credit or subsidy after the first 30 (80 in 2015).[3]

B. The Hawaii Employer Mandate

Hawaii's Prepaid Health Care Act, passed in 1974, requires that all private-sector employers provide health insurance to employees working more than 20 hours per week.[4] Employers must pay at least 50 percent of the premium cost, and the employee contribution is limited to 1.5 percent of an employee's earnings. The Hawaiian mandate exempts coverage of those with very low monthly earnings—less than 87 times the minimum hourly wage, which amounts to earning the minimum wage and working 20 hours per week. Since the minimum wage applies to most workers and primarily exempts disabled people and full-time, non-college-students, very few workers are exempted from Hawaii's employer mandate. As with the ACA, employer-sponsored health insurance must cover minimum benefits, which include inpatient and emergency room hospital care, maternity care, and medical and surgical services.

Hawaiian employers that do not comply with the mandate face penalties and may be shut down (Buchmueller, DiNardo, and Valletta 2011). By requiring coverage of employees working as few as 20 hours per week (versus 30 in the ACA) and by capping

3. These penalty amounts and the maximum percent that employees can pay each year pertain to 2016. Because the penalties are indexed to inflation and the Internal Revenue Service sets the maximum percent that employees can pay each year, all amounts are subject to change.

4. Although Hawaii's health insurance mandate was passed in 1974, its legality was successfully challenged for self-insured employers on the grounds that the federal government regulates employer-sponsored benefit programs under the Employee Retirement and Income Security Act of 1974. In 1983, Congress granted a permanent exception for the Hawaii employer mandate.

employee premium contributions at 1.5 percent of earnings (compared to 9.66 percent of family income), Hawaii's mandate is considerably stronger than the mandate contained in the ACA. Reflecting its strong employer mandate, Hawaii has had much higher levels of employer-sponsored health insurance coverage than other states (Buchmueller, DiNardo, and Valletta 2011).[5] Thus, while other aspects of the ACA affect Hawaii, the ACA employer mandate should not be binding on Hawaiian employers and should have little, if any, effect on their behavior.[6]

Massachusetts is the only other state to have implemented a mandate that employers offer health insurance prior to the ACA. Unlike Hawaiian employers, however, many Massachusetts employers should still have been influenced by the ACA's mandate because the Massachusetts mandate was less stringent than the ACA's on key dimensions: it had a higher hours threshold, lower penalties, and a lower employer premium contribution for low-income workers. In addition, the Massachusetts mandate was repealed in 2013 in anticipation of the ACA taking effect.

C. Related Research on the Part-Time Effects of Health Insurance Mandates for Employers

Two studies have examined how employer mandates in state health insurance reforms have affected part-time employment. Buchmueller, DiNardo, and Valletta (2011) argue that as health insurance costs increased following the passage of Hawaii's employer mandate, employer incentives to pass along the costs of health insurance to workers or to circumvent the mandate by using exempt part-time workers have also grown. Using CPS data for the years 1979–2005, they find evidence that employers increasingly relied on low-hours part-time employees, particularly among workers most affected by the mandate. In a separate study, we estimate the impact of the Massachusetts health insurance reform on part-time employment. Using CPS data and a difference-in-differences strategy, we find that the Massachusetts reform resulted in a significant increase in part-time employment among low-educated Massachusetts workers (Dillender, Heinrich, and Houseman 2016).

In contrast to the studies on state-level reforms, evidence on the early effects of the employer mandate in the ACA on part-time employment has been mixed. Using CPS data, Moriya, Selden, and Simon (2016) examine how part-time employment changed after the ACA and find no evidence of an overall increase in part-time work. They find some evidence of increases in part-time work for low-educated workers, though they argue that the presence of preexisting trends makes this evidence inconclusive. Alternatively, Mathur, Slavov, and Strain (2016) use CPS data for 2008–2014 to examine the

5. In an analysis available upon request, we provide evidence that Hawaii's mandate is binding by using data from the Annual Social and Economic Supplement to the CPS to show that the incidence of employer-sponsored health insurance was high in Hawaii prior to the passage of the ACA. After accounting for job and demographic characteristics, we find that Hawaiian employees were 21 percentage points more likely to receive employer-sponsored coverage than the national average in the six years immediately prior to the ACA.

6. Although we do not intend to draw broad conclusions from anecdotal evidence, we find it reassuring that no Hawaiian employer is on *Investor's Business Daily* list of employers that have reported shifting workers to part-time work because of the ACA's employer mandate. This is consistent with the ACA's employer mandate having minimal effects on Hawaiian employers, while, alternatively, most of the rest of the nation (44 states and the District of Columbia) is represented on the list.

effects of the ACA mandate on part-time work. Using a difference-in-differences strategy, they test whether there was an increase in those working 25–29 hours compared to 31–35 hours in industries and occupations they deemed most likely to be affected by the mandate. They conclude that the ACA mandate had no effect on part-time work, although the coefficients in their models were imprecisely estimated.[7] By these studies' own accounts, an inability to account for time trends related to part-time employment hinders their ability to rule out definitively part-time employment effects of the ACA.

Even and Macpherson (2016) point out that occupations with low levels of employer-sponsored health insurance will be most affected by the employer mandate. Based on this observation, they use data from the CPS (1994–2014) and cross-occupation variation in pre-ACA employer-sponsored health insurance coverage levels to estimate the effect of the employer mandate on part-time employment.[8] They find that involuntary part-time employment has increased more in occupations with the largest share of workers affected by the ACA's employer mandate (an estimated 700,000 additional workers without a college degree) than in occupations with smaller shares of workers affected by the mandate.

The conflicting results of existing studies suggest the need for further research. Our contribution to this nascent literature is twofold. First, we use Hawaii as an already-treated control group, which means we do not have to rely on national trend breaks over time or on cross-occupation comparisons for identification. Second, we focus on the at-risk population, which increases the expected effect size and, in turn, statistical power for identifying any effects of the mandate.[9]

III. Potential Effects of the ACA on Part-Time Employment

A. Applying Theory on Mandated Benefits to the ACA

Although the ACA requires large employers to provide health insurance to employees averaging 30 or more hours per week, employers will not necessarily pay the cost of this benefit, instead shifting some or all of their health insurance costs onto workers. Summers (1989) argues that any mandated benefit will function like a tax at a rate equal to the difference between the employer's cost of providing the benefit and the employee's valuation of it. If employees fully value health insurance benefits, theory predicts that they will bear its costs through lower wages or other forms of compensation, a prediction that has some empirical support (Gruber 1994; Kolstad and Kowalski 2012).

7. Other studies examine employment effects of the ACA's Medicaid expansions or focus on employment changes in areas that had low insurance rates prior to the ACA. For examples, refer to Duggan, Goda, and Jackson (2017); Gooptu et al. (2016); Kaestner et al. (2015); Leung and Mas (2016); and Pinkovskiy (2015). These studies typically have not found that the ACA's Medicaid expansions increased part-time work.

8. In contrast to our approach, they exclude Hawaii and Massachusetts from their estimations, citing the earlier mandates in those states.

9. Identification not coming from other industries and occupations is important for two reasons. First, as other industries are still subject to the mandate, they could increase their use of part-time work as well even if doing so would be less attractive for them. Second, industries that are not as conducive to part-time work will likely not have the same time trend in part-time work as industries that are conducive to part-time work.

While some employers may extend health insurance and reduce wages for some workers, other employers may find offering health insurance to be suboptimal. If employees do not fully value health insurance benefits, employers will be unable to fully pass the health insurance costs onto employees in a competitive market.[10] Even if employees fully value the health insurance benefits, employers' ability to shift their costs onto workers in the form of lower wages may be constrained by minimum wages or union contracts. In addition, in a period of low inflation, as currently exists, employers may need to cut nominal wages to reduce real wages to cover the benefit cost; nominal wage cuts can have significant adverse consequences on worker morale and productivity.

Rather than absorbing health insurance costs or reducing real wages, employers may avoid the mandate, at least for some workers, by altering the way they staff positions. One way, and the empirical focus of our study, is by reducing average weekly hours below the 30-hour threshold for a larger share of the workforce. The likely effect of the mandate on a specific firm's use of part-time employment will vary based on several factors. One factor is the number of workers who are vulnerable to a change in employment arrangement because of the mandate—workers who average 30 or more hours a week and who are employed at firms with at least 50 FTEs, but whose employers did not previously offer them health insurance coverage. Employers with few workers newly covered by the mandate would be unlikely to make major changes to their use of part-time workers. Another factor is the cost of providing health insurance that complies with the mandate, which represents a higher increase in overall compensation in percentage terms for low-income workers. A third factor that likely influences the ACA's effect on part-time employment is industry-level characteristics that affect the ease of using part-time work.

Each of these factors points to the retail and the accommodation and food services industries as being likely to experience the largest impacts of the employer mandate on part-time employment. As Belman, Wolfson, and Nawakitphaitoon (2015) explain, the RAF industries contain disproportionately high shares of low-wage workers. As the vast majority of high-wage workers already had employer-sponsored health insurance prior to the ACA, the economic characteristics of workers in RAF industries suggest that these industries will have relatively high shares of workers affected by the employer mandate. Indeed, in earlier work, we find that while workers in the RAF industries accounted for only 20 percent of wage and salary employment, they accounted for about 42 percent of workers vulnerable to a change in staffing arrangements (Dillender, Heinrich, and Houseman 2015).

In addition to having a high share of low-wage workers affected by the mandate, the RAF industries also have characteristics that make part-time work a particularly attractive way to allocate labor. Firms in RAF industries typically experience fluctuations in customer demand during the day and week and have hours of operation that exceed the full-time working week. Firms in RAF industries thus use part-time staffing to meet peak-load demand, to enhance productivity by employing workers when their marginal productivity is the highest, and to avoid having to pay overtime for full-time workers (Künn-Nelen, De Grip, and Fouarge 2013; Valletta, Bengali, and van der List 2020).

10. Although few studies estimate how much people value health insurance, one by Finkelstein, Hendren, and Luttmer (2015) finds that Medicaid recipients in Oregon only value Medicaid at 20–40 percent of Medicaid's cost, suggesting that many low-income workers may not fully value health insurance.

Consistent with the attractiveness of part-time work in RAF industries, part-time work, including short-hours part-time work, is especially prevalent in the RAF sectors (Künn-Nelen, De Grip, and Fouarge 2013; Valletta, Bengali, and van der List 2020). Many firms in these industries have adopted variable scheduling practices that permit an expansion or contraction of employees' work hours.[11] Survey evidence from the National Longitudinal Survey of Youth shows that three out of the four occupations with the greatest variability in hours were core occupations in these sectors: retail workers, food service workers, and janitors and cleaners (Henly and Lambert 2015; Lambert, Fugiel, and Henly 2014). In other sectors, such as manufacturing, work is organized around full-time shifts. Thus, the incidence of part-time employment is low, and reducing weekly hours to avoid the mandate may be a costly strategy. If manufacturers choose to circumvent the mandate, they might be inclined to use other staffing strategies to do so.[12]

B. Timing of the Effects

The immediate concerns of many employers, the stops and starts of the mandate's implementation, and the 12-month look-back period make it unclear when one might begin to see effects of the employer mandate on part-time work. Although the mandate for employers with 100 or more FTEs was postponed to January 2015 several months before it was to take effect, implementation was originally scheduled for January 2014 and, with 2013 as a look-back period, one might expect to observe the effects of the mandate on part-time employment in 2013. Moreover, because it may be difficult for employers to reduce workers' hours permanently after they are hired, forward-looking firms could have opted to implement more short-hours part-time staffing as early as 2010.

In the years immediately after the ACA's passage, there was also considerable uncertainty about how the employer mandate would be implemented, which could have prompted firms to respond to the mandate before it became binding. In response to the many questions that it received about the employer mandate, in 2012 the Internal Revenue Service (IRS) published answers to frequently asked questions about the mandate. Questions about how full-time work would be defined were included in this list, and the IRS stated that it expected its rule to be that an employee would be considered full-time if the "employee is reasonably expected as of the time of hire to work an average of 30 or more hours per week on an annual basis."[13] According to Jost (2012), the IRS released notices in 2011 and 2012 with implementation rules that indicated that the IRS was "putting a great deal of thought into identifying and blocking stratagems that employers may use to try to evade the requirements of the mandate," which

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11. For example, many firms in these industries use algorithms to vary workers' hours from week to week according to demand. For evidence on such practices, see University of Chicago Work Scheduling Study reports and papers, available at http://ssascholars.uchicago.edu/work-scheduling-study/work-scheduling-study-papers (accessed December 27, 2021).

12. Use of temporary-help employment in manufacturing is high. Although it is beyond the scope of this work, the rise of temporary-help employment since the Great Recession could partly reflect a response to the ACA. Because their weekly hours, when averaged across periodic assignments, are relatively low, most temporary help workers would not be covered by the ACA's employer mandate (Dillender, Heinrich, and Houseman 2015).

13. The list of frequently asked questions can be accessed at https://www.irs.gov/pub/irs-drop/n-12-17.pdf (accessed December 27, 2021).

highlights the uncertainty about how the employer mandate would be implemented and indicates that the employer mandate was likely salient for firms well before it became binding. Immediately after the ACA was passed, benefit consultants began suggesting that firms consider reducing hours as a way to avoid having to offer health insurance coverage, and firms indicated that they were making staffing changes prior to the employer mandate being effective, largely because of uncertainty about how the employer mandate would be implemented (Muller, Isely, and Levin 2015; Singhal, Stueland, and Ungerman 2011).

In addition, an early reaction to the employer mandate may be especially likely because short-hours part-time employment was already high at the time of the ACA's passage because of the Great Recession, so the employer response to the mandate could take the form of little or no change to worker hours. Even in the RAF industries, adjusting workers' hours is not costless, as cutting already-established hours may lower morale, and reconfiguring established staffing patterns may disrupt business operations in the short run. Thus, rather than raising workers' hours and then reducing them again once the employer mandate was implemented, some employers might find it attractive for certain groups of workers to slow or halt the increase in weekly hours to pre-ACA levels while the economy gradually emerged from the Great Recession.[14] Because of the uncertainty over the timing of employers' responses to the mandate, we begin by estimating models with time-flexible coefficients of the effect of the ACA.

IV. Data Description and Trends in Part-Time Work

A. The Basic Monthly CPS

The CPS is a Bureau of Labor Statistics monthly household survey that covers around 60,000 households and is a primary source of data on employment and hours in the United States. In addition to collecting information on workers' demographic and job characteristics, the survey collects information on usual weekly hours worked in up to two jobs. If total usual hours are less than 35 (the BLS definition of part-time employment), workers are asked whether they work part-time for personal (voluntary) or economic (involuntary) reasons. The latter group consists of workers whose employers have reduced their work hours or who indicate that they cannot find full-time work. We classify individuals as part-time based on their usual hours worked in their main job.

Although data on weekly hours include the full range of possible integer values, responses tend to bunch at five-hour increments (for example, 20, 25, 30, 35, 40), suggesting that many respondents round their reported hours. In constructing measures of short-hours part-time employment, we therefore use two alternative cutoffs: work involving less than 30 hours per week in the main job (corresponding to the threshold that applies in the ACA mandate) and work involving 30 hours or less per week in the main job. The latter measure likely includes many who, in practice, work somewhat

14. Although employee turnover is high in RAF industries, some are likely to have a long tenure. Moreover, if firms set hours thresholds for groups of workers, at any point in time, most in the group would be affected by a decision to reduce average work hours.

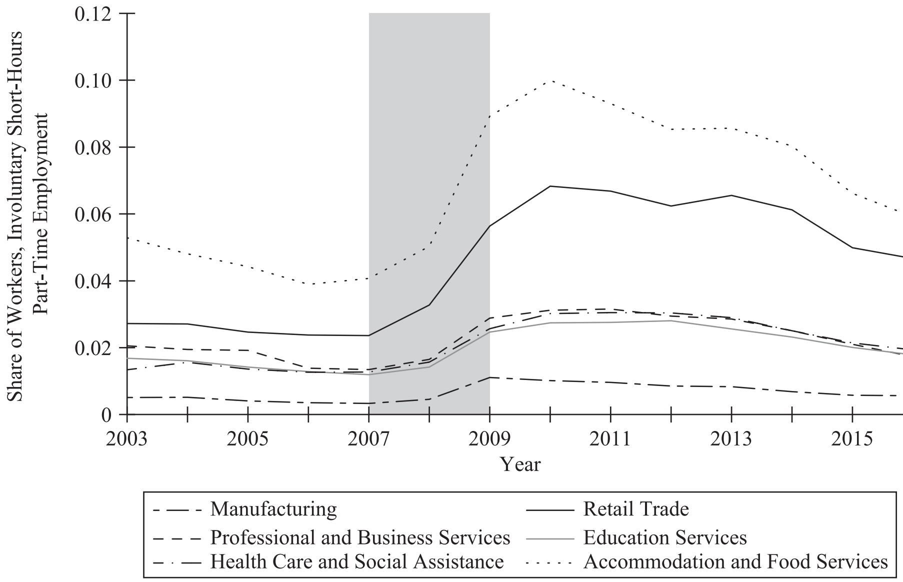

Figure 1

Share of Involuntary Short-Hours Part-Time Employment, Selected Sectors, 2003–2016

Source: Authors' calculations using data from the monthly CPS.
Notes: Gray shading indicates the years of the Great Recession.

below (or above) 30 hours. In addition, we estimate equations with the outcome variable indicating any short-hours part-time work and indicating only involuntary short-hours part-time work.[15]

Our main sample includes wage and salary workers aged 18–64; self-employed workers are excluded from the sample because the employer mandate does not apply to them. Our sample period covers 2003–2016. We use 2003 as the starting period because the CPS made major changes to its industry codes starting in 2003 and because, as Buchmueller, DiNardo, and Valletta (2011) argue, Hawaiian employers could have continued to adjust behavior to the mandate as health insurance premium costs kept rising through the 1990s.[16] Online Appendix Table 1 shows descriptive statistics for our sample separately for Hawaii and for the rest of the nation, both for the RAF sectors and for non-RAF

15. The question concerning the reason for part-time work is only asked of those who were at work during the survey week and who usually work fewer than 35 hours per week in up to two jobs. It is possible that an individual has her hours reduced below 30 hours per week on a main job and picks up a second job that boosts her hours to 35 or above. That individual would not be counted as involuntarily part-time on the main job. The share of individuals with second jobs is very small—about 5 percent. In addition, some workers report that their weekly hours are variable, and they are subsequently asked if their usual hours are part-time or full-time. We classify variable-hours workers who indicate they usually work part-time as short-hours part-time, although excluding them from our sample has no substantive effect on our results.

16. In results available from the authors on request, we have estimated models based on CPS data covering the 1994–2016 period. Comparability of industry definitions over time is less of an issue in the retail and hotel and accommodations than in other sectors. These estimates yield very similar results to ones presented in this paper.

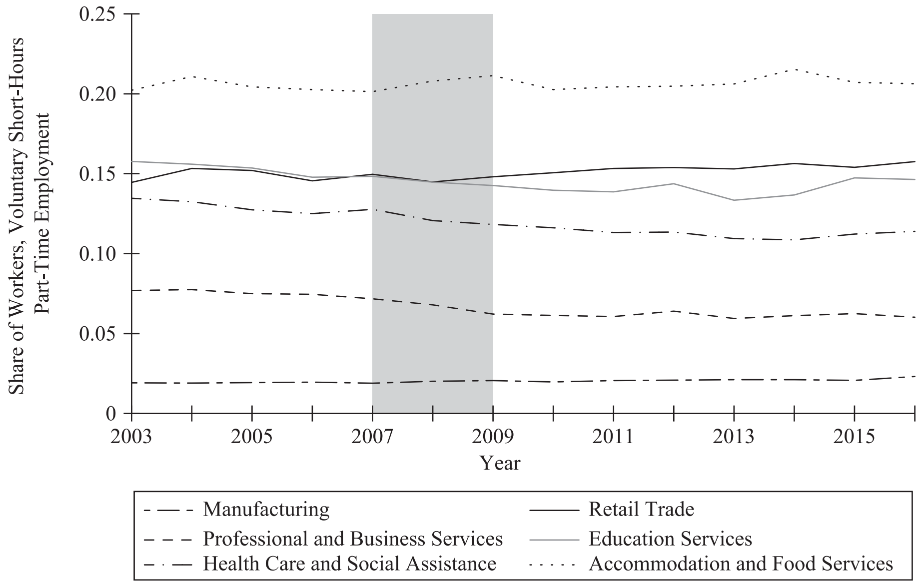

Figure 2

Share of Voluntary Short-Hours Part-Time Employment, Selected Sectors, 2003–2016

Source: Authors' calculations using data from the monthly CPS.
Notes: Gray shading indicates the years of the Great Recession.

sectors. The descriptive statistics indicate that, among the employed, a higher share work in the food and accommodation services industries and a lower share work in retail in Hawaii than in the rest of the nation. Hawaii also has a large Asian population.

B. Trends in Part-Time Work

Several observers have noted that, although the rate of involuntary part-time employment has fallen during the recovery period, it has remained higher than would be expected given the state of the macroeconomy, suggesting that structural factors help to explain why it has a higher incidence today than before the recession.[17] In addition, higher-than-normal involuntary part-time employment has been concentrated in sectors with high baseline rates of part-time employment (Dillender, Heinrich, and Houseman 2015; Robertson and Terry 2014).

Figures 1 and 2 depict the percentage of employees working fewer than 30 hours per week for economic reasons and the percentage working fewer than 30 hours per week

17. Valletta, Bengali, and van der List (2020) report evidence that supply- and demand-side factors related to demographic shifts and shifts in the industry composition of employment over time could account for the increase in involuntary part-time employment observed since the last recession. They acknowledge that cross-industry differences in the effects of the ACA could be related to the substantial role that industry employment shares play in their analysis.

voluntarily, respectively, for selected sectors for the years 2003–2016.[18] No cyclical or secular trends are evident with voluntary part-time employment. The fluctuation in the rate of part-time employment over the business cycle is due entirely to changes in the incidence of part-time employment for economic reasons. The rate of involuntary part-time employment jumped during the recession in 2008 and 2009, particularly in the RAF industries, and has remained stubbornly high.

V. The Impact of the Affordable Care Act on Part-Time Work

A. Estimating Equation and Identifying Assumptions

To estimate the effect of the employer mandate on part-time employment, we implement a difference-in-differences strategy that exploits the fact that, because of Hawaii's prior employer mandate, part-time work in Hawaii should not be affected by the ACA's employer mandate. In our main specification, we restrict our sample to individuals working in RAF industries. Our basic estimating strategy is represented by the following equation:

$$(1) \quad y_{ijst} = \alpha X_{ist} + unemp_{st}\lambda + postACA_t \times NotHI_s\beta + \gamma_t + \phi_s + \theta_j + \epsilon_{ist}.$$

Outcome y is an indicator variable measuring whether individual i working in sector j in state s and time t is in a short-hours part-time job. We estimate models in which the outcome measure includes all short-hours part-time employment, as well as models in which the outcome measure is restricted to part-time employment for economic reasons (that is, involuntary). The regression model controls for demographic and job characteristics (age, age-squared, gender, education, race and ethnicity, occupation), the state monthly unemployment rate, month–year and state fixed effects, and industry fixed effects. In specifications restricted to RAF industries, θ_j is an indicator that distinguishes between the two major sectors (retail or accommodations and food services). The period following the passage of the ACA is interacted with an indicator equal to one for all states other than Hawaii. The state fixed effects account for state differences that are fixed across time, while the month–year fixed effects account for national trends in part-time employment. The coefficient β captures differences between Hawaii and the other states in the probability of being part-time that did not exist prior to the act's passage. The baseline estimates come from estimating Equation 1 using linear probability models (LPM), though we verify the robustness of the results to estimating Equation 1 using probit and logit models. Standard errors are clustered at the state level.

Note that Equation 1 does not assume that part-time employment is the same in Hawaii as in the rest of the nation prior to the ACA, nor does it assume the absence of national trends in involuntary part-time work. Neither of these assumptions would be palatable. Like most states, Hawaii differs from the rest of the United States along many dimensions, as is evident in Online Appendix Table 1. Also, national economic forces, such as

18. The omitted sectors have very low rates of part-time and involuntary part-time employment and were dropped from the figures to improve their readability.

the Great Recession and the recovery that followed, have large effects on involuntary part-time employment during the period we study.

Instead, this approach relies on two weaker but critical assumptions. The first assumption is that the ACA's employer mandate should have virtually no effect on part-time employment in Hawaii. Given that Hawaii's mandate is much stricter than the ACA's and has been in existence for more than 40 years, we feel this assumption is plausible. If this assumption were to be violated, the estimates presented in here would be biased towards zero. The second critical assumption of the estimation strategy is that, controlling for the demographic and occupational characteristics of respondents and for economic conditions as measured by the state unemployment rate, part-time employment in the RAF industries in Hawaii would have trended similarly to part-time employment in these industries in the rest of the nation if not for the ACA. Although we have no reason to doubt this assumption either, its plausibility is not ensured by institutional factors, as is the plausibility of the first assumption, and assessing the direction of any potential bias is not as straightforward. Therefore, after presenting the baseline results, we further consider the validity of this assumption and possible ways to relax it.

B. Descriptive Analysis

To illustrate the identification strategy and as a first approach to considering the plausibility of the parallel-trends assumption, we show annual means of key variables separately for Hawaii and for the rest of the nation in Figure 3. Graph A of Figure 3 plots annual means of low-hours involuntary part-time work, defined as 30 or fewer weekly hours, in Hawaii and the rest of the nation, both for the RAF sector and for all non-RAF sectors. The graph shows that low-hours involuntary part-time employment in the RAF industries in Hawaii and the rest of the nation closely tracked each other during the economic expansion from 2003 to 2007, during the recession in 2008 and 2009, and in 2010, the first full year following the official end of the recession. However, in 2011, the first full year after the ACA's passage, low-hours involuntary part-time employment in the RAF industries in Hawaii and the rest of the nation began to diverge. For non-RAF industries, the annual share of workers in low-hours part-time employment for economic reasons rises somewhat more in Hawaii during the recession and correspondingly declines somewhat more after 2010, though for most of the post-recession period remains slightly higher in Hawaii. In contrast to the situation in the RAF sector, there is no clear break in trend in the rest of the United States relative to Hawaii after the ACA is passed.

Figure 3B shows annual unemployment rates for Hawaii and for the rest of nation from the Local Area Unemployment Statistics. As involuntary part-time employment is procyclical, a differential effect of the Great Recession on unemployment or a differential recovery for Hawaii could hinder our empirical approach. Panel B shows that the unemployment rate is lower in Hawaii than in the rest of the United States for all years of our sample. Panel B also indicates that the unemployment rate rose for both Hawaii and for the rest of the nation during the Great Recession but has steadily returned to pre-recession means. Thus, while there are level differences between Hawaii and the rest of the nation in their unemployment rates, there is no evidence of differential trends over time.

Figure 3C shows the share of workers employed in RAF industries separately for Hawaii and the rest of the nation. Differential changes in the RAF employment shares

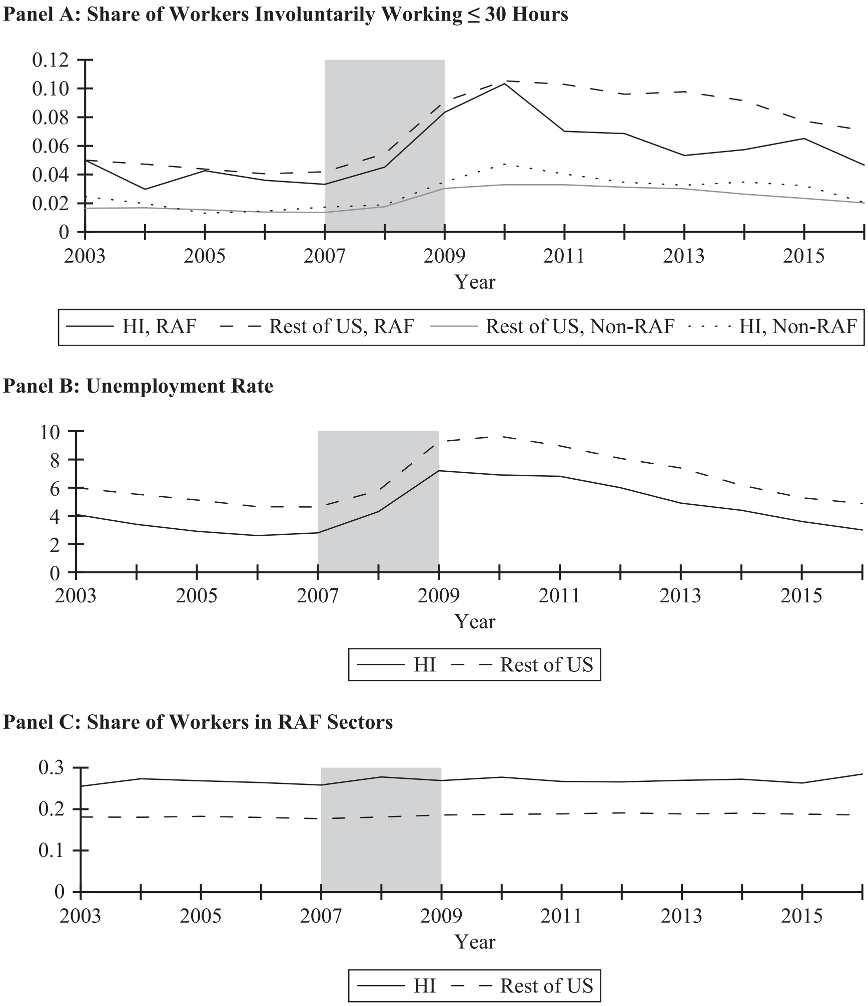

Figure 3

Trends in Key Economic Variables in Hawaii and Rest of the United States

Sources: Panels A and C, authors' calculations using data from the monthly CPS; Panel B, BLS Local Area Unemployment Statistics.

Notes: Gray shading indicates the years of the Great Recession.

between Hawaii and the rest of the nation could be indicative of broader changes to the RAF industry in Hawaii. However, while Hawaii has a higher share of RAF workers than the rest of the United States, the shares for both groups remain steady over time and do not appear to change with the Great Recession or with the passage of the ACA, indicating that other major trends in RAF industries should not be a concern for the analysis.

Overall, this descriptive analysis highlights that there are initial differences between Hawaii and the rest of the nation in the unemployment rate and in the share of workers in RAF industries. However, the analysis also suggests that most outcomes trend in parallel for the entire sample period. The main exception is involuntary part-time employment, particularly for RAF industries. While Hawaii and the rest of the nation have similar trends in involuntary part-time employment prior to the ACA being passed, involuntary part-time employment increased in the RAF sectors in the rest of the United States relative to Hawaii immediately after the ACA was passed. These patterns support the assumption that relevant outcomes would have trended in parallel in Hawaii and the rest of the nation absent the ACA and suggest that the ACA increased part-time employment in the RAF sectors in the rest of the United States relative to Hawaii.

C. Difference-in-Differences Estimates

We next present estimates for variants of Equation 1. As explained in Section III, it is unclear when the employer mandate might begin to affect part-time work, so we start our analysis by setting *postACA*NotHI* in Equation 1 to be a vector of 13 time-varying indicator variables equal to one for all non-Hawaii states for each year, 2004–2016; we restrict our sample to individuals working in retail, accommodations, or food services. The coefficient on each of these interaction variables can be interpreted as how the difference in the outcome variable between Hawaii and the rest of the nation changed in the indicated year relative to the difference in 2003. Figure 4 graphs the estimates for the four main dependent variables. The point estimates and their standard errors are shown in Online Appendix Table 2. The coefficient estimates on the *non-Hawaii–year* interaction terms become large and positive beginning in 2011, indicating that employers began making adjustments ahead of the implementation of the employer mandate, which Pinkovskiy (2015) and Mathur, Slavov, and Strain (2016) also argue is a possibility. In all specifications, tests of joint significance suggest that the probability of being in a short-hours part-time job and of being involuntarily in a short-hours part-time job increased in the rest of the United States relative to Hawaii after the passage of the ACA.

The estimates are similar regardless of whether the dependent variable specifically references being involuntarily part-time, indicating that most of the effect is coming from involuntary part-time employment and so is unambiguously demand driven.[19] If employers increase their demand for short-hours part-time workers in response to the mandate, voluntary part-time employment could also increase by raising the likelihood that workers who want part-time employment can find it. The extent to which increased demand for part-time workers raises involuntary or voluntary part-time employment depends on the relative supply and demand of part-time workers in the economy. Our

19. Estimates from a model in which the dependent variable is specified as voluntary short-hours part-time employment (not reported in Online Appendix Table 2) confirm that coefficient estimates on the non-Hawaiian interaction term during the post-ACA period are not jointly statistically significant. *P*-values are 0.73 and 0.48 for voluntary part-time employment less than 30 hours and for voluntary part-time employment of 30 hours or less, respectively.

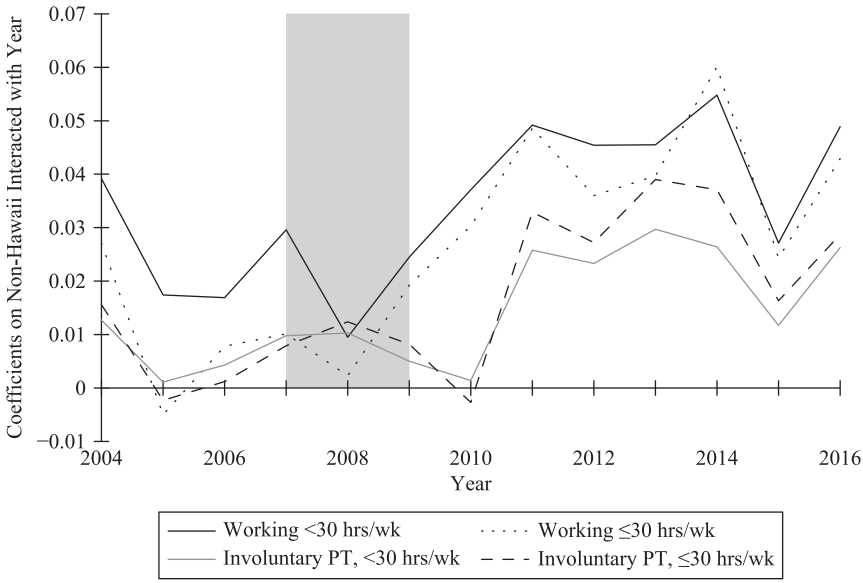

Figure 4

Coefficients on Non-Hawaii Interacted with Year, RAF Industries

Source: Authors' calculations using data from the monthly CPS.
Notes: Gray shading indicates the years of the Great Recession.

finding that the effect of the ACA mandate is largely manifested in involuntary short-hours part-time employment is consistent with Valletta, Bengali, and van der List (2020), who argue that shifts in the demographic and industry composition of the workforce have reduced the supply of part-time workers relative to the demand for part-time workers.

To compute the average effect of the ACA since 2011, we replace the series of non-Hawaii–year interactions with a single indicator equal to one for all states other than Hawaii from 2011 to 2016. Because of concerns that standard errors may be understated in difference-in-differences specifications when the number of treated and untreated states is unbalanced, in addition to reporting standard errors, we estimate a series of placebo tests to assess the statistical significance of our estimates.[20] Specifically, we reestimate Equation 1 50 times, substituting one of the other 49 states or the District of Columbia for Hawaii as the control for the rest of the United States, which gives us a distribution of placebo ACA coefficients. For each model in which the coefficient on β obtained from using Hawaii as the control group is positive, we report the percentage of placebo β coefficients that is larger; in cases where the estimated coefficient

20. See Conley and Taber (2013) and Buchmueller, DiNardo, and Valletta (2011) for discussions of potential biases to standard errors when there are few treated (or in our case control) states. Our approach of generating *p*-values based on a set of placebo estimates follows that in Buchmueller, DiNardo, and Valletta (2011) and Dillender, Heinrich, and Houseman (2016).

Table 1
Effects of ACA on Part-Time Employment: Hawaii Relative to Rest of United States

	Working <30 hrs/wk	Working ≤30 hrs/wk	Involuntary PT, <30 hrs/wk	Involuntary PT, ≤30 hrs/wk
Panel A: The Effect on Retail and Accommodation and Food Services				
Non-Hawaii*post-ACA	0.0232***	0.0304***	0.0182***	0.0250***
	(0.002)	(0.002)	(0.001)	(0.001)
Fraction of placebo estimates larger than when Hawaii is control	0.12	0.10	0.06	0.00
n	1,583,936	1,583,936	1,583,936	1,583,936
Panel B: The Effect on All Other Industries				
Non-Hawaii*post-ACA	–0.0017*	–0.0028**	–0.0009	–0.0001
	(0.001)	(0.001)	(0.001)	(0.001)
Fraction of placebo estimates smaller than when Hawaii is control	0.26	0.28	0.24	0.26
n	7,102,471	7,102,471	7,102,471	7,102,471
Panel C: The Effect on All Industries				
Non-Hawaii*post-ACA	0.0027**	0.0032**	0.0029***	0.0048***
	(0.001)	(0.001)	(0.001)	(0.001)
Fraction of placebo estimates smaller than when Hawaii is control	0.46	0.40	0.28	0.18
n	8,686,407	8,686,407	8,686,407	8,686,407

Notes: The data come from the 2003–2016 monthly CPS. Each column represents a separate regression with the indicated dependent variable, estimated using a linear probability model. For each model, we report the coefficient estimate on the interaction of non-Hawaii states with the post-ACA period (2011–2016). All models also control for month–year and state fixed effects, demographic and job characteristics (age, age-squared, gender, race-ethnicity, education, occupation, industry), and state–month–year unemployment rate. The sample includes wage and salary workers aged 18–64. Panel A restricts the sample to workers in the retail and accommodations and food services industries. Panel B restricts the sample to workers in all other industries. Panel C includes workers from all industries. All regressions are weighted by the CPS final weights, and standard errors are clustered by state. Significance: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

on β obtained from using Hawaii as the control group is negative, we report the percentage of placebo β coefficients that is smaller. We would be concerned about our ability to accurately reject the null hypothesis of no effect of the ACA if many other states yielded estimates of the ACA effect that were larger than the estimated effect from using Hawaii as the control group.

Panel A of Table 1 gives the estimates of the average effect of the ACA on part-time employment in RAF industries since 2011 along with the results of the placebo analysis. The estimates range from a 1.82 percentage point increase in the probability of involuntary part-time less than 30 hours per week to a 3.04 percentage point increase in the probability of working less than or equal to 30 hours per week. As with the time-flexible estimates displayed in Figure 4, the coefficient estimates of the average effect of the ACA are almost as large when the dependent variable measures short-hours involuntary part-time employment as when the dependent variable is any short-hours part-time employment, which suggests that most of the change in part-time employment is coming from the demand side. The placebo analysis, which shows that the overwhelming majority of states experienced an increase in short-hours part-time employment in RAF industries compared to Hawaii, is reassuring. When the dependent variable is working less than 30 hours per week (or working 30 hours or less), six (and five) placebo estimates, respectively, are larger than the estimate when Hawaii is the control state. The placebo tests are somewhat stronger when the dependent variables are the involuntary counterparts—three and zero coefficients, respectively, are larger than the estimates when Hawaii is the control state.

To place these estimates in perspective and to assess their reasonableness, we consider the implied number of workers affected by the mandate and the size of that population relative to the number of workers in RAF industries who were potentially vulnerable to a reduction in their hours when the ACA was enacted. In earlier work, we estimated that about 2.9 million workers in RAF industries were vulnerable to hours reductions, which we defined as employees of firms with at least 50 employees who averaged at least 30 hours per week in their job but who were not offered health insurance (Dillender, Heinrich, and Houseman 2015). An increase of 1.82 percentage points in the share of workers in RAF industries who are involuntarily part-time with less than 30 hours per week implies about a half million affected workers. The somewhat larger estimate of 2.5 percentage points, which includes workers involuntarily part-time and reporting working 30 or fewer hours, translates into almost 700,000 workers.[21] Expressed as a percent of those vulnerable to hours reductions, our estimates suggest that 17 percent, and perhaps as high as 24 percent, of the vulnerable population in RAF industries became involuntarily employed in short-hours part-time jobs. While our prior estimates of the vulnerable population, and thus these calculations, should be taken as rough approximations, they nonetheless suggest that a sizable minority of that population was affected.

In Online Appendix Table 3, we examine where in the weekly-hours-worked distribution the increase in short-hours part-time employment comes from by estimating regressions for RAF workers in which the dependent variables are indicators for working less than 20 hours per week, working 21–30 hours per week, working 31–40 hours per week, and working more than 40 hours per week. The results suggest that most of the increase in part-time employment comes from an increase in people working 21–30 hours per week and a decrease in people working 31–40 hours per week. However, we cannot rule out that the employer mandate has affected hours across all parts of the distribution.

Panel B of Table 1 presents estimates analogous to those in Panel A for all other industries. As explained earlier, we focus on the RAF industries because they have the most workers at risk of being affected by the employer mandate and because they have

21. We use national employment estimates from the Quarterly Census of Employment and Wages (QCEW) program for RAF industries for the years 2011–2016 and subtract out employment estimates in RAF industries for Hawaii. (National Current Employment Statistics estimates are benchmarked to QCEW figures.)

characteristics that make part-time staffing attractive. Although other industries are subject to the mandate, they are not as conducive to expanding the use of part-time work and prior to the mandate had a smaller share of employees working at least 30 hours per week who were not offered health insurance. The point estimates in Panel B are all small, and each estimate has many placebo estimates with larger magnitudes. Although we cannot rule out that the ACA's employer mandate has had some effect on part-time employment in non-RAF industries, we do not find evidence that it has led to major increases in these industries.

Panel C of Table 1 presents estimates for all industries. As the ACA increased part-time employment for RAF industries, the effect on all industries should be positive as well if the ACA did not decrease part-time employment in non-RAF industries. However, as we have discussed elsewhere, the analysis may be underpowered for the full sample of all industries. The estimates in Panel C are positive and provide suggestive evidence that the employer mandate has increased part-time employment for all industries. However, note that while the estimates are statistically significant when assessed using standard errors clustered at the state level, many placebo estimates are larger than the ACA estimates.

VI. Robustness and Heterogeneity

A. Robustness Tests

Our main specifications estimate Equation 1 using a linear probability model. While the LPM has advantages over nonlinear binary choice models in terms of ease of estimation and interpretation, applying LPMs to outcomes with means near zero or one may produce biased estimates.[22] In Table 2, we consider the robustness of the RAF results to estimating Equation 1 using probit and logit models. Table 2 shows the average marginal effects from probit and logit specifications and indicates that the estimates are generally very similar to the main estimates. Online Appendix Tables 4 and 5 show estimates from probit and logit models for the full set of results in Table 1 and verifies that the probit and logit estimates are similar to the full set of LPM estimates presented in Table 1.

The main specification in Table 1 calculated the average effect of the employer mandate since 2011, which combines changes in part-time employment since the employer mandate was implemented with changes in part-time employment since the ACA was enacted but before the employer mandate was implemented. The rationale for estimating this specification is that Figure 4 suggests that employers reacted to the mandate even before it was implemented. An alternative approach, which we implement in Panel A of Table 3, is to estimate models with separate interaction terms for the enactment and implementation periods.

As would be expected based on the analysis shown in Figure 4, the point estimates in Panel A are positive for both the enactment and implementation interactions for all part-

22. Refer to Ai and Norton (2003), Greene (2010), and Puhani (2012) for discussions about interpreting interaction terms in nonlinear models and to Horrace and Oaxaca (2006) for discussions of bias in the LPM. Angrist and Pischke (2009), however, argue that the differences between the LPM and nonlinear binary choice models are likely inconsequential when estimating average marginal effects.

Table 2

Effect of the Employer Mandate on Part-Time Employment in Retail and Accommodations and Food Services: Robustness to Probit and Logit Models

	Working <30 hrs/wk	Working ≤30 hrs/wk	Involuntary PT, <30 hrs/wk	Involuntary PT, ≤30 hrs/wk
Panel A: Average Marginal Effects from Probit Models				
Non-Hawaii*post-ACA	0.0224***	0.0290***	0.0171***	0.0233***
	(0.002)	(0.002)	(0.001)	(0.001)
Fraction of placebo estimates larger than when Hawaii is control	0.12	0.10	0.06	0.04
Panel B: Average Marginal Effects from Logit Models				
Non-Hawaii*post-ACA	0.0214***	0.0285***	0.0169***	0.0232***
	(0.002)	(0.002)	(0.001)	(0.001)
Fraction of placebo estimates larger than when Hawaii is control	0.14	0.12	0.08	0.04
n (for both panels)	1,583,936	1,583,936	1,583,936	1,583,936

Notes: The data come from the 2003–2016 monthly CPS. Each column represents a separate regression with the indicated dependent variable. The sample includes all wage and salary workers aged 18–64 in the retail and accommodation and food services sectors. For each model, we report estimates of the average marginal effect on the interaction of non-Hawaii states with the post ACA period (2011–2016). All models also control for month–year and state fixed effects, demographic and job characteristics (age, age-squared, gender, race-ethnicity, education, occupation, industry), and state–month–year unemployment rate. All regressions are weighted by the CPS final weights, and standard errors are clustered by state. Significance: $*p<0.10$, $**p<0.05$, $***p<0.01$.

time definitions. These results indicate that even if the estimation does not use any variation from the enactment period to identify the effect of the employer mandate, the estimated effect of the mandate on part-time employment is qualitatively similar to the original estimates in Panel A of Table 1. In three out of the four specifications, the point estimates on the implementation interaction are larger than the point estimates on the enactment interaction. For the involuntary part-time variables, they are 11 percent larger (0.0153 vs. 0.0138) and 17 percent larger (0.0211 vs. 0.0180) though the implementation and enactment coefficients are not statistically different in any of the models.

Arguably the strongest assumption that the baseline analysis makes is that, controlling for state unemployment and workers' demographic and job characteristics, part-time employment in RAF industries in Hawaii would have paralleled part-time employment in the rest of the nation if not for the ACA's employer mandate. Even though the pre-trend analysis displayed in Figure 3 suggests no reason to doubt this assumption, we next implement several alternative approaches to relax this assumption and to test its validity, given how vital the assumption is for the analysis.

The period studied includes the Great Recession, and one concern is that the Great Recession could drive our findings if its effect on Hawaiian RAF industries substantially

Table 3
Effect of the Employer Mandate on Part-Time Employment in Retail and Accommodations and Food Services: Robustness to Estimating Separate Enactment and Implementation Effects and to Alternative Approaches to Accounting for Economic Conditions

	Working <30 hrs/wk	Working ≤30 hrs/wk	Involuntary PT, <30 hrs/wk	Involuntary PT, ≤30 hrs/wk
Panel A: Separate Enactment and Implementation Effects				
Non-Hawaii*2010–2013	0.0245***	0.0298***	0.0138***	0.0180***
	(0.002)	(0.002)	(0.002)	(0.002)
Fraction of placebo estimates larger than when Hawaii is control	0.06	0.08	0.10	0.08
Non-Hawaii*2014–2016	0.0240***	0.0336***	0.0153***	0.0211***
	(0.002)	(0.003)	(0.001)	(0.002)
Fraction of placebo estimates larger than when Hawaii is control	0.20	0.18	0.18	0.12
Panel B: Including Hawaii*Great Recession Interaction				
Non-Hawaii*post-ACA	0.0209***	0.0288***	0.0191***	0.0265***
	(0.002)	(0.002)	(0.001)	(0.001)
Fraction of placebo estimates larger than when Hawaii is control	0.16	0.12	0.04	0.00
Panel C: Separate Unemployment Effect for Each State				
Non-Hawaii*post-ACA	0.0225***	0.0293***	0.0183***	0.0249***
	(0.002)	(0.002)	(0.001)	(0.001)
Fraction of placebo estimates larger than when Hawaii is control	0.14	0.10	0.04	0.02
Panel D: Including Control for Indexed State GDP				
Non-Hawaii*post-ACA	0.0231***	0.0303***	0.0181***	0.0249***
	(0.002)	(0.002)	(0.001)	(0.001)
Fraction of placebo estimates larger than when Hawaii is control	0.14	0.08	0.04	0.00

(*continued*)

Table 3 (*continued*)

	Working <30 hrs/wk	Working ≤30 hrs/wk	Involuntary PT, <30 hrs/wk	Involuntary PT, ≤30 hrs/wk
Panel E: Including Control for Indexed State GDP and Indexed State RAF GDP				
Non-Hawaii*post-ACA	0.0300***	0.0382***	0.0201***	0.0276***
	(0.003)	(0.004)	(0.001)	(0.002)
Fraction of placebo estimates larger than when Hawaii is control	0.06	0.04	0.00	0.00
n (for all panels)	1,583,936	1,583,936	1,583,936	1,583,936

Notes: The data come from the 2003–2016 monthly CPS. The sample includes all wage and salary workers aged 18–64 in the retail and accommodation and food services sectors. Each column in Panel A displays two coefficients from a single regression with the indicated dependent variable. Each column in Panels B–E is the non-Hawaii*post-ACA interaction from a separate regression with the indicated dependent variable. All models include month–year and state fixed effects, controls for demographic and job characteristics (age, age-squared, gender, race-ethnicity, education, occupation, industry), and state–month–year unemployment rate. Models in Panel B also control for an interaction of the Great Recession and Hawaii. Models in Panel C include separate month–year-state unemployment controls for each state. Models in Panel D include an index of state real GDP. Models in Panel E include an index for state real GDP and an index for state real GDP in the RAF sector. All regressions are weighted by the CPS final weights, and standard errors are clustered by state. Significance: *$p < 0.10$, **$p < 0.05$, ***$p < 0.01$.

differed from its effect on RAF industries elsewhere in the United States. One way to account for the Great Recession is to exclude variation from the Great Recession in identifying the ACA effect. To remove variation from the Great Recession, we supplement Equation 1 with an indicator variable equal to one for Hawaii during the Great Recession, which the National Bureau of Economic Research dates as lasting from December 2007 to June 2009. This revised specification means that any differences in part-time work between Hawaii and the rest of the nation that arise during the Great Recession will not be used to identify the coefficient on the Hawaii fixed effect, which captures initial differences between Hawaii and the rest of the nation. The results from this revised approach are displayed in Panel B of Table 3 and are similar to those from the baseline analysis.

Another way to account for the Great Recession is to include additional controls to account for economic conditions. Since involuntary part-time employment is cyclical, a possible threat to the empirical strategy would be if Hawaii experienced a faster or more complete recovery from the Great Recession, which is why the main specification controlled for the unemployment rate. That said, the main specification does not account for the possibility that the unemployment rate has a larger effect on involuntary part-time employment in Hawaii than in other states, which might be the case, as involuntary part-time employment is higher relative to the unemployment rate in Hawaii than in the rest of the United States (see Figure 3). To address this concern, Panel C of Table 3 presents results that include a vector of state unemployment rates interacted with a vector of state indicator variables, thus allowing the effect of the unemployment rate to vary by state. This specification accounts for the possibility that involuntary part-time employment rises by a different amount in response to an increase in unemployment in Hawaii. The results in Panel C are similar to the main estimates and provide more evidence that cyclical factors do not confound the analysis.

Panel D further considers the possibility that a differential effect of or recovery from the Great Recession is a threat to the identification by including a control for an index of inflation-adjusted annual state GDP. A dramatic change in results with this additional control might suggest that the unemployment rate control does not adequately capture the influence of the Great Recession. However, the results are very similar to the main specification.

A related concern is that RAF industries in Hawaii may have experienced a state-specific and industry-specific shock that is not captured by our main specification. To consider this possibility, the results in Panel E supplement the models in Panel D with an index of annual state real GDP in the RAF sector. Again, the results are very similar, suggesting that Hawaii did not experience an RAF-specific economic shock that drives our results.

While Table 3 considers ways to overcome possible unobserved shocks the Hawaiian RAF industries may have experienced that coincided with the passage of the ACA, Table 4 displays results from analyses that test for possible confounding economic shocks more directly. To do so, we estimate difference-in-differences models over the 2003–2016 period in which the dependent variable is an indicator of the individual's employment status—employed or employed in the RAF sectors:

$$(2) \quad y_{ist} = \alpha X_{ist} + postACA_t \times NotHI_s \beta + \gamma_t + \phi_s + \epsilon_{ist}$$

Control variables include a vector of demographic characteristics (age, age-squared, gender, race, ethnicity, and education), month–year fixed effects, state fixed effects, and the interaction of the post-ACA period with treated (not Hawaii) states. A negative ACA effect from this regression would indicate that Hawaii, or Hawaii's RAF sector, experienced an economic improvement relative to the rest of the United States that coincided with the ACA's passage.[23] Given that part-time work is counter-cyclical, Hawaii experiencing an economic improvement relative to the United States that coincides with the ACA would cast doubt on the ACA being responsible for the differential changes in part-time work that Hawaii and the rest of the nation experienced beginning in 2011. In contrast, when the dependent variable is any employment, a positive ACA effect would be consistent with Hawaii experiencing weaker economic growth relative to the rest of the United States. For regressions where the dependent variable is employment in the RAF sector, a positive ACA effect on employment, coupled with our earlier estimates showing a negative ACA effect on average hours in RAF, would be consistent with employers hiring more workers to compensate for lower weekly hours induced by the employer mandate.

In the first model, reported in Column 1 of Table 4, the sample includes all CPS respondents ages 18–64 (including nonemployed), and the dependent variable is an

23. The regressions reported in Table 4 do not control for industry or occupation because those who are not employed do not have an industry or occupation and employment in RAF industries is the dependent variable for two of the three specifications. They also do not include the state's unemployment rate, which is mechanically related to the dependent variable, the state's employment rate.

Table 4
Effect of the Employer Mandate on Total Employment

	Employed	Employed in RAF	
Non-Hawaii*post-ACA	0.0030	0.0040	0.0068
	(0.004)	(0.004)	(0.005)
Fraction of placebo estimates larger than when Hawaii is control	0.26	0.22	0.32
Sample	All respondents	All respondents	All workers
n	13,863,934	13,863,934	8,686,407

Notes: The data come from the 2003–2016 monthly CPS. Each column represents a separate regression with the indicated dependent variable. For each model, we report the coefficient estimate on the interaction of non-Hawaii states with the post ACA period (2011–2016). All models include month–year and state fixed effects and controls for demographic characteristics (age, age-squared, gender, race-ethnicity, education). The sample includes people ages 18–64; further restrictions are noted in the table. All regressions are weighted by the CPS final weights, and standard errors are clustered on individuals. Significance: $^{*}p<0.10$, $^{**}p<0.05$, $^{***}p<0.01$.

indicator equal to one if the individual is employed. The point estimate on the ACA effect is positive, though statistically indistinguishable from zero, and 26 percent of the placebo estimates are larger than the estimate obtained from using Hawaii as the control group. In Columns 2 and 3 of Table 4, the dependent variable is an indicator variable equal to one if the individual works in an RAF industry. The sample for the estimates in Column 2 includes the nonemployed, while the sample in Column 3 includes only employed individuals. In both cases, the estimates are small, positive, and statistically insignificant using standard errors clustered at the state level; the estimates also are smaller than a large share of the placebo estimates (22 percent and 32 percent in Columns 2 and 3, respectively).

In sum, the results presented in Table 4 provide no evidence that, following ACA's passage, stronger economic growth in Hawaii as a whole or in Hawaii's RAF industries can explain our findings. If anything, they suggest that some RAF employers may have increased employment in conjunction with lowering average hours, though our results are not definitive.

B. Heterogeneity

We next consider two potential sources of heterogeneous effects of the employer mandate by drawing on information that is available for subsamples of workers in the basic monthly CPS. As previously discussed, we would expect the effect of the employer mandate on part-time employment to be larger for low-wage workers, so the first source of heterogeneity that we consider is based on workers' hourly wages. However, as the employer mandate also has the potential to affect worker compensation, readers should keep in mind that the employer mandate may alter the composition of low-wage workers.

To consider heterogeneity based on hourly wages, we use hourly wage data from the CPS earner study, which covers the subset of CPS respondents who belong to the Outgoing Rotation Group (ORG). Respondents in the ORG are asked about their hourly wages (if they are paid hourly), their usual weekly earnings (if they are salaried), and their usual hours worked. For salaried workers, we impute hourly wages using the earnings and hours variables. We focus on workers who earn an hourly wage of at least \$6.[24] Note that restricting the sample to only ORG respondents reduces the sample size of RAF workers to 363,347 observations.

Panel A.1 of Table 5 shows the estimated effect of the ACA on part-time employment for RAF workers earning less than \$10 per hour, while Panel A.2 shows the estimated effect of the ACA on part-time employment for workers earning at least \$10 per hour. As expected, the point estimates of the effect of the employer mandate on part-time employment are much larger for the low-wage sample. The estimated effects range from 0.0415 to 0.0811. Few of the placebo estimates are larger than the ACA estimates, and none are larger when the dependent variables incorporate working part-time involuntarily. In contrast, when the sample is restricted to people earning at least \$10 per hour, the point estimates become much smaller, and many of the placebo estimates are larger than the ACA estimates.

We next test for evidence of heterogeneous effects of the employer mandate based on RAF workers' firm sizes. As the employer mandate only applies to firms with at least 50 FTE employees, implementing the empirical strategy using only workers at small firms provides a falsification test for the empirical strategy. To test for heterogeneous effects based on firm size, we draw on the March CPS's Annual Social and Economic (ASEC) Supplement, which asks respondents to report the number of people who work at the firm of the job that they held the longest in the previous year.

Before discussing the results, we point out a few caveats to using the ASEC's firm-size information. First, because the ASEC's firm-size categories are not consistent over time, we estimate heterogeneous effects for respondents at firms of fewer than ten employees and for respondents at firms of at least 100 employees. Second, the firm-size variable may be misreported often, either because individuals do not know their firm's size or because individuals incorrectly report establishment size instead of firm size (Callahan and Williams 2013). Third, the sample size of ASEC respondents is much smaller than the sample sizes in the basic monthly CPS and in the ORG subsample. Finally, while the ASEC's firm-size and industry information pertain to the individual's longest-held job in the previous year, the ASEC's questions about hours worked in the previous year pertain to all jobs and, for those who usually worked part-time, do not distinguish whether the individual was working part-time voluntarily or for economic reasons.[25]

The results from using the firm-size information in the ASEC Supplement are shown in Panel B of Table 5. Panel B.1 shows the results for people who report working at RAF firms with fewer than ten employees. For both dependent variables, the point estimates of the effect of the ACA are statistically indistinguishable from zero for small firms

24. Hawaii's lowest minimum wage during our study period is \$6.25, so the small number of workers earning less than \$6 an hour in Hawaii either are exempt from the minimum wage or misreport their hours or earnings.
25. Instead, the March CPS asks people who worked less than 35 hours in *any* week in the previous year why they worked part-time.

Table 5
Effect of the Employer Mandate on Part-Time Employment by Wage Level and Firm Size

	Working <30 hrs/wk	Working ≤30 hrs/wk	Involuntary PT, <30 hrs/wk	Involuntary PT, ≤30 hrs/wk
Panel A: Heterogeneity by Hourly Wage from ORG CPS				
A.1. Hourly Wage <$10				
Non-Hawaii*post-ACA	0.0608***	0.0811***	0.0415***	0.0644***
	(0.003)	(0.004)	(0.003)	(0.003)
Fraction of placebo estimates larger than when Hawaii is control	0.04	0.02	0.00	0.00
n	141,959	141,959	141,959	141,959
A.2. Hourly Wage ≥$10				
Non-Hawaii*post-ACA	0.0078**	0.0073*	0.0083***	0.0110***
	(0.003)	(0.004)	(0.002)	(0.002)
Fraction of placebo estimates larger than when Hawaii is control	0.32	0.40	0.18	0.14
n	221,388	221,388	221,388	221,388
	B.1. Firm Size <10		B.2. Firm Size ≥100	
	Working <30 hrs/wk	Working ≤30 hrs/wk	Working <30 hrs/wk	Working ≤30 hrs/wk
Panel B: Heterogeneity by Firm Size from ASEC CPS				
Non-Hawaii*post-ACA	–0.0042	–0.0047	0.0152***	0.0264***
	(0.006)	(0.007)	(0.004)	(0.005)
Fraction of placebo estimates larger than when Hawaii is control	0.40	0.48	0.24	0.16
n	24,945	24,945	121,863	121,863

Notes: The data in Panel A come from the ORG of the 2003–2016 CPS. The data in Panel B come from the ASEC Supplement of the 2003–2016 March CPS. Each column represents a separate regression with the indicated dependent variable. For each model, we report the coefficient estimate on the interaction of non-Hawaii states with the post ACA period (2011–2016). All models also control for month–year and state fixed effects, demographic and job characteristics (age, age-squared, gender, race-ethnicity, education, occupation, industry), and state–month–year unemployment rate. The sample in Panel A includes all wage and salary workers aged 18–64, while the sample in Panel B includes all wage and salary workers aged 18–64 in the retail and accommodation and food services sectors. All regressions are weighted by the CPS weights for the ORG or ASEC Supplement, and standard errors are clustered by state. Significance: $^{*}p<0.10$, $^{**}p<0.05$, $^{***}p<0.01$.

when assessed using confidence intervals clustered at the state level, and many of the placebo estimates are larger than the ACA estimates.

When we focus on people who work at RAF firms with at least 100 employees in Panel B.2, the point estimates of the effect of the ACA on part-time employment are positive. The point estimate of the effect on usually working less than 30 hours per week is 0.0152, while the point estimate of the effect on usually working 30 hours per week or less is 0.0264. These estimates are statistically different from zero when significance is assessed using standard errors clustered at the state level, though many of the placebo estimates are larger than the ACA estimates. As such, we view this analysis as providing suggestive rather than conclusive evidence of heterogeneous effects based on firm size.

VII. Conclusion

The ACA's passage in 2010 led to considerable speculation that there would be a widespread reduction in weekly scheduled hours below the 30-hour threshold by employers seeking to avoid the employer mandate. Those predictions, however, ignore the fact that most workers who were not receiving offers of health insurance coverage from their employer prior to the ACA were not covered by the employer mandate. Estimates suggest that, at the time of the act's passage, at most 5 percent of wage and salary workers would be vulnerable to a reduction in hours (Dillender, Heinrich, and Houseman 2015; Glied and Solis-Roman 2014). Moreover, we emphasize that although compliance with the mandate could result in a substantial percentage increase in compensation for some—particularly low-wage workers—employers would have to weigh any increased costs associated with offering health insurance against the costs of reorganizing work around short-hours schedules. In addition, employers wishing to circumvent the mandate have other options, such as using more contract, temporary, and on-call workers for certain tasks.

We focus our study on the retail and accommodation and food service industries, which have large concentrations of workers affected by the mandate, as well as characteristics that make part-time employment particularly attractive. By exploiting the fact that the employer mandate was not binding in Hawaii to implement a difference-in-differences strategy, we find strong evidence that the ACA increased short-hours part-time employment in these industries. We estimate that short-hours part-time employment increased by 1.8 to 3.0 percentage points, and that almost all of the increase was the result of an increase in involuntary part-time employment, indicating that employer demand, not worker supply, was the cause of the growth in short-hours part-time employment. These estimates indicate that 17–24 percent of at-risk RAF workers became involuntarily employed part-time, which is a sizable share of a low-wage population that is often the focus of labor and social safety net policies. We find no evidence of increases of part-time work in other industries.

Although our results indicate that the ACA has increased part-time work for a vulnerable population, they suggest that fears of a part-time nation were overblown. In total, our estimates indicate an increase in involuntary, short-hours part-time employment of 500,000–700,000 workers, which represented 0.4–0.5 percent of the aggregate wage and salary workforce during our period of study.

Our results contrast with those of most existing studies of the effects of the ACA on part-time employment. This, we suspect, is because, in aggregate, the effects will be small as a percentage of the workforce and because most prior studies are insufficiently powered or targeted to detect effects in sectors where they are likely concentrated. A notable exception is the study of Even and Macpherson (2016), who compare how involuntary part-time employment changed across occupations after the passage of the ACA. Their finding that the ACA has shifted roughly 700,000 workers to involuntary part-time work is at the upper range of our estimates, but they estimate that only about 250,000 were in the retail and accommodation and food services industries. Without relying on the assumption that occupations are good control groups for each other, we find that the increase in involuntary part-time work is concentrated in industries that are particularly conducive to part-time work. Our estimates of the part-time-employment effect of the ACA's employer mandate are smaller than the estimates of the part-time-employment effect of Hawaii's mandate from Buchmueller, DiNardo, and Valletta (2011), which is consistent with the fact that the ACA's mandate is not as strong as Hawaii's mandate.

Readers should keep in mind that our study identifies causal effects resulting from employers' initial reactions to the mandate and that these effects could change over time. During much of the period of our analysis, the regulations governing the mandate were being developed, and there was considerable confusion among employers over their costs. The period was also marked by strong opposition to the mandate among many employers, which may have been associated with some exaggeration of potential costs. As employers learn about the actual cost implications of the mandate in their circumstances, behavior may change. In addition, it is important to recognize that the effects of the mandate on part-time employment are not independent of economic conditions. In times when the economy is improving and labor markets are tightening, more employers may find it in their interests to offer longer work hours and health insurance benefits. Increases in inflation could also increase the probability that employers offer workers health insurance, by making it easier for them to pass along premium costs in the form of lower real wages without cutting nominal wages. There is a need, therefore, for future studies of the effects of the ACA mandate on part-time employment, as well as on other staffing arrangements, under varying economic conditions, which have not been addressed in this or other research.

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