The Affordable Care Act and the Growth of Involuntary Part-Time Employment*

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Abstract

This study tests whether the employer mandate under the Affordable Care Act (ACA) increased involuntary part-time (IPT) employment. Using data from the Current Population Survey between 1994 and 2015, we find that IPT employment in 2015 was higher than predicted based on economic conditions and the composition of jobs and workers in the labor market. More importantly, using difference-in-difference methods, we find that the increase in the probability of IPT employment since passage of the ACA has been greatest in the occupations with a larger share of workers affected by the mandate. Our estimates suggest that up to 700,000 additional workers without a college degree between the ages of 19 and 64 are in IPT employment as a result of the ACA employer mandate.

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1. Introduction.

At its passage, the 2010 Affordable Care Act (ACA) required that firms with 50 or more employees provide health insurance for their full-time workers or be subjected to penalties beginning in 2014. Many analysts argued that the law created incentives for large firms to shift from full-time to part-time workers to escape the penalties and the cost of providing health insurance.

While there have been numerous media reports of firms shifting to part-time work because of the ACA, there are only a small number of studies that use nationally representative data to determine whether the effects are quantitatively or statistically significant. While several researchers claim that the share of workers engaged in part-time work rose sharply during the Great Recession and remains above pre-recession levels, there is disagreement about why the number remains so high – particularly for the share that are employed part-time involuntarily (i.e., those who have part-time jobs but who would prefer full-time jobs). Some believe that the threat of the ACA reduced employers' willingness to hire full-time workers during the recovery. Others believe that the high level of part-time work is an aftermath of the recession and structural changes in the labor market unrelated to the ACA.

This paper attempts to sort out the ACA effects on involuntary part-time (IPT) employment from cyclical or other structural changes in the labor market. To isolate the effect of the ACA, we use Current Population Survey (CPS) data between 1994 and 2015 to examine factors that may have contributed to the growth in IPT in recent years. Our empirical analysis shows that IPT employment is significantly higher in 2015 than one would forecast given economic conditions and the occupational, industrial, and demographic structure of the labor market. More importantly, we show that the rate of growth in IPT employment since passage of the ACA has been greatest in the occupations where the ACA mandate binds the most. We also provide evidence that the effect is most pronounced in low wage occupations where employers would have the greatest difficulty shifting the cost of the mandated health insurance coverage to their workers. The analysis provides several tests for the robustness of the results and considers alternative explanations for the patterns observed in the data. Overall, we believe that there is compelling evidence that the ACA employer mandate has contributed to an increase in involuntary part-time work in the United States.

While other research has examined the impact of the ACA or other state-specific employer mandates on part-time work, our study goes beyond the earlier work. Earlier studies use difference-in-difference methods that compare the rate of growth in part-time work across states based on whether they had passed a state-specific employer mandate (e.g., Dillender, Heinrich, and Houseman, 2016a, b). Ours is the first study to use a difference-in-difference method comparing the rate of growth in part-time work across occupations based on the extent to which the mandate is likely to be binding. With this approach, we can avoid relying on the identifying assumption that part-time employment trends would be similar across states had the ACA not been passed. Of course, we are forced to rely upon a different identifying assumption – namely, controlling for economic conditions, trends in part-time employment would be identical across occupations without passage of the ACA. It is important to note that we control for a wide range of economic conditions as well as structural changes in the economy that could have differentially affected labor demand across occupations. We also allow the demographic and cyclical determinants of part-time employment to differ across occupation.

2. Background.

Because of non-discrimination laws in the provision of fringe benefits that have existed since 1978, if an employer has a self-insured medical plan, the plan must be offered to all full-time employees who are classified as non-highly compensated. Because of this rule, some employers have used part-time workers or independent contractors to avoid the costs of providing health insurance for low wage workers (Houseman 2001). It is important to note, however, that these rules only applied to employers with self-insured medical plans and, as of 2015, only 39 percent of private sector establishments had a self-insured health plan (Fronstin 2016) though the percentage is greater at larger firms. Consequently, many employers could exclude full-time employees from their health insurance plans. Moreover, there was no penalty for offering no health insurance whatsoever.

The passage of the Affordable Care Act changed the rules considerably by imposing penalties on employers who do not offer health insurance to all full-time workers. To understand how the rules changed, we provide a brief review of the penalties and mandate.

With initial passage of the ACA in 2010, "applicable large employers" (ALEs) would be assessed an "Employer Shared Responsibility" (ESR) penalty if they did not provide "minimal essential coverage" for their full-time employees and dependents at an "affordable price." A firm is considered an "applicable large employer" (ALE) if, based upon employment in the prior year, it employed an average of at least 50 full-time equivalent employees (FTEs). Any worker averaging at least 30 hours per week (or 130 hours per month) counts as one FTE. For part-time workers averaging less than 30 hours per week, total weekly hours of all part-time workers are divided by 30 to calculate FTEs (or monthly hours by 130). For example, if an employer has 40 employees who each work more than 30 hours per week and 60 part-time employees who each average 10 hours per week, total FTEs would be 40 + (600/30) = 60 FTEs.[1] Any employer with less than 50 FTEs is not an ALE and is exempt from ESR penalties.

An employer's insurance plan provides minimal essential coverage if the plan covers at least 60 percent of expected health care costs. The plan is deemed affordable if the employee's cost of coverage does not exceed 9.5% of the employee's household income. Since employers may not know their worker's household income, they can instead rely on one of three safe harbors for determining the maximum cost for affordable coverage.[2]

If any of an ALE's full-time employees purchase health insurance in the private market and receive a federal premium subsidy, the employer may be subject to an ESR penalty. To be eligible for a premium subsidy, a worker's income must be less than 400% of the federal poverty level, and the worker cannot have access to minimal essential health insurance coverage that is affordable through the workplace. Given the federal poverty lines in 2014, workers with one household member would be eligible for a subsidy if they earned less than $46,680 annually. A worker from a household with four members would be eligible if the household earned less than $95,400.

If a full-time employee of an ALE receives a federal subsidy, the employer could be assessed an ESR penalty. The size of the penalty depends on whether the firm provides affordable and minimal essential coverage to at least 95% of its full-time workers. If the firm meets this requirement, the penalty is $3,000 per full-time worker receiving a subsidy, but no

[1] There is a seasonal worker exception exempting employers that average more than 50 full-time workers for 120 days or less during a year.

[2] The three safe-harbors define affordability based on either the worker's wages reported on form W2, the worker's rate of pay at the beginning of the period, or the federal poverty line for a single individual in the relevant calendar year.

more than $2,000 times the number of full-time employees less 30. If an ALE provides coverage to fewer than 95% of its full-time workers, its penalty will be $2,000 per full-time employee (minus up to 30).

When the initial ESR penalties were scheduled to begin in the 2014 calendar year, the number of employees and their full-time status were determined by the number of employees and their work hours in 2013.[3] As a result, if a firm wished to reduce exposure to ESR penalties that started in 2014, it would have to make adjustments during 2013.

After initial passage, some of the implementation dates for the employer penalties and the workers with mandatory coverage were adjusted. Table 1 summarizes some of the key dates. In July 2013, the administration announced that the implementation date for employer penalties would be pushed back to 2015. In February 2014, the implementation date for mid-sized employers (50-99 employees) was pushed back to 2016. Transition relief was also provided for large employers (100+) by reducing the minimum coverage rate to 70 percent of employees in 2015 and deferring the implementation of the 95 percent coverage rate until 2016. Although the implementation date was pushed back, some employers reported making adjustments to avoid the penalties detailed in the new law as early as 2012. For example, Investor's Business Daily collected reports of workers' hours being reduced to avoid the ACA penalties from over 450 employers with some reports filed as early as 2011.

While transition relief pushed the implementation of the ESR payments back to 2015, large firms that made the adjustments in the first half of 2013 may have kept them in place knowing that with implementation in 2015, exposure to ESR penalties would be determined by 2014 employment levels. The mid-sized firms (50-99) may have been slower to adjust since they would not be subject to ESR payments until 2016.

There are a few ways that an ALE could reduce exposure to ESR penalties. First, the firm could provide affordable coverage to its full-time workers. According to data provided by the Kaiser Family Foundation, in 2012 the average premium for employer-based health

[3] To determine the number of employees and their classification as a full- or part-time worker, each firm had to define three time periods: the measurement period, and the subsequent administrative and stability periods. To determine whether a firm was an "applicable large employer" in 2014, it would use a measurement period of between three and 12 months in the prior calendar year to determine which employees were full-time and part-time. Beginning in 2014, any continuing worker who was classified as a full-time worker during the measurement period must be treated as a full-time worker for the "stability period." In subsequent years, the firm could define a measurement period followed by an "administrative period," which would allow the firm to determine those workers classified as full-time for the subsequent stability period.

insurance in the United States was $5,384.[4] For a full-time worker (2,000 hours per year), this is equivalent to an hourly wage increase of $2.69. The cost would be higher for older workers and would vary across the states.

Firms increasing coverage for their full-time workers could try to shift the cost to the newly covered workers. If the firm is to avoid ESR penalties, however, the firm cannot require a worker eligible for a subsidy (i.e., those with household income below 400 percent of the poverty line) to contribute more than 9.5% of household income for the insurance. Thus, for example, if a full-time worker (2,000 hours annually), earned $10 per hour ($20,000 per year) and had no other household income, the firm could not require that the worker contribute more than $1,900 for the health insurance. On the other hand, if the worker earned $25 per hour, the firm could require the worker to contribute $4,750 per year for the plan and shift most of the cost to the employee. Since employers are less able to shift the cost of adding coverage to low-wage workers, they will have a greater incentive to find ways to avoid providing coverage for them.

If an ALE is unable to shift the cost of the health insurance to the worker by requiring that the employee pay a large share of the premiums, it could instead cut the worker's hourly wage. For example, if a worker previously earned $10 per hour and received no health insurance, rather than require the worker to pay the equivalent of $2 per hour for the health insurance (which would violate the rule on affordability), the firm could cut the worker's wage by $2 per hour and require no employee contribution. This strategy is not feasible if the wage cut would put the worker's hourly wage below the minimum wage. Cutting the hourly rate would also limit a firm's ability to attract low-wage workers who place minimal value on the health insurance.

In sum, providing health insurance coverage is costly and firms have incentives to shift the cost of the new coverage to workers by requiring employee contributions or cutting wages. Since such strategies are least effective for low-wage workers, employers are motivated to find alternative means to avoid the coverage requirement for low-wage employees. Several such strategies are available. First, firms with close to 50 FTE employees could try to keep employment below that threshold and escape the mandate entirely. Employers could accomplish this by reducing the number of full-time employees while increasing hours worked per employee. For example, a firm with 40 employees who average 50 hours per week would

[4] http://kff.org/other/state-indicator/single-coverage/

be exempt whereas a firm with 50 employees who average 40 hours per week would not be exempt. Alternatively, a firm could keep its FTEs below the limit of 50 by hiring independent contractors or hiring temporary workers.

Second, a firm could reduce the size of the penalty and the cost of providing affordable coverage by shifting from full-time to part-time (<30 hours) workers. Failing to provide part-time workers with coverage does not result in a penalty. Also, because a firm can compute a worker's full-time status by averaging hours worked over as many as 12 months, hiring short-term employees would reduce the number of workers that must be provided coverage.

While the threat of ESR penalties created incentives for large firms to switch to part-time employment, in a 2014 report, the Congressional Budget Office (CBO) emphasized the impact of the health insurance subsidies on the supply side of the labor market. Mulligan (2014) provides a detailed analysis of the subsidy formula and concludes that, with the health insurance subsidies, millions of workers could yield more disposable income with a part-time than a full-time schedule. In fact, the Congressional Budget Office (2014) estimated that the ACA subsidies will reduce the number of full-time-equivalent (FTE) employees by about 2.0 million by 2017. The CBO projects that virtually all of the reduction will be due to workers wanting to reduce hours or drop out of the labor market entirely because of the new health insurance subsidies.

When there is an increase in part-time employment from the supply side, the expectation is that the number of workers who are working part-time voluntarily would rise. On the other hand, if there is an increase in the demand for part-time workers (perhaps due to the ACA mandate), some workers with full-time jobs may be forced to accept a part-time job involuntarily. As Ehrenberg and Smith (1988) note, a smaller number of workers who prefer voluntary part-time employment could increase the number who are forced into part-time work. As an example, if the number of workers who prefer part-time work decreases, employers may respond by cutting back on the hours of full-time workers. For example, teenagers have been a major source of part-time workers, and as teenage employment rates have dropped, some employers may cut hours for workers who prefer full-time jobs.

Several studies examine recent trends in part-time employment with a focus on differential trends in voluntary (VPT) and involuntary part-time (IPT) employment. The Bureau of Labor Statistics defines IPT workers as those who work part-time, prefer a full-time job, but are unable to find a full-time job because of economic conditions. VPT workers have part-time jobs but do not want a full-time job, or are unable to take a full-time job because of personal reasons (e.g., schooling, family responsibilities, phased retirement).

Valletta and Bengali (2013) show that VPT work has been trending downward over the past few decades and was largely unchanged during the Great Recession. On the other hand, Valletta and List (2015) point out that IPT employment rose sharply during the Great Recession and dropped slowly since the recovery. Their analysis suggests that cyclical factors played an important role in the rapid rise in IPT work during the Great Recession, but structural factors have also been important. Structural factors that may have led to increases in IPT employment include employment shifts toward service industries that make greater use of part-time workers (e.g., wholesale trade and leisure and hospitality sectors) as well as demographic changes in the labor market resulting in fewer workers preferring part-time work. The high level of IPT work since the Great Recession is also documented by Cajner, Mawhirter, Nekarda, and Ratner (2014), and Canon et al. (2014). These studies also show that the growth of the service industry contributed to increases in IPT employment.[5]

Overall, there is a good deal of evidence that IPT employment remains unexpectedly high after the Great Recession. The existing research attributes the persistently high levels of IPT employment to a combination of cyclical and structural factors. While numerous media reports document how employers are switching to part-time workers to avoid the pending ACA penalties, there are only a few empirical studies that specifically address the role of the ACA.

Garrett and Kaestner (2015) use 2000-2014 Current Population Survey data to test whether the health insurance subsidies and expansion of Medicaid had negative supply side effects. Their empirical analysis finds virtually no adverse effect of either the Medicaid expansion or ACA policies on labor force participation, employment, or usual hours worked per week. They do, however, report a modest increase in part-time employment (less than 30 hours per week).

Mathur, Slavov and Strain (2015) use 2008-2014 CPS data to examine the effect of the ACA mandate on part-time work without any separate analysis of the effects on VPT and IPT work. Their analysis examines the effect of the mandate on the ratio of people working 25-29 hours versus 31-35 hours. Their estimated effect of the mandate is based on a difference-in-

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[5] Yellen (2014) also discusses the potential role of a rising share of employment in the service sector as a potential explanation for increases in part-time employment.

difference methodology that compares the change in the ratio after 2010 for a treatment and a control group. They consider three treatment groups defined on the basis of occupation, industry, or wage and find no evidence that the ratio of workers in the 25-29 versus 31-35 hours group grew faster for the treatment groups. They conclude that the ACA did not cause employers to switch to part-time work. Their study differs from ours in several respects. First, every worker is defined as either "treated" or "untreated" based on either their occupation, industry, or wage. Second, they define the "post-treatment" period as 2010-2014. Since the ACA was not passed until 2010 and the mandates didn't officially go into effect for several years, their definition of the post-treatment period may contribute to an underestimate of the impact of the ACA on part-time work.

A 2014 survey of employers by the ADP Research Institute found that 38 percent planned to adjust worker hours in response to the ACA employer mandate.[6] Despite this fact, the authors found no change between 2013 and 2014 in the distribution of hours worked. The authors did not, however, control for changes in economic conditions that may have offset the effect of the ACA on the average level of part-time work. They were also unable to focus on the types of workers that were most likely to be impacted by the mandate – low skill workers at large firms that did not previously have health insurance coverage.

Dillender, Heinrich and Houseman (2016a) use a difference-in-difference approach to estimate the effect of the ACA on part-time work. Their analysis treats Hawaii as a control group since the state had previously passed legislation that would make passage of the ACA employer mandate unimportant. Consistent with expectations, the results suggest that involuntary part-time employment grew at a slower rate in Hawaii than in the other states, particularly in retail, accommodations, and food services – industries with a significant share of workers who are low skill and where part-time work has historically been common. They estimate that the ACA led to an increase in involuntary part-time employment of between one-half and one million workers in retail, accommodations, and food services. This is the first study that uses a nationally representative data set and finds an effect of the ACA mandate on part-time work.

Dillender, Heinrich and Houseman (2016b) examine the effect of the 2007 employer mandate in Massachusetts on part-time employment using a difference-in-difference approach

[6] See ADP Research Institute (2015).

comparing changes in part-time employment in Massachusetts to other states. They find that the Massachusetts employer mandate increased part-time employment among workers without a college degree. They find similar results for a variety of different control groups.

Overall, the extant literature finds that part-time work (particularly involuntary) remains higher than predicted after the Great Recession. There is some evidence that the ACA mandate increased involuntary part-time work using nationally representative data, but it relies critically on the assumption that Hawaii is an appropriate control group.

Our research differs from the earlier work by testing whether the growth in IPT across occupations varies systematically with the share of workers affected by the mandate. Also, we show that the effects have been largest where theory predicts – among low wage workers. While the earlier studies have merit, we believe that this alternative approach provides more identifying variation allowing a more precise estimate of the impact of the ACA.

3. Data and Empirical Methods.

To investigate the effects of the ACA on IPT employment, we use data from the Current Population Survey (CPS). The monthly CPS has the requisite information on hours worked, earnings, and worker characteristics, but it does not routinely collect information on firm size or health insurance coverage. To obtain this information, we use the March Annual Demographic Supplements to the CPS which provide firm size and health insurance information.

Unfortunately, prior to 2011, the firm size categories in the CPS do not allow us to precisely distinguish firms above and below 50 employees. As a result, we are forced to rely on a cut-off of 100 employees for a time-consistent definition of the firms that are most likely to be affected by the ACA. Even this definition is not perfect because the March data on firm size is based on the number of employees rather than the number of FTEs as defined by the ACA. The primary hypothesis we wish to test is that the threat of the ESR penalties established by the ACA caused large employers to shift from full-time to part-time workers and increased the share of workers with IPT jobs. The challenge is sorting out the effect of the business cycle or structural changes in the labor market from the ACA.

To examine the trend in part-time work, we use the monthly CPS data to compute the annual average of the percentage of workers aged 19-64 that are in VPT and IPT employment.

We exclude workers aged 65 and over because they are eligible for Medicare and not eligible for a health insurance subsidy. We exclude 16-18 year olds since many of them are in school and would not work enough weeks in the year to be classified as a full-time worker by the ACA. We also restrict our sample to wage and salary workers since self-employed workers are not directly affected by the mandate. Finally, we exclude workers residing in Massachusetts and Hawaii because these states had employer mandates prior to passage of the ACA. We also exclude workers from the District of Columbia due to a lack of data on economic conditions.

Consistent with the Bureau of Labor Statistics (BLS) definition of IPT (or "part-time for economic reasons"), we define IPT workers as those who are working less than 30 hours per week, want a full-time job, but are working part-time because of either slack work or unfavorable business conditions, an inability to find full-time work, or seasonal declines in demand.[7]

Figure 1 presents annual estimates of the share of workers engaged in VPT and IPT employment between 1994 and 2015. Separate estimates are provided for workers with less than 16 years of education and those with at least 16 years of education. The estimates make a few points clear. First, the share of workers engaged in VPT work has been relatively stable over the past 20 years and increased only slightly during the two recessions since 2000. The share in IPT employment is more volatile and rose during both recessions since 2000, especially during the Great Recession and particularly for the less educated work force. Finally, while IPT employment has fallen below the record levels seen in 2011, it remains above any level experienced between 1994 and the start of the Great Recession for both education groups.

As discussed previously, it is possible that cyclical or structural factors could be responsible for the persistence of unusually high levels of IPT since the Great Recession. To investigate this possibility, we estimate linear probability models of both IPT and VPT employment controlling for economic conditions as well as individual worker characteristics and the industries and occupations of employment. The controls for state economic conditions and demographic characteristics are monthly in frequency and include the state unemployment rate and its 6, 12, 18, and 24 month lags, the square of the state unemployment rate, the state-specific

[7] It is important to note, however, that BLS tabulations of part-time work are based on a work week of less than 35 hours whereas we use a cut-off of 30 hours since that is the relevant cut-off for the ACA employer mandate. In 2014, reducing the cut-off for IPT employment from 35 to 30 hours per week reduces the number of IPT workers by approximately one-third.

coincident economic indicator and its 6, 12, and 18 month lags, the ratio of the state minimum wage to the 25th percentile wage, and the teen share of the state labor force.[8] We also control for the worker's industry and occupation,[9] month of year, and a wide range of personal characteristics that influence a worker's preference for part-time employment.[10] We estimate the regressions separately for education groups with and without a college degree.[11] Rather than choose a single year for a reference period, we force the sum of the year effects to equal zero and include a constant term in all the regressions. This effectively makes the intercept capture the average intercept over the 1994-2015 time period and the coefficients on the individual year dummies represent deviations from the 1994-2015 average.[12]

The results for IPT and VPT are presented in figures 2 and 3. Since we only report the year regression coefficients, a brief summary of the other results is in order. Economic conditions (unemployment rates, coincident index, minimum wage ratio, and teen labor force share) are jointly statistically significant (at the .001 level) on both VPT and IPT.[13] Also, there is statistically significant variation in the probability of VPT and IPT across groups formed by age, education, race, marital status, industry, and occupation. Hence, changes in the state of the economy, or the demographic, occupational or industrial composition of the labor market contribute to changes in both IPT and VPT.

Figure 2 shows how the probability of IPT has varied over time. The unadjusted plot shows how the probability has varied over time without controlling for changes in labor market conditions or the demographics of the workers. As in figure 1, it shows that the probability of IPT rose sharply during the great recession and has fallen since. The plot of the adjusted year effects reveal that for the less educated group of workers, controlling for economic conditions

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[8] With the exception of the teen share of the labor force and the occupation-specific unemployment rate which is measured annually, all of the controls for economic conditions are measured monthly. The minimum wage to 25th percentile wage ratio is calculated separately for the low and high education groups. The coincident economic indicators are from the Federal Reserve Bank of Philadelphia: http://www.phil.frb.org/research-and-data/regional-economy/indexes/coincident/

[9] The models include controls for 13 major occupations and 22 major industries.

[10] The personal characteristics include state of residence; age (5 categories); race (3 categories); female, marital status; female interacted with marital status; education; and female interacted with number of children aged 0 to 5 and aged 6 to 17.

[11] In an earlier version of the paper, we split the sample according to whether workers had a high-school degree. Our overall estimates of the change in part-time employment were similar to those reported here, but subsequent analysis revealed that the group of workers with some college was more similar in terms of part-time work trends than those in the group of workers with at least a bachelor's degree.

[12] See Suits (1984) for a discussion of this approach.

[13] Unless noted otherwise, for the remainder of this paper, the term "statistically significant" implies that the effect is significant at the .05 level of significance.

and the demographics of the work force, the probability of IPT was significantly above (at the .05 level) the 1994-2015 average for the first time beginning in 2013. For the more educated workers (16+ years of education), the adjusted probability of IPT has been relatively flat since 2010 and the adjusted probability is never significant above the average for the reference period.

Figure 3 presents the results for VPT. While the unadjusted probability of VPT has been relatively flat over time, the adjusted probability has been dropping steadily over time for both the high and low education groups. This suggests that some factor that we are unable to control for is VPT to rise over time. Given that this trend was apparent from 2000-2010, we conclude that it is not driven by the ACA.

To add some perspective on the size of the IPT employment effects in 2015, estimates from the CPS imply that 2.3 million of the 77.6 million workers without a college degree represented by our restricted sample are engaged in IPT employment. In this group, the level of IPT employment is approximately 358,000 higher than expected relative to the average 1994-2015 level. In other words, about one of six workers engaged in IPT employment in 2015 is the result of IPT being at higher than expected levels relative to the reference period.[14]

While IPT employment since 2013 was significantly higher than predicted based on economic conditions and labor market structure, it does not necessarily indicate that the ACA is responsible. Other structural changes in the economy that are not controlled for in our model might explain an increase in employer demand for part-time workers or a decrease in the supply of workers wanting part-time jobs.

To assess whether the ACA is an important reason for the rise in IPT employment, we estimate a model that allows the growth in the probability of IPT employment to vary depending on the extent to which a particular job is likely to be impacted by the ACA employer mandate.[15] In particular, we test whether the increase in IPT employment since the onset of the ACA is greatest in the occupations where the share of workers that would be affected by the mandate is greatest. To test whether this is the case, we first use March CPS data from 2003-2007 to

[14] For college graduates, approximately 520,000 of the 43.2 million workers represented by our restricted sample are engaged in IPT employment. IPT employment is approximately 50,000 higher than expected relative to the reference period but this difference is not statistically significant at the .05 level.

[15] Card (1992) uses a similar approach to estimate the effect of the April 1990 increase in the minimum wage on the employment of teenagers. The size of the treatment effect varied across states according to the fraction of teens earning less than the new minimum wage.

estimate the share of workers in each industry most likely to be affected by the ACA penalties.[16] For each 3-digit occupation, we calculate the proportion of workers who are employed at firms with 100 or more workers, work 30 or more hours per week, and do not have health insurance.[17] We choose the 2002-2006 sample period to exclude the effects of the Great Recession on IPT work and to get an estimate of the percent affected under more normal economic conditions.[18] The affected share is computed separately for our low and high education groupings.

Figure 4 presents histograms of the proportion of workers affected across 3-digit occupations by education group. In the low education group, 16.3 percent of the workers are likely to be affected by the ACA employer mandate (i.e., employed at a large firm, working 30 or more hours per week, and without employer provided health insurance). In the high education group, 11.4 percent are identified as potentially affected. While the percent affected is higher for those without a college degree, a surprisingly large share of those with a college degree are in the potentially affected group. The histogram illustrates substantial variation in the percent affected across the 3-digit occupations. It is this variation in the affected share that we use to identify the effect of the ACA mandate.

Among the less educated workers, a sample of occupations with the affected share in the top 5 percent include cashiers, kitchen workers, and stock and inventory clerks. At the other extreme, occupations in the bottom 5 percent of those affected include dental assistants, dental hygienists, musicians and composers, and clergy.

To determine whether the probability of IPT/VPT has grown the most in the occupations with the largest share of workers affected by the ACA mandate, we pursue a two-step process. First, for each 3-digit occupation, we estimate a linear probability model for VPT and IPT employment as a function of the same controls used in figures 2 and 3. The models also include a dummy variable for each year from 1994 forward and the sum of the year coefficients is

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[16] To match the share affected from the 2003 to 2007 March Current Population Survey to the monthly 1994-2015 Current Population Survey data, all 3-digit occupation codes were cross-walked into 1990 3-digit occupation codes using crosswalks available from Flood et al. (2015). 1990 Census occupations with less than 500 observations in the 1994-2015 sample for each of the two education groups were collapsed into the largest 3-digit occupation within its detailed occupation (46). This affected less than 0.5 percent of workers in the sample.

[17] We exclude observations with imputed industry, firm size, occupation, weekly hours, employer health insurance coverage, class of worker, weeks worked, or age. The weights are adjusted to account for deletions using the approach outlined in Bollinger and Hirsch (2006).

[18] The 2003-2007 March CPS reports the data for the 2002-2006 period.

constrained to equal zero. That is, we estimate first-stage regressions for each occupation of the following form:

$$(1)\ PT_{ijt} = \theta_{jt} + X_{it}\beta_j + \varepsilon_{ijt}$$

where PT_{ijt} is a dummy variable for either VPT or IPT employment; i indexes the individual; j indexes the 3-digit occupation; and t indexes year. X_{ijt} includes controls for economic conditions in period t in the person's state of residence and the worker's personal characteristics. The controls for economic conditions and personal characteristics are the same as those used in figures 2 and 3 except that controls with no within occupation variation must be dropped.

An important feature of the first stage regression is that it allows the effect of economic conditions and worker characteristics to differ by occupation. Hirsch (2005) presents evidence that the occupation-specific skill requirements tend to be lower in jobs where part-time work is most common. Moreover, theory suggests that when there is a downturn in demand for a firm's product, the employer is more likely to cut hours per worker instead of the number of workers when the firm has a large investment in the worker's training (Rosen 1968). Hence, we believe it is important to allow the relationship between part-time employment and economic conditions to differ by occupation.

After estimating the year effects for each occupation ($\hat{\theta}_{jt}$), we use a second stage regression to determine whether the change in the probability of part-time employment over time varies consistently with the share of workers affected by the ACA. That is, we estimate the following second stage regression:

$$(2)\ \hat{\theta}_{jt} = \alpha_j + \mu_t + affect_j * \beta_t + occ\_unemp_{jt} * \delta + occ\_unemp_{jt} * affect_j * \gamma + u_{jt}\ ,$$
$$\mathrm{t} = 1994, \ldots\ 2015$$

where $\hat{\theta}_{jt}$ is the estimated year-specific intercept for occupation j in year t obtained in the first stage regression; α_j is an occupation specific fixed effect; μ_t is a year fixed effect, $affect_j$ is the estimated proportion of workers affected by the ACA in occupation j, and occ_unemp_{jt} is a vector containing the year-specific unemployment rate in the 3-digit occupation and its one and two year lag. As in our earlier analysis, we restrict the sum of the year-specific effects (μ_t) and the sum of the year-specific effects of $affect_j$ (β_t) to be equal to zero. This makes the relevant reference period 1994-2015. The regression model is estimated with weighted least squares with the weights being the sample size used to calculate the proportion affected from the March 2003-2007 Current Population Survey. The standard errors used to calculate the t-statistics are bootstrapped (2,000 replications) to account for the estimation of the affected share and clustering of residuals by occupation.[19]

The second stage regression allows us to determine how much of the change in IPT employment across time is shared by all occupations (changes in a_t), and how much of the change in IPT varies according to the affected share by the ACA (changes in b_t). The estimates can be used to obtain a "difference-in-difference" estimate of the impact of the ACA by comparing groups with varying degrees of the treatment. To illustrate, the above regression model implies that, controlling for other factors, the growth in the probability of part-time (PT) employment in occupation j between, 2015 and the 1994-2015 period, is:

$$(3)\quad \Delta\theta_{j,2015} \quad = \hat{\mu}_{2015} + \hat{b}_{2015} * affect_j$$

Differencing the rate of growth across two occupations j and k, yields:

$$(4)\ \Delta\theta_{j,2015} = (affect_j - affect_k) * \hat{b}_{2015}$$

This expression shows how the differential rate of growth in PT employment varies according to the share affected by the mandate. If the ACA led to a shift toward PT employment, the coefficient on the affected share should rise when employers begin responding.

Relying on differential growth rates in PT employment to identify the effect of the ACA rests upon the assumption that, absent the ACA, the rate of growth in PT employment would not have been related to the percent affected by ACA. If this assumption is violated, our method of identification would lead to a biased estimate of the ACA. For example, if there is evidence that

[19] In the first stage of the bootstrap, we randomly draw with replacement, from within each occupation in the March CPS to estimate the share affected. In the second stage, we randomly draw (with replacement) entire clusters of occupations from the monthly CPS data. The first stage generates variation in the share affected by occupation. The second stage corrects for potential clustering of residuals by occupation (see Bertrand, Duflo, and Mullainathan 2004).

PT employment was growing more rapidly in affected occupations even prior to passage of the ACA (i.e., $\hat{b}_t$ was rising over time), a continuation of that trend would lead to an over-estimate of the effect of the ACA.

To investigate the validity of the identifying assumptions, figures 5 and 6 present the coefficient on the affected share ($\hat{b}_t$ from equation 2) between 1994 and 2015 from the models using PTI and PTV as the dependent variable in the first stage. The evidence is that, despite controlling for numerous measures of economic conditions, $\hat{b}_t$ tends to rise as the economy moves into recession and fall thereafter. This is true for both the high and low education groups. What is striking, however, is that after the Great Recession, $\hat{b}_t$ fell to a level that matched the previous low since 1994. This implies that, by 2011/12, the adjusted gap in the probability of IPT for affected and unaffected jobs had returned to pre-recession lows. However, by 2014 and 2015, a gap in the probability of IPT had emerged. For the first time since 1994, the probability of IPT was significantly higher in the highly affected occupations among the less educated work force. For the more educated workers, there is no statistically significant evidence that the adjusted probability of IPT rose faster in the more highly affected occupations.

In figure 6, the same information is provided for the relationship between the percent affected and the probability of VPT employment. Like IPT, there is no evidence that b_t rose over time. Unlike IPT, however, b_t does not appear to have counter-cyclical behavior. Relative to the 1994-2015 average, b_t is significantly different from zero in only one year (2006) in the low education group, and three years (1995, 2008-2009) in the high education group. Hence, we find no differential trends in VPT based on the extent to which occupations would have been affected by the ACA.

Since the only strong relationship between the affected share by the ACA and the probability of PT work is for workers without a college degree and the relationship exists only for involuntary, not voluntary, part-time work. For this reason, the remainder of the paper will focus on links between involuntary part-time work among workers without a college degree and passage of the ACA.

To provide some sense of the quantitative significance of these effects, the 90-10 difference in the affected share for workers without a college degree is 0.17. Given that the coefficient on affected share is .059 in 2015, this implies that, relative to the base period of 1994-

2015, the probability of IPT grew by 0.010 more among workers in a highly affected occupation (90th percentile) compared to a worker in occupation that is only slightly affected (10th percentile). While a 1.0 percentage point difference in the growth of the probability of IPT may seem small, the average probability of IPT employment in 2015 for a worker without a college degree is only 3.0 percentage points. This is strong evidence that, relative to the base period of 1994-2015, the probability of IPT employment grew most for the workers who were employed in occupations that were most likely to be impacted by the ACA.

One limitation of this two-step analysis is its assumption that, for a given education group, the effect of the mandate on part-time employment varies only with the percentage of workers affected. In fact, we expect that the effect would vary across workers depending on the nature of the job. The ACA limits the employee contribution for health insurance to 9.5% of household income for those receiving a subsidy. Thus, if the wage rate of a worker is higher, the firm is more able to pass on the cost of health insurance to the worker. Consequently, workers in high-wage industries would likely be less threatened with a switch to part-time employment as a result of the mandate.

To test this prediction, we pursue the following approach. First, we estimate the 10th percentile of the real hourly wage in 2014 dollars (Wage_10) by 3-digit occupation and education group using the 2010-2015 Outgoing Rotation Groups of the CPS. Second, we re-estimate the second stage regressions of the occupation-specific change in the probability of IPT employment on the affected share by adding Wage_10, and an interaction between Wage_10, the affected share, and all of the year dummies. The sum of the coefficients on these interaction terms is restricted to zero making 1994-2015 the reference period. The expectation is that the mandate would be more likely to increase IPT employment in occupations with lower hourly wages. Thus, we expect a negative coefficient on the interaction term between the wage rate and affected share after the ACA mandate takes effect.

The results are presented in table 2 for workers without a college degree. To facilitate the interpretation of coefficients, the affected share and Wage_10 are converted to deviations from sample means. For simplicity, we first focus on the results for workers without a college degree. Since the control variables are measured as deviations from means, the estimates of the intercept reveal the change in the probability of IPT for the occupation with the average 10th percentile wage and average affected share. The fact that the intercept is positive and statistically significant for 2013-2015 implies that the probability of IPT has been rising gradually relative to 1994-2015 average. The positive coefficient on the affected share variable interacted with the 2013-2015 year dummies implies that the probability of IPT rose most for workers in occupations where a large share would be affected by the mandate.

The negative coefficient on the interaction between affected share and Wage_10 in 2014 and 2015 is consistent with expectations. The positive association between the share of workers affected by the ACA and the growth of IPT employment is strongest in low wage occupations. While the coefficients for the other years are not presented for these interaction terms, it is worth noting that, other than in 2014 and 2015, the coefficient is never significantly different from zero at the .10 level. Hence, the more rapid growth in IPT in highly affected occupations with low wages did not start until 2014.

As an illustration of the size of these effects, compare a high and low wage occupation where Wage_10 is either $5 above or $5 below the average value. The marginal effect of the affected share by the ACA is close to zero for workers in the high wage occupation. On the other hand, for workers in occupations where Wage_10 is $5 below the average, the effect of the ACA is about 80% higher than that for a worker in an occupation where Wage_10 matches that for the average occupation.[20]

While we have presented strong evidence that the growth in the probability of IPT has been greatest in the occupations where a large share of workers would be affected by the ACA, it is possible that our results are simply spurious correlation between structural changes in the labor market and the affected share. For example, Autor and Dorn (2013) discuss the growth of low-skill service jobs that was partly driven by technological change displacing workers in occupations that performed routine tasks. Canon et al. (2014) provide evidence that IPT employment has become more common since 2010, and that IPT is most common in non-routine manual occupations (typically low-skill service). They suggest that the effect of technological change on the polarization of jobs could be partly responsible for the rise in IPT.

If our occupation-specific estimate of the affected share by the ACA varies systematically with the impact of technological change, our estimated effect of the ACA could be biased and

[20] This comparison is based on the derivative of the probability of IPT employment with respect to %affected implied by specification (1) in table 3. In 2015, this derivative is 0.046-.0075*Wage_10. Recall that the occupation- specific wage measure is converted to a deviation from means in the regression so that Wage_10 is +5 (-5) in the high and low wage occupations, respectively. Similar results are obtained when the 25th percentile wage is used in place of the 10th percentile, but Wage_10 has a slightly higher level of significance.

capturing what is actually the effect of technological change. To test for this, we merge Autor and Dorn's occupation specific measure of "routine task intensity" (RTI) to our data and repeat the second stage regressions reported in specification 1 of table 2. We include RTI as well as interactions between RTI and year dummies. The interactions allow for the possibility that the impact of technological change on IPT is changing over time. If technological change explains the patterns observed in our analysis, controlling for RTI should reduce or eliminate the explanatory power of the affected share by the ACA for the growth in IPT employment. The results, provided as specification (2) of table 2, reveal that the addition of RTI to our model has virtually no impact on our primary conclusions. Namely, even after allowing the growth in the probability of IPT to differ depending on an occupation's exposure to the effects of technological change (RTI), the growth in the probability of IPT employment since 2013 was greatest in occupations with low wages and had a larger share of workers affected by the ACA employer mandate.

To provide a sense of the quantitative importance of our results, we use the models in figure 5 and table 2 to compute the implied change in IPT employment due to the ACA for workers without a college degree. In particular, we compute the predicted change in the probability of IPT employment that would occur if the occupation specific changes observed in 2014 and 2015 had not occurred. This is estimated as the predicted change in the probability of IPT associated with switching the coefficients on the year*affect and year*affect*Wage_10 coefficients to zero (presented in table 2, specification 2) in 2014 and 2015. We then multiply this occupation-specific increase in the probability of IPT employment by the level of employment in that occupation in the relevant year.[21] We compute standard errors for our estimates by using the 2,000 bootstrapped coefficients for each regression to generate 2,000 estimates of the aggregate change in IPT employment in each year.

It is worth noting that our approach is using the 1994-2015 average as the counterfactual for what would have happened had the ACA not been passed. This is a relatively conservative approach since the estimated effects would be larger if we had used one or more of the years shortly before the mandate began to bind (e.g. 2010-2012).

[21] An alternative approach is to use the 2015 level of employment rather than the actual level of employment in a given year. This method removes any effects of shifts in the occupation distribution of employment on IPT. This approach changes the results very little.

Table 3 presents our estimated changes in IPT employment due to the ACA ds. The estimated increase in IPT relative to the base-period of 1994-2015 is fairly robust to the model chosen. In all three models, the impact of the ACA starts at around 200,000 in 2013, rises to approximately 700,000 in 2014, and then is either flat or falls slightly. Given that the employer mandate did not assess fines for a lack of coverage until 2015, one might be surprised that the effects did not rise between 2014 and 2015 – though the change is not statistically significant. There are several plausible explanations for the moderation in the effect. First, while the penalties did not take effect until 2015, the definition of who must be covered in 2015 was based on 2014 employment data. As a consequence, employers needed to adjust employee hours in 2014 if they were to avoid the cost of providing coverage in 2015. Second, if employees are forced into involuntary part-time employment because of the ACA, they will seek out a full-time job elsewhere. Employers who wish to continue with part-time workers will eventually move toward those who voluntarily choose part-time employment. For similar reasons, the effect of the ACA on IPT may diminish over time as workers who are involuntarily employed part-time switch to other jobs and employers use workers who prefer part-time work or other ways to avoid providing coverage for low wage workers. For example, large employers may outsource low skill jobs or replace wage and salary workers with self-employed workers. We leave it for future researchers to test whether large firms are pursuing such strategies in response to the ACA.

Robustness of Results.

While the above empirical results are largely consistent with the theory that the ACA mandate caused IPT employment to rise, it is important to test the robustness of the results to alternative specifications. First, we consider two alternative versions of the occupation specific measure of the affected share. As noted above, the ACA will apply to employers with 50 or more employees, though the implementation date for firms with 50-99 employees was postponed by one year.

Prior to 2011, the March CPS allows us to identify workers at firms with 100 or more workers, but not exactly 50. From 2011 forward, we can identify workers at employers with 50-

99 workers. Using information from the 2011-2015 March CPS, we estimate the share of workers at firms with less than 100 employees that have between 50-99 employees. We then use this to create a probability that each worker employed at a firm with less than 100 workers in the 2003-2007 March CPS would be at a firm with 50-99 workers. This information is then used to estimate the share of workers in a given occupation that are at a firm of 50 or more employees without health insurance coverage.

This alternative measure of the affected share causes a significant increase in the average share of workers affected by the ACA. For workers without a college degree, switching the definition of affected workers to include those at firms of 100+ to 50+ employees increases the affected share from .16 to .25. Using this alternative measure of the affected share could either increase or decrease the estimated impact of the ACA. It could decrease the estimated effect because the mid-sized firms (50-99) were not subjected to the penalties until 2016 and thus may display a muted response. It may also decrease the estimated effect because this estimate of the affected share is noisier since the 2003-2007 data does not have the 50-99 employer size category precisely defined. On the other hand, it could increase the estimated effect since it will lead to a more accurate classification of workers according to whether they will eventually be affected by the mandate.

A comparison of the results is presented in table 4. While the estimated coefficients on the key parameters is somewhat muted by using the alternative measure of the affected share, the statistical significance remains for the primary parameters of interest. Namely, the growth in the probability of IPT is significantly greater in occupations with a larger share of affected workers, regardless of whether we use a firm size of 100 or 50 to define who is affected. While the coefficients are the affected share by the ACA are slightly lower, it is important to keep in mind that the average affected share is higher when using a firm size of 50 as the cut-off.

Another issue in our measure of the affected share is the problem of "heaping" in survey responses to questions regarding hours worked. That is, many workers will round their answers to 20, 30, or 40 hours. Since the 30-hour cut-off is used in defining who is affected by the ACA, it is likely that some people working slightly less than 30 hours round their answers and report 30 and we would misclassify the worker as affected if she did not have health insurance coverage.

To determine whether this rounding error creates a problem, we repeated all of our analysis by excluding workers who reported exactly 30 hours of work in determining the affected share and in the regressions used to estimate the determinants of IPT. The results are presented in specifications (3) and (6) of table 4. Our key results are robust to this change in the measure of affected share. The growth in IPT employment is greatest in occupations where the share of workers affected by the ACA is largest.

Who are the affected workers?

Table 5 presents a list of occupations with the largest and smallest increases in the probability of IPT due to the ACA relative to the reference period. The list is restricted to occupations with at least 300,000 workers without a college degree in 2015. The table also lists the average 10th percentile wage and affected share by the ACA for each of the occupations.

The eight occupations with the largest increase in the probability of IPT relative to the reference period include, not surprisingly, low wage occupations with a relatively large share of workers affected by the mandate. The list includes cashiers, cooks, waiters and waitresses, and kitchen workers. All of these occupations have 10th percentile wages averaging less than \$9 per hour.

On the other extreme, the eight occupations where the ACA caused the smallest increase in the probability of IPT include relatively high wage occupations (all have a 10th percentile of wages above \$11 per hour) and where the affected share is quite low. Occupations with small increases in IPT due to the ACA include registered nurses, police officers, electricians, and machinists. In separate analysis, we examined the change in the probability of IPT due to the ACA for each worker and found that women, racial minorities, and younger workers were more likely than others to see a shift to IPT work as a result of the ACA.

4. Summary and Conclusions.

Starting in 2015, the Affordable Care Act imposed penalties on large employers that do not provide affordable health insurance to their full-time employees. This study investigates whether employers have shifted from full-time to part-time workers to avoid these penalties. Our estimation of models of part-time employment that control for changing economic conditions, worker characteristics, and occupational and industrial composition suggest that IPT employment between 2013 and 2015 was higher than predicted based on a reference period of 1994-2015. More importantly, we find that the increase in IPT employment over the past few years was greatest in the occupations that employed the largest percentage of workers who would be affected by the mandate -- those employed at a large firm without health insurance coverage. This result persists after we control for the state of the economy, the demographics of workers, and control for other factors (such as technological change) that might contribute to elevated levels of involuntary part-time work. We believe this provides strong evidence that the ACA penalties have caused a shift to IPT employment.

The empirical analysis indicates that the ACA employer mandate had a quantitatively important impact on IPT employment. Our estimates suggest that as many as 700,000 workers without a college degree are engaged in IPT employment because of the ACA. It is conceivable these effects will gradually diminish as employers find other ways to avoid the mandate (e.g., outsource low skill work to small firms) and IPT workers gradually relocate where they can obtain full-time jobs.

Overall, we believe this study presents robust evidence that the ACA employer mandate is at least partially responsible for the elevated levels of involuntary part-time work since 2013. While it is conceivable that there are other structural changes at work that explain differential growth in IPT employment across occupations, we are unaware of any such structural change that would have grown in importance in the occupations where the ACA is most likely to be binding.

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Table 1. Key Affordable Care Dates	
Legislation Dates:	
Passed Senate	December 24, 2009
Passed House	March 21, 2010
Signed into Law by President	March 23, 2010
Upheld by Supreme Court	June 28, 2012
Original Dates in Legislation:	
Start of Health Insurance Exchange	October 1, 2013
Employee Fines Start for No Coverage	January 1, 2014
Subsidies Made Available	January 1, 2014
Revised Dates:	
Delay of Employer Fines Until 2015	July 2, 2013
Delay of Employer Fines Until 2016 for 50-99 Employee Firms	February 10, 2014
Coverage Requirement for 100+ Employee Firms changed to 70 percent for 2015 and 95 percent by 2016	February 10, 2014

Note: The legislation dates are from U.S. Senate, Committee on Finance (http://www.finance.senate.gov/issue/?id=32be19bd-491e-4192-812f-f65215c1ba65). The implementation dates for Affordable Care Act are from Kaiser Family Foundation (http://kff.org/interactive/implementation-timeline/).

Table 2. Impact of ACA on Involuntary Part-Time Employment for Workers with Less Than 16 Years of Education.

Explanation: Coefficients below are from the second step in the two-stage process described in the text. The coefficients below represent the relationship between the control variables and the occupation specific increase in involuntary part-time employment. In first stage, using data from 1994 through 2015, a linear probability model of part-time employment is estimated for each 3-digit occupation with controls for worker characteristics and economic conditions. The coefficients on the year dummies represent the change in the probability of part-time employment relative to the selected reference period of 1994-2015. In the second stage, the occupation specific coefficients on the year dummies from the first stage are regressed on year dummies, the affected share, affected share * 10th percentile wage, and the controls listed below. The regressions are weighted by the occupation-specific sample size used to estimate the percent of workers affected. T-statistics are based on bootstrapped standard errors (2,000 replications) correcting for the estimation of percent affected and clustering by occupation. Additional details are included in the text.

	(1)	(2)
Year:		
2013	0.002	–
	(2.08)	(2.04)
2014	0.003	0.003
	(3.67)	(3.58)
2015	0.003	0.003
	(4.28)	(4.09)
Year Interacted with Affected share		
2013 * Affected share	0.015	0.015
	(0.83)	(0.80)
2014 * Affected share	0.043	0.043
	(3.09)	(2.99)
2015 * Affected share	0.046	0.042
	(3.42)	(3.36)
Year interacted with Affected share and Wage_10		
2013 * Affected share * Wage_10/100	0.055	0.055
	(0.13)	(0.13)
2014 * Affected share * Wage_10/100	-0.747	-0.743
	(-2.00)	(-1.97)
2015 * Affected share * Wage_10/100	-0.750	-0.795
	(-1.58)	(-1.70)
Year Dummies for 1994-2012 Included	Y	Y
Year Dummies 1994-2012 * Affected Share Included	Y	Y
Year Dummies for 1994-2012 * Affected Share * Wage Included	Y	Y
Dummies for each occupation included	Y	Y
Occupation unemployment rate included	Y	Y
Occupation unemployment * Affected Share included	Y	Y
Controls for RTI included	N	Y
Sample Size	5,918	5,918

Table 3. Estimated Increase in Involuntary Part-Time Employment due to ACA for Workers with Less than 16 Years of Education.

Explanation: The models presented in Figure 5 and table 2 are used to compute the implied change in involuntary part-time employment due to the ACA for workers with less than 16 years of education. We compute the predicted change in the probability of part-time employment from changing the occupation-specific share of workers affected by the ACA from the observed value to zero. We then multiply this occupation-specific increase in the probability of part-time employment by the level of employment in that occupation in the relevant year. Standard errors are in parentheses. All employment estimates are in 1,000s.

	Involuntary Part-Time Employment		
2013	234,251	229,970	230,065
	(247,518)	(247,702)	(255,701)
2014	705,312***	661,207***	659,570***
	(194,289)	(191,701)	(198,587)
2015	711,705***	669,055***	623,499***
	(228,390)	(202,500)	(186,472)
Controls for Wage	N	Y	Y
Controls for RTI	N	N	Y

Standard errors are given in parentheses and are estimated with 2,000 bootstrap replications. *, **, *** indicate the change in employment is statistically significant at the .10, .05, or .01 level.

[1] RTI is an occupation specific measure of a routine task index designed to control for the rising impact of technological change on the displacement of low skill workers. The models with RTI include interactions between the year dummies and the RTI index to allow the impact of RTI on part-time employment to change over time.

Table 4. Relationship between Occupation Specific Measure of Share Affected by ACA and Growth in Involuntary Part-Time Employment Among Workers without a College Degree.

Explanation: The models below replicate the regressions in figure 2 and table 2 but with different measures of the share of an occupation affected by the ACA. Specifications (1) and (4) define affected workers as those working 30 or more hours at a large firm without health insurance coverage where large is based on a firm size of 100 or more. Specifications (2) and (5) define large as a firm size of 50 or more. Specifications (3) and (6) exclude all workers reporting exactly 30 hours of work to remove observations that could be rounded and therefore misclassified in terms of whether the ACA would affect them.

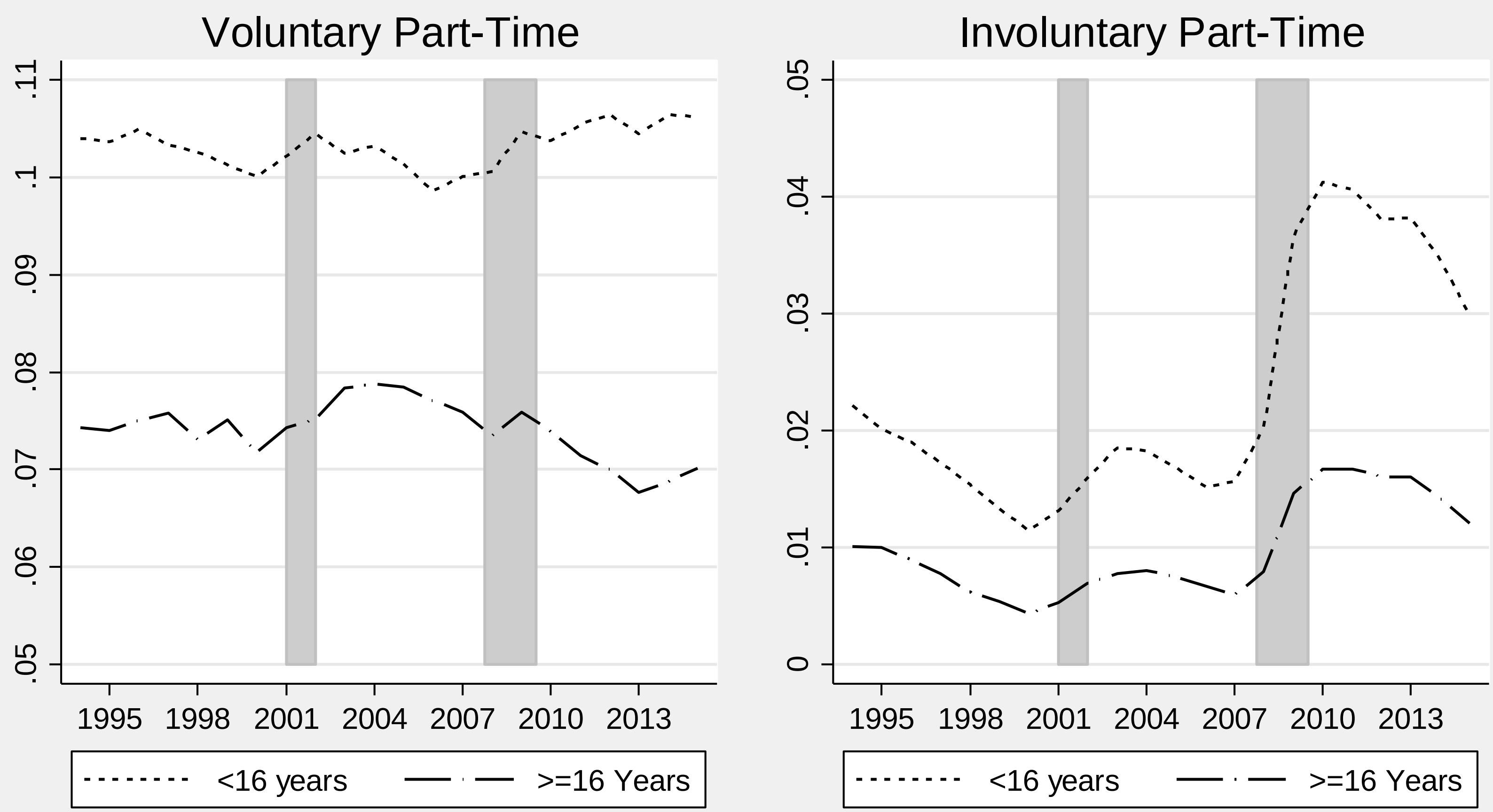
	Without Wage * Affect Interactions			With Wage* Affect Interactions		
	(1) Affected based on employers of 100+ workers	(2) Affected based on employers of 50+ workers	(3) Affected based on employers of 100+ workers deleting reports of 30 hours	(4) Affected based on employers of 100+ workers	(5) Affected based on employers of 50+ workers	(6) Affected based on employers of 100+ workers deleting reports of 30 hours
Year dummies:						
2013	0.002	0.002	0.002	0.002	0.003	0.002
	(2.39)	(2.92)	(2.15)	(2.08)	(2.80)	(1.85)
2014	0.003	0.004	0.003	0.003	0.003	0.003
	(4.94)	(5.25)	(4.49)	(3.67)	(3.87)	(3.32)
2015	0.003	0.004	0.004	0.003	0.004	0.003
	(5.49)	(5.73)	(5.33)	(4.28)	(4.98)	(4.18)
Year dummies * Affected Share						
2013 * Affected	0.015	0.004	0.016	0.015	0.010	0.017
	(0.75)	(0.29)	(0.68)	(0.83)	(0.74)	(0.76)
2014 * Affected	0.056	0.031	0.063	0.043	0.026	0.047
	(3.66)	(2.70)	(3.18)	(3.09)	(2.42)	(2.86)
2015 * Affected	0.059	0.032	0.071	0.046	0.031	0.050
	(3.29)	(2.24)	(3.05)	(3.42)	(2.67)	(3.26)
Year dummies * Affected Share * Wage_10						
2013 * Affected * Wage_10/100	--	--	--	0.055	0.341	0.030
				(0.13)	(1.25)	(0.05)
2014 * Affected * Wage_10/100	--	--	--	-0.747	-0.261	-0.985
				(-2.00)	(-0.99)	(-1.76)
2015 * Affected * Wage_10/100	--	--	--	-0.750	-0.052	-1.248
				(-1.58)	(-0.16)	(-1.72)
Other variables included:						
Year dummies 1994-2012	Y	Y	Y	Y	Y	Y
Year dummies 1994-2012 * Affected	Y	Y	Y	Y	Y	Y
Occupation dummies	Y	Y	Y	Y	Y	Y
Occupation unemployment rate	Y	Y	Y	Y	Y	Y
Occupation unemployment rate * Affected	Y	Y	Y	Y	Y	Y
Sample Size	5,918	5,918	5,918	5,918	5,918	5,918

Table 5. Occupations Ranked by Change in Probability of Involuntary Part-Time Work due to ACA.

Note: Sample is restricted to workers without a college degree, aged 19-64, and employed in 3 digit occupations with at least 300,000 workers in 2015. Estimated changes are based upon specification (4) in table 4. Change in probability is based on a comparison of 2015 with the reference period of 1994-2015.

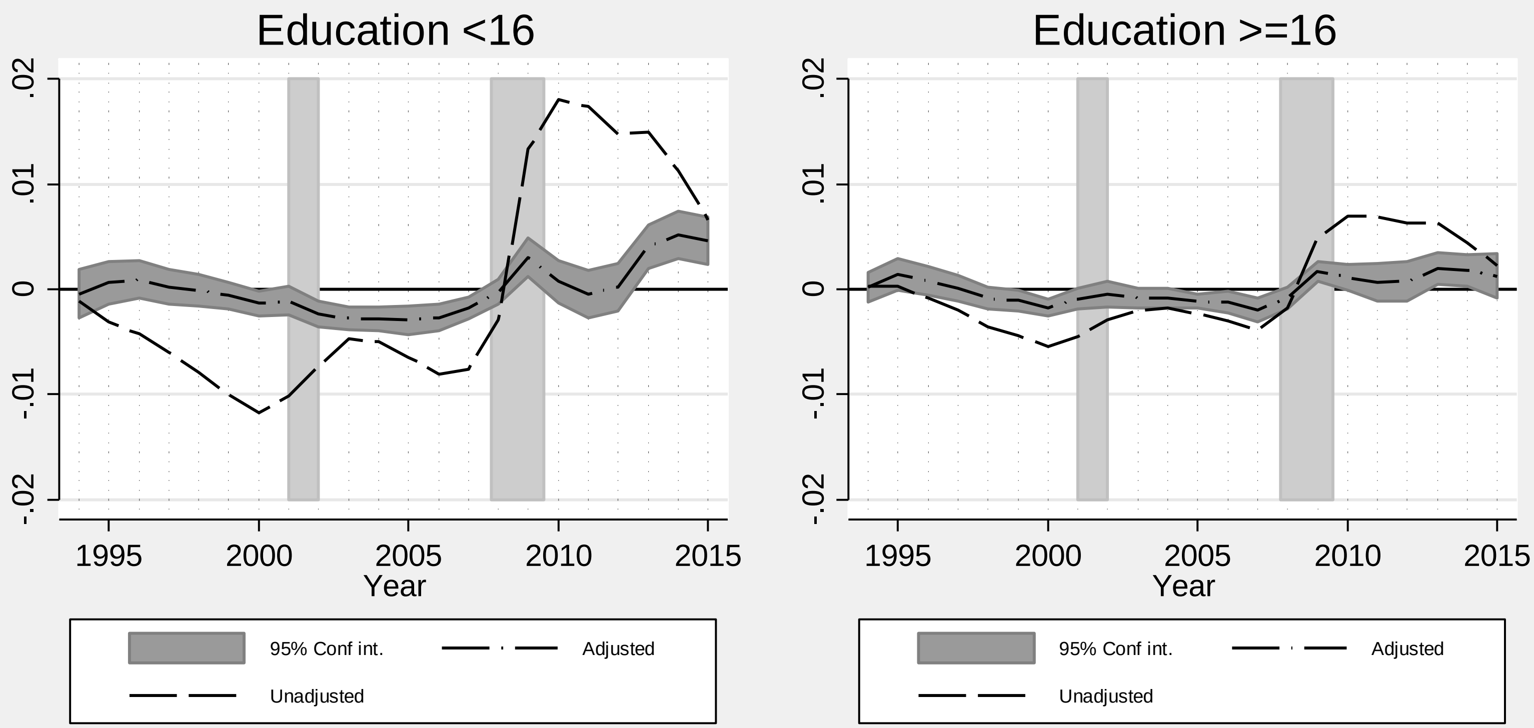
	Increase in Probability of PTI due to ACA	Number of Additional Workers employed in PTI jobs due to ACA	Total employment in 2015	Average 10th percentile wage in occupation	Affected share in Occupation
	8 occupations with Largest increase in probability of PTI				
Kitchen workers	2.1%	6,365	302,884	$7.48	32.1%
Cashiers	2.0%	45,917	2,279,389	$7.58	31.2%
Stock and inventory clerks	1.8%	23,139	1,251,611	$8.00	30.1%
Waiter/waitress	1.8%	28,225	1,555,891	$5.41	22.5%
Packers and packagers by hand	1.8%	7,263	400,704	$7.88	29.0%
Cooks, variously defined	1.7%	40,990	2,388,638	$7.74	27.0%
Guards, watchmen, doorkeepers	1.6%	10,167	617,870	$8.69	29.1%
Misc food prep workers	1.5%	12,900	849,814	$7.62	23.6%
	8 occupations with smallest increase in probability of PTI.				
Supervisors of construction work	0.1%	420	486,257	$14.84	7.7%
Registered nurses	0.1%	1,086	1,070,374	$15.47	15.4%
Computer software developers	0.1%	411	344,809	$15.00	11.9%
Police, detectives, and private investigators	0.1%	740	527,925	$13.72	7.2%
Electricians	0.2%	1,053	654,796	$12.98	6.5%
Machinists	0.2%	633	302,501	$11.94	6.4%
Managers and administrators, n.e.c.	0.3%	6,971	2,785,824	$12.70	9.3%
Industrial machinery repairers	0.3%	977	370,925	$12.49	9.2%

Figure 1
Fraction Employed Part-Time by Education Level



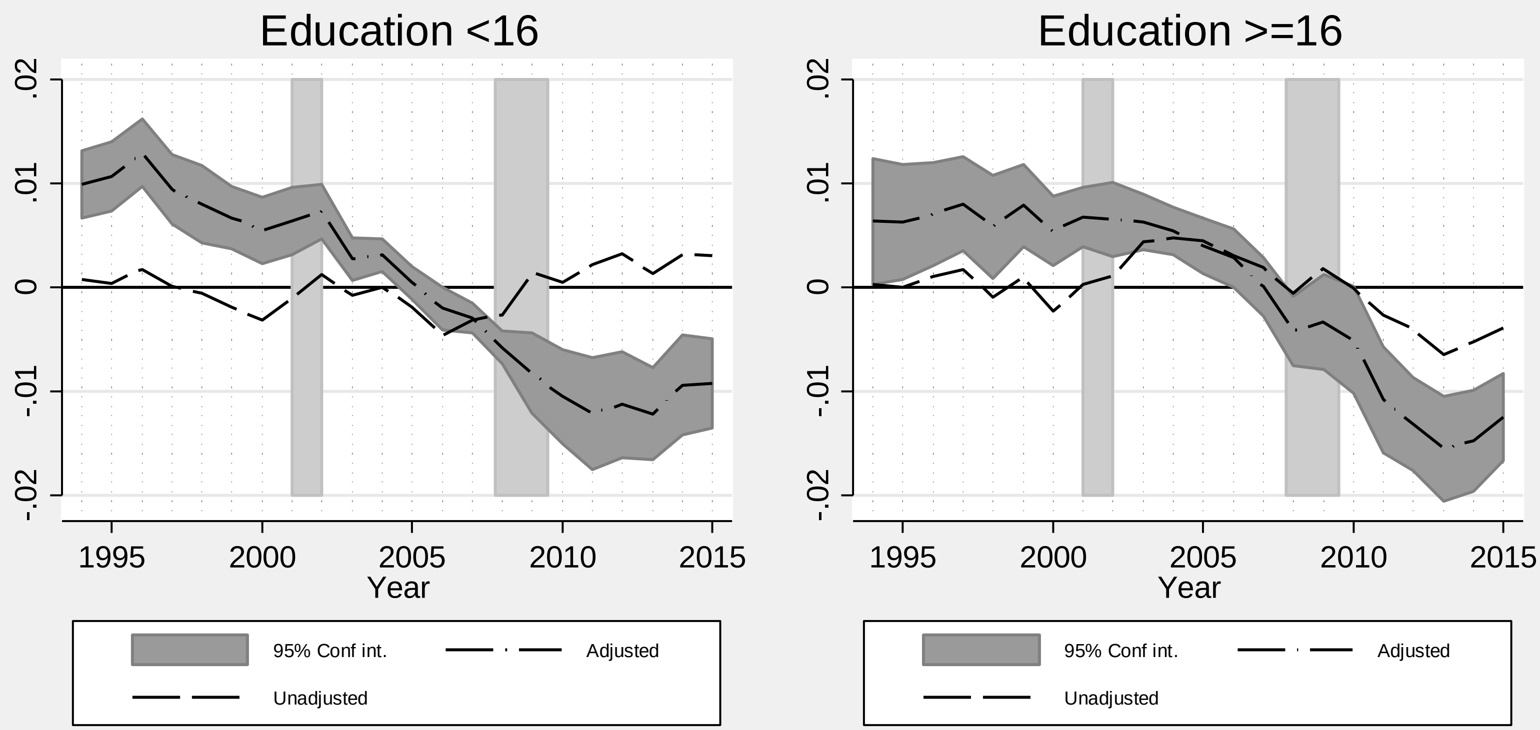
The sample includes wage and salary workers aged 19-64 not residing in District of Columbia, Hawaii, or Massachusetts. Part-time is defined as working less than 30 hours per week. Shaded areas represent recessions.

Figure 2
Variation in Probability of Involuntary Part-Time: 1994-2015



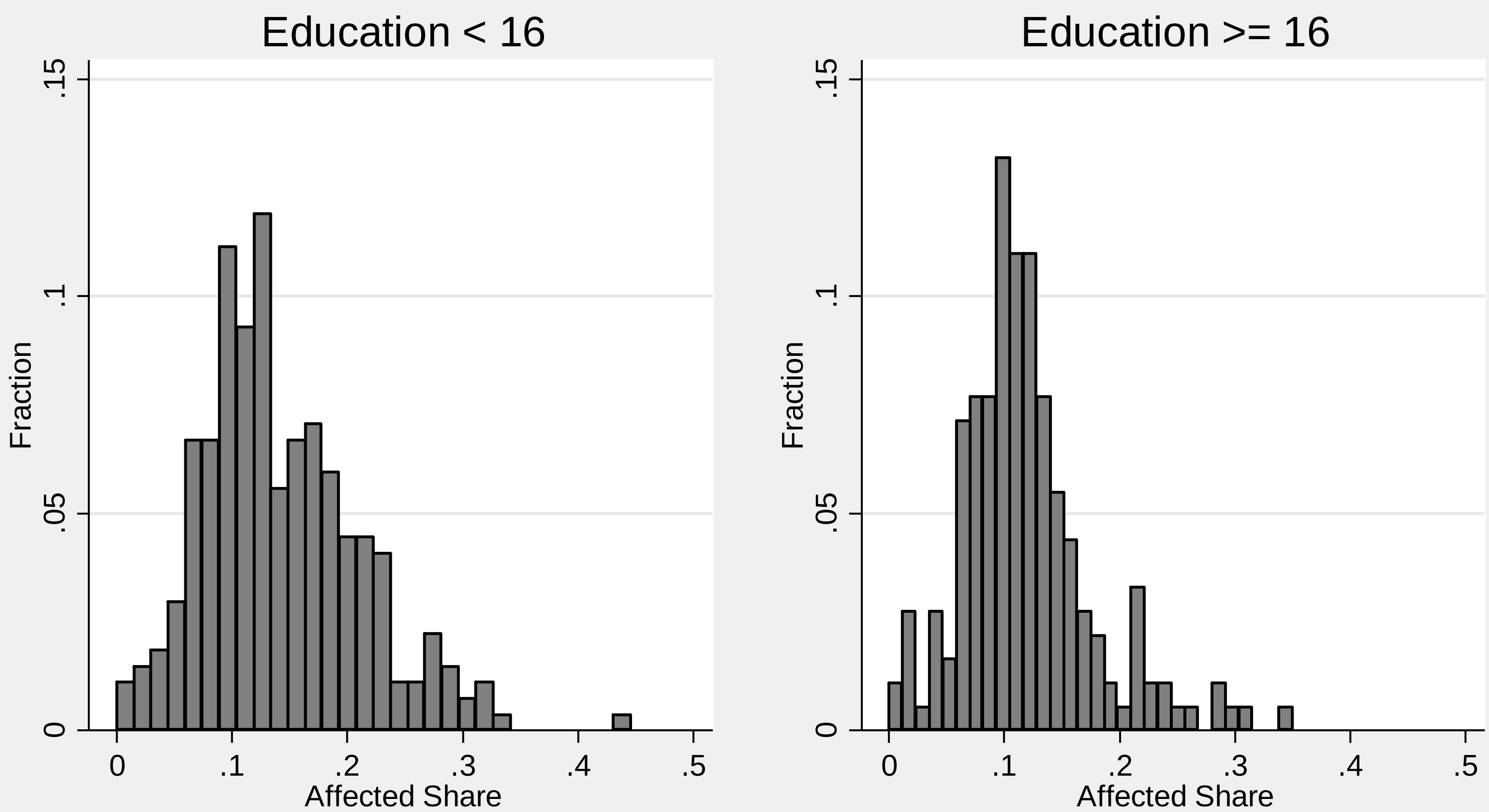
The plots represent deviations in the probability of involuntary part-time relative to the 1994-2015 average. The adjusted rate is from a regression model with controls for state economic conditions and demographic characteristics, and personal characteristics. The unadjusted rate is from a regression model with no controls. The confidence intervals are based on bootstrapped standard errors correcting for clustering by state. The year effects are forced to sum to zero. See the text for details. Shaded areas represent recessions.

Figure 3
Variation in Probability of Voluntary Part-Time: 1994-2015



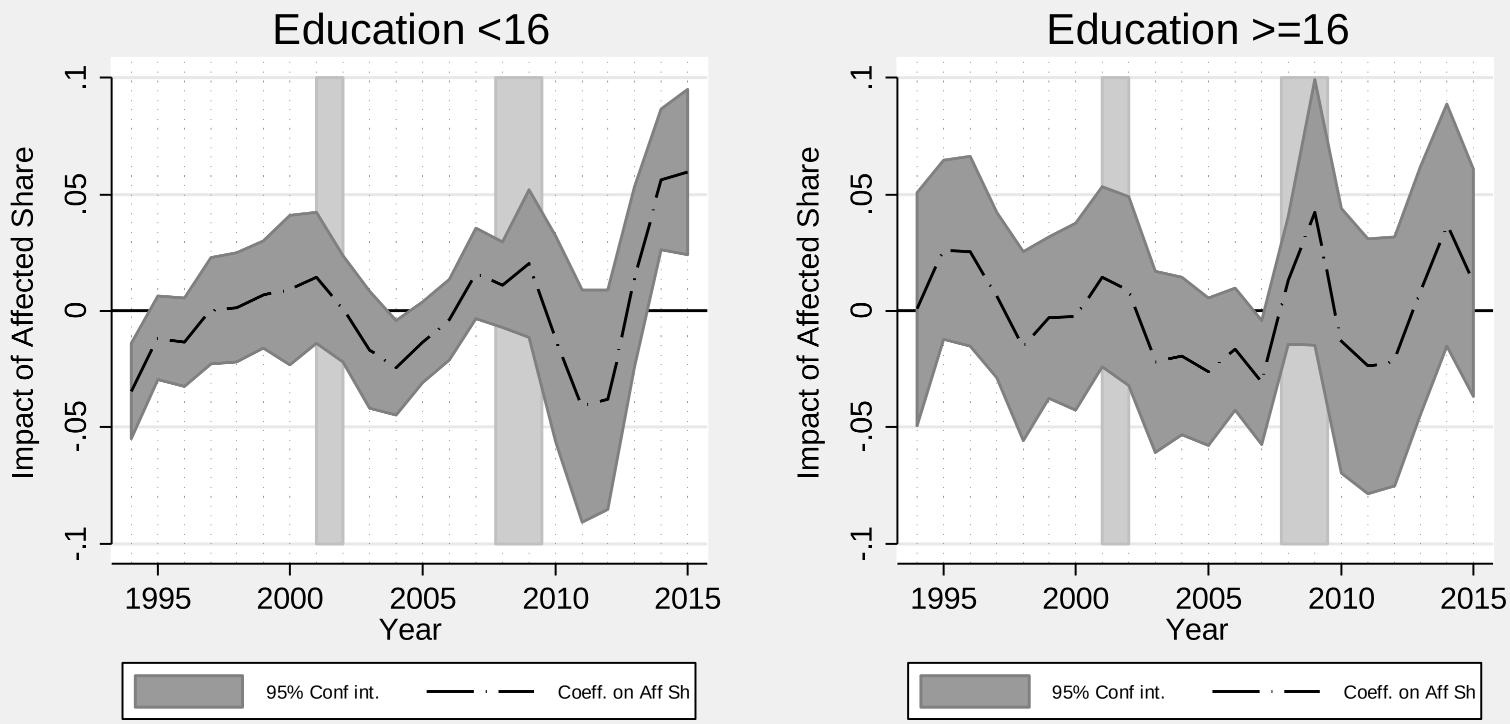
The plots represent deviations in the probability of voluntary part-time relative to the 1994-2015 average. The adjusted rate is from a regression model with controls for state economic conditions and demographic characteristics, and personal characteristics. The unadjusted rate is from a regression model with no controls. The confidence intervals are based on bootstrapped standard errors correcting for clustering by state. The year effects are forced to sum to zero. See the text for details. Shaded areas represent recessions.

Figure 4
Share of Workers Affected by ACA Across Occupations



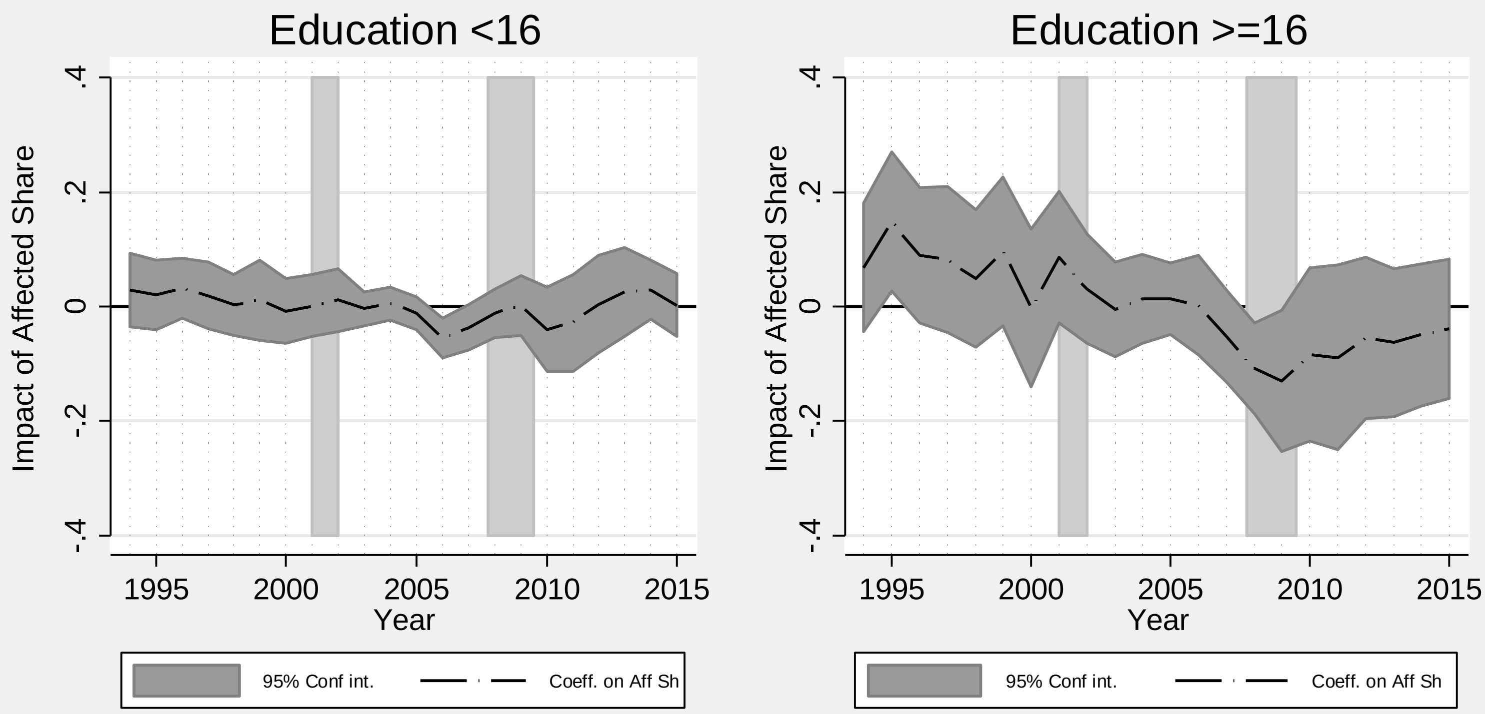
Affected workers are those at a firm with 100 or more workers, work 30 or more hours per week, and don't have health insurance. Employment weighted mean share affected in 2015 is 0.163 for education <16 and 0.114 for education>=16.

Figure 5
Impact of Share Affected by ACA on Involuntary Part-Time



Plotted coefficient represents the estimated effect of the occupation-specific share affected by the ACA on the year-specific increase in the probabiity of involuntary part-time employment. The regression model includes controls for state economic conditions and demographic characteristics, personal characteristics, occupation unemployment rate and its interaction with the affected share. The confidence intervals are based on bootstrapped standard errors correcting for the estimation of share affected and clustering by occupation. The year effects are forced to sum to zero. See the text for details. Shaded areas represent recessions.

Figure 6
Impact of Share Affected by ACA on Voluntary Part-Time



Plotted coefficient represents the estimated effect of the occupation-specific share affected by the ACA on the year-specific increase in the probabiity of voluntary part-time employment. The regression model includes controls for state economic conditions and demographic characteristics, personal characteristics, occupation unemployment rate and its interaction with the affected share. The confidence intervals are based on bootstrapped standard errors correcting for the estimation of share affected and clustering by occupation. The year effects are forced to sum to zero. See the text for details. Shaded areas represent recessions.

Appendix Table 1: Impact of Share Affected by ACA on Involuntary Part-Time Employment.

Explanation: For each 3-digit occupation, a linear probability model of involuntary part-time employment is estimated with controls for personal characteristics of workers and economic conditions. The sample period is 1994 through 2015. The coefficients on the year dummies represent the change in the probability of part-time employment relative to the base period (1994-2015). The regressions are weighted by the occupation-specific sample size used to estimate the percent of workers affected. T-statistics are based on bootstrapped standard errors (2,000 replications) correcting for the estimation of percent affected and clustering by occupation Additional details are included in the text.

	Less than 16 years of education	16 or more years of education
Year:		
1994	0.000	0.000
	(0.04)	(0.19)
1995	0.000	0.001
	(0.78)	(1.38)
1996	0.001	0.001
	(2.33)	(0.88)
1997	0.001	0.000
	(1.14)	(0.36)
1998	0.000	-0.000
	(0.46)	(-0.42)
1999	-0.000	-0.001
	(-0.40)	(-1.22)
2000	-0.001	-0.001
	(-1.28)	(-1.54)
2001	-0.001	-0.000
	(-1.13)	(-0.58)
2002	-0.002	-0.001
	(-4.14)	(-0.90)
2003	-0.002	-0.001
	(-3.78)	(-1.85)
2004	-0.002	-0.001
	(-4.49)	(-1.38)
2005	-0.002	-0.001
	(-5.84)	(-1.72)
2006	-0.002	-0.001
	(-4.09)	(-1.40)
2007	-0.001	-0.001
	(-2.45)	(-3.28)
2008	0.000	-0.000
	(0.50)	(-0.53)
2009	0.002	0.001
	(3.13)	(1.23)
2010	-0.000	-0.000
	(-0.01)	(-0.25)
2011	-0.001	-0.000
	(-1.25)	(-0.42)
2012	-0.001	-0.000
	(-0.80)	(-0.11)

	Less than 16 years of education	16 or more years of education
2013	0.002	0.002
	(2.39)	(1.76)
2014	0.003	0.002
	(4.94)	(2.30)
2015	0.003	0.002
	(5.49)	(2.06)
Year Interacted with Affected Share		
1994 * Affected Share	-0.034	0.001
	(-3.29)	(0.03)
1995 * Affected Share	-0.012	0.026
	(-1.26)	(1.33)
1996 * Affected Share	-0.013	0.025
	(-1.39)	(1.23)
1997 * Affected Share	-0.000	0.007
	(-0.00)	(0.37)
1998 * Affected Share	0.001	-0.015
	(0.11)	(-0.74)
1999 * Affected Share	0.007	-0.003
	(0.58)	(-0.17)
2000 * Affected Share	0.009	-0.003
	(0.55)	(-0.12)
2001 * Affected Share	0.014	0.015
	(1.00)	(0.74)
2002 * Affected Share	0.001	0.008
	(0.06)	(0.41)
2003 * Affected Share	-0.017	-0.022
	(-1.31)	(-1.11)
2004 * Affected Share	-0.025	-0.019
	(-2.37)	(-1.13)
2005 * Affected Share	-0.014	-0.026
	(-1.55)	(-1.62)
2006 * Affected Share	-0.004	-0.016
	(-0.45)	(-1.22)
2007 * Affected Share	0.016	-0.031
	(1.61)	(-2.26)
2008 * Affected Share	0.011	0.013
	(1.20)	(0.94)
2009 * Affected Share	0.020	0.042
	(1.25)	(1.45)
2010 * Affected Share	-0.012	-0.013
	(-0.54)	(-0.44)
2011 * Affected Share	-0.041	-0.024
	(-1.61)	(-0.85)
2012 * Affected Share	-0.038	-0.022
	(-1.59)	(-0.80)
2013 * Affected Share	0.015	0.008
	(0.75)	(0.31)

	Less than 16 years of education	16 or more years of education
2014 * Affected Share	0.056	0.037
	(3.66)	(1.38)
2015 * Affected Share	0.059	0.012
	(3.29)	(0.48)
Occupation dummies included	Y	Y
Occupation unemployment rate included	Y	Y
Occupation unemployment * affected share included	Y	Y
Sample Size	5,918	3,982