THE LABOR MARKET EFFECTS OF EXPANDING OVERTIME COVERAGE*

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Abstract

Despite the prevalence of overtime regulations, estimating their impact has been hindered by a lack of policy variation and high-quality data. This paper evaluates the labor market impact of recent expansions in overtime coverage for salaried workers in the US. While overtime was historically intended to raise employment by encouraging firms to substitute more workers for fewer hours, a competing theory predicts that employers would instead reduce base pays to offset the cost of the overtime premium. Leveraging recent changes in state salary thresholds for overtime coverage of salaried employees, in conjunction with high-frequency administrative payroll data, I find evidence inconsistent with both views. Rather than reducing base pays, firms raised salaries above the new eligibility threshold to keep jobs exempt from overtime. On the other hand, expansions in overtime coverage had no significant impact on employment or workers' hours. I find the effects are consistent with a labor demand model whereby employers determine workers' hours.

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I Introduction

In nearly every OECD country, basic employment standards are regulated through at least two major rules: a minimum wage to set a hard floor on workers' earnings, and an overtime premium to set a soft ceiling on workers' hours (OECD, 2021; ILO, 2021). While interest in wage and hour regulations has sparked a large literature on the economic impacts of the minimum wage (Autor et al., 2016; Harasztosi and Lindner, 2019; Cengiz et al., 2019), far less is known about the effects of overtime despite its potentially large impact on workers' hours, earnings, and employment. In 2019, while only 2% of workers in the U.S. earn the federal minimum wage (BLS, 2019), over half of U.S. workers are guaranteed a 50% overtime wage premium for each hour worked above 40 per week (DOL, 2019a). The broad extent of overtime coverage induces a larger transfer from employers to workers each year, with firms paying nearly twice as much in overtime compensation in 2019 as they do in taxes to fund the entire unemployment insurance system (DOL, 2019b).

Despite the prevalence of overtime regulations, research on its effects have been hindered by a shortage of policy changes and the unavailability of data to accurately measure their impacts. In the U.S., expansions in overtime coverage to additional industries and demographic groups over the past 80 years often coincided with expansions in the minimum wage. As a result, previous papers are confronted with the challenge of separately identifying the effects of overtime on income and employment. Apart from a shortage of clean variation, few datasets in the U.S. distinguish between workers' base pay and overtime pay, and those that do often lack the sample size or panel structure to precisely estimate changes in aggregate employment. In light of these empirical challenges, a review by Brown and Hamermesh (2019) concludes that "no study presents estimates of effects [of overtime coverage] on employment, and none offers evidence on all outcomes: [wages, earnings, and hours]".

Even from a theoretical perspective, the labor market impacts of overtime coverage are ambiguous. In models of labor supply, workers would increase their hours in response to the overtime premium (Idson and Robins, 1991). On the other hand, a labor demand model predicts firms would reduce weekly hours and bunch them at 40. Historically, policymakers believed that employers would substitute for the reduction in hours by hiring more workers, leading to a "work-sharing" effect (Ehrenberg and Schumann, 1981). However, an opposing scale effect implies that the employment response is ambiguous even within the labor demand model. Lastly, combining demand and supply in a compensating differentials framework implies that base wages would simply fall in equilibrium to fully offset the cost of overtime, leading to no real effects on hours, earnings, or employment (Trejo, 1991).

My paper empirically tests the competing models of overtime, and provides the first

causal estimates of its employment and income effects in the United States. I overcome prior empirical limitations by leveraging anonymous administrative payroll data that covers over one tenth of the U.S. labor force to evaluate the effect of recent federal and state expansions in overtime eligibility for low-income salaried employees. Unlike hourly workers, overtime eligibility for salaried employees is determined by their base pay relative to a legislated "overtime exemption threshold." All workers who earn below this salary threshold are guaranteed overtime protection, whereas white-collared salaried workers who earn above it are legally exempt. Between 2014 and 2021, there were two federal rule changes and nineteen state reforms that raised the overtime exemption threshold.[1]

To estimate the labor market effects of raising the overtime exemption threshold, I implement an event-study difference-in-difference design that compares the evolution of employment and earnings in states that increased their thresholds to those that did not. My empirical strategy assumes that absent the reforms, outcomes would have evolved similarly between both groups treatment and control states. To validate my methodology, I find no differential pre-trends between states that did and did not expand overtime protection, and any responses to the policy were immediately apparent on the exact month of the reform. Moreover, I estimate the impact of the policy along the entire pay distribution and, as a placebo test, show that the threshold changes had no effect outside the salary interval directly targeted by the policy (Cengiz et al., 2019). My results are thus robust to even state-specific shocks, unless those shocks occurred at precisely the month of the policy change and targeted only jobs between the old and new overtime exemption thresholds.

I report four set of results. First, on the exact month that the overtime exemption threshold increases, the number of salaried workers earning between the old and new thresholds falls by 21% (s.e. 6.1%). Three responses explain the dramatic change in the pay distribution for salaried jobs. 1) Three quarters of the decrease in jobs below the new threshold is accounted for by an increase in jobs right above it. This bunching in the distribution reflects firms' decision to raise workers' base pay above the new cutoff to keep them exempt from overtime. 2) Half of the remaining missing mass were jobs reclassified from salaried to hourly. Individuals in these jobs no longer receive a fixed salary, but are paid per hour of labor and qualify for overtime protection. 3) The remaining missing mass was lost due to a statistically insignificant reduction in employment.

Second, the expansion in overtime protection for salaried workers had spillover effects onto hourly jobs with comparable wages. Similar to the impact on the distribution of salaried jobs,

[1]Recently, the federal Department of Labor announced in August 2023 a new proposal to raise the threshold to $1,059 per week ($55,068 per year) - a policy that is expected to impact 3.6 million salaried workers (DOL Wage and Hour Division, 2023).

I find that firms raised the wages of some hourly workers such that their weekly base pay, defined as 40 times their wage, is bunched right above the overtime exemption threshold. Desipte the pay increase, I find no decrease in employment of hourly workers.

Third, incomes increased by about 1.3% for salaried workers targeted by the policy and 0.8% for hourly workers impacted by the spillovers. Although the policies aimed to expand overtime coverage, I find that the majority of the rise in earnings is attributed to an increase in base pay, rather than an increase in overtime compensation. Given that firms would only raise base salaries to keep workers exempt from overtime, it appears that the benefit of the reforms is unevenly distributed across workers and is concentrated primarily among a small group of workers who get bunched right above the overtime exemption threshold. Under the assumption that bunched workers are the only ones to receive an increase in base pay, I show that the salaries of this group increased by approximately 8.3%. At the same time, using data from the Current Population Survey, I find zero impact of raising the overtime threshold on workers' hours. Taken together, the evidence suggests that expansions in overtime coverage benefit workers by raising average salaries with no significant impact on either their hours or employment.

Fourth, to understand the mechanisms driving the labor market effects of overtime coverage, I use the empirical results to test three general theories of wage and hour determination and show that the evidence is consistent with a labor demand model whereby employers unilaterally set workers' hours. If wages and hours are determined jointly by the market, then compensating differential forces would reduce base pays to fully offset the costs of overtime coverage, leading to no real effects on the price or quantity of labor (Trejo, 1991). Due to a lack of policy changes, prior support for the compensating differentials model have relied on a negative cross-sectional correlation between wages and overtime hours (Trejo, 1991; Barkume, 2010). In contrast, using the threshold changes as a natural experiment, I find no evidence that firms cut salaries to negate the effects of overtime. Alternatively, if workers select their own hours, then a labor supply model predicts that no person eligible for overtime would ever choose to work exactly 40 hours per week (Idson and Robins, 1991). However, consistent with previous work, I find significant bunching in the distribution of hours at precisely 40 per week for hourly workers, suggesting there exists hour constraints within firms (Estevão et al., 2008; Goff, 2020).[2] Indeed, I find a labor demand model where firms choose hours and employment to maximize profits can explain the bunching, reclassi-

[2]Constraints by firms on workers' hours has also been proposed to explain the dynamics of hours among job-switchers (Altonji and Paxson, 1992), differences between micro and macro labor supply elasticities (Chetty et al., 2011), and the responsiveness of hours to tax reforms (Labanca and Pozzoli, 2022).

fication, and employment effects observed in the data. Within that framework, expansions in overtime coverage increase the cost of salaried jobs paid below the exemption threshold, thereby leading firms to reallocate away from those positions.

From a policy perspective, the empirical findings provide evidence against the historical intent that overtime coverage would encourage firms to create new jobs to substitute for a reduction in hours (Ehrenberg, 1971).[3] Given limitations in variation and data, tests of the work-sharing hypothesis of overtime in the U.S. have focused solely on its effects on workers' hours (Costa, 2000; Hamermesh and Trejo, 2000; Johnson, 2003; Trejo, 2003). Outside the U.S., rather than expansions in coverage, papers have instead examined policies that shortened the length of the standard workweek, and found a mix of zero and negative employment effects (Hunt, 1999; Crépon and Kramarz, 2002; Skuterud, 2007; Chemin and Wasmer, 2009; Sanchez, 2013). However, studies of reforms that shortened the workweek often lack adequate data to compute precise elasticities or examine the distributional impacts of overtime. To my knowledge, my paper provides the first causal estimates of the employment and income effects of overtime coverage in the U.S., and the first estimate of their implied elasticity in any context.

The remainder of this paper is organized as follows. In section II, I explain the institutional details governing U.S. overtime regulations and the specific policies that expanded coverage for salaried workers. Section III outlines the predictions of the competing models of overtime. In section IV, I describe the administrative payroll data from ADP. Sections V reports my main results on the aggregate employment and income effects of overtime. In section VI, I discuss the policy implications of raising the overtime exemption threshold for salaried workers. I conclude in section VII by discussing the implications of my findings and areas for future research.

II Federal and State Overtime Regulation

The Fair Labor Standards Act (FLSA) requires employers to pay workers one and a half times their regular rate of pay for each hour worked above 40 in a week.[4] While the overtime

[3]The significance of the work-sharing mechanism continues to be debated in recent policy proposals to introduce overtime protection for agricultural workers, with some arguing that a reduction in hours would actually decrease workers' earnings (Hsu and Bustillo, 2023).

[4]For hourly workers, the regular rate of pay is simply their wage. For salaried workers, the regular rate of pay is defined as their weekly salary divided by the number of hours for which the salary is intended to compensate (29 C.F.R. § 778.113). In practice, firms typically calculate salaried workers' regular pay rate as their weekly salary divided by 40. For example, a worker paid a salary of \$450 per week has an implied wage of $\$11.25 = \frac{450}{40}$. If the worker is covered for overtime, she would receive $\$16.88 = 1.5 \cdot 11.25$ for each hour above 40 that she works in a given week, in

premium applies to nearly all hourly workers in the U.S., the FLSA exempts a large group of salaried workers who are considered executive, administrative, or professional employees. To exempt a salaried employee, an employer must show that the worker performs primarily white-collared duties, and earns a salary equal to or greater than the "overtime exemption threshold" set by the Department of Labor (DOL).[5] Since the FLSA's overtime exemption threshold is not adjusted for inflation, the share of salaried workers earning less than the threshold, and thereby guaranteed overtime coverage, fell from over 50% in 1975 to less than 10% in 2016 (see Appendix Figure A.1).[6] In an effort to restore overtime protection to low-income salaried workers, such as managers at fast food restaurants and retail stores, Departments of Labors at both the federal and state levels have recently raised their overtime exemption thresholds. My paper uses these rule changes in the exemption threshold as natural experiments to study the effects of overtime coverage.

At the federal level, there have been three major policy proposals to increase the FLSA's overtime exemption threshold. First, the Department of Labor announced in May 2016 that it would more than double the federal exemption threshold from \$455 per week (\$23,660 per year) to \$913 per week (\$47,476 per year) effective December 1, 2016. The new rule would effectively raise the threshold from the 10th percentile of the salaried income distribution to the 35th percentile. However, to employers' surprise, a federal judge imposed an injunction on the policy on November 22, 2016, stating that such a large increase in the threshold oversteps the power of the DOL and requires Congressional approval. Following the retraction of the 2016 rule change, the DOL debated a smaller increase to the FLSA overtime exemption threshold and eventually raised the threshold to \$684 per week on January 1, 2020. Most recently, the federal Department of Labor announced in August 2023 that it plans to increase the threshold to \$1,059 per week (\$55,068 per year). Given that the federal policies affect all states simultaneously, I drop these reforms from my analysis.[7]

My analysis uses an event-study design to evaluate the 19 state-level increases to the overtime exemption threshold between 2014 and 2021.[8] Similar to the minimum wage, multiple

addition to her regular salary of \$450.

[5]The law also makes exceptions for special occupations such as teachers and outside sale employees. For a detailed overview of all exemptions, refer to Face Sheet #17A published by the DOL.

[6]In appendix figure A.2, I show that over the same time period, the share of salaried workers who say they would be paid for working more than their usual hours per week dropped from 27% to 12%.

[7]In a separate paper, I use the retraction of the 2016 reform as a natural experiment to study the role of fairness norms in determining persistence wage responses to policy changes. Although the 2016 federal proposal was never binding, I nevertheless find evidence that firms raised incumbents and new hires' salaries above the overtime exemption threshold, and this impact persisted even a year after the court injunction (Quach, 2020).

[8]Although the data ends in 2021, California, Colorado, and Washington state have continued to

states impose their own exemption thresholds that exceed the one set by the FLSA. I present in Figure I all state and federal thresholds from 2005 to 2021, along with the invalidated proposal in 2016.[9] My state-level analysis uses variation from six states: California, New York, Colorado, Washington, Alaska, and Maine. With the exception of Colorado, all six states define their overtime exemption thresholds as a multiple of their minimum wage. Thus, each time these states raise their minimum wages, the overtime exemption thresholds simultaneously increases following a known formula. While these policies happen concurrently, they target very different population groups. For example, in New York and California, which are the states with the most policy variation, the overtime exemption threshold is set at 75 and 80 times the minimum wage, respectively. Thus, a minimum wage worker would have to work at least 75 hours in New York to earn the overtime exemption threshold. I will show empirically that the overtime exemption threshold is high enough such that the segment of the income distribution affected by changes in the threshold does not interact with changes in the minimum wage, even after accounting for potential spillovers.

In addition to the increases in the overtime exemption thresholds, the nature of the regulation also provides another source of variation that can be used as a placebo check. Since the rule changes only directly affect salaried workers earning between the old and new thresholds, they should have little effect on workers with incomes much higher in the pay distribution.

III Theoretical Predictions

To guide my empirical analysis, I present a framework of wage and hour setting that nests three prominent models of overtime developed in the literature: a labor supply model, a labor demand model, and a compensating differentials model. I show in my framework that the effects of overtime differ depending on whether workers choose their own hours, employers set workers' hours, or earning-hour profiles are determined jointly by the market. I use my framework to generate testable predictions of each case.

III.a Set-up

Firm. The employer's production function depends on both the number of workers n and the number of hours per worker h. Following standard parametrizations (Ehrenberg, 1971;

increase their overtime exemption thresholds. For instance, Washington passed a law to gradually raise its threshold to $1,780 per week by 2028, which is equivalent to a salary of $92,560 per year.

[9] I exclude from my event study the five most recent rule changes in Alaska, which cumulatively increased the exemption threshold by only $47 to adjust for inflation. I also exclude the January 2014 event in New York due to missing data.

Hart, 2004), I characterize the firm's profit function by

$$\pi = xn^{\alpha}h^{\beta} - Y(h)n \tag{1}$$

where the parameter α introduces decreasing returns to scale, β allows for nonlinear returns to long work weeks, x is the price of the firm's output, and $Y(h)$ is the cost of hiring a worker for h hours.

The cost structure of a job depends on whether it is a salaried or hourly position:

$$Y(h) = \begin{cases} w(h + p(h-40)) + F & \text{if hourly} \\ S(1 + p\frac{h-40}{40}1[S < \bar{S}]) + R & \text{if salaried} \end{cases} \tag{2}$$

Hourly workers are paid a wage w for each hour of labor and receive a premium $p = 0.5 \cdot 1[h > 40]$ for each hour worked above 40 in a week. Salaried workers receive a fixed base pay S regardless of their hours and are only eligible for overtime compensation if their base pay is below the exemption threshold $\bar{S}$. I introduce the fixed costs F and R to capture the reduced-form value of classifying a job as salaried or hourly. For example, R encompasses the benefits (e.g. more flexibility, no need to monitor hours, etc.) and costs (e.g. easier to shirk, less control, etc.) of paying a worker by salary.[10] I assume the firm chooses whether a job is salaried or hourly.

Workers. There are M equally productive workers who differ only in their reservation wages, ordered by $r_1 < \cdots < r_M$. Workers' payoff from a job with weekly earnings Y and hours h equals the utility of consuming their income minus the disutility from working:

$$U(Y,h) = Y(h) - a^{-\frac{1}{\epsilon}}\frac{h^{1+\frac{1}{\epsilon}}}{1+\frac{1}{\epsilon}} \tag{3}$$

where ϵ represents the intensive labor supply elasticity and $Y(h)$ is the earnings from working h hours. I define extensive labor supply to be the number of workers for whom utility at (Y,h) is at least equal to their reservation wage: $N^s(Y,h) = j$ such that $r_j \leq U(Y,h) < r_{j+1}$.

[10] The salaried-hourly decision can be formally motivated by an agency problem where firms choose an occupation's pay classification depending on whether the number hours worked is informative of workers' effort and output (Fama, 1991). I take a reduced-form approach to focus on the costs of overtime.

Equilibrium. A market equilibrium is characterized by a set of hours, salaries, and wages (h^*, S^*, w^*) such that

1. *Hours:*
   - In the labor supply model, workers treat (S^*, w^*) as given and choose hours to maximize utility (i.e. $\frac{\partial U}{\partial h} = 0$).
   - In the labor demand model, firms treat (S^*, w^*) as given and choose hours to maximize profits (i.e. $\frac{\partial \pi}{\partial h} = 0$).
   - In the compensating differentials model, firms and workers view salaries as a function of hours (i.e. $(S(h), w(h))$). Hours satisfy both parties' first order conditions.
2. *Labor demand:* $N^d(Y^*, h^*)$ satisfies the firms' first order condition (i.e. $\frac{\partial \pi}{\partial n} = 0$).
3. *Salaried/hourly status:* The firm cannot increase profits by switching salaried/hourly classification.
4. *Profit Maximizing Salary:* The firm cannot increase profits by raising salaries (i.e. $\pi(S^*, h^*) \geq \pi(S, h^*)$ for every $S > S^*$). If decreasing base salaries reduces profits, then demand can exceed supply. Otherwise, the market clears: $N^d(Y^*, h^*) = N^s(Y^*, h^*)$.

The first three conditions ensure that firms and workers have no incentive to deviate hours, employment, or job classification in equilibrium. The last condition determines the equilibrium salary level by moving along the labor demand curve until either the market clears, or it is no longer profitable for firms to cut base salaries, which may occur if bunching base pays at the overtime exemption threshold is cheaper than paying overtime. I now examine how the equilibrium outcomes respond to an expansion in overtime coverage under three different assumptions about the manner in which hours are determined.

III.b Labor Supply Model

Suppose workers unilaterally sets hours as in traditional labor supply models (Blundell and MaCurdy, 1999). Taking wages and salaries as given, workers choose weekly hours to satisfy the first order condition implied by equation 3:

$$a^{-\frac{1}{\epsilon}} h = \begin{cases} w(1+p) & \text{if hourly} \\ \frac{S}{40} p 1[S < \bar{S}] & \text{if salaried} \end{cases}$$

Proposition 1. *If workers set their own hours, then workers eligible for overtime would never work exactly 40 hours per week.*

Proof. See Appendix B.a. ■

By introducing a kink in employees' budget constraint, workers are never indifferent to working exactly 40 hours per week. Although the policy variation targets salaried jobs, the first proposition nevertheless provides a clear null hypothesis that can be tested simply from cross-sectional data on hourly workers.

Proposition 2. *If workers set their own hours, an exogenous increase in p for salaried jobs will increase hours and reduce base pay. Effects on employment are ambiguous.*

Proof. See Appendix B.b. ■

An expansion in overtime coverage for salaried workers can be interpreted as an increase in the premium p from 0 to 1.5 for all jobs with base pays less than $\bar{S}$. Since workers choose hours, the larger overtime premium incentivizes them to work more hours. Moreover, workers now receive higher net earnings for any combination of base pay S and hours h, leading to an increase in extensive labor supply. On the other hand, labor demand falls due to both a scale effect from the increased cost of labor and a substitution effect from the increase in hours. The combination of an increased labor supply and decreased labor demand implies a fall in base salaries and ambiguous employment responses.

III.c Labor Demand Model

Next, suppose firms unilaterally set hours. This setting would be an extreme interpretation of the growing evidence that workers face hours constraints within firms (Altonji and Paxson, 1992; Chetty et al., 2011; Labanca and Pozzoli, 2022). Taking wages as given, the firm simultaneously chooses hours and employment to maximize equation 1. In the case of salaried workers, Appendix B.c shows that the first order conditions imply

$$n^D(S, h^*(S,p), p) = \left[\frac{x\alpha h^*(S,p)^\beta}{S(1 + p\frac{h^*(S,p)-40}{40}1[S \leq \bar{S}]) + R}\right]^{\frac{1}{1-\alpha}} \tag{4}$$

where

$$h^*(S,p) = \frac{40\beta(S(1 - p1[S \leq \bar{S}]) + R)}{(\alpha - \beta)Sp1[S \leq \bar{S}]} \tag{5}$$

Proposition 3. *If firms set hours and p exogenously increases for salaried jobs paying less than $\bar{S}$, then*

1. *Weekly hours will decrease*

2. *Employment will increase for sufficiently small β and decrease otherwise*

3. *Salaries will bunch at $\bar{S}$, with the missing mass coming from jobs paying directly below the exemption threshold*

4. *Jobs will be reclassified from salaried to hourly*

Proof. See Appendix B.c. ■

Following equation 5, increasing the marginal hourly cost of labor incentivizes employers to cut hours. Aside from increasing the absolute cost of labor, overtime coverage also distorts relative costs, making it relatively cheaper to hire workers than raise hours. From equation 4, the opposing scale and substitution effects lead to ambiguous employment responses.

$$\frac{dn^d}{dp} = \underbrace{\frac{\partial n^d}{\partial p}}_{\text{Scale Effect}<0} + \underbrace{\frac{\partial n^d}{\partial h}\frac{\partial h}{\partial p}}_{\text{Substitution Effect}>0}$$

Historically, when overtime was introduced during the Great Depression, policymakers had intended for the policy to raise employment, implicitly believing that the substitution effect exceeds the scale effect (Ehrenberg, 1971; Trejo, 2003). Appendix B.c shows that a sufficient condition for employment to rise is for β to be small. In effect, employers engage in work-sharing only if there is little value to employing workers for long hours so that it is inexpensive to reallocate hours across workers.

In addition to the possibility of work-sharing, the labor demand model also generates two additional predictions: bunching and reclassification. First, if paying workers a salary at the overtime exemption threshold (i.e. $\bar{S}$) is cheaper than paying them the market clearing rate plus overtime (i.e. $S^*(1 + 1.5\frac{h^*-40}{40})$), then firms would raise workers' salaries to exactly the threshold, leading to a bunching mass in the distribution of base pays. Second, since overtime increased the cost of salaried workers relative to hourly ones, jobs on the margin of being salaried are reclassified to hourly. Neither the bunching nor reclassification effects would occur in the labor supply model because if workers can set their own hours, salaried employees would choose to work the minimum number of hours to keep their jobs. Thus, firms would actually prefer that salaried jobs are covered for overtime to incentivize them to work longer.

Note that the reduction in hours apply specifically to salaried workers who do not get bunched above the threshold. Bunched workers may actually be required to work longer hours. Since whether a worker is bunched or not is an endogenous decision, I will only be able to test the aggregate impact on workers' hours, averaged over both groups. Nevertheless,

the prediction that hours decrease generates a useful corollary that allows me to distinguish between the labor demand and labor supply models using only cross-sectional data for hourly workers.

Proposition 4. *If firms set hours, then there would be a bunching mass in the distribution of weekly hours at precisely 40 per week.*

Proof. See Appendix B.d. ■

III.d Compensating Differentials Model

A criticism of the preceding models is that hours are unilaterally determined by one party. In reality, the preferences of workers and firms likely both play a role in setting hours. However, the discussion of the labor supply and labor demand models shows that regardless of the base wage, there is no weekly hours for which both the worker and firm are indifferent - workers would never choose a 40 workweek whereas employers would bunch hours at 40.

To reconcile the preferences of both parties, the literature has argued through the lens of a compensating differentials model that overtime would have no real labor market effects (Trejo, 1991). If the government mandates that employers pay workers a premium for each hour worked above 40 in a week, firms and workers could simply renegotiate a lower base pay such that after accounting for the overtime premium, total earnings remain unchanged. As a result, overtime would have no effect on either hours or employment.

In appendix B.e, I replicate the predictions of the compensating differentials model within the framework developed in this section. Unlike the labor supply and labor demand models wherein jobs are defined in equation 2 by either a fixed salary S or wage w, I assume the market determines a general earnings-hour profile $Y(h)$. Taking $Y(h)$ as given, firms and workers solve equations 1 and 3, giving the following first order conditions:

$$\frac{d\pi}{dh} = x\beta n^{\alpha} h^{\beta-1} - Y'(h)n = 0 \tag{6}$$

$$\frac{d\pi}{dn} = x\alpha n^{\alpha-1} h^{\beta} - Y(h) = 0 \tag{7}$$

$$\frac{dU}{dh} = Y'(h) - a^{-\frac{1}{\epsilon}} h^{\frac{1}{\epsilon}} = 0 \tag{8}$$

Together with the market clearing condition $n^d(Y,h) = n^s(Y,h)$, the four equations define an equilibrium characterized by four outcomes: employment n^*, hours h^*, earnings $Y(h^*)$, and wage $Y'(h^*)$.

Proposition 5. *If the market determines a wage-hour profile $Y(h)$, then expansions in overtime coverage would decrease base salaries but have no effect on total earnings, hours, or employment.*

Proof. See Appendix B.e. ■

An insight from the demand-supply framework relative to the compensating differentials model is that it provides equilibrium conditions under which neither workers nor firms have an incentive to deviate from the agreed upon weekly number of hours. To illustrate, suppose prior to the rule change, a job was working 55 hours per week for a wage of \$15 per hour. While the compensating differentials model predicts that wages would fall to entirely offset the new overtime premium, it does not define the conditions under which such an equilibrium would satisfy each party's marginal decision per hour. In the above example, it is often assumed that wages would decrease to \$13.2 per hour, so that gross earnings are constant (i.e. $15 \cdot 55 = 13.2 \cdot 40 + 1.5 \cdot 13.2 \cdot (55 - 40)$). However, 55 hours per week is no longer incentive compatible given a wage of \$13.2 per hour. Since the marginal cost of the 55th hour has increased to \$19.8, firms have an incentive to cut hours whereas workers have an incentive to increase hours. In the demand-supply framework, I show that for the same pair of weekly hours and earnings to be incentive compatible, the market will adjust the earnings-hour profile $Y(h)$ such that not only are hours and earnings held constant relative to the pre-policy values, but the implied wage at the equilibrium hours $Y'(h^*)$ is held constant too. To accomplish that while also satisfying the requirements of the law, firms could reduce wages by a factor of 1.5 and then pay a lump sum bonus to compensate for the reduction in wages. In the above example, wages would fall to \$10 per hour and firms would give workers a \$200 bonus for each week of labor (i.e. $15 \cdot 55 = 10 \cdot 40 + 1.5 \cdot 10 \cdot (55 - 40) + 200$).

III.e Testing the predictions

I summarize the testable predictions of the three models in Table I. In all cases, an expansion in coverage mechanically increases overtime pay. In response to this, if wages and hours are determined jointly, then base pay would decrease to fully offset the costs of overtime. Base pay would likewise fall in the labor supply model, but I would also expect a missing mass at 40 hours per week among hourly jobs. In contrast, if firms control hours, then I expect an increase in the overtime exemption threshold to lead to a bunching mass at the threshold coming from jobs that would otherwise be paid right below it, and reclassification of jobs from salaried to hourly. Employment may rise or fall in this model depending on the marginal productivity of long hours.

IV ADP Data

I use anonymized monthly administrative payroll data provided by ADP LLC, a global provider of human resource services that helps employers manage their payroll, taxes, and benefits. Their matched employer-employee panel allows me to observe monthly aggregates of anonymous individual paycheck information between May 2008 and July 2021. The data contains detailed information on employee's salaried/hourly status, income, hours, pay frequency (i.e. weekly, bi-weekly, or monthly), and state of employment for over a tenth of the U.S. labor force.[11] Since ADP manages both W-2 and 1099-MISC tax forms for employers, the data contains data on both regular employees and independent contractors. I aggregate across both types of workers to estimate the effect of overtime coverage on total employment and average earnings across all workers.

A significant advantage of the ADP data over traditional survey data and other administrative datasets is that it precisely records each worker's standard rate of pay, separate from other forms of compensation. This variable enables me to precisely calculate the measure of weekly base pay that the DOL uses to determine employees' exemption status. Following the DOL's guidelines, I compute salaried workers' weekly base pay as the ratio between their salary per pay-period and the number of weeks per per-period.[12] As a simple benchmark to compare the rate of pay for workers who transition between salaried and hourly status, I define the weekly base pay of hourly jobs as 40 times their wage.

In addition to workers' pay rate, the data also records employees' monthly overtime pay and monthly gross pay. For a given worker-month, the gross earnings variable is defined as the total pre-tax remuneration paid over all paychecks issued to the worker in that month, including overtime pay, bonuses, cashed-out vacation days, and reimbursements. To express gross pay and overtime pay in the same weekly denominator as base pay, I scale them by the number of paychecks received each month and the number of weeks per pay-period. Appendix C provides more detail on how I construct the two variables. For my analysis, I restrict the sample to a balanced panel of employers because the entry and exit of firms in the data reflect both real business formations and the decision of existing firms to partner with ADP.

While the data also records workers' total number of hours worked per month, employers

[11] For observations prior to 2016, I use workers' state of residence to proxy for their state of employment. This approximation is often implicitly assumed in papers that use the Current Population Survey. Testing the validity of this assumption in the post-2016 ADP data, I find that 95.5% of workers work in the same state that they live.

[12] For example, a salaried worker with a statutory pay of \$3000 per month would have a weekly base pay of $\$3000 * \frac{12}{52}$=\$692.31.

only accurately track this information for hourly employees. Since employers are not required to record the hours of salaried workers who are not covered for overtime, this limitation is likely endemic to all administrative employer datasets, including Census or IRS data. To overcome this limitation, I will supplement my analysis with survey data from the CPS Outgoing Rotation Group, which asks workers their weekly earnings, usual weekly hours, and the number of hours they worked in the previous week. I distinguish between salaried and hourly workers in the CPS ORG based on respondents' answer to whether or not they are paid by the hour.

In Appendix D, I explore the characteristics of firms affected by the changes in the overtime exemption threshold.[13] I show three descriptive results. First, I find that these expansions in overtime coverage affects workers across all industries, but primarily impacts large firms. Employers with at least one salaried worker in the interval of base pay targeted by the rule changes are about twice as large as the average firm in the sample. Second, plotting the distribution of weekly base pay, I find clear evidence of employers bunching workers' salaries right at the overtime exemption threshold, thereby keeping them exempt from overtime. Third, the probability that salaried workers receive overtime pay is highly responsive to the overtime exemption threshold. In California and New York, salaried employees earning right below the threshold have a 10-20% probability of earning overtime pay. However, the probability of receiving OT pay discontinuously drops by half for workers right above the cutoff, suggesting that employers are complying with the overtime regulations. In the next section, I use the state policy changes in the thresholds to identify the causal impact of raising the overtime exemption threshold on employment, earnings, hours, and salaried/hourly classification.

V Results

V.a Employment, Bunching, and Reclassification Effects

Empirical Strategy.

To start, I estimate the impact of the labor reforms on the distribution of weekly base pay. My analysis uses an event-study difference-in-difference design that leverages the 19 state rule changes. Intuitively, I compare the evolution of employment and earnings between states that raised their overtime exemption thresholds to states that did not. Formally, I estimate

[13] For a detailed analysis of the representativeness of the ADP data in general, refer to Grigsby et al. (2021). They find that while the data closely matches the demographics of workers in the Current Population Survey, it under-represents employment in firms with over 5000 employees relative to the Business Dynamic Statistics.

the following stacked difference-in-difference regression:

$$n_{jstkv} = \sum_{\substack{t=-6 \\ t\neq -1}}^{5} \sum_{k=-6}^{15} \beta_{kt} \cdot I_{stk} + \alpha_{skv} + \delta_{ktv} + \varepsilon_{jstkv} \tag{9}$$

where n_{jstkv} is the number of workers employed in firm j from state s at event-time t, with base pay in bin k for event v. I define each bin as a \$40 interval of weekly base pay, normalized to 0 at the new threshold. The treatment dummy I_{stk} equals 1 for the treatment state at event time t and bin k. I omit the month before the policy change as a reference period. My benchmark specification includes state-bin-event (α_{skv}) and month-bin-event (δ_{tkv}) fixed effects to control for state-specific differences in the base pay distribution and nationwide changes in inequality, respectively. Standard errors are clustered by state.

There are three features of the regression that differ from traditional difference-in-difference designs. First, rather than aggregating employment at the firm level, I follow recent advancements in the minimum wage literature and measure employment at the firm-bin level (Cengiz et al., 2019; Derenoncourt and Montialoux, 2019; Harasztosi and Lindner, 2019; Gopalan et al., 2020). The extra level of disaggregation allows me to estimate changes in employment along the entire income distribution, which provides two useful advantages over a single aggregate statistic: 1) I am able to test whether employers bunch workers' salaries right above the overtime exemption threshold, and 2) I can use jobs at the right tail of the income distribution, where I expect no effect of the policy, as a placebo check of the parallel trends assumption.

Second, instead of a two-way fixed effects model, I estimate a stacked difference-in-difference model by creating a distinct dataset for each of the 19 state reforms. Each dataset comprises of the treated state along with all the control states that have never increased their overtime exemption threshold yet up to the that point in time. I then append the 19 datasets together to estimate equation 9. By organizing the data in this way and interacting each of the fixed-effects with an event indicator, equation 9 is equivalent to estimating 19 individual differences-in-differences and then taking a weighted average of the treatment effects to compute β_{kt}. The advantage of estimating a stacked difference-in-difference relative to a traditional two-way fixed-effects model is that I avoid biasing my estimates from heterogeneous treatment effects over varying time periods (Goodman-Bacon, 2021; Baker et al., 2022; Roth et al., 2023; Freedman et al., 2023).

Third, for each event, I rescale the distribution of base pay for each state in the control group to exactly match the distribution in the treatment group in the month before the threshold change. Since employment is measured in levels rather than logs, even if firms

grow at the same rate in both groups, the state with the largest population will nevertheless gain more jobs simply because it had higher baseline employment levels. To account for this, I apply the following transformation to the observations in the control group:

$$\tilde{n}_{jstkv} = n_{jstkv} \cdot \frac{\bar{n}_{s=treat,t=-1,k,v}}{\bar{n}_{s,t=-1,k,v}}$$

where $\bar{n}$ is the average employment across all firms within a state. Given that the transformation is only applied to the control states and uses baseline characteristics, it is uncorrelated with the change in employment levels post-reform and will therefore not bias my results.

I estimate equation 9 separately for salaried and hourly workers. The aggregate change in employment is simply the sum of the β_k estimates across both class of workers. For interpretability, unless otherwise noted, I scale all estimates by the number of "affected salaried workers", which I define as the number of salaried workers between the old and new thresholds in the baseline month.

My identification strategy relies on the assumption that absent the state threshold changes, the distribution of base pay in the treated states would have evolved the same as the control states. I test my identification assumption in two ways. First, I show that treated and control states do not exhibit differential trends prior to the reform. Second, I show that treated and control states exhibit similar trends *post-reform* for workers paid well above the overtime exemption threshold who are unlikely to be affected by the policy. Given these two validation checks and the high frequency nature of the data, my estimates would only be biased if there is a shock that occurs at precisely the month of the policy change, in precisely the treated states, and affects solely jobs within the interval targeted by the overtime threshold, but not jobs higher in the income distribution.

Estimates of the Employment Effect Along the Distribution of Base Pay.

Figure II plots the estimates of the treatment effect, β_k, separately for the distribution of salaried and hourly workers. Panel (a) shows that immediately in the month a new threshold goes into effect, there is a decrease in the number of salaried employees below the threshold and a spike in workers right above it.[14] As a placebo check, I find no effect on any bins of base pay above the new threshold. To further validate my empirical strategy, panel (b) plots the change in the number of salaried workers paid below and above the new threshold over time. Examining the figure from left to right, three features stand out. First, there is little evidence of a pre-trend prior to the policy change, indicating that the parallel trends

[14]The largest impact occurs within -80 to 40 above the new threshold simply because the majority (15 out of 19) of the policy changes raised the overtime exemption threshold by at most $80 weekly base pay.

assumption holds. Second, there is a sharp drop in the number of jobs below the threshold and a sharp increase in the number of jobs above it at precisely the month of the rule change, consistent with the bunching from the cross-sectional estimates. Third, the magnitude of the decrease in employment below the threshold is visibly larger than the increase in employment above it.

Plotting analogous figures for hourly workers, I find that the base pay distribution for hourly jobs responded in a qualitatively similar fashion. Panel (c) of Figure II shows that raising the state overtime exemption threshold cut hourly jobs earning between the old and new thresholds, and increased the number of jobs above it. In panel (d), I confirm that the bunching effect is statistically significant by plotting the effect on hourly employment over time. Mirroring the estimates for the salaried distribution, I find no significant pre-trends, followed by a sharp divergence in the number of jobs below and above the threshold at exactly the month of the rule change. However, the number of jobs below the threshold appear to trend upwards in the months following the reform. The bunching of hourly employees who were unaffected by the policy is consistent with growing evidence of relative pay concerns within firms (Card et al., 2012; Dube et al., 2019; Quach, 2020).

Contrary to the compensating differentials model of overtime, I find no evidence that employers reduced workers' base pay to offset the costs of overtime. The majority of the policy reforms only increased the overtime exemption threshold by $80 of weekly base pay, and yet I find no increase in the number of salaried workers even $240 below the new threshold. For hourly workers, I do see a small change in the number of jobs earning in the left tail of the distribution in Figure II, but this can be attributed to changes in the minimum wage. In particular, if a state increased its overtime exemption threshold by $80 and the new minimum wage is within $160/week (i.e. $4 per hour) of the old threshold, then the effect of the minimum wage would be reflected in the figure. To avoid any possible contamination with the minimum wage and other differences in state trends, I only aggregate the estimates from -$160 to $80 of normalized base pay when computing the employment effects.

In the appendix, I present two pieces of evidence that my preferred estimate of the employment effect is not confounded by changes in the minimum wage. First, appendix figure A.3 shows the distributional impacts of the overtime rule after omitting all states with a new minimum wage within $200 of the old overtime exemption threshold.[15] As expected, I no longer observe a change in the left tail of the hourly workers' distribution. Nevertheless, the sum of my estimates are very similar to the earlier results with the full sample. Second, appendix figure A.4 extends the estimates further to the left and right of the pay distribution.

[15] This restriction excludes 5 events and is equivalent to dropping Colorado, Maine, and Washington from the data.

Similar to the truncated results, I find no statistically significant impact at any point above the new overtime exemption threshold, suggesting that the treatment and control states are following similar post-trends after the reform. In contrast, there is clear bunching from the minimum wage in the lower end of the pay distribution, particularly for hourly jobs. While the effect of the minimum wage is multiple times larger than the impact of the overtime reforms, its impacts are concentrated primarily in the far left tail of the distribution, away from the bounds affected by the change in the overtime exemption threshold.

Table II summarizes my results for the employment effects separately for salaried and hourly jobs. The estimates in column 1 are computed from equation 9 and correspond to those presented earlier in figure II. Panel (a) reports the change in the number of jobs along the income distribution for each affected salaried worker. I find that the number of salaried jobs in the affected interval of base pay falls by 20.9% (s.e. 6.1%). Among those jobs, 15.8 p.p. (s.e. 5.6) or about three-quarters of the missing mass can be accounted for by employers bunching salaried workers right above the new overtime exemption threshold. Similarly, I find that for each affected salaried worker, the policy reduces the number of hourly jobs paid between the old and new thresholds by 0.07 (s.e. 0.026) workers, all of which can be accounted for by bunching above the cutoff.

Panel (b) reports two measures of the aggregate employment effect based on the four estimates in panel (a). First, the sum of the above estimates imply that employment falls by 2.3 (s.e. 2.3) jobs for every 100 salaried workers earning between the old and new thresholds at baseline. However, while the number of affected salaried jobs is a useful denominator for thinking about the share of salaried workers that receive a raise above the threshold, it makes more sense to include hourly workers in the denominator when calculating a labor demand elasticity. There are at least three reasons for this distinction. First, it is clear from Figure II that the policy affected the wages of not only salaried workers, but also hourly ones. Second, the change in aggregate employment (i.e. the numerator in $\frac{\Delta \text{Employment}}{\text{Employment}_0}$) is calculated as not only the change in salaried employment, but also hourly. Lastly, previous studies in the minimum wage literature also consider both salaried and hourly workers as treated, so to make my results comparable, I use a similar denominator.

Given the above rationale, panel B also reports the change in employment as a percentage of "all affected workers", which I define to be the total number of salaried and hourly workers with base pays between the old and new thresholds. The increase in the overtime exemption threshold had no statistically significant impact on aggregate employment, and the 95% confidence bounds can rule out employment losses greater than 1.5% and employment gains above 0.5%.

To assess the robustness of my results, I report a series of alternative specifications

that control for geographically localized shocks and variation in firm composition across states. In column (2), I show that the estimate of the employment effect is similar even if I compare states within the same Census division. This specification eliminates any bias from spatial shocks that target only specific regions of the U.S. (Dube et al., 2010). In column (3), I introduce firm-interacted fixed effects to compare only firms that operate in both the treated and control states. This stricter specification controls for differences in time-trends across states that may arise due to industry or even firm-specific shocks. While the sign of the employment effect flips, the estimates nevertheless suggests only minor changes in employment at most. In column (4), I restrict the sample to only policy changes where the old overtime exemption threshold is at least $200 weekly base pay above the new minimum wage. The sample restriction thereby eliminates concerns regarding interactions between the two policies. Column (5) imposes all the restrictions and controls from columns (2)-(4). Lastly, column (6) estimates a traditional two-way fixed effects model where the outcome variable is simply the total number of jobs between -$160 and +$80 weekly base pay relative to the new threshold. Overall, the culmination of evidence suggests that the aggregate employment was small relative to the total number of affected workers.[16]

Decomposing the Changes in the Number of Workers.

The results thus far are consistent with the predictions of the labor demand model whereby firms set workers' hours. However, multiple questions remain. First, aside from the bunching effect, the model predicts that firms would also reclassify workers from salaried to hourly - did that happen? Second, it takes time for firms to adjust the stock of employment. Did firms instead change their hiring or separation rates?

To answer these questions, this section studies the impact of increasing the overtime exemption threshold on the flow of workers. Intuitively, changes in the stock of workers within a bin of base pay for a particular hourly/salaried classification can be decomposed into three margins of response: flows across bins within the same classification (i.e. wage growth), flows across hourly/salary status (i.e. reclassifications), and flows in/out of employment. I measure each of these flows directly using the employer-employee panel structure of the data, and use the number of such flows at the firm level as the outcome variable to equation 9.

Figure III plots the effect of the state policies on net employment flows (i.e. hires minus separations) and net reclassification flows. In panel (a), I find a decrease in employment

[16] Since the stacked event study is a weighted sum of the individual difference-in-difference estimates, I also test whether the results are driven by any outlier events. Appendix figure II plots the difference-in-difference estimates separately for each policy change. I drop Maine from the figure since the estimates are very noisy given the small sample size in the state. While there appears to be one outlier in New York 2016 with a very large negative response, the other estimates tend to hover around 0, with no particular pattern towards a negative or positive employment effect.

flows among salaried workers paid between the old and new thresholds at precisely the month of the reform, and an increase in flows above the threshold. Interestingly, the change in employment flows does not persists over time. Turning to hourly jobs, the estimates in panel (b) are far noisier and exhibit a downward pre-trend. In either case, there does not appear to be a persistent change in employment in the month of the reform, nor afterward. The lack of an effect on salaried employment flows suggests that the decline in salaried jobs observed earlier in figure II is due to the reclassification of jobs from salaried to hourly.

To directly measure the reclassification effect, panels (c) and (d) plot the impact of the overtime reform on the flow of jobs out of and into the salaried distribution, respectively. Consistent with the labor demand model, I observe a sharp increase in the number of jobs being reclassified from salaried to hourly at precisely the month that the new overtime exemption threshold goes into effect. This impact solely affects jobs paid below the threshold. On the other hand, there does not appear to be a large impact on the number of jobs being reclassified from hourly to salaried, which usually occurs following a promotion. However, jobs moving from hourly to salaried are now more likely to be paid above the overtime exemption threshold. Overall, I find clear evidence of a reclassification effect that was obscured when examining the stock of workers.

Table III summarizes the effect of raising the overtime exemption threshold on employment and reclassification flows. Given that the bulk of the flow responses occurred at precisely the month of the policy reform, I report the β_{tk} estimates from equation 9 for $t = 0$. All estimates are scaled by the number of affected salaried workers. In columns (1)-(3), I present the change in hires, separations, and net employment flows, respectively. The first row finds that there is a precisely estimated decrease in net employment flows to salaried jobs paid below the new overtime exemption threshold, with a confidence bound between -1 to -0.2 jobs per 100 affected salaried position. Half of the reduction is driven by a decrease in hires and half is from an increase in separations. However, the second row shows that about two-thirds of the drop in employment flows below the threshold can be account for by new jobs above it.

As for hourly jobs, the third and forth rows report a negative impact on hourly employment concentrated primarily among jobs paid below the new threshold. The response appears to be driven more by an increase in separations than a decrease in hires. However, as noted in the figures, there is a already pre-trend prior to the policy reform so I am cautious to interpret the change in hourly employment flows as causal. Nevertheless, column 3 shows that on net, the estimates imply a cumulative decrease in employment flow of approximately 0.021 (s.e. 0.028) jobs per affected salaried worker. Despite using a different method to compute the employment effect, my estimate from examining changes in employment flows is

very close to the estimate in table II calculated using changes in the stock of workers.

Relative to the employment effects, the change in reclassification flows are very precisely estimated even for small responses. The first two rows correspond to panel (c) and (d) from figure III where I observe a net flow of jobs from salaried to hourly primarily among jobs being paid below the new threshold. The magnitude of the estimates suggest that employers reclassify approximately 2.2% of affected salaried workers from salaried to hourly. While reclassifying workers has no impact on their overtime eligibility, it enables employers to reduce workers' pay if they work less than 40 hours per week instead of simply paying a fixed weekly salary. In the third and fourth rows, I show that there is a corresponding increase in the number of hourly positions both below and above the threshold. This suggests that newly reclassified workers are paid a weekly base pay similar to what they earned in their previous title.

Overall, the evidence from the analysis on flows suggests that raising the overtime exemption threshold has no significant impact on employment, but leads to a reclassification of jobs from salaried to hourly. In terms of magnitudes, the reclassification effect is much smaller than the bunching effect. Taken together with the results from table II, the estimates suggest that three quarter of the reduction in affected salaried jobs below the threshold is due to bunching, half of the remainder is accounted for by reclassification, and the residual noise is due to employment loss.

V.b Income Effects: Base Pay and Overtime Pay

In this section, I estimate the effect of raising the overtime exemption threshold on workers' incomes using a difference-in-difference design. As in section V, my identification strategy compares workers in states that raised their overtime exemption thresholds to similar workers in states that did not. My baseline regression is

$$y_{isvt} = \sum_{t=T_0}^{T_1} \beta_t \cdot I_{st} + \alpha_{vs} + \delta_{vt} + \varepsilon_{isvt} \tag{10}$$

where y_{isvt} is the weekly earnings of worker i in state s at event time t for event v, and I_{st} is an indicator that equals 1 at month t for workers in the treatment group. I control for event-state (α_{vs}) and event-month (δ_{vt}) fixed effects. My identification strategy assumes that the wages of workers in the treated states would have evolved at the same rate as the control states absent the policy. As before, I validate my empirical strategy by checking that the parallel trend assumption holds pre-reform and that my results are robust to a series of alternative specifications. Standard errors are clustered by state.

I make two sample restrictions. First, to focus on workers directly targeted by the policies, I restrict the sample for each event to workers who are paid a weekly base pay between the old and new thresholds in the month prior to the policy change. I make the same restriction for the control groups to create a comparable counterfactual. Second, since I do not observe the wages of workers in firms not using ADP's software, I restrict the sample to workers who are continuously employed at the same firm in all months of the event window. I acknowledge that the sample restriction may introduce selection bias if the policies had an impact on the separation of workers. Given that I did not find strong employment effects, this is likely a minor issue. Nevertheless, I will show that despite the potential differences in composition, workers in the treated and control states still had very similar pre-trends.

Figure IV plots the difference-in-difference estimates for base and overtime pay, separately by whether the worker was salaried or hourly at baseline. I estimate all regressions in levels rather than logs to be able to compare the dollar increase in weekly base pay relative to weekly overtime pay. Starting with salaried workers, three key features stand out from this analysis. First, consistent with the parallel trends assumption, both weekly base pay and overtime pay were trending similarly between the treatment and control states prior to the announcement of the new rule. Second, workers experience a sharp jump in base pay and a minor increase in overtime pay at precisely the month that the threshold increases, and this raise remains fairly stable afterwards. Lastly, workers experience a far larger increase in base pay than overtime pay, suggesting that bunched workers gained the most from the policy. The rise in salaries further rejects the prediction of the compensating differentials model that firms would cut workers' base pay to nullify the costs of overtime.

Turning to hourly workers, I find a similar pattern of parallel pre-trends, a sharp increase in earnings at the month of the reform, and continued elevated earnings afterwards. However, there are two differences to note from salaried workers. First, the increase in base pay is far smaller on average for hourly workers than salaried workers. This is to be expected as not only did fewer hourly workers receive a raise above the threshold, but there are also more hourly workers than salaried workers at baseline. Second, the increase in overtime pay is more volatile than for base pay. However, the steady elevated level of overtime pay post reform suggests that despite the volatility, the overtime reforms had a persistent small positive impact on hourly workers' overtime earnings.

Table IV summarizes the estimates of the income effects from expanding overtime coverage. All estimates are computed from equation 10, and reported for the first month of the reform to ease computational burden. In the first column, I show that the base pay impact is an order of magnitude larger than the increase in overtime pay, suggesting that most of the benefits of the policy for workers is realized through the bunching effect. Summing the

effects on base and overtime pay, then dividing by baseline incomes, the overtime reforms increased average total income of salaried workers by 1.4% (s.e. 0.2%). In the last row, I show that results are similar if I simply estimate the difference-in-difference using log total incomes as the outcome.

Column (2) reports analogous estimates for hourly workers. In this case, I find a smaller increase in weekly base pay, but a larger increase in overtime compensation. Overall, the earnings of hourly workers between the old and new thresholds increased by about 0.9% (s.e. 0.2%). Although the magnitude of the income effect for hourly workers is smaller than that of salaried workers, it nevertheless highlights that raising the overtime exemption threshold has strong spillover effects onto hourly workers too.

I implement three tests of the robustness of my results. First, in columns (3) and (4), I control for Census-division interacted with time fixed effects, thereby only comparing states within the same Census division over time. Second, columns (4) and (5) control for event-firm-month and event-firm-state fixed effects. This specification compares workers who work at the same firm but reside in different states. Lastly, columns (7) and (8) drops events where a state simultaneously raised the minimum wage at the same time it increased the overtime exemption threshold, and the new minimum wage is within \$200 weekly base pay of the old threshold. Across all specifications, I find a clear increase in salaried workers' earnings by around 1%. I find a slightly smaller increase in hourly workers' earnings, and the estimate shrinks significantly once I compare workers within the same firm. That perhaps suggest that the initial spike in overtime in panel (d) of figure IV is driven by firm-specific changes and the composition of firms across states.

Taken together, the evidence shows that expanding overtime coverage increases the earnings of both salaried and hourly workers. However, the increase in base pay is far larger than the increase in overtime pay, suggesting that employers would rather pay to keep workers exempt from overtime than simply pay them overtime compensation. I conduct a simple calculation to understand the magnitude of the bunching effect and the share of workers who are benefit from such a raise. In theory, the only reason firms would raise base pays in response to the overtime policies is to pay workers above the new overtime exemption threshold. Assuming that the rise in base pay accrues entirely to bunched workers, then the previous analysis in table II shows that about 16% of affected salaried workers drive the entire \$11.54 increase in weekly base pay. Together, the estimates imply that the weekly base pay of bunched workers increase by $\frac{\$11.54}{0.16} = \$72.13$. Given a baseline weekly salary of \$870, the simple calculation suggests that the earnings of the 16% of affected workers increased by approximately 8.3%, whereas remaining affected workers only experienced minor increases to their earnings from overtime pay.

V.c Hours Effects

In this section, I analyze the effect of raising the overtime exemption threshold on workers' hours. My analysis proceeds in two parts. First, I explore the distribution of hours in the ADP data. However, given that employers are not required to record the hours of salaried workers exempt from overtime, this preliminary analysis with the ADP data is limited to only hourly workers. Second, I supplement my analysis using data from the CPS Outgoing Rotation Survey, which asks respondents' their usual weekly work hours and salaried/hourly status.

To start, I use the ADP data to test one of the fundamental distinctions between the labor demand and labor supply models of overtime: bunching in the distribution of weekly hours. Similar to overtime pay and gross pay, the variable for hours is aggregated over all hours worked in a month and I convert it to an average weekly value following the procedure in appendix C. Appendix Figure A.6 plots the distribution of average weekly hours for all hourly workers in the baseline month. Despite the imprecision from converting monthly hours to an average weekly value, there is a clear bunching mass at precisely 40 hours per week. The bunching is consistent with recent evidence from Goff (2020) that finds a similar result using administrative payroll data at the weekly level. Such bunching at the overtime kink point is contrary to the labor supply model and adds to growing evidence that employers play a pivotal role in determining workers' hours (Labanca and Pozzoli, 2022).

While the cross-sectional distribution of hours provides a useful test of the competing models of hours determination, it does not contain any information on the impact of raising the overtime exemption threshold. In theory, the response can be in either direction. Newly covered salaried workers may see a decrease in hours leading to a bunching mass at 40. However, those who receive a raise to above the overtime exemption threshold may actually be asked to work more hours to compensate. Given the theoretical ambiguity, I next test for the hours response in the CPS. Specifically, I estimate the following regression:

$$\log(Y_{it}) = \alpha_{iv} + \alpha_{tv} + \sum_{q=-4}^{3} \beta_q I_{sq} + \varepsilon_{it} \tag{11}$$

where Y_{it} is the outcome variable for respondent i at month t. I estimate the stacked difference-in-difference regression controlling for event-individual (α_{iv}) and event-month (α_{tv}) fixed effects. The treatment interacted with event-time indicators (I_{sq}) are aggregated at the quarterly level, q. For statistical power, I expand my sample period to include all 30 state overtime threshold changes from 2008 to 2023. I cluster estimates at the state level.

Note that by controlling for individual fixed effects, my preferred specification follows

the same individuals over time. As such, the regression is identified from individuals who are employed both pre and post reform. Since the CPS only surveys respondents over 16 months, some individuals may not exist throughout the entire panel. Moreover, respondents in the CPS are asked their salaried/hourly status and weekly earnings only in the 4th and 16th month they are surveyed. As such, the sample of workers used to identify each β_q changes each quarter. Workers are however asked their weekly hours in every wave of the survey, so the sample changes less often when examining the effect of the policy on hours. In either case, β_q is interpreted as the percent change in Y_{it} within-individual between the treated states relative to the control states.

My main specification restricts the sample to workers who reported working a salaried job during the 4th wave of the survey, with weekly earnings between the old and new overtime exemption thresholds. As a placebo check, I also estimate the difference-in-difference for workers earning up to $100 per week above the new threshold.

Figure V presents my estimates of equation 11 for two outcome variables: log weekly earnings and log weekly hours worked last week. First, as a validity check on the quality of the data, I replicate my earlier analysis of the income effect of expanding overtime coverage for salaries workers. One concern with the CPS is that the weekly earnings variable is measured with self-reporting error, so I cannot perfectly identify workers directly affected by the reforms. Despite the measurement error, panel (a) shows similar pre-trends between the treated and control states, and an increase in average weekly earnings immediately following the expansion in overtime coverage. As a placebo check, panel (b) finds no such immediate impact for salaried workers already earning above the overtime exemption threshold. The successful replication of the income effect, although imprecisely estimated, builds confidence that the earlier results in the ADP data are externally valid.[17]

After confirming that the CPS data appears adequate enough to identify workers affected by the overtime reforms, I next estimate the effect on workers' hours. Panel (c) shows that weekly hours increase in the 2nd and 3rd quarters after the expansion in overtime coverage. However, while hours increased relative to the quarter immediately before the policy change, it is not significantly different from even earlier quarters. Hours also return back to their baseline levels 4 quarters after the expansions in overtime. Overall, I find weak evidence of an increase in hours, but nothing persistent. Panel (d) shows that hours did no change for the placebo group relative to the year before the policy change.

Table V assesses the robustness of my analysis of the CPS data. First, the estimates in column (1) are analogous to figure V except I aggregate the dynamic effects using a simple

[17] Given the relatively small sample size of the CPS, I do not have enough statistical power to replicate the bunching results.

indicator for the treated states post-reform. Consistent with the figure, I find a 3.7% (s.e. 1%) increase in weekly earnings among the treatment group, no impact on income of the placebo group, and no effects on hours for either sample. The estimates of the hours effect are precise enough that I can rule out any positive increase in hours above 1.2%. Column (2) adds demographic controls for gender, race, age, and age-squared. Lastly, in column (3), I interact the month fixed effects with the Census division of the survey respondent. This stricter specification accounts for region-specific time trends so that I am comparing states in the same Census division over time. I find that the magnitude of the estimates all remain fairly stable over the three specifications. Taken together, the evidence suggests that raising the overtime exemption threshold increases workers' earning, without significantly impacting workers' hours.

VI Policy Implications

To benchmark the costs and benefits of raising the overtime exemption threshold relative to other labor market policies, table VI computes the ratio of the employment and income effects. The elasticity is computed by jointly estimating the employment and income effects, and then calculating their ratio using the delta method.

In column (1), I report the estimates from my baseline firm-level regression. The elasticity in panel (A) is equivalent to taking the ratio of the employment effect with respect to affected salaried workers in column 1 of table II and the income effect for salaried jobs in column 1 of table IV. The ratio suggests that for each percent increase in affected salaried workers' earnings, the number of affected jobs falls by 1.9% (s.e. 2.4%). However, this statistic likely overstates the percent change in employment because it assumes that the entire employment loss comes from salaried jobs. As evident from the earlier analysis, there is significant spillover effects onto hourly jobs in terms of both income and employment. To account for such spillovers, panel (B) reports the estimates once I treat all workers between the old and new thresholds as affected by the policy, regardless of their salary/hourly status. As expected, both the percent change in employment and income fall significantly. Taken together, the estimates imply that the number of affected jobs fall by -0.59% (s.e. 0.55%) for each percent increase in workers' earnings.[18] My 95% confidence bound can rule out elasticities smaller than -1.66 and larger than 0.48.

The remaining columns of table VI assess the robustness of my estimate to a series of

[18] If overtime coverage also imposes administrative costs in additional the increased payroll costs, then the ratio of the employment and income effects alone does not represent a labor demand elasticity. Nevertheless, this ratio is still a policy-relevant statistic for gauging the cost and benefits to workers. I thank Steve Trejo for this clarification.

stricter specifications checks. Focusing on panel (B), I find that the magnitude of the elasticity is stable to including Census-division fixed effects, dropping events where a concurrent change in the minimum wage risked interacting with the change in the overtime exemption threshold, or simply estimating a stacked two-way fixed effects model. However, the elasticity flips signs and becomes much noisier if I only use within-firm variation. Overall, the point estimates of the elasticity appear fairly small, ranging from -0.75 to 1.1.

To my knowledge, this is the first study to compare the employment effects of overtime to its income effects. As such, there are no estimates in the literature with which I can make a one-to-one comparison. Given the bunching in the pay distribution, the most similar benchmark for reference would be from studies of the minimum wage. A meta-analysis by Dube (2019) finds a median elasticity of employment with respect to own wage of -0.17 across 36 studies of the minimum wage in the U.S, with a range of -3 to 1.5. Relative to these studies, I examine a higher-income population where the implied hourly wage of affected salaried workers is about $21.75/hour. As such, there is no reason why the overtime exemption policy should have comparable effects. Nevertheless, I find small elasticities within a similar range of those in the minimum wage literature.

VII Discussion and Conclusion

This paper presents new facts about the labor market effects of expanding overtime coverage and informs policy debates surrounding recent initiatives to raise the overtime exemption threshold. In this section, I summarize my findings by comparing the estimates of the effects of the 2016 FLSA policy to the predictions in the Department of Labor's cost-benefit assessment. To generate these predictions, the DOL conducted a thorough review of the literature on overtime and used existing labor demand elasticities to infer from the Current Population Survey the expected effects of their proposed reform.

My empirical results differ from the conclusions of the Department of Labor in four ways. First, the DOL believed that by increasing the marginal cost of labor per hour, "employers have an incentive to avoid overtime hours worked by newly overtime-eligible workers, spreading work to other employees" (U.S. Department of Labor, 2016). In contrast, I estimate that expansions in overtime coverage actually have little effect on employment. Second, while the DOL accurately predicted that average weekly earnings would rise, they calculated an income effect of only 0.7%, whereas I show that earnings increased by nearly twice that amount for salaried workers. I also show that this positive income effect was not uniformly distributed across the range of affected base pays, and primarily benefited a small group of workers who receive a raise above the threshold. Third, drawing from previous studies of the

compensating differentials model of overtime, the DOL calculated that 18% of workers would experience a decrease in base pay to partially offset the increase in overtime pay. However, I find no evidence that firms reduced base pays in response to overtime coverage. Fourth, the DOL considered the reclassification effects of the policy negligible given the available evidence at the time. In contrast, I find that approximate 2.6% of affected workers are reclassified from salaried and hourly, and the policy appears to have noticeable spillover effects on the income and employment of hourly workers.

Although my paper offers the most comprehensive evaluation of the overtime exemption policy to date, there are many avenues for future research that are beyond the scope of this study. Similar to the minimum wage, overtime can potentially have important implications for consumers (Harasztosi and Lindner, 2019), inequality (Lee, 1999; Autor et al., 2016), and productivity (Coviello et al., 2021). Unlike the minimum wage though, there did not even exist estimates of the employment effects of overtime eligibility prior to my study (Brown and Hamermesh, 2019). This paper adds to the scarce literature on overtime coverage by providing the first causal estimates of its employment and income effects. In future endeavors, it would be fruitful to explore the effects of overtime coverage on additional margins of responses.

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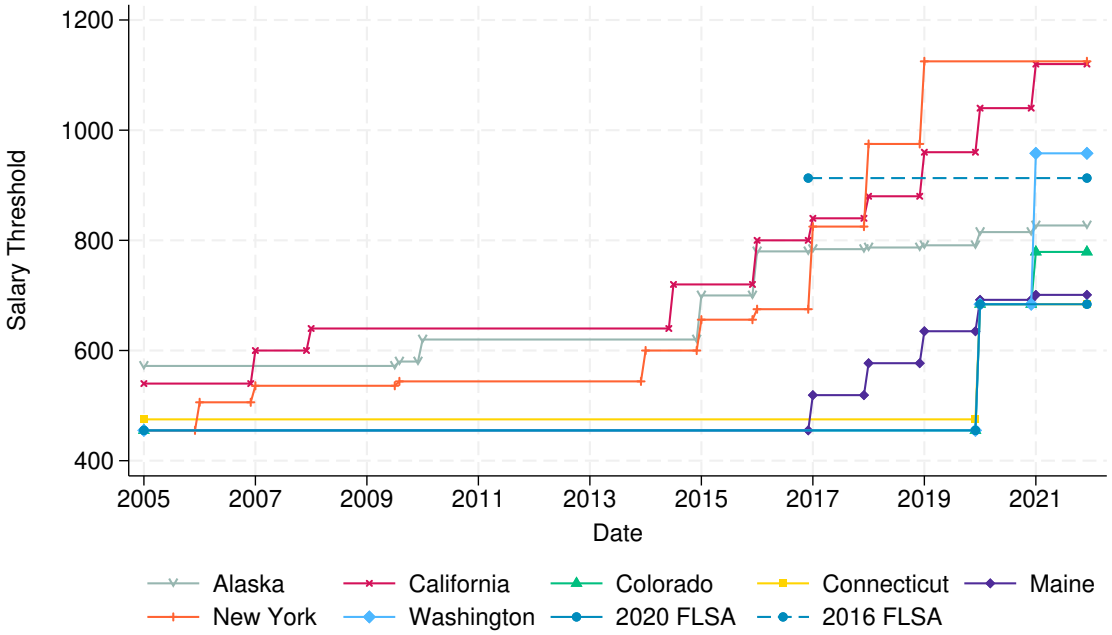
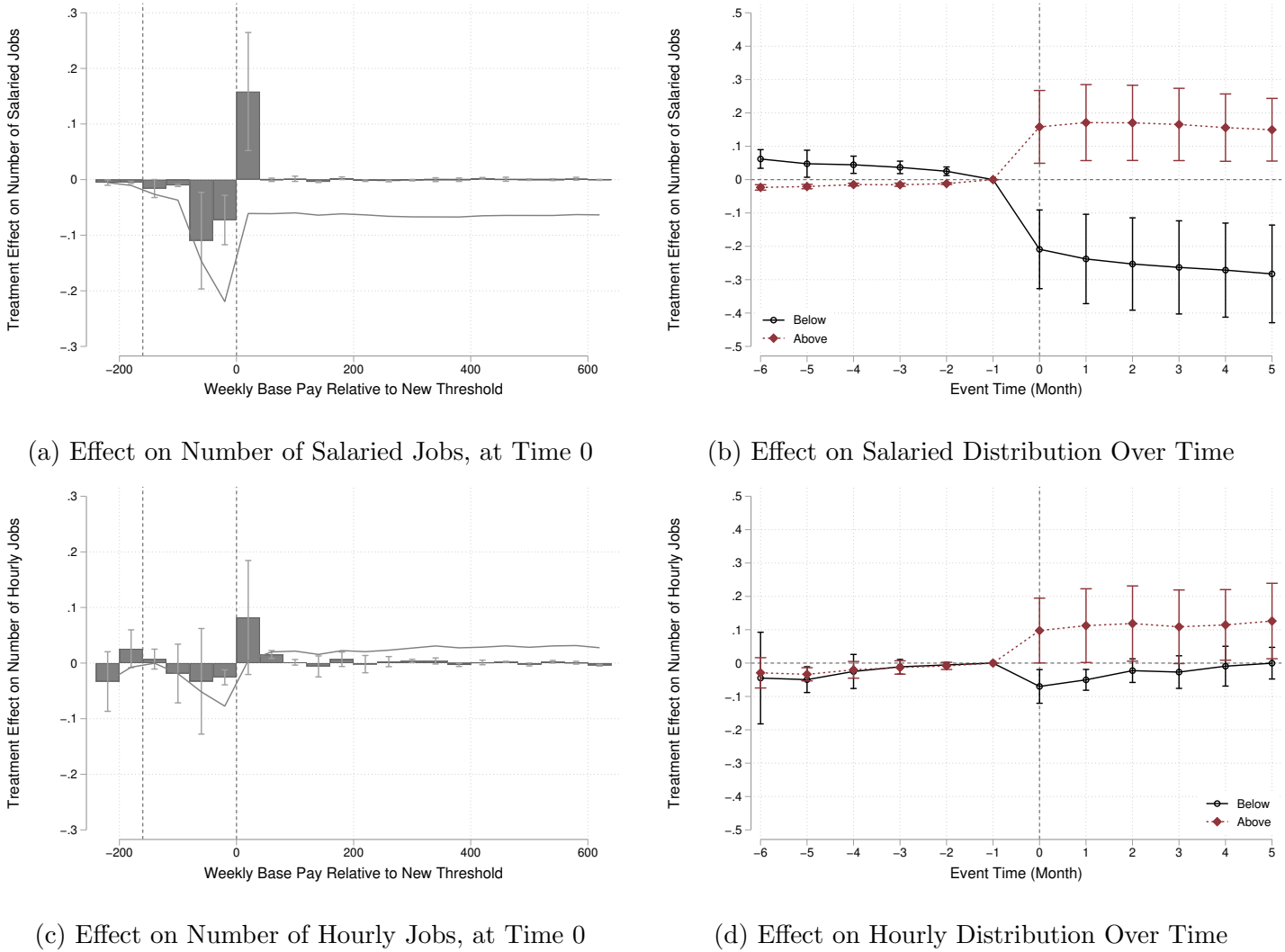


Figure I
Variation in Overtime Exemption Thresholds Across States

Notes. This figure shows the binding overtime exemption threshold in each state between 2005 and 2020. All states not explicitly included in the graph are covered by the Fair Labor Standards Act (FLSA). The line "2016 FLSA" represents the federal threshold that was supposed to go into effect on December 1, 2016 but was nullified in November 2016. In Alaska and California, the threshold equals 80 times the state minimum wage. In New York, the threshold equals 75 times the minimum wage. In Washington, the overtime exemption threshold equals 70 times the minimum wage in 2021. In Maine, the threshold equals 3000/52 times the minimum wage. Colorado's overtime exemption threshold is not pegged to the minimum wage. Starting in January 2017, the minimum wage and threshold varies by firm size in CA, and county and firm size in NY. When the threshold varies within-state, I plot the highest threshold faced by any employer in the state.



(a) Effect on Number of Salaried Jobs, at Time 0

(b) Effect on Salaried Distribution Over Time

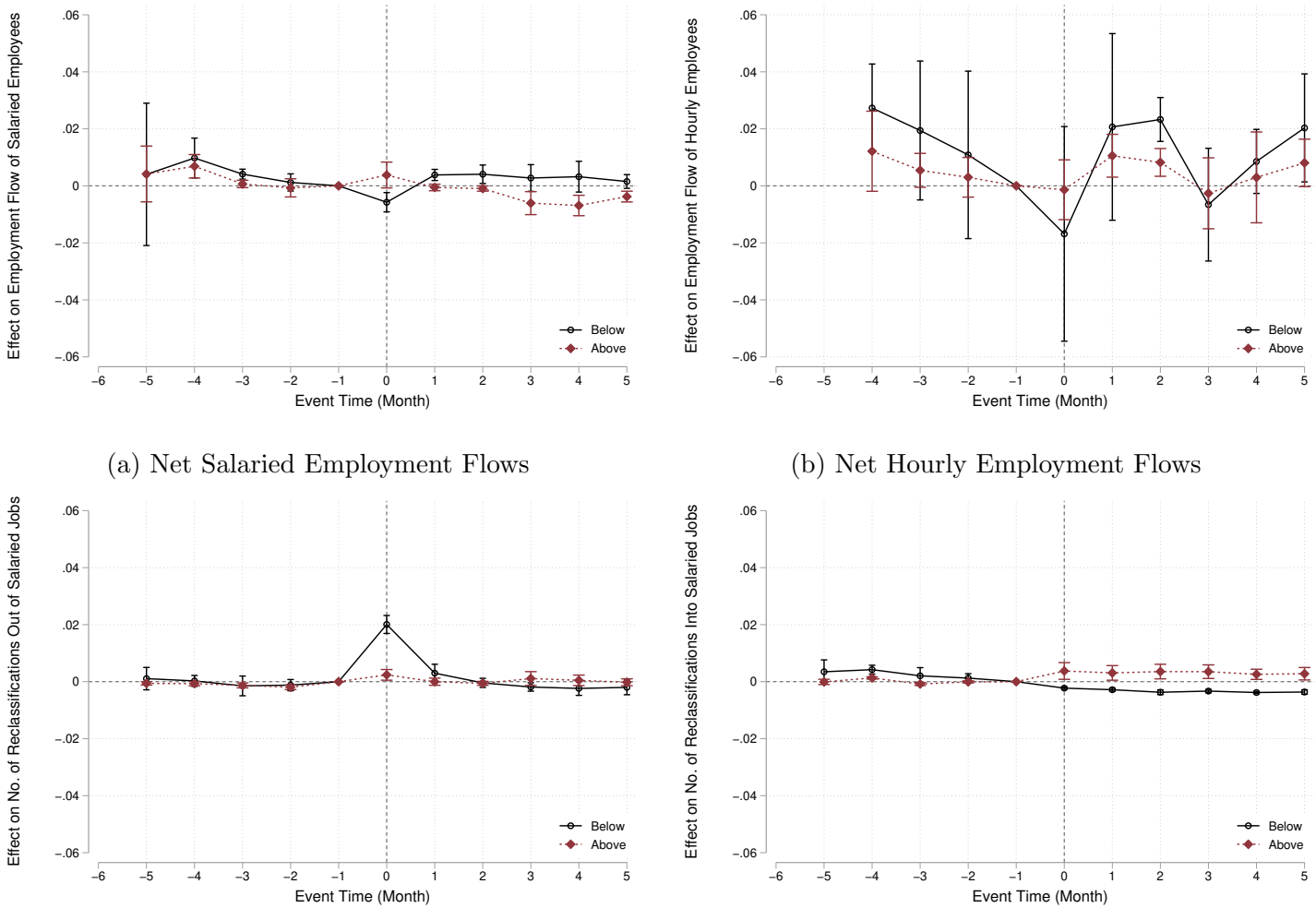
(c) Effect on Number of Hourly Jobs, at Time 0

(d) Effect on Hourly Distribution Over Time

Figure II

Effect of Raising States' OT Exemption Thresholds on the Frequency Distribution of Base Pay

Notes. Panel (a) shows the event study estimates from equation 9. The height of each bar indicates the effect of raising the OT exemption threshold on the number of salaried jobs in each \$40 bin of base pay on the month that the new threshold becomes binding, scaled by the baseline number of salaried workers between the old and new thresholds. The solid line is the running sum of these estimates. The bins are normalized so that the new threshold for each event is 0. The left vertical dashed line is set at the smallest baseline threshold across all the events. Panel (b) shows the sum of the estimates over time, separately for bins between the old and new thresholds and bins up to \$80 above the new threshold. Panels (c) and (d) are analogous to Panels (a) and (b) for the distribution of hourly jobs. For each estimate, I show the 95% confidence interval using standard errors clustered by state.



(a) Net Salaried Employment Flows

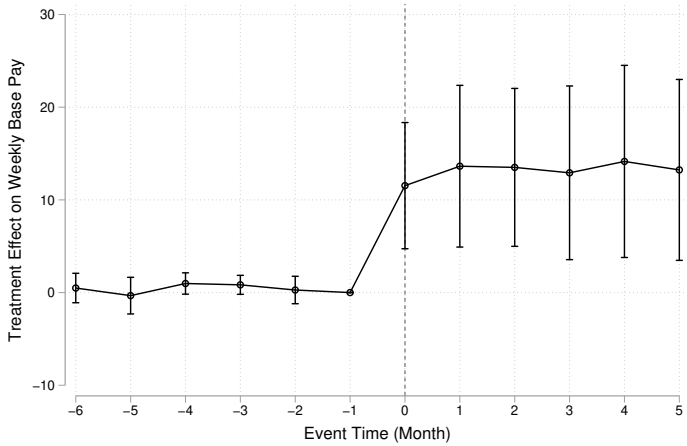
(b) Net Hourly Employment Flows

(c) Reclassification Flows Out of Salaried

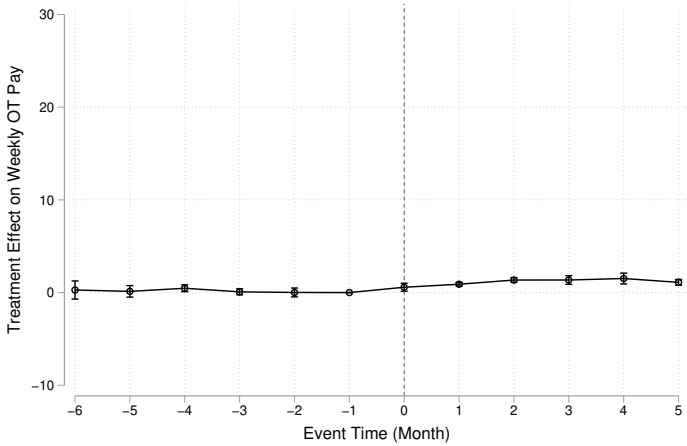
(d) Reclassification Flows Into Salaried

Figure III
Effect of State Threshold Changes on the Flow of Workers Into, Out of, and Within Firms

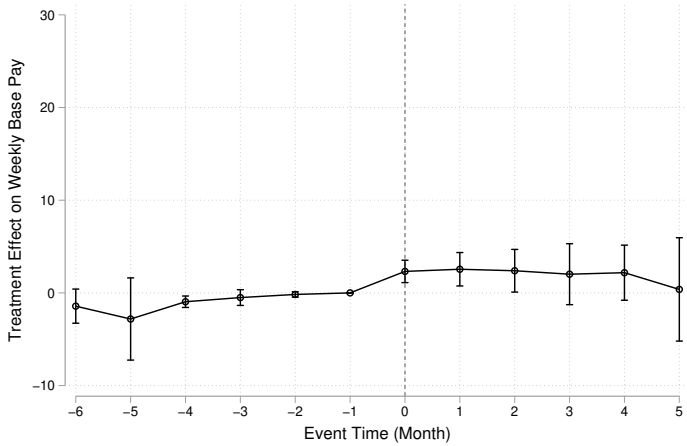
Notes. Panel (a) plots the effect of the state threshold changes on the net employment flow of salaried employees for each month since the threshold increased, scaled by the baseline number of salaried workers between the old and new thresholds. Panel (b) plots the analogous figure for net employment flows of hourly employees. Panel (c) plots the effect on the number of workers being reclassified from salaried to hourly each month and Panel (d) plots the effect on the number of employees that were reclassified from hourly to salaried. All estimates are computed using equation 9, and aggregated separately for bins between the old and new thresholds (circles) and bins up to $80 above the new threshold (diamonds). For each estimate, I show the 95% confidence interval using standard errors clustered by state.



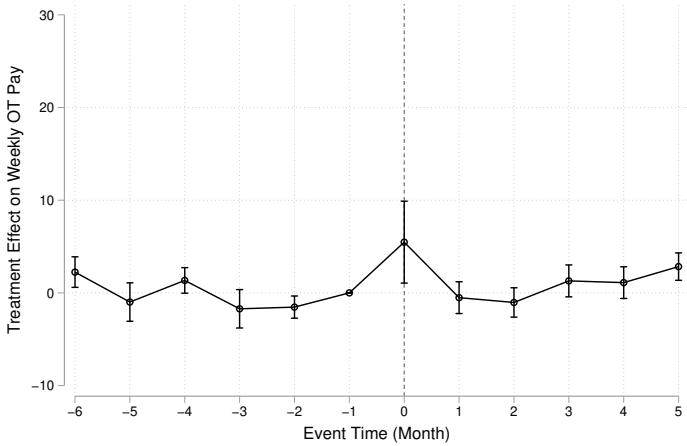
(a) Salaried: Weekly Base Pay



(b) Salaried: Weekly OT Pay



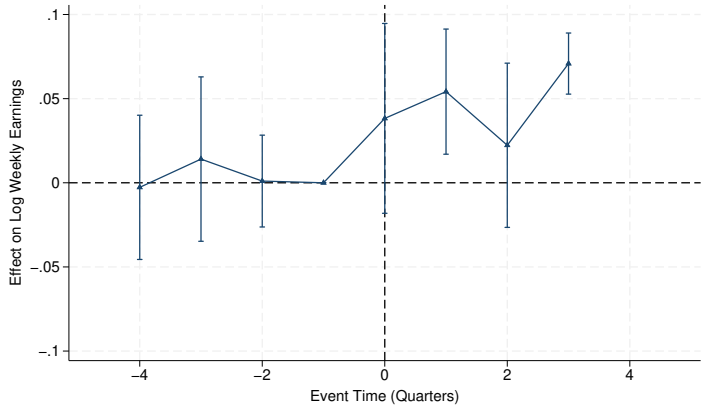
(c) Hourly: Weekly Base Pay



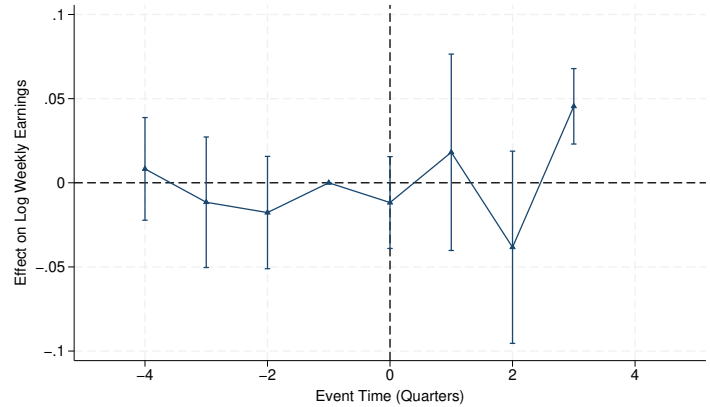
(d) Hourly: Weekly OT Pay

Figure IV
Difference-in-Difference Estimates of the Income Effect of Raising the OT Exemption Threshold

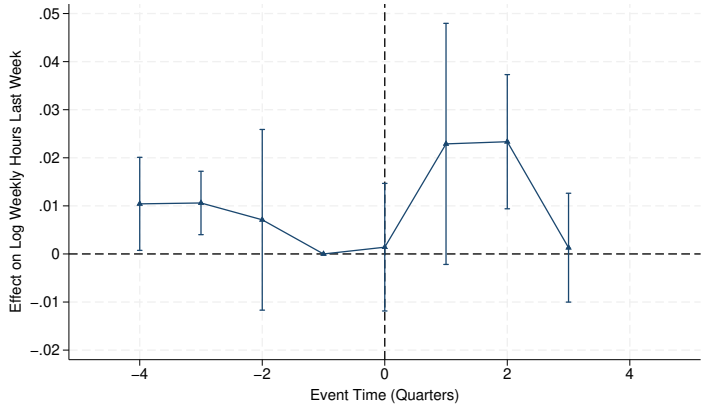
Notes. Panels (a)-(d) show the effect of raising the overtime exemption threshold on base pay and overtime pay, separately for salaried and hourly workers initially earning between the old and new thresholds. All estimates are computed from equation 10. Standard errors are clustered by state.



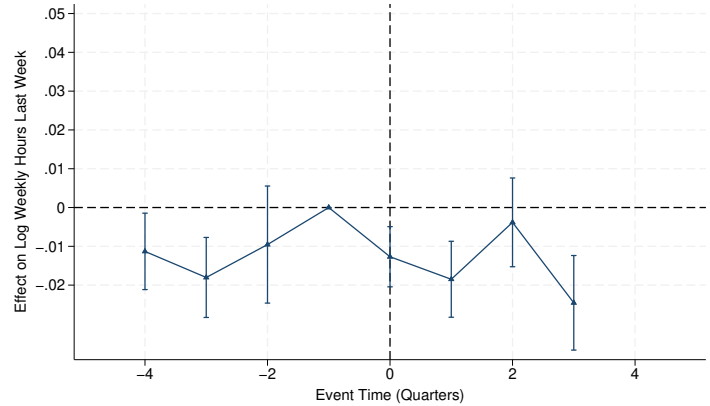
(a) Treated: Log Weekly Earnings



(b) Placebo: Log Weekly Earnings



(c) Treated: Log Weekly Hours



(d) Placebo: Log Weekly Hours

Figure V
Difference-in-Difference Estimates of the Income and Hours Effect using the CPS

Notes. This figure shows the estimates of equation 11, which compares the outcomes of CPS respondents in states that raised the overtime exemption threshold to respondents in states that did not. Panels (a) and (c) uses the sample of salaried workers with reported weekly earnings between the old and new thresholds. Panels (b) and (d) uses salaried workers with reported weekly earnings up to $100 above the new threshold.

Table I
Summary of Theoretical Predictions

Prediction	Compensating Differentials	Labor Supply	Labor Demand
Base Pay	↓	↓	Bunching
Overtime Pay	↑	↑	↑
Employment	-	?	?
Pay structure	-	-	Reclass from salaried to hourly
Hours	-	Missing mass at 40	Bunching at 40

Notes. This table summarizes the predictions of the three models of overtime discussed in Section III. The first four rows refer to the effect on each outcome from an expansion in overtime coverage for salaried workers. The last row refers to the effect of overtime among hourly workers. In the labor demand model, the employment effect is positive if the marginal productivity of additional hours diminishes sufficiently quickly, and negative otherwise.

Table II
Effect of Raising the Overtime Exemption Threshold on the Pay Distribution

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A						
Salaried Jobs						
Below Threshold	-.209***	-.149***	-.203***	-.216***	-.209***	
	(.061)	(.032)	(.063)	(.056)	(.062)	
Above Threshold	.158***	.107***	.157***	.164***	.153**	
	(.056)	(.03)	(.057)	(.054)	(.062)	
Hourly Jobs						
Below Threshold	-.07***	-.042*	-.028	-.077***	-.009	
	(.026)	(.023)	(.021)	(.023)	(.017)	
Above Threshold	.098**	.065**	.088**	.102**	.08*	
	(.05)	(.026)	(.042)	(.049)	(.046)	
Panel B						
Emp. w.r.t. Aff. Salaried	-.023	-.019	.015	-.027	.015	-.024
	(.023)	(.013)	(.018)	(.025)	(.035)	(.024)
Emp. w.r.t. All Affected	-.005	-.004	.003	-.006	.003	-.005
	(.005)	(.003)	(.004)	(.005)	(.008)	(.005)
Number of Affected Salaried	1.075	1.075	1.096	1.15	1.198	1.075
Number of Affected Hourly	4.212	4.212	4.272	4.233	4.378	4.212
Event-Bin-Month FE	Y	Y	Y	Y	Y	
Event-Bin-State FE	Y	Y	Y	Y	Y	
Event-Bin-Division-Month FE		Y			Y	
Event-Bin-Month-Firm FE			Y		Y	
Event-Bin-State-Firm FE			Y		Y	
Event-State						Y
Event-Month						Y
Number of event-firms (treatment)	237,285	237,285	184,560	213,356	162,454	237,285
Number of event-firms (control)	5,124,260	5,124,260	2,225,276	3,770,451	1,854,267	5,124,260

Notes. Rows (1) and (2) report the effect of raising the OT exemption threshold on the number of salaried jobs below and above the new threshold, respectively, scaled by the number of affected salaried jobs. Affected jobs are defined as salaried employees with base pay between the old and new threshold on the month before the announcement of the rule change. Rows (3) and (4) report similar estimates for the number of hourly jobs. Row (5) reports the sum of rows (1) to (4), and represents the effect on the total number of jobs, scaled by the number of affected salaried jobs. Row (6) scales the change in total employment by the total number of jobs (i.e. salaried and hourly) with base pays between the old and new thresholds before the rule change, instead of just salaried jobs.

Column (1) reports the effects of the state threshold increases on the exact month of the rule change, estimated using equation 9. Column (2) estimates the effects using within-census division variation. Column (3) uses variation within firms that operate in both the treatment and control groups. Column (4) drops the five events where the old overtime exemption threshold was within $200 of the new minimum wage. Column (5) estimates a saturated model with the full sample. Column (6) aggregates employment across bins of base pay and then estimates a two-way fixed-effects model. Standard errors are clustered by state. *10%,** 5%,*** 1% significance level.

Table III
Event-Study Estimates of Employment Flow and Reclassification Effects

	Employment			Reclassification		
	Hires	Separations	Net	Into	Out of	Net
Salaried Below	-.003***	.003	-.006***	-.002***	.02***	-.022***
	(.001)	(.003)	(.002)	(0)	(.002)	(.002)
Salaried Above	.006***	.002**	.004**	.004***	.002**	.001
	(.001)	(.001)	(.002)	(.001)	(.001)	(.001)
Hourly Below	-.001	.017*	-.018	.019***	.002**	.018***
	(.01)	(.01)	(.019)	(.002)	(.001)	(.002)
Hourly Above	.003	.005**	-.001	.005***	0	.005***
	(.005)	(.002)	(.005)	(.001)	(0)	(.001)
Cumulative	.005	.027**	-.021	.026***	.024***	.001
	(.015)	(.014)	(.028)	(.004)	(.003)	(.001)
Treatment Group						
Number of Affected Salaried	1.075	1.075	1.075	1.075	1.075	1.075
Number of Affected Hourly	4.212	4.212	4.212	4.212	4.212	4.212
Event-Bin-Month FE	Y	Y	Y	Y	Y	Y
Event-Bin-State FE	Y	Y	Y	Y	Y	Y
Number of event-firms	237,285	237,285	237,285	237,285	237,285	237,285

Notes. The first column reports the effect of raising the state OT exemption threshold on the number of new hires among salaried jobs paying within \$160 below the new threshold (row 1), salaried jobs paying within \$80 above the new threshold (row 2), hourly jobs within the same ranges (rows 3 & 4), and the sum of salaried and hourly jobs within those ranges (row 5). Each estimate is scaled by the number of affected salaried jobs, reported in row 6. Column (2) reports estimates for the effect on separations by the same groups. Column (3) is the difference between columns (1) and (2). Columns (4) and (5) report the effect on the number of reclassifications into and out of each group, respectively. Column (6) is the difference between columns (4) and (5). All values are estimated from equation 9, and robust standard errors in parentheses are clustered by state. $^{*}p < .1,^{**}p < .05,^{***}p < .01$

Table IV
Income Effect of Raising the OT Exemption Threshold

	Main Spec		Within-Census		Within-Firm		No Minimum Wage	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Weekly Base Pay	11.536***	2.319***	7.359***	1.528***	9.602***	1.551***	11.405***	2.379***
	3.473	.616	1.851	.299	2.38	.239	3.619	.677
Weekly OT Pay	.574***	5.474**	.282***	4.647***	.285*	1.485*	.564**	5.395**
	.222	2.254	.099	1.557	.162	.849	.232	2.467
Weekly Total Pay	12.11***	7.793***	7.641***	6.175***	9.887***	3.036***	11.968***	7.774***
	3.68	1.69	1.88	1.272	2.444	.793	3.836	1.836
% Δ Total Pay	.014***	.009***	.009***	.007***	.011***	.003***	.014***	.009***
	.004	.002	.002	.001	.003	.001	.004	.002
Log Total Pay	.013***	.008***	.008***	.006***	.01***	.003***	.012***	.007***
	.004	.001	.002	.001	.003	.001	.004	.002
Baseline Weekly Base Pay	870	815	870	815	878	820	874	821
Baseline Weekly OT Pay	7	79	7	79	10	86	7	81
Event-Month FE	Y	Y	Y	Y	Y	Y	Y	Y
Event-State FE	Y	Y	Y	Y	Y	Y	Y	Y
Event-Division-Month FE			Y	Y			Y	Y
Event-Month-Firm FE					Y	Y	Y	Y
Event-State-Firm FE					Y	Y	Y	Y
N (treatment)	189,241	708,375	189,241	708,375	97,002	402,593	92,796	362,043
N (control)	2,322,226	9,195,386	2,322,226	9,195,386	1,447,181	6,711,793	1,172,638	4,307,095
Sample	Salaried	Hourly	Salaried	Hourly	Salaried	Hourly	Salaried	Hourly

Notes. Rows (1) and (2) report the effect of raising the OT exemption threshold on continuously employed workers' base pay and overtime pay, respectively. Row (3) equals the sum of rows (1) and (2). Row (4) scales row (3) by the average baseline income of the treatment group. Row (5) reports the estimate of the policy's effect on log total pay.

Columns (1) reports the effect of increasing the state overtime exemption threshold on salaried workers' earnings, estimated from equation 10. The sample consists of salaried workers who were earning between the old and new threshold prior to the rule change. Column (2) reports the income effect for hourly workers with weekly base pay within the affected salary interval. Columns (3) and (4) report analogous estimates comparing workers within the same Census division. Columns (5) and (6) compare workers within the same firm, across states. Columns (7) and (8) drops the five events where the old overtime exemption threshold was within $200 of the new minimum wage. Robust standard errors in parentheses are clustered by state. *10%,** 5%,*** 1% significance level.

Table V
Effect of Raising the OT Exemption Threshold on Workers' Earnings and Hours

	(1)	(2)	(3)
Treatment (n=161,804)			
Log Weekly Earnings	.037***	.038***	.049***
	(.01)	(.011)	(.016)
Log Weekly Hours	.004	.003	-.011
	(.004)	(.003)	(.003)
Placebo (n=188,576)			
Log Weekly Earnings	.009	.009	.011
	(.005)	(.005)	(.011)
Log Weekly Hours	.001	.001	
	(.002)	(.002)	(.004)
Event-Individual	Y	Y	Y
Event-Month	Y	Y	
Event-Month-Region			Y
Controls		Y	Y

Notes. This table presents stacked difference-in-difference estimates comparing states that raised their overtime exemption thresholds to states that did not. Column (1) reports the estimates for two outcome variables (i.e. log weekly earnings and log weekly hours) separately for a treatment and placebo sample constructed from the CPS. Column (2) includes controls for gender, race, age, age-squared, and marital status. Column (3) adds fixed effects for Census division-event-month fixed effects. The treatment panel restricts the data to only salaried workers who report earning between the old and new overtime exemption threshold before the policy change. The placebo panel restricts the sample to those earning between the new threshold and $100 per week above it. I report besides each group name the number of observations over the two year event window for all events. All robust standard errors in parentheses are clustered by state. *10%,** 5%,*** 1% significance level.

Table VI
Ratio of Employment and Income Effects

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A						
% Δ Emp. w.r.t. Aff. Salaried	-.023	-.019	.015	-.027	.015	-.024
	(.023)	(.013)	(.018)	(.025)	(.035)	(.024)
% Δ Salaried Income	.013	.008	.01	.012	.004	.013
	(.004)	(.002)	(.003)	(.004)	(.001)	(.004)
Elasticity w.r.t. Aff. Salaried	-1.855	-2.25	1.452	-2.152	3.503	-1.905
	(2.357)	(1.688)	(1.49)	(2.654)	(7.335)	(2.453)
Panel B						
% Δ Emp. w.r.t. All Affected	-.005	-.004	.003	-.006	.003	-.005
	(.005)	(.003)	(.004)	(.005)	(.007)	(.005)
% Δ All Income	.008	.006	.004	.008	.003	.008
	(.001)	(.001)	(0)	(.001)	(.001)	(.001)
Elasticity w.r.t. Income	-.592	-.622	.745	-.722	1.097	-.608
	(.545)	(.393)	(.93)	(.611)	(2.487)	(.571)
Number of Affected Salaried	1.075	1.075	1.096	1.15	1.198	1.075
Number of Affected Hourly	4.212	4.212	4.272	4.233	4.378	4.212
Event-Bin-Month FE	Y	Y	Y	Y	Y	
Event-Bin-State FE	Y	Y	Y	Y	Y	
Event-Bin-Division-Month FE		Y			Y	
Event-Bin-Month-Firm FE			Y		Y	
Event-Bin-State-Firm FE			Y		Y	
Event-State						Y
Event-Month						Y
Number of event-firms	5,361,545	5,361,545	2,409,836	3,983,807	2,016,721	5,361,545
Number of event-workers (sal)	2,679,244	2,679,244	1,676,655	2,256,403	1,435,846	2,679,244
Number of event-workers (sal + hr)	9,388,099	9,388,099	6,453,514	6,571,030	4,502,730	9,388,099

Notes. Row (1) reports the change in employment scaled by the number of affected salaried workers, defined as salaried workers earning between the old and new overtime exemption thresholds in the month before the policy change. Row (2) reports the percent change in income of affected salaried workers. Row (3) reports the ratio of the estimates in rows (1) and (2). Rows (4)-(6) repeats the same analysis, but defines the treated group as all jobs (i.e. both salaried and hourly) with base pays between the old and new thresholds before the rule change.

Column (1) reports the employment effect from column (1) of table II, the income effect in column (1) of table IV, and their ratio. Column (2) estimates the effects using within-census division variation. Column (3) uses variation within firms that operate in both the treatment and control groups. Column (4) drops the five events where the old overtime exemption threshold was within $200 of the new minimum wage. Column (5) estimates a saturated model with the full sample. Column (6) aggregates employment across bins of base pay and then estimates a two-way fixed-effects model. Standard errors are clustered by state. *10%,** 5%,*** 1% significance level.

Appendix: For Online Publication

Appendix A. Additional figures and tables

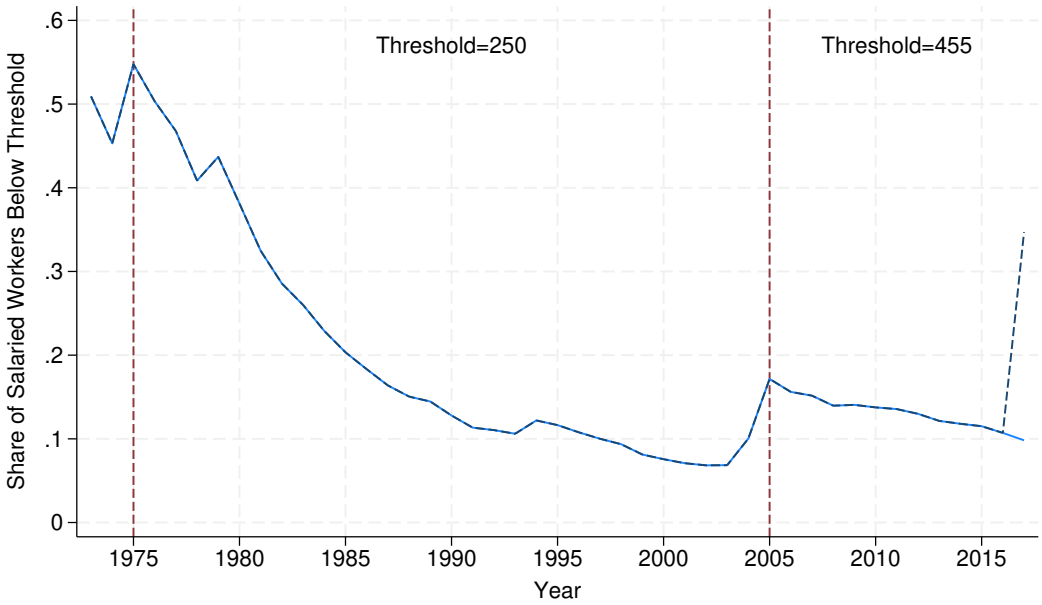


Figure A.1
Percent of Salaried Workers Earning Below the FLSA OT Exemption Threshold

Notes. The figure shows the share of all salaried workers in the May extracts of the CPS who report usual weekly earnings below the effective FLSA overtime exemption threshold from 1973 to 2017. The threshold increased from $200 per week to $250 per week in January 1975, and then to $455 in August 2004. The dotted blue line shows the percent of salaried workers with usual weekly earnings below the $913 per week threshold announced in the 2016 policy.

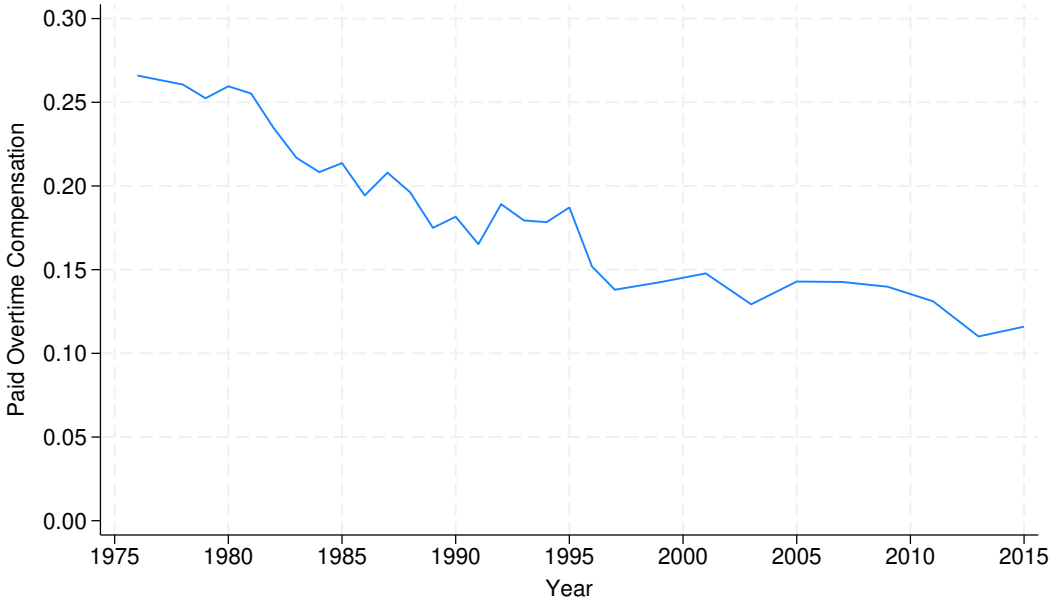
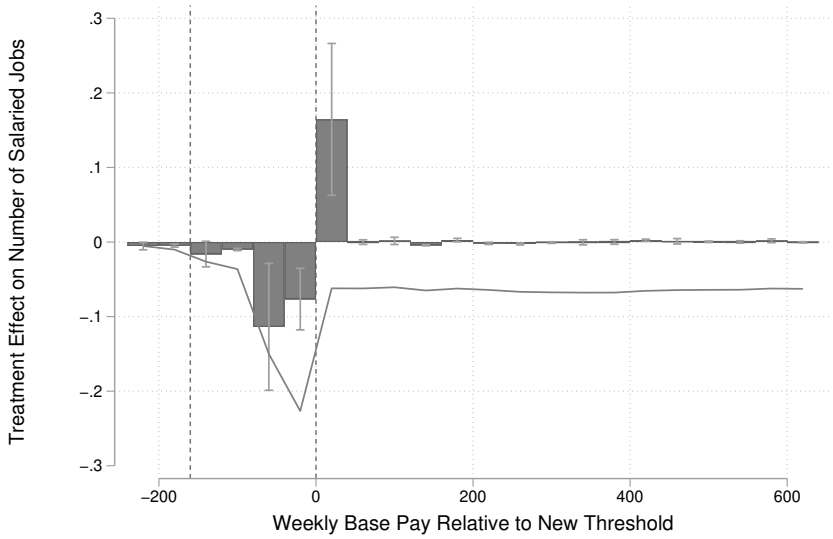
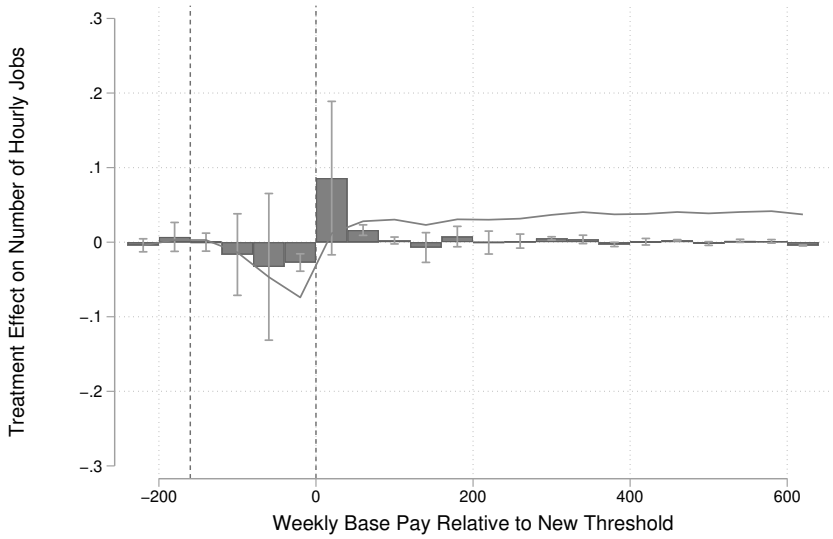


Figure A.2
Percent of Salaried Workers Eligible for Overtime

Notes. This figure shows the percent of salaried workers in the PSID who work at least 40 hours a week and respond yes to the question "If you were to work more hours than usual during some week, would you get paid for those extra hours of work".



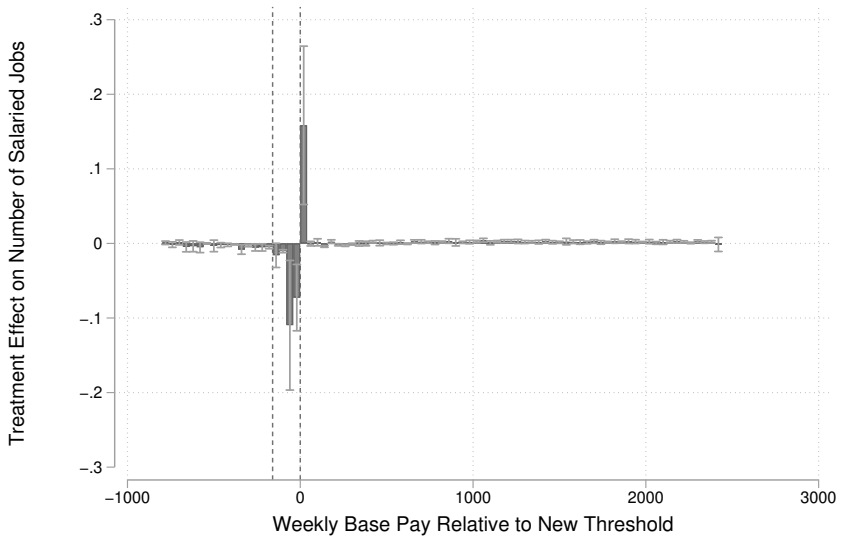
(a) Effect on Number of Salaried Jobs, at Time 0



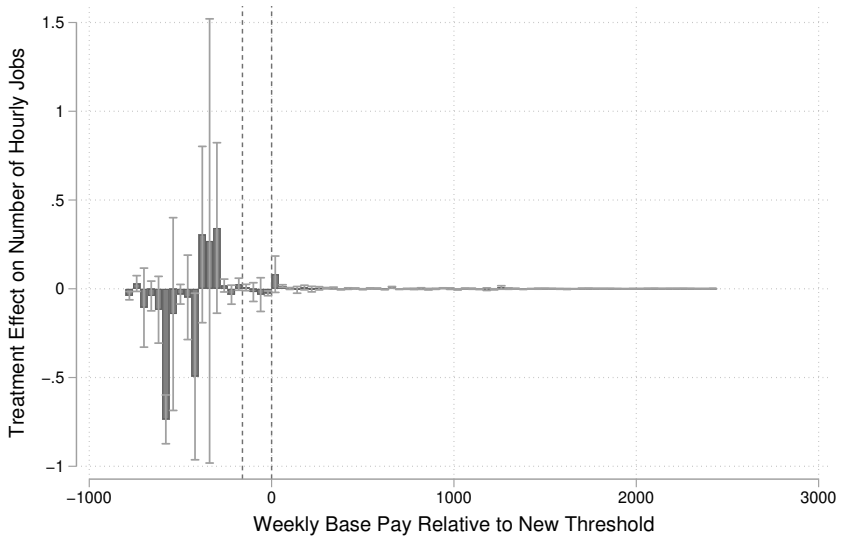
(b) Effect on Number of Hourly Jobs, at Time 0

Figure A.3
Effect of Raising States' OT Exemption Thresholds on the Distribution of Base, No Interaction with Minimum Wage Pay

Notes. The figures show the event study estimates from equation 9, separately for salaried and hourly jobs. The sample is restricted to the 14 events where the the old overtime exemption threshold is at least $200 weekly base pay above a new minimum wage. The height of each bar indicates the effect of raising the OT exemption threshold on the number jobs in each $40 bin of base pay on the month that the new threshold becomes binding, scaled by the baseline number of salaried workers between the old and new thresholds. The solid line is the running sum of these estimates. The bins are normalized so that the new threshold for each event is 0. The left vertical dashed line is set at the smallest baseline threshold across all the events. For each estimate, I show the 95% confidence interval using standard errors clustered by state.



(a) Effect on Number of Salaried Jobs, at Time 0



(b) Effect on Number of Hourly Jobs, at Time 0

Figure A.4
Effect of Raising States' OT Exemption Thresholds on the Distribution of Base Pay

Notes. The figures show the event study estimates from equation 9, separately for salaried and hourly jobs. The height of each bar indicates the effect of raising the OT exemption threshold on the number jobs in each $40 bin of base pay on the month that the new threshold becomes binding, scaled by the baseline number of salaried workers between the old and new thresholds. The bins are normalized so that the new threshold for each event is 0. The left vertical dashed line is set at the smallest baseline threshold across all the events. For each estimate, I show the 95% confidence interval using standard errors clustered by state.

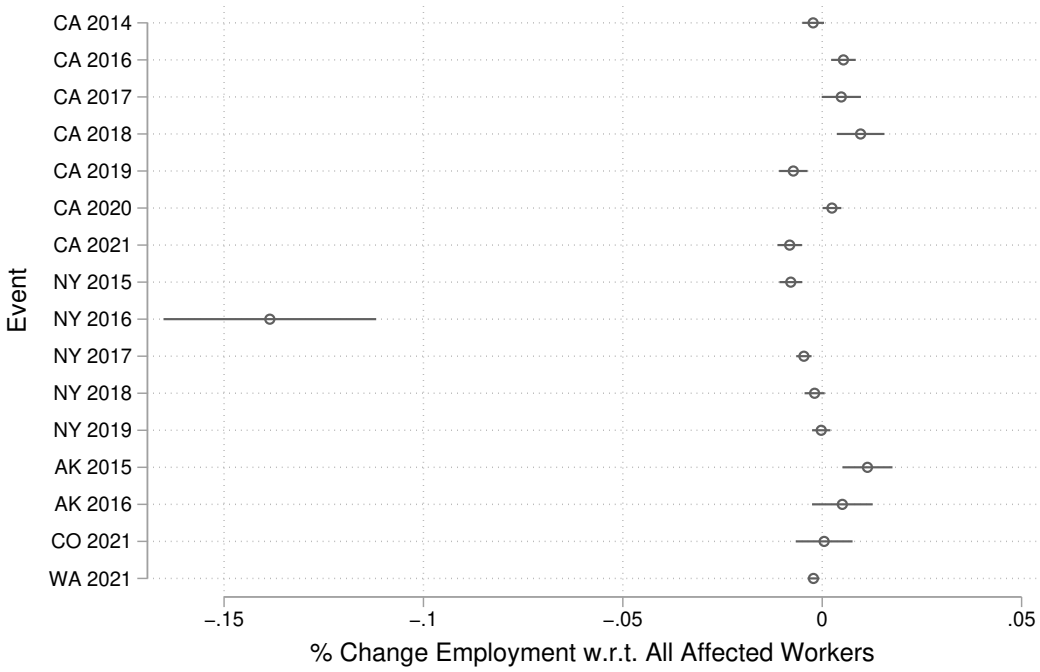


Figure A.5
Employment Effect Scaled by Affected Workers, by Event

Notes. The figure shows the difference-in-difference estimates from equation 9, separately for 16 events. All regression estimates are scaled by the total number of salaried and hourly workers between the old and new overtime exemption threshold in the month before the policy change. The bars around each estimate represent 95% confidence intervals. Standard errors are clustered by state.

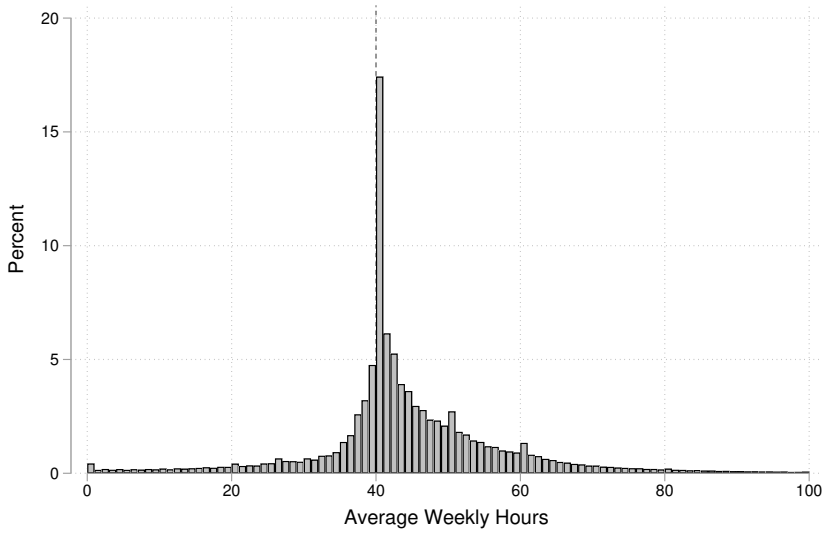


Figure A.6
Distribution of Average Weekly Hours Among Hourly Workers in April 2016

Notes. The figure shows the distribution of average workweeks among hourly workers with base pays between the old and new overtime exemption threshold in the month before a policy change. Average weekly hours is imputed from the total hours worked in a month following the methodology in Appendix C. Each bin is a one hour increment. The left vertical line is at 40 hours per week.

Appendix B. Derivation of the Conceptual Framework

B.a Proof of Proposition 1

Proof. To determine h, hourly workers solve

$$\max_h U(h) = w[h + p(h - 40)] - a^{-\frac{1}{\epsilon}} \frac{h^{1+\frac{1}{\epsilon}}}{1 + \frac{1}{\epsilon}}$$

The first order conditions $w(1 + p) - a^{-\frac{1}{\epsilon}} h^{\frac{1}{\epsilon}} = 0$ imply that

$$h^* = a[w(1 + p)]^{\epsilon}$$

The overtime premium $p = 0.5 \cdot 1[h > 40]$ creates a kink in the budget constraint, causing workers with $40 = aw^{\epsilon}$ to increase their hours until $h^* = a[1.5w]^{\epsilon} > aw^{\epsilon} = 40$. ■

B.b Proof of Proposition 2

Proof. To determine h, salaried workers solve

$$\max_h U(h) = S(1 + p\frac{h - 40}{40} 1[S < \bar{S}]) - a^{-\frac{1}{\epsilon}} \frac{h^{1+\frac{1}{\epsilon}}}{1 + \frac{1}{\epsilon}} \tag{1}$$

Suppose p increases from 0 to 1.5. First, I show that hours increase with respect to p. The first order condition to the worker's problem in equation 1 is

$$\begin{aligned} \frac{dU}{dh} &= 0 \\ \Rightarrow \frac{Sp}{40} 1[S \leq \bar{S}] &= a^{-\frac{1}{\epsilon}} h^{\frac{1}{\epsilon}} \\ \Rightarrow h^* &= a\Big[\frac{Sp}{40} 1[S \leq \bar{S}]\Big]^{\epsilon} \end{aligned} \tag{2}$$

Hours is thus an increasing function of the overtime premium p.

Second, I show that increasing p decreases labor demand. Given h^*, the firm solves

$$\max_n \pi = xn^{\alpha} h^{*\beta} - [S(1 + p\frac{h^* - 40}{40} 1[S \leq \bar{S}]) + R]n$$

giving the first order condition

$$\frac{d\pi}{dn} = x\alpha n^{\alpha-1}h^{\beta} - [S(1+p\frac{h-40}{40}1[S \leq \bar{S}]) + R] = 0$$

Rearranging for labor demand:

$$n^D(S, h^*(S,p), p) = \Big[\frac{x\alpha h(S,p)^{\beta}}{S(1+p\frac{h(S,p)-40}{40}1[S \leq \bar{S}]) + R}\Big]^{\frac{1}{1-\alpha}}$$

To determine the change in labor demand, take the derivative of $n^D(S, h^*(S,p), p)$ with respect to p:

$$\frac{dn^d}{dp} = \underbrace{\frac{1}{1-\alpha}\Big[\frac{x\alpha h^{\beta}}{S(1+p\frac{h-40}{40}) + R}\Big]^{\frac{\alpha}{1-\alpha}}}_{>0} \Big[\frac{x\beta\alpha h^{\beta-1}\frac{\partial h}{\partial p}[S(1+p\frac{h-40}{40}) + R] - S[\frac{h-40}{40} + \frac{p}{40}\frac{\partial h}{\partial p}]x\alpha h^{\beta}}{[S(1+p\frac{h-40}{40}1[S \leq \bar{S}]) + R]^2}\Big]$$

where $1[S \leq \bar{S}]$ is attached to every p but omitted for brevity.

The numerator in the second term simplifies to

$$x\alpha h^{\beta-1}\Big[\big(\underbrace{\beta[S(1-p)+R] - (1-\beta)\frac{ph}{40}S]}_{<0 \text{ because } p=1.5}\big)\underbrace{\frac{\partial h}{\partial p}}_{>0} - \frac{h-40}{40}Sh\Big]$$

The term is always negative, implying that labor demand decreases. Intuitively, an increase in the overtime premium induces a scale and substitution effect. In this case, since workers are increasing hours, both these effects point in the same direction, leading to a negative impact on labor demand.

$$\frac{dn^d}{dp} = \underbrace{\frac{\partial n^d}{\partial p}}_{\text{Scale Effect}<0} + \underbrace{\frac{\partial n^d}{\partial h}\frac{\partial h}{\partial p}}_{\text{Substitution Effect}<0}$$

Third, raising p increases labor supply since for any given S, the worker can receive more for the same h. The market clearing salary S^* equates demand and supply: $n^d(S^*) = n^S(S^*)$. A drop in labor demand and an increase in labor supply imply that wages will fall, and the employment effect is ambiguous.

Forth, there will be no bunching at the overtime exemption threshold. Since workers control hours, they will work as few hours as possible if exempt from overtime, leaving firms with zero surplus from the employment relationship. As such firms have no incentives to raise salaries above the market clearing rate to the exemption threshold. ■

B.c Proof of Proposition 3

Proof. The firm take salary S as given. It chooses hours and employment to maximize profits:

$$\max_{(n,h)} \pi = xn^{\alpha}h^{\beta} - [S(1 + p\frac{h-40}{40}1[S \leq \bar{S}]) + R]n \tag{3}$$

If $p = 0$, then for any S, the firm will have workers work as many hours as possible because the marginal cost per hour of labor is zero. However, increasing hours will reduce the number of workers willing to supply labor to the firm. Given the firm's ability to control hours, it will therefore always be on the worker's extensive labor supply curve. As such, when $p = 0$, the firm behaves as a monopsonist, maximizing equation 3 subject to the following constraint:

$$U(S,h) = S(1 + p\frac{h-40}{40}1[S \leq \bar{S}]) - a^{-\frac{1}{\epsilon}}\frac{h^{1+\frac{1}{\epsilon}}}{1+\frac{1}{\epsilon}} \geq r_n \tag{4}$$

Rearranging equation 4 for S, substituting it into equation 3 and taking first order conditions, the firm's decision solves

$$\frac{d\pi}{dh} = x\beta n^{\alpha}h^{\beta-1} - a^{-\frac{1}{\epsilon}}h^{\frac{1}{\epsilon}} = 0$$
$$\frac{d\pi}{dn} = x\alpha n^{\alpha-1}h^{\beta} - [r_n + a^{-\frac{1}{\epsilon}}\frac{h^{1+\frac{1}{\epsilon}}}{1+\frac{1}{\epsilon}} + R] - r'(n)n = 0$$

The first order conditions implicitly define n and h, which can be substituted back into the worker's labor supply function to solve for S.

Next, suppose p increases from 0 to 1.5. The market for a specific job can respond in one of three ways, depending on the fundamental parameters of the job (i.e. α, β, R):

Case 1: Firms cut hours (i.e. $\frac{\partial h}{\partial p} < 0$). If β is small, then employment increases and effect on salary is ambiguous. If β is large, then salaries fall and employment is ambiguous.

First, I show that firms cut hours as p increases. The first order conditions to equation 3

are

$$\frac{d\pi}{dh} = x\beta n^{\alpha} h^{\beta-1} - \frac{Sp}{40} 1[S \leq \bar{S}] n = 0$$
$$\frac{d\pi}{dn} = x\alpha n^{\alpha-1} h^{\beta} - [S(1 + p\frac{h-40}{40} 1[S \leq \bar{S}]) + R] = 0$$

Ratio of the first order conditions:

$$\frac{\beta}{\alpha h} = \frac{Sp1[S \leq \bar{S}]}{40[S(1 + p\frac{h-40}{40} 1[S \leq \bar{S}]) + R]}$$

Solve for hours:

$$h^* = \frac{40\beta(S(1 - p1[S \leq \bar{S}]) + R)}{(\alpha - \beta)Sp1[S \leq \bar{S}]}$$

Hours is a decreasing function of the premium p.

Second, I show that the effect on labor demand is ambiguous and depends on the magnitude of β. Intuitively, an increase in p induces opposing scale and substitution effects on labor demand:

$$\frac{dn^d}{dp} = \underbrace{\frac{\partial n^d}{\partial p}}_{\text{Scale Effect}<0} + \underbrace{\frac{\partial n^d}{\partial h}\frac{\partial h}{\partial p}}_{\text{Substitution Effect}>0}$$

To determine labor demand, n^D, rearrange $\frac{d\pi}{dn} = 0$:

$$n^D(S, h(S,p), p) = \Big[\frac{x\alpha h(S,p)^{\beta}}{S(1 + p\frac{h(S,p)-40}{40} 1[S \leq \bar{S}]) + R}\Big]^{\frac{1}{1-\alpha}}$$

To determine the change in labor demand, I follow the same argument as in proof B.b. Take the derivative of $n^D(S, h(S,p), p)$ with respect to p:

$$\frac{dn^d}{dp} = \quad = \underbrace{\frac{1}{1-\alpha}\Big[\frac{x\alpha h^{\beta}}{S(1 + p\frac{h-40}{40}) + R}\Big]^{\frac{\alpha}{1-\alpha}}}_{>0} \Big[\frac{x\beta\alpha h^{\beta-1}\frac{\partial h}{\partial p}[S(1 + p\frac{h-40}{40}) + R] - S[\frac{h-40}{40} + \frac{p}{40}\frac{\partial h}{\partial p}]x\alpha h^{\beta}}{[S(1 + p\frac{h-40}{40} 1[S \leq \bar{S}]) + R]^2}\Big]$$

where $1[S \leq \bar{S}]$ is attached to every p but omitted for brevity.

The numerator in the second term simplifies to

$$x\alpha h^{\beta-1}\Big[\underbrace{\big(\beta[S(1-p)+R]-(1-\beta)\frac{ph}{40}S]\big)\frac{\partial h}{\partial p}}_{\text{Substitution Effect}>0}\underbrace{-\frac{h-40}{40}Sh}_{\text{Scale Effect}<0}\Big]$$

Since p=1.5, the substitution effect is positive for all β. On the other hand, the scale effect is negative. To show that the sum can be positive or negative, I show that it is positive at $\beta = 0$ and strictly decreasing in β. If $\beta = 0$ then $\frac{dn}{dp} \propto -[p\frac{\partial h}{\partial p} + h - 40]\frac{Sh}{40}$, implying that work-sharing would occur if 1.5 times the reduction in hours exceeds the amount of overtime hours. This must be true if $\beta = 0$ since firms would simply choose 0 hours considering that it is irrelevant to production. As a result, if β is sufficiently small, then prior to coverage, firms will extract as many hours from workers as possible but after coverage, they will reduce all overtime hours. The derivative of the sum of the substitution and scale effects is $[S+R+\frac{pS(h-40)}{40}]\frac{\partial h}{\partial p} < 0$ so the employment effect is strictly decreasing in β. At a sufficiently large β, the employment effect will therefore be negative.

Third, I show that in equilibrium, one of the effects on employment and salary is ambiguous, depending on what happened to labor demand. Equilibrium (n^*, S^*, h^*) are characterized by the firm's two first order conditions and workers' extensive labor supply equation. In equilibrium, salary equates extensive labor demand and supply, $n^D(S, h(S,p), p) = n^S(S, h(S,p), p)$. An increase in p decreases h and rewards workers for hours of overtime. As a result, extensive labor supply $N^S(S) = n^S(S, h(S,p), p)$ has increased for each value of S. As shown before, labor demand may either increase or decrease depending on the magnitudes of the scale and substitution effects. If labor demand increases, then equilibrium employment must rise, and the effect on base salary is ambiguous. If labor demand falls, then equilibrium employment is ambiguous while base salaries will fall.

Case 2 (Bunching): Some jobs receive a raise to exactly the overtime exemption threshold $\bar{S}$. Jobs initially paid close to the threshold are more likely to receive a raise.

The equilibrium outcome (n^*, S^*, h^*) must satisfy $\pi(n^*, S^*, h^*) \geq \pi(n^*, S, h^*)$ for every salary $S > S^*$, otherwise the firm would increase salaries. Since only employees earning less than $\bar{S}$ are covered for overtime, then at the value of n^* and h^* in case 1, there exists a $\underline{S}$ such that

$$\underline{S}(1 + 1.5\frac{h^* - 40}{40}) = \bar{S}$$

That is, it costs the firm as much to pay each worker at a base pay of $\underline{S}$ with overtime as

it does to simply pay them $\bar{S}$. In this case, no $S \in (\underline{S}, \bar{S})$ can exist in equilibrium because firms would simply raise salaries for these jobs to the threshold. For these jobs, a stable equilibrium is characterized by $S = \bar{S}$, and (n, h) are determined by the firm's first order conditions as functions of $\bar{S}$. At this equilibrium, firms have no incentives to deviate hours or employment. Even though workers are willing to accept a lower base salary to work, that is not incentive compatible for the firm. Similar to a minimum wage, employment is no longer on the worker's extensive labor supply curve.

Note that by raising salaries, the bunching effect leads firms to reduce employment relative to baseline levels. Nevertheless, it is still true that for small values of β, employment would rise. If firms do not value long work hours, they would simply reduce hours to 40. In that case, raising salaries to the threshold would only increase labor costs.

Case 3 (Reclassification): Some jobs are reclassified from salaried to hourly.

For jobs with large values of R, the value of classifying the job salaried was initially only marginally better than having the job hourly (i.e. $\pi^{sal} - \pi^{hr} = \epsilon > 0$). After p increases, the cost of overtime reduces the profitability of salaried jobs so that it is now more profitable to classify these jobs as hourly: $\pi^{sal} - \pi^{hr} < 0$. ■

B.d Proof of Proposition 4

Proof. The firm takes wage w as given. It chooses hours and employment to maximize profits:

$$\max_{(n,h)} \pi = xn^{\alpha}h^{\beta} - [w(h + p(h - 40)) + F]n \tag{5}$$

The first order conditions are

$$\frac{d\pi}{dh} = x\beta n^{\alpha}h^{\beta-1} - w(1+p)n = 0$$
$$\frac{d\pi}{dn} = x\alpha n^{\alpha-1}h^{\beta} - [w(h + p(h - 40)) + F] = 0$$

Taking the ratio of the first order conditions implies

$$\frac{\alpha}{\beta}h = \frac{w(h + p(h - 40)) + F}{w(1+p)}$$

Isolating h, I show that hours is a decreasing function in p, which is the premium that goes into effect for each hour worked above 40 per week:

$$h = \frac{\beta}{\alpha + \beta} \frac{F - 40wp}{w(1+p)}$$

■

B.e Proof of Proposition 5

Proof. **Market:** Suppose the market determines a earnings-hour profile $Y(h)$ where $Y'(h) > 0$ for every h.
Firm: The firm takes the function $Y(h)$ as given and solves

$$\max_{(n,h)} \pi = xn^{\alpha}h^{\beta} - Y(h)n$$

The first order conditions are

$$\frac{d\pi}{dh} = x\beta n^{\alpha}h^{\beta-1} - Y'(h)n = 0 \tag{6}$$

$$\frac{d\pi}{dn} = x\alpha n^{\alpha-1}h^{\beta} - Y(h) = 0 \tag{7}$$

Worker: Workers take the function $Y(h)$ as given and solve

$$\max_{h} U(h) = Y(h) - a^{-\frac{1}{\epsilon}} \frac{h^{1+\frac{1}{\epsilon}}}{1+\frac{1}{\epsilon}}$$

The first order condition is

$$\frac{dU}{dh} = Y'(h) - a^{-\frac{1}{\epsilon}}h^{\frac{1}{\epsilon}} = 0 \tag{8}$$

Equilibrium: Equilibrium consists of values of h^*, n^*, Y^*, and $Y'(h^*)$ such that

1. Both parties are indifferent to switching their hours: h^*, n^* and $Y'(h^*)$ satisfy equations 6 and 8. Together, the two first order conditions imply

$$h^*(n) = \left[\frac{a^{\frac{1}{\epsilon}}x\beta}{n^{1-\alpha}}\right]^{\frac{1}{1+\frac{1}{\epsilon}-\beta}}$$

2. The firm is indifferent to hiring more or fewer workers: n^*, h^*, and Y^* satisfy equation 7. Substituting $h^*(n)$ into equation 7 and solving for n^* implies the following labor demand function:

$$n^D(Y^*) = \left[a^{\frac{1}{\epsilon}} x\beta\right]^{(1-\alpha)(1+\frac{1}{\epsilon})} \left[\frac{x\alpha}{Y^*}\right]^{\frac{1+\frac{1}{\epsilon}-\beta}{(1-\alpha)(1+\frac{1}{\epsilon})}}$$

   Subbing $n^D(Y^*)$ back into $h^*(n)$ gives hours as a function of Y^*

3. The labor market clears: $n^D(Y^*) = n^S(Y^*)$ where extensive labor supply $n^S(Y, h)$ can be written only in terms of Y^* and model primitives. Given the Y^* that clears the market, I can write n^* and h^* in terms of model parameters, and then the worker's first order condition defines $Y'(h^*)$.

Comparative Statics:

Case 1: Hourly jobs

To start, suppose earning-hour profiles are defined in terms of only an hourly wage contract. In that case, without overtime coverage, $Y'(h)$ is constant for every h. An increase in the overtime premium from $p = 0$ to $p = 0.5$ imposes a new constraint on the earnings-hour profile. Namely, suppose $Y'(h) = w$ is constant for every $h < 40$, then $Y'(h) \geq 1.5w$ for every $h > 40$. Initially, the equilibrium outcomes are independent of the overtime premium. Thus, as long as the earnings-hour profile adjusts in a way that the new constraint holds without violating any of the previous first order conditions, the equilibrium will not change.

First, I show that the initial equilibrium cannot be maintained by simply reducing the wage rate $Y'(h)$. Suppose the equilibrium was initially at $h^* > 40$, n^*, Y^*, and $Y'(h^*)$. Also suppose $Y'(h) = w$ for every h. In that case, I can express weekly earnings Y^* as follows:

$$Y^* = \int_0^{h^*} w dh \tag{9}$$

After overtime coverage, the law requires that

$$Y^* = \int_0^{40} Y'(h) + \int_4 0^{h^*} 1.5Y'(h)dh \tag{10}$$

However, equation 10 cannot hold if the slope of the earnings-hour profile still satisfies the first order conditions. For the FOC to be satisfied, the slope at h^* must remain the same

as it was in the initial equilibrium: $Y'(h^*) = w$. That leads to a contradiction:

$$\begin{aligned} Y^* &= \int_0^{40} Y'(h) + \int_4 0^{h^*} 1.5Y'(h)dh \\ &= \int_0^{40} \frac{w}{1.5} + \int_4 0^{h^*} wdh \\ &< \int_0^{h^*} wdh \\ &= Y^* \end{aligned}$$

In order for Y^* to stay the same, the wage rate at h^* must be higher than its initial value.

Second, I present a simple solution to satisfy the overtime regulation for hourly jobs, while also maintaining the initial equilibrium. Let $D = \int_0^{h^*} wdh - \int_0^{40} \frac{w}{1.5} + \int_4 0^{h^*} wdh$ be the difference between the total earnings at baseline and the total earning if the same wage rate remained constant at the point h^*. To satisfy the law while keeping earnings and wages constant at h^*, the new earnings profile can offer hourly workers a wage $\frac{w}{1.5}$ for the first 40 hours, a bonus D for working above 40 hours, and then a wage w for all hours above 40.

Case 2: Salaried jobs

Now consider the case of salaried jobs. I interpret $Y(h)$ as a *market* level earnings-hour profile. That is, jobs that work more hours are paid a higher salary. Within a job, salaried workers simply work hours h for a salary S that implicitly depends on h at the market level, but does not explicitly depend on hours in the employment contract. Define the implied wage $w = \frac{S}{40}$, be the salary divided by 40. The law requires that $Y'(h) = 1.5w$ for every $h > 40$. Given the notation, the same proof that showed wages would decrease for hourly jobs would show that implied wages, and by tension salaries, would decrease for salaried jobs. However, in this case, each hour-salary pair should be interpreted as a job, so that to change hours, a worker needs to switch employment. ■

To determine labor demand n^d, solve

$$\max_n \pi = xn^\alpha h^\beta - [w[h + p(h - 40)] + F]n$$

$$\frac{d\pi}{dn} = x\alpha n^{\alpha-1} h^\beta - [w[h + p(h - 40)] + F] = 0$$

$$\Rightarrow N^d(w, h, p) = \left[\frac{x\alpha h^\beta}{w[h + p(h - 40)] + F}\right]^{\frac{1}{1-\alpha}} \tag{11}$$

In equilibrium, select w^* to satisfy $N^d(w, h(w,p), p) = N^s(w, h(w,p), p)$ and then input w^* into the formula for h and n.

Comparative statics: Suppose p increases from 0 to 0.5

Proposition 6. *Hours will increase, and no one will work exactly 40 hours per week.*

Proof. This follows from equation III.b and the fact that the overtime premium p only applies for hours greater than 40. ■

Proposition 7. *Extensive labor supply rises. If there is diminishing returns to scale (i.e. $\alpha < 1$) and the fixed cost per worker F is small relative to the wage bill, then labor demand will fall. As such. wages will decrease.*

Proof. For a given w, workers' utility $U(h,p)$ increases with p so $N^s(w)$ increases.

To understand the effect on labor demand, take the derivative of equation 11 with respect to p,

$$\frac{dn^d}{dp} = \underbrace{\frac{\partial n^d}{\partial p}}_{\text{Scale Effect}<0} + \underbrace{\frac{\partial n^d}{\partial h}\frac{\partial h}{\partial p}}_{\text{Substitution Effect}} \tag{12}$$

$$= \underbrace{\frac{1}{1-\alpha}\Big[\frac{x\alpha h^\beta}{w[h+p(h-40)]+F}\Big]^{\frac{\alpha}{1-\alpha}}}_{>0}\Big[\frac{\beta x\alpha h^{\beta-1}\frac{\partial h}{\partial p}[w[h+p(h-40)]+F] - w[\frac{dh}{dp}+(h-40)+p\frac{\partial h}{\partial p}]x\alpha h^\beta}{[w[h+p(h-40)]+F]^2}\Big]$$

The numerator in the second term simplifies to

$$\begin{aligned}
&x\alpha h^{\beta-1}\Big[\big(\beta w[h+p(h-40)]+\beta F - wh[1+p]\big)\frac{\partial h}{\partial p} - wh(h-40)\Big]\\
=&x\alpha h^{\beta-1}\Big[\big(\beta wh[1+p]-\beta 40wp+\beta F - wh[1+p]\big)\frac{\partial h}{\partial p} - wh(h-40)\Big]\\
=&x\alpha h^{\beta-1}\Big[\underbrace{\big(\beta(F-40wp)-(1-\beta)wh[1+p]\big)\frac{\partial h}{\partial p}}_{\text{Substitution Effect}}\underbrace{-wh(h-40)}_{\text{Scale Effect}}\Big]
\end{aligned}$$

The scale effect is always negative. If we assume $F < 40wp$ (i.e. the majority of the cost of labor is via their wages) then the substitution effect is in the opposite direction as $\frac{\partial h}{\partial p}$. In the case where workers select hours, $\frac{\partial h}{\partial p} > 0$ so $\frac{dn^d}{dp} < 0$. Given that supply increases and demand fall, to satisfy the equilibrium condition $N^d(w) = N^s(w)$, it must be that w decreases.[19] ■

[19]Note that w will not decrease enough to offset the effects of overtime. From the labor supply

equation, to achieve the same hours, w would need to decrease to $\frac{w}{1+p}$, but then workers' level of utility fell relative to when $p = 0$ because they are being paid less per hour before 40. At this point, while N^d does not change because hours are constant, N^s has decreased, creating upward pressure on w.

Appendix C. Defining the Compensation Variables

C.a Overtime Pay

In this subsection, I present the procedure I use to determine each individual's overtime pay from the "OT earnings" variable, when available. There are two challenges to inferring workers' overtime pay from the ADP data.

First, firms are not required to input a value into the "OT earnings" field. Although the ADP data contains four separate earnings variables and four corresponding hours variables, each capturing a different component of gross compensation, firms are only required to report employees' gross pay and standard rate of pay. Thus, it is uncertain whether a missing value for overtime earnings means that the firm does not record the value or the worker did not receive any overtime pay. To test how often firms separately record workers' overtime pay, I calculate the sum of workers' four components of pay and find that it matches the measure of gross pay 99.8% of the time. This suggests that most employers are indeed properly recording the multiple aspects of individuals' incomes.

The second challenge with measuring workers' overtime pay is that the type of compensation included in the "OT earnings" field is at the discretion of the firm. Thus, some employers may use the field to record other forms of compensation than overtime pay. To account for this, I impute overtime pay following the methodology described by Grigsby et al. (2021). First, I define an implied overtime wage as the ratio between the "OT earnings" and "OT hours" variables. Next, I divide the implied wage by workers' actual wage to compute an implied overtime premium (i.e. $\frac{\text{OT earnings}}{\text{OT hours}\cdot\text{wage}}$), where a salaried worker's "wage" for overtime purposes is defined by the Department of Labor as $\frac{\text{weekly base pay}}{40}$. I consider the "OT earnings" variable to represent true overtime pay if the implied overtime premium is less than or equal to 2. I find that the distribution of the implied overtime premium exhibits significant bunching at 1.5, and 2, indicating that the variable usually captures true overtime earnings. Among workers with non-missing "OT earnings", 94% of hourly workers and 86% of salaried workers have implied overtime premiums within either 1.4-1.6 or 1.9-2.1. To validate my measure of overtime for salaried workers, appendix figure D.2 plots the probability that a salaried worker receives overtime as a function of their weekly base pay, and finds clear discontinuity at the overtime exemption threshold.

C.b Computing Weekly Measure of Income

While the measure of base pay that the Department of Labor uses to determine overtime eligibility is denominated at the weekly level, workers' gross pay and overtime pay are recorded at the monthly level in the data. In this section, I explain the procedure I use to standardize these two key measures of compensation to the weekly level. Table C.1 shows the share of workers with each pay frequency in April 2016, and the formula used to compute their weekly base pay, gross pay, and overtime pay.

Appendix Table C.1
Normalizing Compensation to Weekly Level, by Pay Frequency

	Share of Workers			
Pay Frequency	Hourly	Salaried	Base Pay	Gross & OT Pay
Weekly	.276	.079	S	$\frac{1}{N}$ Y
Biweekly	.653	.570	$\frac{1}{2}$S	$\frac{1}{2N}$Y
Semimonthly	.067	.311	$\frac{24}{52}$S	$\frac{12}{52}$Y
Monthly	.004	.04	$\frac{12}{52}$S	$\frac{12}{52}$Y
All workers	.762	.238		

Notes. The first column shows the four frequencies at which individuals can receive their paycheck. Columns 2 and 3 show the share of hourly and salaried workers with each pay frequency, respectively, in the baseline month across all 19 events studied in the paper. Column 4 shows the formula to normalize salaried workers' standard rate of pay, denoted by S, to weekly base pay for each pay frequency. Column 5 shows the formula to normalize monthly gross pay and overtime pay, denoted by Y, to an average weekly amount conditional on receiving N paychecks in the month.

To derive workers' weekly base pay from their standard rate of pay, I follow the rules set by the Department of Labor and scale each worker's standard rate of pay by their pay frequency (i.e. $\frac{\text{standard pay}}{\text{week}} = \frac{\text{standard pay}}{\text{paycheck}} \cdot \frac{\text{paycheck}}{\text{week}}$). For workers paid weekly or biweekly, I simply multiply the standard rate of pay by 1 and 0.5, respectively, to compute their weekly base pay. For workers paid semimonthly or monthly, the DOL's formula makes the approximation that each month is 1/12 of the year and each year has 52 weeks. Thus, weekly base pay equals standard rate of pay times $\frac{24}{52}$ for workers paid semimonthly, and standard rate of pay times $\frac{12}{52}$ for workers paid monthly.

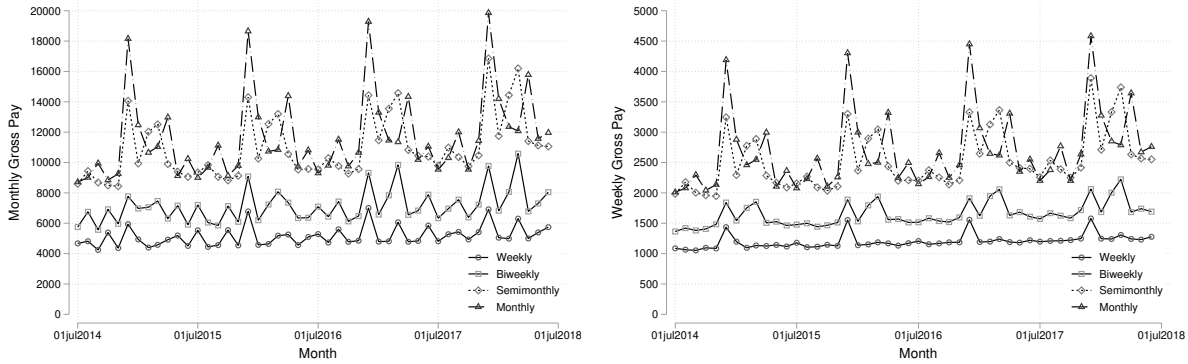
To express the monthly gross and overtime pay variables at the weekly level, I normalize it by the number of paychecks they receive each month and the number of weeks covered

per paycheck:

$$\frac{\text{OT pay}}{\text{week}} = \frac{\text{OT pay}}{\text{month}} \Big/ \left(\frac{\text{paychecks}}{\text{month}} \cdot \frac{\text{weeks}}{\text{paycheck}}\right)$$

This scaling calculation is simple to compute for observations after 2016 since I observe the number of paychecks per month, and the term $\frac{\text{paycheck}}{\text{weeks}}$ is equivalent to the scaling factor used to translate the standard rate of pay to weekly base pay. For observations prior to 2016 though, I have to impute the number of paychecks per month.

I define $\frac{\text{paychecks}}{\text{month}} = 1$ for workers paid monthly and $\frac{\text{paychecks}}{\text{month}} = 2$ for workers paid semi-monthly. For weekly and biweekly paid workers, the number of paychecks received each month depends on both the day of the week that each worker gets paid, and the number of times that day appears in the month. For instance, if a worker gets paid on a Thursday every two weeks, then the worker's gross pay includes 3 paychecks in December 2016 when there were 5 Thursdays, but only 2 paychecks in April 2016. To illustrate this problem, I plot in figure C.1a the monthly gross pay for all continuously employed workers from July 2014 to July 2018, by their pay frequency. Not only do biweekly and weekly paid workers experience spikes in their gross pay, the peaks and troughs do not occur on the same months between years. In contrast, monthly and semi-monthly paid workers only experience a large spike in December of each year, likely reflecting bonuses.



(a) Monthly Gross Pay

(b) Normalized Weekly Gross Pay

Appendix Figure C.1
Gross Income, by Pay Frequency

Notes. Panel (a) shows the average monthly gross pay for a balanced panel of workers from July 2014 to July 2018. Panel (b) shows the average weekly gross pay for the same panel of workers.

While different workers may receive an extra paycheck in different months, employees of the same firm tend to receive a paycheck on the same day of the month, conditional on their pay frequency. To impute the number of paychecks per month that each firm issues in

a month, I apply the following algorithm:

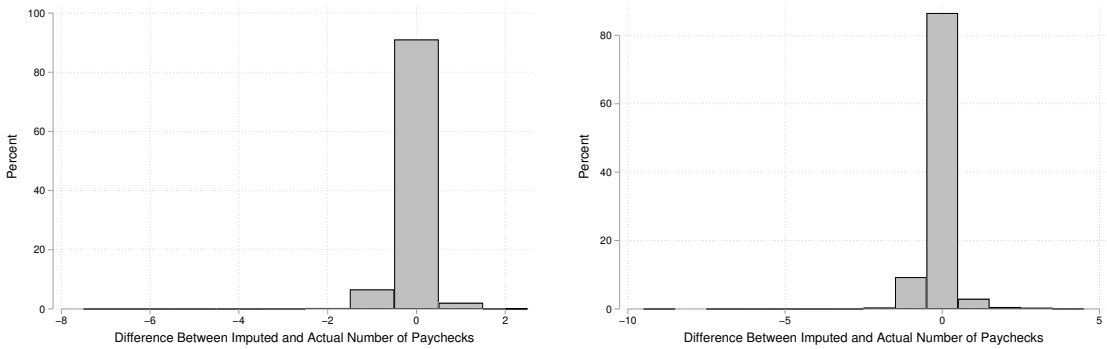
1. Compute the average gross pay across all workers of the same pay frequency within each firm-month.

2. Within each year, for each firm-frequency, compute the median of the 12 average monthly gross pays computed in step 1.

3. Record biweekly workers as receiving 3 paychecks in months where the average gross pay in their firm-frequency exceeds 1.25 times the firm's median gross pay in that year, and 2 otherwise.

4. Record weekly workers as receiving 5 paychecks in months where the average gross pay in their firm-frequency exceeds 1.25 times the firm's median gross pay in that year, and 4 otherwise.

In effect, I assume a firm issued one additional paycheck in the month if the average worker's gross pay for that month far exceeds the median pay in the year. I implement two validation tests of my imputation strategy. First, plotting workers' weekly gross pay implied by their imputed number of paychecks, I show in figure C.1b that the periodic spikes in gross pay among biweekly and weekly paid workers completely disappear. Second, I compare the imputed number of paychecks per month to the actual number of paychecks per month using data post-2016 (see figure C.2). I find that I am able to match the actual number of paychecks for nearly 90% of biweekly paid worker-months and 85% of weekly paid worker-months.



(a) Biweekly (b) Weekly

Appendix Figure C.2
Impute Number of Pay Checks, by Pay Frequency

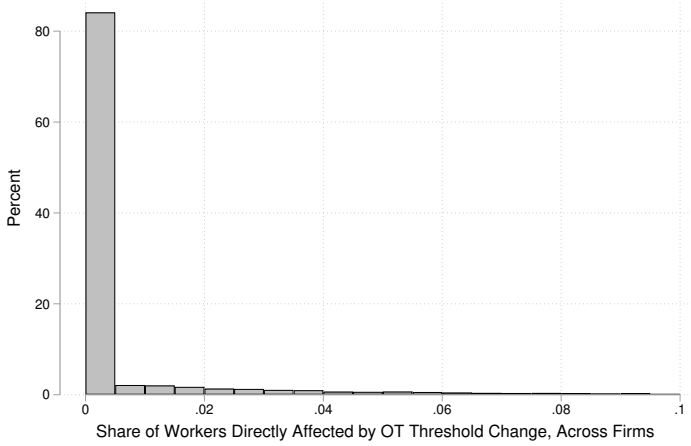
Notes. Panel (a) shows the distribution of the difference between imputed and actual number of paychecks per month, for all worker-months in between July 2016 to July 2018 where the worker is paid biweekly. Panel (b) shows a similar distribution for workers who are paid weekly.

Appendix D. Descriptive Statistics

In this section, I describe the characteristics of the firms and workers in the ADP data.

D.a Directly Affected Firms vs. Entire Sample

Figure D.1 plots the distribution of firms by their share of salaried employees with base pays between the old and new state overtime exemption thresholds in the month before the policy change. The first feature to note is that the majority of firms had less than 0.5% of salaried workers with base pays between the old and new thresholds. Nevertheless, firms with no directly affected workers can still respond to the policy through changes in hiring decisions or spillovers from the reallocation of jobs between firms. Thus, I keep the full sample of firms in my main analysis.



Appendix Figure D.1
Distribution of Share Directly Affected by the State Overtime Policy

Notes. The figure shows the distribution of firms by the share of workers who are paid by salary, and earn between the old and new overtime exemption threshold at baseline.

Table D.1 describes in more detail the characteristics of firms and workers affected by each of the increases in the overtime exemption threshold. In column (1), I record the size distribution, industry mix, and worker composition among firms in the treated states. I find that the sample comprises primarily of small and medium size firms across a range of industries, albeit the data has a smaller share in non-tradeable and a larger share in "other" relative to representative statistics (Mian and Sufi, 2014). Of the firms in the treated states, 17% hire only hourly workers. In column (2), I restrict the sample to only firms with at least one salary worker between the old and new thresholds at baseline. Relative to the

average employer, directly affected firms are over twice as large and have a greater share of salaried workers, but follow a similar industry mix. The observation that larger firms are more susceptible to reforms in the exemption threshold follows from purely a probabilistic standpoint - firms with more employees are more likely to have at least one worker paid within any fixed interval of base pay. Given that the direct response to the rule changes is driven by large firms, there may be concern about the representativeness of my estimates. However, it should be noted that although directly affected firms only make up 21% of the sample, they employ nearly half of all workers. Thus, the response of these large firms is highly relevant to the evaluation of the policy. One final point to note is that the majority of the treated firms are located in California and New York, so my firm-level analyses primarily capture the response of firms located in these states.

Columns (3) and (4) repeat the summary statistics for firms in the states that did not raise their overtime exemption thresholds. I included Colorado and Washington in the control group for all events prior to 2021 as these states simply followed the FLSA's overtime exemption threshold at that time. A "firm" in the control state should be understood as a "firm-state" combination. Overall, I find very similar industry and worker distribution as firms in the treated states. To account for any differences in trends by industry, the main analysis will test for the robustness of my results to keeping only firms that operate in both treated and control states.

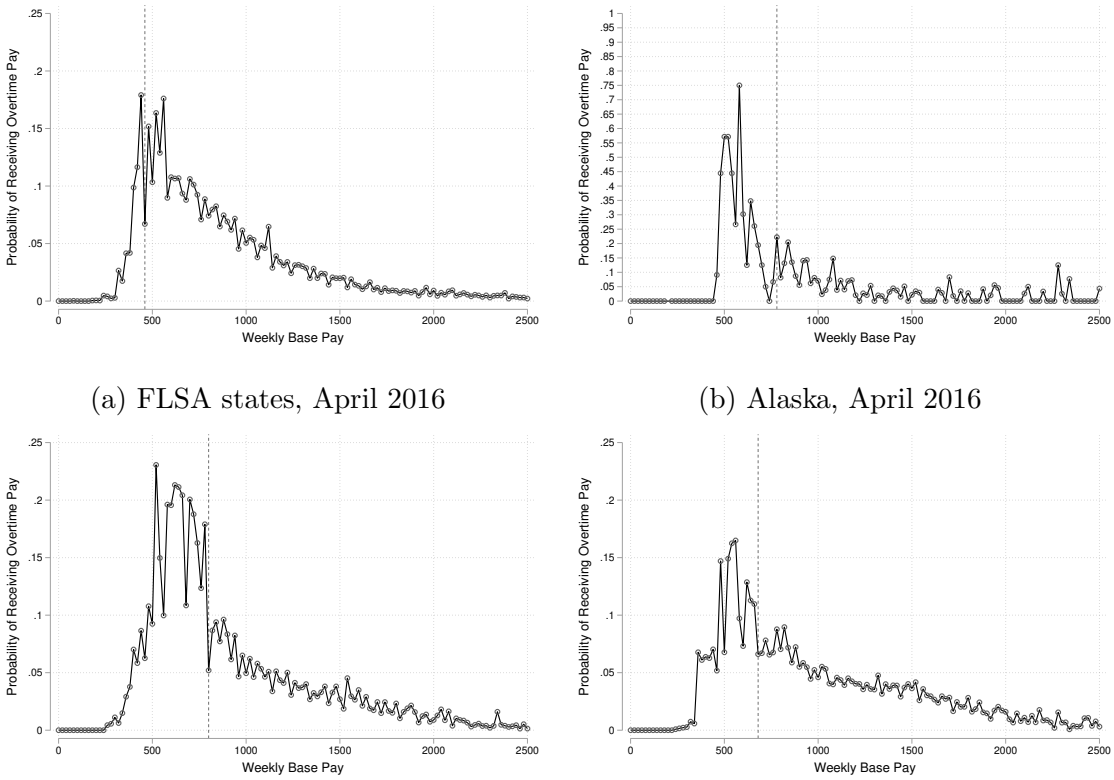
D.b Cross-sectional Evidence of Compliance

Next, I explore whether employers comply with the overtime rules. Figure D.2 plots the probability that a salaried worker receives overtime pay as a function of their weekly base pay. Consistent with compliance to the overtime regulation, salaried workers earning less than the overtime exemption threshold are far more likely to receive overtime pay compared to those earning above it. In particular, the probability of receiving overtime in California and New York in April 2016, exhibits a discontinuous drop at exactly the overtime exemption threshold. The discontinuity likely represents compliance to the policy and not simply the selection of workers who get bunched above the threshold. For example, firms likely prefer to bunch worker who work long hours above the threshold to avoid paying them overtime. In that case, those below the threshold should actually have lower hours than average, and should be *less* likely to earn overtime. Despite the bias against finding a discontinuity, I nevertheless observe that workers paid below the overtime exemption threshold are more likely to earn overtime in California and New York. However, this composition effect could explain the lack of a discontinuity in FLSA states in April 2016.

Appendix Table D.1
Firms Affected by Changes in the Overtime Exemption Threshold

	Treated States		Control States	
	(1)	(2)	(3)	(4)
Firms				
Average Size	87.135	208.169	42.258	151.144
% Firm Size: < 50	.703	.366	.842	.463
% Firm Size: 50-499	.266	.549	.144	.476
% Firm Size: 500-999	.019	.051	.009	.041
% Firm Size: 1000-4999	.011	.031	.004	.019
% Firm Size: ≥ 5000	.001	.002	0	.001
Tradeable	.168	.129	.191	.148
Nontradeable	.044	.061	.04	.063
Construction	.098	.085	.093	.09
Other	.65	.693	.643	.671
Restaurant	.023	.028	.016	.026
Retail	.048	.065	.045	.07
Manufacturing	.168	.135	.189	.156
Share Salaried	.521	.514	.545	.527
Only Salaried	.297	.125	.388	.169
Only Hourly	.174	0	.233	0
Both Salaried and Hourly	.529	.875	.379	.831
Share Treated	.017	.08	.016	.123
Variation				
California	.553	.568	0	0
New York	.333	.385	0	0
Alaska	.007	.002	0	0
Maine	.036	.005	0	0
Colorado	.036	.011	.023	.018
Washington	.036	.029	.024	.022
No. Event-Firms	253,139	52,918	5,698,548	754,045
No. Event-Workers	22,057,267	11,015,864	240,811,385	113,969,262
Sample	All	≥ 1 aff. worker	All	≥ 1 aff. worker

Notes. The table reports the characteristics of firms and workers in the baseline month prior to each threshold change, separately for all firms and for only firms that employed at least one salary worker directly affected by the reform. Columns (1)-(2) report these statistics for firms in states that raised their overtime exemption threshold. Columns (3)-(4) reports similar statistics for firms in the control states. The first three group of rows report the distribution of firm sizes, industry mix, and worker composition of firms. The fourth group of rows captures the distribution of firms across the states that ever raised the overtime exemption threshold from 2014-2021.



(a) FLSA states, April 2016

(b) Alaska, April 2016

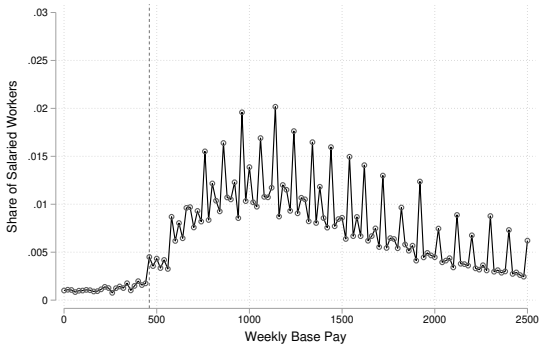
(c) California, April 2016

(d) New York, April 2016

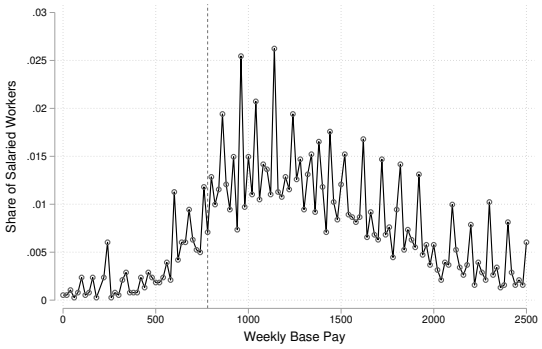
Appendix Figure D.2
Probability of Receiving Overtime Pay, Conditional on Base Pay

Notes. Each graph shows the probability that salaried workers receive non-zero overtime pay in April 2016, as a function of their weekly base pay. The sample in figure (a) is restricted to salaried workers not living in California, New York, Maine, or Alaska. The sample in figure (b) is restricted to salaried workers in Alaska, figure (c) is restricted to California, and figure (d) is restricted to New York. The dotted vertical black line is the overtime exemption threshold in effect for each respective sample.

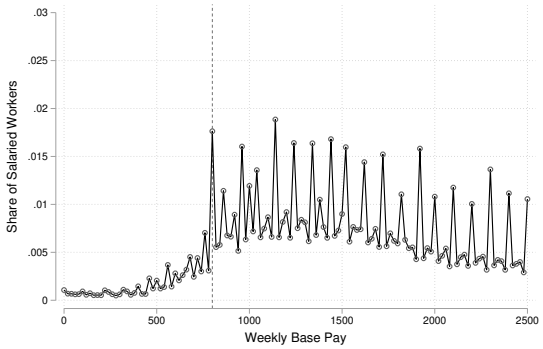
To provide preliminary evidence of a bunching effect, figure D.3 finds clear spikes in the distribution of weekly base pay at the overtime exemption threshold. Since I divided the distribution into $20 bins of weekly base pay, the distribution exhibits spikes at regular intervals that represent annual salaries in multiples of $5,000. Nevertheless, there is evidence of bunching at precisely each state's overtime exemption threshold. and the bunching does not exist at the same salary bin in states where the threshold is not binding. One feature worth pointing out is that there are rarely any workers to the left of the FLSA threshold. That not only provides evidence that few workers are covered for overtime, but also explains why there is no discontinuity in overtime pay at the cutoff in appendix figure D.2a.



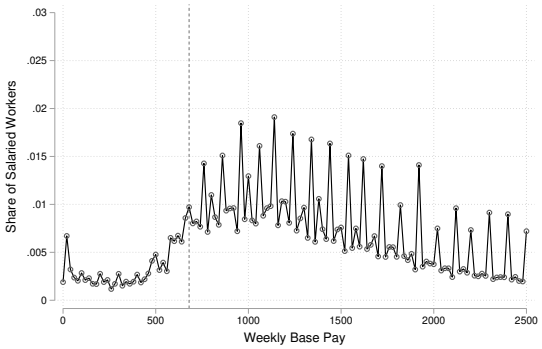
(a) FLSA states, April 2016



(b) Alaska, April 2016



(c) California, April 2016



(d) New York, April 2016

Appendix Figure D.3
Distribution of Weekly Base Pay

Notes. Each graph shows the distribution of weekly base pay for salaried workers in April 2016. The sample in figure (a) is restricted to salaried workers not living in California, New York, Maine, or Alaska. The sample in figure (b) is restricted to salaried workers in Alaska, figure (c) is restricted to California, and figure (d) is restricted to New York. The dotted vertical black line is the overtime exemption threshold in effect for each respective sample.