Taxes Today, Benefits Tomorrow*

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Abstract

How distortive are labor taxes that entitle workers to social benefits? I provide new empirical evidence of bunching in U.S. earnings distribution that show how workers draw a link between taxes and benefits. I find that partially unemployed workers value future preserved benefits when they bunch at the kink of the unemployment insurance benefit-withdrawal schedule. I then extend the bunching formula of Saez (2010) to a dynamic setting that accounts for the value of future benefits tied to taxation. This yields new tests of tax-benefit linkage based on bunching heterogeneity. I verify in quasi-experiments that UI extension programs that decrease the value of future benefits lead to more bunching and to lower labor supply. Last, a quantification exercise of the dynamic bunching formula provides extra support for a strong tax-benefit linkage.

Keywords: tax-benefit linkage, bunching, unemployment insurance

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1 Introduction

Social insurance contributions - for old-age pensions, health, and unemployment insurance - represent 9.2% of 2016 GDP in OECD countries, and this rate doubled over the past 50 years.[1] In many countries social insurance contributions are levied as payroll taxes on wages and represent more than a quarter of overall taxes. By introducing a wedge between workers' net wages and firms' labor cost, payroll taxes may distort labor market allocation and lead to inefficiently low levels of employment. However, the size of tax distortions depends on the workers' valuation of social insurance and the link they draw between their payroll taxes and the future benefits. If workers value social insurance and if benefit eligibility is tied to taxed employment, workers are willing to accept lower net wages internalizing that the overall compensation package includes social insurance. As in the classical example of mandated benefits (Summers, 1989), these conditions dampen distortionary effects of taxation. In this paper, I study whether workers draw the link between taxes paid today and social insurance benefits received later on.

The paper advances the literature on several dimensions. First, I provide evidence of bunching in earnings distribution that directly support the existence of tax-benefit linkage. Second, on the theory side, I extend the bunching formula of Saez (2010) to a dynamic setting that accounts for the value of future benefits tied to taxation. Third, and informed by the theory, I find in the data that policy shocks that decrease the future value of benefits while holding taxation constant lead to higher bunching and to lower labor supply.

I analyze the tax-benefit linkage embedded in the partial unemployment insurance rules in the U.S. (McCall, 1996). Partial UI allows unemployment insurance claimants to keep some unemployment benefits while they work in low-earnings jobs – usually part-time or temporary work. In most U.S. states, weekly labor earnings below a certain threshold, termed the *disregard*, do not trigger any reduction in current benefits. Then, for every dollar earned above the disregard level, current weekly benefits are reduced on a dollar-per-dollar basis. The benefit withdrawal schedule then implies a 100% marginal tax rate on earnings above the disregard level, drawing a large kink in the claimant budget set. Importantly, the reduction in current benefits leads to future entitlement, i.e. preserved benefits can be paid later in the claiming spell. In other words, intertemporal benefit transfers delay the potential benefit exhaustion date. I study whether claimants bunch at the disregard level in reaction to the kink in the benefit withdrawal schedule, and whether bunching depends on the expected value of preserved benefits.

[1]The OECD publishes aggregate statistics on social insurance at the following link: https://data.oecd.org/tax/social-security-contributions.htm

I compute bunching estimates using UI administrative data from four U.S. states: Idaho, Louisiana, New Mexico and Missouri. The data come from the Continuous Work and Benefit History (CWBH) project.[2] I find substantial bunching at the disregard level. In Idaho and Louisiana, the excess mass of workers at the disregard is five times the population density of workers that would earn this amount absent the kink. I also observe that a significant fraction of claimants have earnings above the disregard amount. This observation is consistent with claimants reacting to an effective marginal tax rate that is lower than the static 100% benefit-withdrawal rate.[3] Two empirical facts rule out the alternative view where *myopic* claimants react to the 100% benefit-withdrawal rate only, but face frictions preventing them from avoiding their strictly dominated earnings region just above the disregard. Under this alternative view, the fraction of workers with earnings above the disregard should decrease over time as workers gradually overcome frictional constraints. I find the opposite pattern in the data: the fraction of claimants above the disregard increases over the claim (holding the sample of claimants constant over time). In addition, bunching patterns follow within a few months changes in the disregard levels triggered by a reform in Louisiana, which further suggests low adjustment costs or small frictions in the market of low-earnings jobs.

Building on these suggestive empirical evidence, I develop a job-search model with frictionless labor supply while on claim, which shows how tax-benefit linkage affects claimants' behavior. In the model, job seekers work in low-earnings jobs while they search for permanent jobs ineligible for partial UI. They make their labor supply decisions in the low-earnings market based on an effective *dynamic* marginal tax rate that accounts for the present value of benefits preservation. The dynamic marginal tax rate depends not only on the discount factor, but also on the claimants' expected probability to find a permanent job and on the horizon over which preserved benefits are rolled over. First, if the claimant expects to rapidly find a permanent job and to exit the UI registers, then she is less likely to profit from the benefit-preservation mechanism and her dynamic marginal tax rate is larger, closer to the static benefit-withdrawal rate. Second, if the claimant is entitled to many weekly payments of unemployment benefits (long potential benefit duration), preserving benefits is less valuable. She is less likely to profit from a delay in the date when her benefit payments exhaust, as she would just have found a permanent job before her claim ends. Thus, claimants with longer potential benefit duration have larger effective marginal tax rate. I then show that excess bunching at the kink of the partial-UI schedule equals the product

[2] I thank Camille Landais for sharing the data. See Moffitt (1985) and Landais (2014) for more details about the data.

[3] With 100% benefit-withdrawal rate, net income is flat above the disregard level.

of the earnings elasticity to the net-of-tax rate and of the change in the effective dynamic marginal tax rate at the kink. This extends the bunching formula of Saez (2010) to contexts where taxes entitle workers to future benefits.

Consistent with the model-based implications of tax-benefit linkage on effective tax rates, I find in my data that claimants with longer potential benefit duration - i.e. initial number of benefit weeks if totally unemployed - bunch more. This is confirmed when focusing on within-individual variation in potential benefit duration across claims, or when focusing on exogenous variations in potential benefit duration initiated by triggers of emergency UI extension programs. This provides quasi-experimental evidence on the existence of tax-benefit linkages, building on discontinuities in tax schedules à la Saez and on well-identified policy shocks.[4] I also verify in the cross-section that bunching estimates are larger for claimants with a low propensity to remain on the UI registers, such as claimants expecting to be recalled by their previous employer (Katz and Meyer, 1990b).

Lastly, I quantify whether the tax-benefit linkage mechanism may rationalize claimants' observed labor supply. I then assume that claimants have rational expectations about their permanent job finding rate. I estimate a hazard model of exiting the UI registers that depends on workers' socio-demographics (age, gender, education), their claim characteristics (benefit level, potential benefit duration) and their recall expectations. This enables to predict the expected probability of exiting the UI registers for each claimant and to compute their model-based dynamic marginal tax rate. I find that the average effective tax rate amounts to 55%, significantly lower than the 100% static benefit-withdrawal rate. The bunching formula then identifies the earnings elasticity to the net-of-tax rate, whose estimate lies between 0.1 and 0.2. This estimate is in line with the consensus estimates in the literature (see the review of quasi-experimental estimates in Chetty (2012) or Chetty et al. (2011b)). Overall this confirms the explanatory power of perfect tax-benefit linkage under rational expectations.

My paper contributes to the literature on the effects of benefit taxation. I provide new empirical evidence of tax-benefit linkage leveraging bunching estimates and policy shocks. Close to my paper, Liebman et al. (2009) find that workers are more likely to retire when the marginal effect of labor supply on future social security benefits is low. Their identification uses non-linearities in the value of benefits only, while I combine both sharp discontinuities in the current taxation schedule and quasi-experimental variations in the value of future

[4] Moffitt and Nicholson (1982); Farber et al. (2015); Johnston and Mas (2018) use triggers of emergency UI extension programs, either in the 80s or during the Great Recession, as a source of exogenous variation in the U.S. unemployment insurance generosity.

benefits.[5] My result on tax-benefit linkage suggests that social contributions are not pure taxes and their incidence would fall on workers as found in Gruber (1997b), Anderson and Meyer (1997a, 2000) and Bozio et al. (2020).[6] My results also suggest that entitlement effects matter when assessing the dynamic effects of unemployment insurance (Mortensen, 1977; Hamermesh, 1979; Kuhn and Riddell, 2010), of welfare programs with time limits (Grogger and Michalopoulos, 2003), and of social security programs (Coile and Gruber, 2001), especially their earning tests (Friedberg, 1998, 2000; Haider and Loughran, 2008; Gelber et al., 2019).[7]

I contribute to the bunching literature (see Saez (2010) and Chetty et al. (2013), and the review in Kleven (2016)). Kleven (2016) writes that "Extending the bunching approach to dynamic settings is still in its infancy." (p.13) The bunching literature on labor supply adopts a static framework, where workers do not link taxes and future benefits. This assumption is questionable when assessing the earnings elasticity to payroll taxation. I show how the bunching strategy and the Saez (2010) formula can be extended to account for these dynamic aspects. The closest papers in the bunching literature are Brown (2013) and Manoli and Weber (2016). They both study bunching at the legal retirement age taking into account that delaying retirement increases either severance payments or future annuities. In this context, they adopt either a static model of lifetime labor supply, or a non-stochastic dynamic model of annual labor supply. The dynamic approach of my paper allows for *stochastic* events generating benefit payments, which are essential features of any social insurance linked to payroll taxation.[8]

My paper also fills a literature gap on the effect of unemployment insurance. I study partial unemployment insurance programs that are widespread in OECD countries. In 2017, 11%

[5]In the context of Liebman et al. (2009), uncertainty smooths discontinuities in the social security benefit rule, which prevents them from adopting a standard regression discontinuity approach. Discontinuities in the schedule of current taxation do not suffer from uncertainty smoothing, so that I can adopt a bunching approach.

[6]Note that payroll tax incidence on equilibrium wages does not only depend on workers' tax-benefit linkage, but incidence also depends on labor supply and labor demand elasticities and on wage rigidities induced by labor market institutions. Even the "full shifting [of payroll tax shocks to wages] can either be due to elastic demand, inelastic supply or due to full tax/benefit linkages" (p.S77 Gruber, 1997b). Consequently, tax incidence is an indirect test of tax-benefit linkage.

[7]Gelber et al. (2019) discuss intertemporal aspects of the U.S. Social Security Annual Earnings Test. Reductions in current benefits can lead to increases in future scheduled benefits (i.e. benefit enhancement mechanism). However, benefit enhancement is triggered only when a sufficient amount of current benefits is reduced. Thus there is no difference between the static benefit-reduction rate and the dynamic marginal tax rate at the kink in the SSAET schedule. Consequently, bunching at the SSAET kink, studied in Friedberg (1998, 2000) and Gelber et al. (2019), is not informative about tax-benefit linkage.

[8]le Maire and Schjerning (2013) also consider dynamic aspects in income tax schedule, but they specifically model income shifting by the self-employed.

of UI claimants in OECD countries work while on claim.[9] In the U.S., McCall (1996) and the early contributions of Holen and Horowitz (1974) and of Kiefer and Neumann (1979) document the behavioral response at the extensive margin, i.e. whether claimants take up low-earnings jobs when partial UI is more generous.[10] I provide the first evidence on significant behavioral response to partial UI rules at the intensive margin.[11] Conditional on working in low-earnings jobs, I show how labor earnings react to changes in partial UI generosity.

The paper is organized as follows. In Section 2, I describe the U.S. partial unemployment insurance program. In Section 3, I introduce the data and show some descriptive statistics. In Section 4, I provide visual evidence of bunching patterns and their dynamics. In Section 5, I develop a job-search model of claimants working while on claim. In Section 6, I test the tax-benefit linkage using bunching heterogeneity in quasi-experiments. In Section 7, I show that perfect tax-benefit linkage quantitatively explains bunching patterns with reasonable earnings elasticity to the net-of-tax rate and under the rational expectation hypothesis. Section 8 concludes.

2 Institutional background

In the U.S., when unemployment insurance (UI) claimants work while on claim, they are eligible for partial unemployment benefits, provided that they do not earn more than a maximum amount of labor income per week.[12] This maximum amount is usually set as a function of the weekly benefit amount (WBA), which is the unemployment benefits (UB) payment when claimants do not work, i.e. total unemployment benefits.[13] Except for this maximum amount, there is no other specific eligibility condition for partial UI. Claimants must only meet the usual UI eligibility requirements (such as active search for jobs, see Appendix A). Partial-UI claimants are allowed to work for any employer, including their

[9] See OECD data for national shares of partial-UI claimants, such as 33% in Sweden, 22% in Finland and 6% in Portugal (OECD, 2020). See Kyyra (2010) for older figures.

[10] In Europe, Kyyra (2010), Caliendo et al. (2016), Kyyra et al. (2013), Fremigacci and Terracol (2013) and Godoy and Roed (2016) study the effects of partial-UI jobs on regular employment.

[11] Following a previous version of my work, Lee et al. (2019) find significant fiscal externalities of the partial UI program when pooling together the intensive- and extensive-margin responses to a small increase in the benefit reduction rate (by 9 percentage point). Their estimates on labor earnings are statistically consistent with the intensive-margin elasticity estimated in this paper.

[12] All components of labor income are considered in the computation, except payments for jury service in some states like New Mexico.

[13] Appendix Table A1 displays the partial UI parameters for the four states covered by my dataset.

past employers; claimants who are temporarily laid off are also eligible for partial UI. [14]

Partial-UI claimants are paid their weekly benefit amount when their weekly earnings are below some state-specific *disregard* thresholds. When partial-UI claimants earn between the disregard level and the maximum eligibility amount, their current benefits are reduced by their earnings minus the disregard. The static marginal benefit-reduction rate is then 100%.

Figure 1 illustrates the partial-UI schedules for the four states and for the time period covered by my dataset: Idaho (ID), Louisiana (LA), New Mexico (NM) and Missouri (MO) in the late 70s and early 80s. As partial-UI rules have hardly changed since then, the discussion of the institutions refers to both current rules and rules of the late 70s and early 80s, except when I discuss nominal amounts.[15] I plot the weekly net income (earnings plus UB payments) against the weekly earnings while on claim.[16] I normalize earnings and UB payments by the WBA, as the maximal amount and the disregard are expressed as a fraction of the WBA for three of the four states (see summary in Appendix Table A1). The figure illustrates that the schedule is kinked at the disregard amount. Intuitively, one expects claimants to bunch at the disregard level. In addition, from a *static* point of view, there are no incentives to work for wages right above the disregard, as the net income is essentially a plateau above that level.

The graphics also illustrate the notches at the maximal eligibility amount in Louisiana, New Mexico and Missouri (see Munts (1970) for an early discussion on notches in the U.S. partial-UI schedule). Notches generate even stronger disincentives to work than kinks, as claimants lose income when they work above the notch threshold (Kleven and Waseem, 2013). Because of data limitations, I will not analyze the claimants' behavior around notches. The incentives to claim drop discontinuously at the notch value, so that workers above the notch should leave the UI registers, and hence my data. There are no such data concerns around the kink.[17]

As will become clear below, I take advantage of the absence of kinks in Missouri at the

[14] Also, individuals whose hours have been reduced at their current workplace are eligible for partial UI, as long as they can file a claim based on this reduction in hours worked. Claimants with reduced hours represent only a small share of partial-UI claimants. In the CWBH data used in this paper, I can only distinguish between claimants taking up new jobs and claimants with reduced hours (Short Time Compensation) in Louisiana since 1982. From 1982 to 1984, only 15.7% of partial-UI weeks in Louisiana concerned claimants with reduced hours.

[15] See Chapter 3 of the 2019 DOLE booklet "Comparison of the State Unemployment Laws"

[16] The plots and my analysis abstract from income taxes. I expect partially unemployed claimants' labor income to be lower than minimum taxable income thresholds. Similarly, only unemployment benefits over some minimal thresholds are subject to income taxes.

[17] Just above the disregard level, claimants still have strong incentives to remain on the UI rolls. If they leave, their total income drops.

disregard level prevailing in Idaho and Louisiana ($0.5 \times WBA$) to perform a first placebo exercise. In a second placebo exercise, I will conduct a difference-in-difference analysis around a policy shock on the disregard amount in Louisiana.

I now turn to the dynamic aspects of the partial-UI rules. At the beginning of each claim, the UI administration computes the claimant's weekly benefit amount (WBA) and potential benefit duration (PBD), which both depend on past earnings. The PBD typically varies between 10 and 26 weeks (see Appendix A for more details). The product of the WBA and the PBD is called the (total) benefit entitlement which I denote B_0. The benefit entitlement can be thought of as a kind of UB capital that depreciates over time with UB payments. If claimants are totally unemployed all along their claim, they receive each week their WBA, and their benefits lapse after $PBD = B_0/WBA$ weeks. When claimants are only paid part of their WBA in a given week, the unpaid amount is rolled over to a later week in the claim and the UB capital depreciates at a slower pace. Let me take the example of a claimant entitled to a WBA of 300 dollars over a PBD of 10 weeks. Under total unemployment, the claimant receives 300 dollars every week from week 1 to week 10. If in week 5 the claimant takes up a low-earning job that reduces her UB payment by 150 dollars in that week, she is entitled to a 150 dollar payment in week 11.

The partial UI rules clearly draw a link between benefit withdrawal/taxes above the disregard level and future benefits. Intuitively, the benefit preservation mechanism provides incentives to earn weekly wages above the disregard level and should lower the amount of bunching at the disregard level. The theoretical model in Section 5 shows formally how the tax-benefit linkage affects bunching. In words, working while on claim, with earnings above the disregard level, is thus a way to delay the benefit exhaustion date.

In principle, there is one limitation to the possibility to delay exhaustion, as any remaining UB capital is lost one year after the first claiming week, defined as the *benefit year*. However, in the data, almost all claimants exhaust their benefits or find a regular job before the end of the benefit year. Consequently, I abstract in the remainder from any horizon effects of the benefit-year rule.

The partial-UI rules described above remain in place when additional UI programs are triggered because of tough labor market conditions. During the late 70s and early 80s, there were two additional programs in place - the Extended Benefit (EB) program (Tier II) and the Federal Supplemental Compensation (FSC) program (Tier IV) - which both increased the potential benefit duration of claimants (see more details in Appendix A). I leverage the quasi-experimental variation in PBD generated by these policy shocks to test for the tax-benefit linkage in Section 6.

3 Data

I use individual panel data from the Continuous Wage and Benefit History project.[18] The CWBH project collected weekly claims for a random subsample of UI claimants in the U.S., and the resulting dataset has the unique advantage of including the *weekly* earnings that claimants report to the UI administration and the consecutive UB payments.[19] I can thus characterize whether claimants are partially unemployed. The data cover four U.S. states – Idaho, Louisiana, Missouri and New Mexico – during the late 70s and early 80s.[20] The data include all relevant information about the claim: weekly benefit amount, total entitlement, and pre-unemployment earnings. Socio-demographics characteristics are also available, such as gender, age, education, ethnicity, past firm sector and industry, and past occupation. In addition, the data set includes survey information about recall expectations for a subsample of claimants.

Table 1 reports descriptive statistics of partial UI claimants by state. I select unemployment spells where workers are partially unemployed for at least one week. This amounts to 32.1% of the initial sample of spells. The share of men varies between 50% to 70% across states. Claimants are in their early 30s, with around 11 years of education on average. Manufacturing is the most common industry of the pre-unemployment firms except in Louisiana where construction is as important. The weekly benefit amount (WBA) is around $100 (current dollars) and the average replacement rate is between 40% and 50%. The potential benefit duration (PBD) is greater than 26 weeks (the maximal PBD in Tier 1) as the early 80s is a period of high unemployment, and of UI extensions. The average claiming duration is around four months.

4 Distribution of earnings while on claim

In this section, I document the bunching patterns found in the unemployment insurance administrative records, and explain why they are unlikely to arise from a pure static behavior of claimants.

[18]The CWBH data have been analyzed in Katz and Meyer (1990b), Katz and Meyer (1990a), Anderson and Meyer (1993), Anderson and Meyer (1997b) and Landais (2014), among others. To the best of my knowledge, partial UI has never been analyzed in the CWBH data.

[19]One concern is that claimants manipulate their earning reports to become eligible for partial UI. However the UI administration takes action to limit false statements. The UI administration performs random audits of claimants' declarations. The UI administration currently cross-checks W-2 and new hires declarations of employers with claimants reported earnings. If fraud is detected, it can be severely punished as a Class VI Felony. Criminal action may result in up to 2 years in prison and fines up to $150,000 for each false statement.

[20]The CWBH project collected data for other states but weekly earnings were missing for these.

4.1 Bunching patterns

Figure 2 displays the weekly earnings density reported by UI claimants together with the empirical partial-UI schedule state by state. In line with the partial-UI rules in Idaho, Louisiana and New Mexico, I normalize the weekly earnings by the weekly benefit amount. The empirical schedules (dashed blue lines), which describe the actual total weekly income (unemployment benefits plus earnings) as a function of weekly earnings, closely follow the theoretical schedules displayed in Figure 1. The upper panels - Idaho and Louisiana - clearly display bunching at the level of the disregard (50% of the weekly benefit amount). In Louisiana (upper right-hand panel), there is also a sharp drop in the density at the weekly benefit amount, when claimants are no longer eligible for partial UI. This may be related to the notch in the schedule, but it can also be due to the fact that individuals have no incentives to stay registered above this "exit" level. In New Mexico, where the disregard level is only 20% of the weekly benefit amount, bunching is less salient (lower left-hand panel). The lower right-hand panel illustrates a placebo test. In Missouri, the level of disregard is \$10, so that the schedule is totally flat when earnings amount to $0.5 \times WBA$. There is indeed no bunching at this placebo level. Thus the bunching observed in Idaho or Louisiana at $0.5 \times WBA$ is unlikely to be an artifact of other labor legislations or norms, or of hour constraints according to which claimants take some part-time jobs that provide roughly one fourth of their previous wages.[21]

I quantify the extent of bunching along the lines of Chetty et al. (2011a). I fit a polynomial on the earnings density of partial-UI claimants, taking into account that there may be bunching in a bandwidth around the disregard, and that the bunching mass comes from the earnings distribution above the disregard. I report the details of the estimation procedure in Appendix B. Appendix Figure B1 also shows that the procedure fits well the earnings distribution in each state.[22] Table 2 reports the results of the bunching estimation for each state. I find that in Idaho and Louisiana, the mass bunched at the disregard level is around five times in excess to the mass that would have been at the disregard level, had the kink disappeared. Excess bunching is highly statistically significant. In New Mexico, the excess bunching mass amounts to 1.2 times the counterfactual density at the disregard level, but is not statistically significant. In Missouri, there seems to be a missing mass of claimants at the placebo threshold level, confirming the visual placebo test above.

[21] $0.5 \times WBA$ amounts to one fourth of previous wages, as the weekly benefit amount is around 50% of previous wages (see replacement rates in Table 1).

[22] The procedure fits a polynomial of degree 7. The bandwidth around the kink threshold is such that $-\underline{R} = -5\$$ and $\bar{R} = 2$.

4.2 Workers above the disregard level

Another important feature of the earnings distribution in Figure 2 and in Appendix Figure B1 is the substantial fraction of claimants working for earnings above the disregard level. This observation is inconsistent with *myopic* claimants reacting to the static 100% benefit-reduction rate only. Indeed, myopic claimants have no incentive to work above the disregard level. In this paper, I argue that forward-looking claimants have incentives to supply labor above the disregard level as they remain entitled for the withdrawn benefits later on. Before moving on to modeling this behavior, I consider an alternative explanation whereby myopic claimants face *frictions in the market for low-wage jobs* that prevent them from working below or at the disregard level. Arguably, under the assumption of labor market frictions, myopic claimants should gradually leave the dominated region above the disregard level. As time passes, they receive and accept wage offers closer to their notional labor supply (below or at the disregard level). I test this implication using a sample of claimants with multiple weeks of partial unemployment. I select from the Idaho and Louisiana samples claimants with at least eight weeks of partial unemployment and I study their labor supply over these first eight partial-UI weeks to avoid composition effects. Figure 3 plots the share of claimants working above the disregard level over time. As the share rather increases over the spell, claimants do not leave the earnings region that would be dominated for myopic claimants. This does not provide support for the alternative view.

I further test the importance of adjustment cost or labor market frictions leveraging the policy reform in Louisiana (see Gelber et al., 2019, for a formal exposition of adjustment cost test with bunching patterns). In April 1983, Louisiana changed UI rules. The change in partial UI affected both the stock of individuals registered in April 1983 and new inflows after that point in time.[23] The disregard level was reduced from $0.5 \times WBA$ to \$50 for all claimants whose WBA is more than \$100. I select claimants with over \$100 WBA who claim between October 1982 and September 1983, from six months before the reform to 6 months after, and study how the bunching location changes following the policy shock. Under the assumption of frictions in the market for low-wage jobs, bunching should remain at the pre-reform disregard level just after the reform, and should only gradually move to the post-reform disregard level. Figure 4 plots the density of the weekly earnings divided by the weekly benefit amount for the pre-reform subsample (blue solid line) and for the post-reform subsample (dashed red line). As in Figure 2, the pre-reform density shows a large amount of bunching at the pre-reform disregard level (red vertical line). The fraction of

[23] There was also a reduction in the maximum number of entitlement weeks from 28 to 26 weeks. This could have affected the amount of bunching, but not its location.

workers bunching at the pre-reform level shrinks significantly in the post-reform density. Pre-reform bunching workers reduce their labor supply as the disregard level decreases. Figure 5 plots the monthly evolution of the bunching estimates at the pre-reform disregard. The pre-reform bunching pattern disappears within a few months after the reform, which suggests that adjustment costs in the market for low-wage jobs are small.[24]

The presence of small labor market frictions in the partial UI context contrasts with the existing evidence from Chetty et al. (2011a) and Kleven and Waseem (2013). One important difference is that we focus on *weekly* labor supply in low-earning jobs, while the other papers analyze *yearly* earnings in any jobs whatever their hourly rates and hours worked. The frequency and job types may explain our difference in findings.

Overall, these empirical patterns of the earnings distribution provide motivation for the model assumptions in the next section. Namely, I model forward-looking claimants who are aware of the partial UI rules and do not face large frictions when supplying labor in the market for low-wage jobs eligible for partial UI. Note that the empirical patterns in this section do not rule out career concerns as drivers of labor supply. However, we discuss below how the bunching heterogeneity tests are robust to the career concern channel, also called the stepping-stone effects.

5 Theoretical model

In this section, I develop a model of job seekers working while on claim that incorporates the dynamic aspects of the partial-UI program. The objective is to derive the implications of tax-benefit linkage on claimants' labor supply. I show that claimants make their labor supply decision based on a dynamic marginal tax rate, which is lower than the static marginal benefit-reduction rate, because job seekers value the expected benefit transfers generated by their work while on claim. I derive a modified bunching formula à la Saez (2010) that links bunching, effective marginal tax rate, and earned income elasticity. The modified bunching formula allows me to test for tax-benefit linkage.

5.1 Setup

I consider an infinitely lived individual i, claiming benefits from period 0 on. Following UI rules, periods are weeks in my model. Until she finds a *permanent* job, the job-seeker may

[24] In Appendix E, I further document that the new location of bunching is exactly at $50. I also use claimants whose WBA is less than $100 as a control group in a difference-in-difference analysis. Results show that bunching is clearly related to the kink in the partial UI schedule.

work in a *low-earnings* job, corresponding to short-term or part-time work eligible for partial UI. In the remainder, I also refer to these low-earnings jobs as *low-wage* jobs. I assume that low-wage and permanent jobs are different types of jobs. The market for low-wage jobs is tight, and there are no search frictions. On the contrary, the market for permanent jobs features search frictions.

The job-seeker's earnings in the low-wage job in period t are denoted z_t. In line with Saez (2010), I do not make any distinction between wage rates and hours as those different components are not observed in the data.[25] The per period utility $u_i(c_t, z_t)$ of job-seeker i depends on consumption c_t and on labor earnings in low-wage jobs z_t - the latter dependence captures disutility of labor. The individual heterogeneity in preferences is smoothly distributed in the population, so that earnings z_t would also be smoothly distributed in the absence of any kinks in the benefit reduction schedule. This is the key assumption of the bunching identification strategy.[26] In the baseline model, I assume that the job-seeker is risk-neutral.[27] Then it is convenient to parametrize the period utility function as follows:

$$u(c_t, z_t; n_i) = c_t - \frac{n_i}{1 + 1/e} \left(\frac{z_t}{n_i} \right)^{1+1/e} \tag{1}$$

where n_i is an individual ability or preference parameter - smoothly distributed in the population - and e is the parameter capturing the earnings elasticity to the net-of-tax rate. As discussed in Saez (2010), the identification argument also holds with more general utility function as long as the individual heterogeneity is smoothly distributed. Such a parametrized utility function is convenient, as the heterogeneity parameter n_i then equals the earnings level in the absence of any benefit-reduction (as derived below). n_i may capture both individual taste for work and ability, I choose to refer to n_i as individual ability in the remainder.

At each date $t > 0$, the job-seeker may find a permanent job with probability $(1 - p)$. Then she leaves the unemployment registers. Permanent jobs yield the expected intertemporal utility W, which is assumed to be greater than the continuation value of unemployment at any period. Therefore, claimants never decline permanent job offers. In the baseline model, I assume that the probability to find a permanent job does not depend on the amount of

[25] Alternatively, one can think of the wage rate as being fixed and that the job-seeker chooses the number of hours worked.

[26] When turning to estimation, we further assume a polynomial shape for the smooth distribution of individual heterogeneity. As discussed in Blomquist et al. (2017), this is an important assumption. We test the robustness of the main estimates when varying the degree of the polynomial.

[27] In Appendix C.2, I provide a model extension with risk-averse workers and discuss how risk-aversion affects bunching.

earnings in low-wage jobs. This follows the empirical literature that finds small effects of partial-UI jobs on permanent employment (McCall, 1996; Kyyra, 2010; Caliendo et al., 2016; Kyyra et al., 2013; Fremigacci and Terracol, 2013; Godoy and Roed, 2016).[28] I introduce individual heterogeneity in p in Section 5.4. Consistent with the view that the markets for low-wage and permanent jobs are separated, I assume that the utility derived from permanent job W is not related to the individual ability in low-wage jobs.

At the beginning of her claim, the job-seeker has a UB capital - total benefit entitlement - equal to B_0. Weekly benefit payments are deducted from the UB capital, so that B_t, the current entitlement at the beginning of period t, decreases over the spell. At each period that she does not work at all (total unemployment), the job-seeker receives an amount b of unemployment benefits, or the remaining entitlement B_t if her current UB capital is not large enough to pay b. If the job-seeker does not work at all along her unemployment spell, she receives benefits during $\bar{t}^{Utot} = B_0/b$ periods.[29] When she takes up a low-wage job with earnings z_t in a given week, she receives an amount $b - T(z_t)$ of unemployment benefits, where $T(z_t)$ is the reduction in benefits. This reduction in benefits $T(z_t)$ is then "transferred" to a later period within the claim. When benefits are exhausted, the job-seeker leaves the unemployment registers, but she still looks for a permanent job and she may still work for a low-wage job. The partial-UI schedule $T(.)$ is defined as:[30]

$$T(z) = \begin{cases} 0 & \text{if } z < z^* \\ z - z^* & \text{if } z \in (z^*, z^* + \min(b, B_t)) \\ \min(b, B_t) & \text{if } z > z^* + \min(b, B_t) \end{cases} \tag{2}$$

where z^* is the amount of disregard. The partial-UI schedule feature two kinks: the marginal benefit reduction rate jumps from 0% to 100% at the disregard level z^*, and comes back to 0% at the maximum earnings level. Except at the very end of the claim, the remaining UB capital B_t is greater than the weekly benefit amount b and the maximum earnings level equals $z^* + b$. As explained in Section 2, I will abstract from the second kink as data limitation prevents me to analyze behaviors around the maximal earnings amount.[31] This

[28] I discuss in Appendix C.3 the model extension to potential stepping-stone effects or job-search crowding-out effects of low-wage jobs.

[29] The model parameters b and $\bar{t}^{Utot}$ correspond to the following institutional parameters: Weekly Benefit Amount (WBA) and Potential Benefit Duration (PBD).

[30] I assume here that current benefit reduction can reach the actual weekly benefit amount b, as in Idaho. In other states, the maximal amount earned by partial-UI claimants is smaller. However this simplification does not affect the identification as the focus is on earnings close to the kink.

[31] I observe earnings reported to the UI administration. When individuals earn more than the maximal amount, there are no incentives to remain on the UI register and report earnings.

second kink does not affect my identification strategy that is local and around the first kink.

Let me define $U(B_t; n_i)$ the value of unemployment of job-seeker i when the UB capital is B_t. At each date, the job-seeker with discount factor β, maximizes the following program:

$$U(B_t; n_i) = \max_{c_t, z_t} u(c_t, z_t; n_i) + \beta \left[pU(B_{t+1}; n_i) + (1-p)W \right] \tag{3}$$

such that

$$\begin{cases} c_t & = z_t + \min(b, B_t) - T(z_t) \\ B_{t+1} & = B_t - \min(b, B_t) + T(z_t) \\ B_{t+1} & \geq 0 \end{cases}$$

The first constraint of the program is the current budget constraint. I assume that workers cannot save or borrow, as UI claimants are likely to be low-skilled workers who are credit-constrained. The second constraint captures the endogenous entitlement reduction (or UB capital depreciation). The UB capital is reduced by the UB payment $\min(b, B_t) - T(z_t)$. The last constraint states that job-seekers cannot borrow UB entitlement from the UI administration.[32]

5.2 Model solution

I focus on the case where the UB capital is strictly decreasing, and I define $\bar{t} < \infty$ the exhaustion date, i.e. the first date when $B_t = 0$.[33] The exhaustion date is endogenous, as it depends on the solution path of z_t. I describe the model solution for a period when $B_t > b$. This is more relevant to the empirical analysis, as most of my observations are in this case. The case $B_t < b$ is reported in Appendix C.1. For all z_t, such that $T(.)$ is differentiable at z_t, the first order condition is:

$$\underbrace{u_c(c_t, z_t; n_i) \left(1 - T'(z_t)\right)}_{(I)} + \underbrace{\beta p T'(z_t) U'(B_{t+1}; n_i)}_{(II)} = -u_z(c_t, z_t; n_i) \tag{4}$$

where u_c is the marginal utility of consumption and u_z the marginal disutility of work. Equation (4) equates the marginal gains of work (on the left-hand side) with the marginal cost of effort (or disutility of work). The marginal gains have two components. The first

[32] For the sake of simplicity, I do not model the fact that any remaining entitlement at the end of the benefit year is lost, as in the data, almost all job-seekers find permanent jobs or exhaust their UB entitlement before that date.

[33] I show, in Appendix C.1, that such a focus is relevant when studying the behavior of claimants around the disregard level.

term on the left-hand side is the current marginal utility of consumption due to one extra dollar of earnings, which is taxed at the marginal benefit-reduction rate $T'(z)$. The second term is the marginal value of an increase in future UB capital due to one extra dollar of earnings. It is scaled by the discount factor β and the survival rate p.

From the envelope theorem - used at every future period -, it is possible to compute the marginal value of UB capital. Computation details are reported in Appendix C.1. The second term of Equation (4) then simplifies to:

$$\beta p T'(z_t) U'(B_{t+1}; n_i) = T'(z_t) \beta^{\bar{t}-t-1} p^{\bar{t}-t-1} u_c(c_{\bar{t}-1}, z_{\bar{t}-1}; n_i). \tag{5}$$

Using Equation (5) and the parametrization of the utility in Equation (1), Equation (4) simplifies to:

$$1 - T'(z_t)\tau_t = \left(\frac{z_t}{n_i}\right)^{1/e} \tag{6}$$

where τ_t is the wedge between the static marginal tax rate $T'(z_t)$ and the dynamic marginal tax rate $\tau_t T'(z_t)$:

$$\tau_t = 1 - \beta^{\bar{t}-t-1} p^{\bar{t}-t-1}. \tag{7}$$

Note that, if there was no benefit reduction at all, all individuals would supply $z_t = n$. The ability n_i of individual i can thus be interpreted as her potential earnings in low-wage jobs. The actual partial-UI schedule features a kink at the disregard level: the marginal benefit-reduction rate jumps from 0% to 100% (see Equation 2). Such a kink implies that some claimants bunch at the disregard amount.

To describe the bunching behavior, I define a first threshold at ability $n^* = z^*$, i.e. the disregard level. The FOC implies that all individuals with $n < n^*$ earn $z_t = n$, as $T'(z_t) = 0$ below z^*. I define another threshold of ability $n^* + \delta n$, such that all individuals with ability strictly above $n^* + \delta n$ earn strictly more than the disregard z^*. Such individuals have their current benefits reduced and their earnings in low-wage jobs are $z_t = n(1-\tau_t)^e$, as $T'(z_t) = 1$. Using the FOC, the upper threshold then verifies:

$$z^* = (n^* + \delta n)(1 - \tau_t)^e. \tag{8}$$

Equation (8) illustrates that the upper threshold depends on the dynamic marginal tax rate and consequently on its determinants, such as the time period. We highlight this dependence by denoting the ability gap between first and last buncher as $\delta n(t)$. More fundamentally, the dynamic marginal tax rate depends on the time to exhaustion, which is endogenous. As job seekers with an ability just above the upper threshold preserve a small amount of benefits and thus delay their exhaustion date by only one period, their benefit exhaustion date $\bar{t}$ is equal to their potential benefit duration plus one.[34] Finally, all individuals with $n \in (n^*, n^* + \delta n(t))$, earn exactly the disregard amount $z_t = z^*$: they bunch at the kink point of the schedule.

To summarize, the earnings density function $g_t(z)$ at period t verifies:[35]

$$g_t(z) = \begin{cases} f(z) & \text{if } z < z^* \\ \int_{n^*}^{n^*+\delta n(t)} f(n)dn & \text{if } z = z^* \\ f\left(\frac{z}{(1-\tau_t)^e}\right)\frac{1}{(1-\tau_t)^e} & \text{if } z > z^* \end{cases} \tag{9}$$

where $f(n)$ is the ability density of claimants, assumed smoothly distributed.

At this stage, I review in formal terms the discussion in the previous section 4.2 on claimants with earnings above the disregard level. If claimants are myopic and do not link taxes and benefits ($\tau_t = 1$), Equation (6) simplifies to the usual static condition $1 - T'(z_t) = \left(\frac{z_t}{n_i}\right)^{1/e}$. This generates corner solutions when $T'(z_t) = 1$. As the partial-UI schedule features 100% marginal tax rate above the disregard and 0% below, all claimant with ability above z^* bunch at the disregard level. There are no claimants with earnings above the disregard level. On the contrary, when forward-looking claimants link taxes and benefits, there are claimants above the disregard level and the bunching mass is related to the effective tax wedge. I clarify this relation in the next section.

5.3 What bunching identifies

Suppose that the earnings distribution $g_t(z)$ is identified in the data. This yields the bunching mass at the disregard level $g_t(z^*)$ and the left limit of the earnings density at the disregard level $g_t^-(z^*)$. The ratio of these two quantities corresponds to the excess bunching at period t, denoted $\mathcal{B}_t$, which is equal to:

$$\mathcal{B}_t \equiv \frac{g_t(z^*)}{g_t^-(z^*)} = \frac{1}{f(n^*)}\int_{n^*}^{n^*+\delta n(t)} f(n)dn \simeq \delta n(t) \tag{10}$$

[34]From a theoretical point of view, there could be other bunching masses at the earnings levels where the theoretical exhaustion date increases by one period. Because the corresponding changes in the dynamic marginal rate are small, especially at the beginning of the spell, I expect the resulting bunching to be small as well. Indeed, I find none in the data and thus abstract from those further kinks.

[35]$g_t(z)$ is a density with respect to $\lambda + \delta(z^*)$ where λ is the Lebesgue measure and $\delta()$ is the Dirac measure.

where the first equality is obtained thanks to Equation (9) and the second equality uses an approximation of the integral of a continuous function. The excess bunching thus identifies the difference in ability between the first job-seeker bunching from below and the last job-seeker bunching from above: $\delta n(t)$.

Using Equation (10) and the definition of the lower ability n^*, a first-order approximation of Equation (8) yields the following expression for the earnings elasticity:[36]

$$e = \frac{\mathcal{B}_t}{z^* \tau_t}. \tag{11}$$

The main difference between the static bunching formula of Saez (2010) and the above expression is the definition of the marginal tax rate. When taxes above the kink are linked to differed benefits, the bunching formula holds with *effective* marginal tax rates. In my setting, the effective dynamic marginal tax rate depends on the discount factor and the probability to exhaust the initial benefit entitlement. This shows that provided that the discount factor and the probability to exhaust the initial benefit entitlement are identified, the bunching formula allows to identify the earnings elasticity.

Last, I aggregate bunching over time. Let me define $\mathcal{B} = \frac{1}{f(z^*)} \int_t \int_{n^*}^{n^*+\delta n(t)} f(n) dn dG(t) = \int_t \mathcal{B}_t dG(t)$ where $G(t)$ is the cumulative distribution of time spent claiming. Using Equations (10) and (11), I obtain the aggregate bunching formula:

$$e = \frac{\mathcal{B}}{z^* \int_t \tau_t dG(t)} \tag{12}$$

where $\int_t \tau_t dG(t)$ is the marginal tax rate that new claimants expect.

5.4 Heterogeneity in bunching

The bunching formula (11) implies that bunching decreases as the probability to find a permanent job and to leave the UI rolls decreases. Intuitively, forward-looking job-seekers with a higher propensity to keep claiming have higher expected returns to partial UI: they are more likely to profit from benefit transfers later in the claim.[37] I then have the following comparative statics result:

[36] Assuming $\delta n << z^*$, I obtain $e = -\frac{\mathcal{B}_t}{z^* \ln(1-\tau_t)}$. Assuming $\tau_t << 1$, I obtain the formula in the main text. I check below that the estimation results are robust when I do not assume $\tau_t << 1$. Indeed, I estimate below that on average $\tau_t = 0.6$.

[37] Formally, this holds for forward-looking claimants with $\beta > 0$.

Proposition 1 (Bunching across p-strata)**.** *At any given period t, excess bunching decreases with the probability to remain claiming p.*

Proof: The dynamic marginal tax rate decreases with p: $\frac{d\tau_t(p)}{dp} = -(\bar{t}-t) \times (\beta p)^{\bar{t}-t}/p < 0$.

The bunching formula (11) also implies that workers with longer potential benefit duration are more likely to bunch. This yields the following Proposition:

Proposition 2 (Bunching effect of potential benefit duration)**.** *Bunching increases with potential benefit duration.*

Proof: Bunching depends on the potential benefit duration $\bar{t}^{Utot}$ through the dynamic marginal tax rate. For bunchers, $\tau_t = 1 - \beta^{\bar{t}^{Utot}-t} p^{\bar{t}^{Utot}-t}$, which increases with $\bar{t}^{Utot}$: $\frac{d\tau_t}{d\bar{t}^{Utot}} = -\log(\beta p) \times (\beta p)^{\bar{t}^{Utot}-t} > 0$.

The bunching heterogeneity, highlighted in the previous propositions, is a direct consequence of the tax-benefit linkage. If workers do not internalize benefit preservation (no tax-benefit linkage), the tax rate in the bunching formula does not depend on survival probability, nor on potential benefit duration. I test below bunching heterogeneity in the U.S. data.

Before that, I discuss the robustness of the bunching heterogeneity test to two important assumptions. First, Propositions 1 and 2 do not rely on workers having rational expectations. Whatever the process generating the workers' expectations about their job finding rate, Proposition 1 holds. Choosing the relevant measures of job finding expectation is rather an empirical issue that we discuss below. Regarding Proposition 2, the *sign* of the effect of potential benefit duration (PBD) on bunching does not depend on the expected job finding rate p. Testing the null hypothesis of no relation between PBD and bunching is thus a robust test of the tax-benefit linkage.

Second, Propositions 1 and 2 are robust to allowing for stepping-stone/crowding-out effects. The intuition for this result is that the career concerns above and below the earnings disregard have similar importance. As long as the expected gains (or costs) of working in low-earnings jobs on future career are a smooth continuous function of earnings at the disregard level, career concerns do not contribute to bunching. The detailed proof is in Appendix C.3.

6 Bunching heterogeneity in the data

In this section, I implement direct empirical tests of the tax-benefit linkage. I test for heterogeneity in bunching along the lines of Propositions 1 and 2 above. I estimate excess bunching, i.e. the quantity $\mathcal{B}$ in Equation (10), following the procedure in Chetty et al. (2011a), as in Section 4. To maximize statistical power, I jointly analyze US states that share the same disregard level ($0.5 \times WBA$): Idaho and Louisiana. I test bunching heterogeneity by potential benefit duration first, and then by expected survival rate (expected time before finding permanent job).

6.1 Potential benefit duration

According to Proposition 2, claimants with longer potential benefit duration (PBD) bunch more. First, I report the correlation between bunching and PBD, within the cross-section. Second, I exploit within-worker variation in PBD, namely across multiple claiming spells. Third, I leverage policy shocks on PBD arising from emergency triggers.

First, Figure 6 plots bunching estimates by groups of initial potential benefit duration in Tier I (that is before any triggered extensions). Here, I analyze Idaho and Louisiana separately to account for their different PBD distributions. They have different rules converting past work history into PBD. Whatever the state, bunching is significantly greater when claimants have longer potential benefit durations. In Idaho (left-hand panel), excess bunching for claimants with PBD below 17 weeks amounts to around four times the mass of claimants who would have worked at the disregard level, had the kink disappeared. For claimants with 26 weeks of PBD, excess bunching reaches six. I compute the slope of the relation between bunching and PBD for Idaho: it is statistically significant with a p-value of 0.01. Similarly, excess bunching is statistically different across the two PBD groups in Louisiana with a p-value of 0.02. Of course, this comparison may be confounded by other factors correlated with potential benefit duration. For example, it is well-established that longer potential benefit durations cause higher survival rates (see early contributions in Katz and Meyer (1990a) and Lalive et al. (2006), and the review in Schmieder and von Wachter (2016)). Higher survival rates tend to decrease bunching (see Proposition 1), so that Figure 6 likely underestimates the positive relation between bunching and potential benefit duration. Another example of confounder - with less obvious direction of bias - is heterogeneous earnings elasticity across workers.

To control for worker heterogeneity in earnings elasticity and other time-invariant factors, I restrict the sample to workers with two claiming spells. I compute the difference in PBD between the first and second spell, and split the sample between workers experiencing a drop in PBD of more than 3 weeks, workers with small changes in PBD (between a 2-week drop or 2-week increase), and workers experiencing an increase in PBD of more than 3 weeks. Figure 7 plots the across-spell change in bunching for these three groups merging data from Idaho and Louisiana. The change in bunching is positively correlated with the change in PBD. Going from one PBD-change group to the next increases bunching change by 1.5 with bootstrapped standard errors equal to 0.26 (p-value less than 0.01). I further scale the bunching change by actual PBD change across group: one week increase in PBD leads to a 0.15 increase in excess bunching with standard errors of 0.027. This is also statistically significant at the 1% level. This provides further evidence consistent with Proposition 2.

The third strategy leverages state-wide increases in potential benefit duration triggered by emergency plans. These policy shocks are automatically triggered when the state-level unemployment rate reaches certain activation levels. I compute bunching quarter by quarter separately for Idaho and Louisiana together with their respective average potential benefit duration (PBD). I then regress bunching on PBD using extension emergency programs as an instrument for PBD. More precisely, the instrument T_{sq} is the number of triggered extension programs in state s in quarter q. It is equal to one if either EB or FSC is available in the quarter, and to two if both are available. My main regression specification writes:

$$\mathcal{B}_{sq} = \alpha + \beta PBD_{sq} + \delta_s + \gamma_q + \epsilon_{sq} \tag{13}$$

where $\mathcal{B}_{sq}$ is excess bunching in state s quarter q, PBD_{sq} is the average potential benefit duration of claimants in state s and quarter q (assuming they exhaust all benefits in each tier, i.e. standard Tier I and supplemental programs Tier II to IV), δ_s are state fixed effects and γ_q are quarter fixed effects. Controlling for quarter and state fixed effects identifies β as in a difference-in-difference design. As $\mathcal{B}_{sq}$ are estimated quantities, I bootstrap the whole estimation process - both bunching estimation and regression (13) - to compute the standard errors for the coefficient of interest β. Table 3 reports the effect of potential benefit duration on excess bunching, i.e. the coefficient β. It is statistically significant and robust across columns. According to Column (2) which corresponds to my main specification above (Equation 13), a 10-week increase in PBD yields 0.7 excess bunching. This amounts to around 14% of average bunching reported in Table 2. In Column (3), I replace the quarter fixed effects of regression (13) by the state quarterly unemployment rates (U_{sq}). Identification of β then relies on discontinuities in the rule between extended potential benefit duration and local unemployment rate. The point estimate is smaller than in Columns (1) and (2), but still statistically significant. Overall, this last piece of empirical evidence combines both discontinuities in benefit-withdrawal schedule and policy shocks for credible identification. It shows that workers supply labor accounting for the expected value of preserved benefits.

6.2 Survival expectations

I further provide an empirical test of Proposition 1 that leverages survey answers of claimants to direct questions on expectations. In the survey available for a subsample of the CWBH dataset, workers answer whether they expect to be recalled to their previous employers. Katz and Meyer (1990b) find that job-seekers who expect to be recalled have shorter unemployment duration, i.e. they have a lower probability to remain claimants. Figure 8 shows that, consistent with Proposition 1, claimants expecting to be recalled bunch significantly more. The bunching mass at the disregard level is 50% larger. Note that, at this stage, I do not assume that workers have rational expectations on their job finding rate, as I rely on workers' own answers to expectations questions in a survey.[38]

Overall, bunching heterogeneity in the data provides strong support for the tax-benefit linkage. In the next section, I perform a quantitative exercise that relies more extensively on the theoretical model. Namely, I test whether model-based effective tax rates allow to estimate reasonable earnings elasticity to the net-of-tax rate.

7 Estimates of effective tax rate and earned income elasticity to the net-of-tax rate

In this section, I estimate the model-based effective marginal tax rate under the assumption of rational expectations. Together with the bunching estimates, this identifies the earning elasticity to the net-of-tax rate (as shown in Equation 11). I compare the elasticity estimates to those found in the literature. This comparison tests whether the data support both the tax-benefit linkage and the rational expectations assumption. It complements the direct bunching-heterogeneity test of the tax-benefit linkage in the previous section.

[38]Appendix Figure F1 provides further evidence that supports Proposition 1. Under the assumption of rational expectations for claimants' job finding rates, I predict for each worker her expected survival rate (see next section for more details). I then compare bunching across the quartiles of the predicted survival rate distribution. Overall, bunching tends to decrease from the first to the fourth quartile, confirming Proposition 1. The relation between bunching and predicted survival rate is significantly negative: the test of zero slope is rejected with p-value 0.03.

7.1 Dynamic effective marginal tax rate

The theoretical model shows that claimants working just above the disregard in period t react to a dynamic effective marginal tax rate τ_t defined as:

$$\tau_t = 1 - (\beta p)^{\bar{t}^{Utot} - t - 1}.$$

The different components of the effective tax rate are pinned down as follows. I first calibrate the weekly discount factor β to 0.9996, corresponding to an annual discount rate of 4%. Second, I compute for each individual the potential benefit duration under total unemployment: $\bar{t}^{Utot}$. Third, I compute the expected survival rate p taking into account observed individual heterogeneity. More precisely I estimate a proportional hazard model of exiting the UI registers $h_i = h_0 \exp(\beta X_i)$. I include in the proportional hazard model various characteristics of claimants (gender, age, education and ethnicity), claim characteristics (WBA, PBD, recall expectations) and year fixed effects. The model is estimated on the sample of totally unemployed claimants (i.e. without any benefit withdrawal while on claim). Detailed estimation results are reported in Appendix D. The weekly hazard rates vary between 3% and 4% across states (reported in Table 2). I then predict hazard rates at each date, and compute the expected probability of UB exhaustion taking into account the remaining number of entitlement weeks: $(p)^{\bar{t}^{Utot} - t - 1}$. By using predicted rates, claimants are assumed to have rational expectations about their compensated unemployment duration.

Table 2 reports the estimates of the average dynamic marginal tax rate, over all individuals and weeks. The average estimate is around 55% in Idaho and Louisiana; it is larger in New Mexico, where it amounts to 60%. It is far below the 100% static benefit-withdrawal marginal tax rate above the disregard level.

7.2 Earned income elasticity to the net-of-tax rate

I use the identification relation (11) to compute the earned income elasticities to the net-of-tax rate. The standard errors of the elasticity estimates are obtained by the delta method. All components of Equation (11) - bunching, and tax wedge - are already quantified in Table 2. I obtain statistically significant elasticities in Idaho and Louisiana, respectively 0.19 and 0.13. The elasticity in New Mexico has a similar magnitude (0.1), but it is not statistically significant. Elasticity estimates remain between 0.1 and 0.2, when I vary the bunching estimation window and the polynomial degree in the estimation procedure.[39] When I do

[39]Robustness results are reported in Appendix Table F1.

not use first-order approximation of the logarithm of the net-of-tax rate, elasticity estimates are slightly lower, but still around 0.1.[40]

As a last robustness check, I discuss how the introduction of risk-aversion in the model affects the elasticity estimates derived from bunching. The detail of the bunching formula with risk-aversion is reported in Appendix Section C.2. Intuitively, risk-averse job-seekers have an extra incentive of transferring benefits to future periods: they insure themselves against future consumption drops. Abstracting from this extra-incentive leads to over-estimate the dynamic marginal tax rate and thus to under-estimate the earnings elasticity e. In Appendix C.2, I obtain a first-order approximation of the ratio of the elasticity estimates with or without risk-aversion: $1 - \sigma\frac{\Delta c}{c_t}\frac{1}{\tau_t}$ where σ is the coefficient of relative risk-aversion and $\Delta c / c_t$ is the relative change in consumption between the current period and the last week of claim before exhaustion. Gruber (1997a) finds that UI claimants experience a 10% drop in consumption when they become unemployed and a further 12% when the replacement rate of UI benefits goes to zero. Then I consider that consumption drops by around 12% between the current period and the last week of claim before exhaustion. Choosing σ between 1 and 2 (Chetty, 2006; Hendren, 2017) and τ_t around 0.5, the ratio of elasticity estimates is between 1.24 and 1.48. The average elasticity in Table 2 is around 0.15, so that the elasticity taking into account risk-aversion would be between 0.186 and 0.222.

My elasticity estimates - between 0.1 and 0.2 - are broadly consistent with the estimates of the intensive labor supply elasticity found in previous micro empirical work (see the review of quasi-experimental estimates in Chetty (2012) or Chetty et al. (2011b)). Closer to my context, Gelber et al. (2019) find bunching for both wage-earners and self-employed individuals at the kinks of the Social Security Annual Earnings Test (for workers over the national retirement age). Their estimate of the average earnings elasticity, not taking into account adjustment cost, is 0.23, which is in the range of my estimates.

Overall, the elasticity-comparison test supports tax-benefit linkage. The observed bunching patterns are quantitatively consistent with rational expectation claimants with standard earnings elasticity supplying labor according to an effective marginal tax rate that fully internalizes the benefit preservation mechanism.

I see the weekly frequency of benefit withdrawal in the US has an important contextual feature leading to *salient* taxation. The short horizon - within a year - over which benefits are rolled over is also likely to contribute to the *salience* of the benefits tied to taxation. All in all, the salience of both taxation and benefit entitlement may explain why partial UI claimants fully internalize the tax-benefit linkage in the US. Further research could apply

[40]Robustness results are reported in Appendix Table F2.

my approach to other contexts to test the extent to which salience contributes to the tax-benefit linkage.

8 Conclusion

This paper provides empirical evidence that workers are willing to pay taxes today in order to become entitled to benefits tomorrow. In the US context, unemployment insurance claimants accept to work more intensively in low-earning jobs and see their current benefits withdrawn, as they value benefit preservation for future payments. The tax-benefit linkage is identified through the consequences of kinks and plateau in the withdrawal schedule and their interactions with policy shocks. To the best of my knowledge, this is the first direct quasi-experimental evidence of tax-benefit linkage.

Tax-benefit linkage has important consequences for the design of optimal benefit taxation, and for the capacity of payroll tax changes to identify labor supply elasticity. This paper provides a first framework to identify earnings elasticity from bunching in the presence of benefit preservation rules. My approach could be directly applied to study partial-UI schedules in other OECD countries (e.g. Germany features kinks in the partial-UI schedule) or to analyze any social insurance with benefit transfers across periods, such as old-age pensions. From a broader perspective, gathering evidence on these dynamic programs would enhance our understanding of intertemporal decisions under uncertainty.

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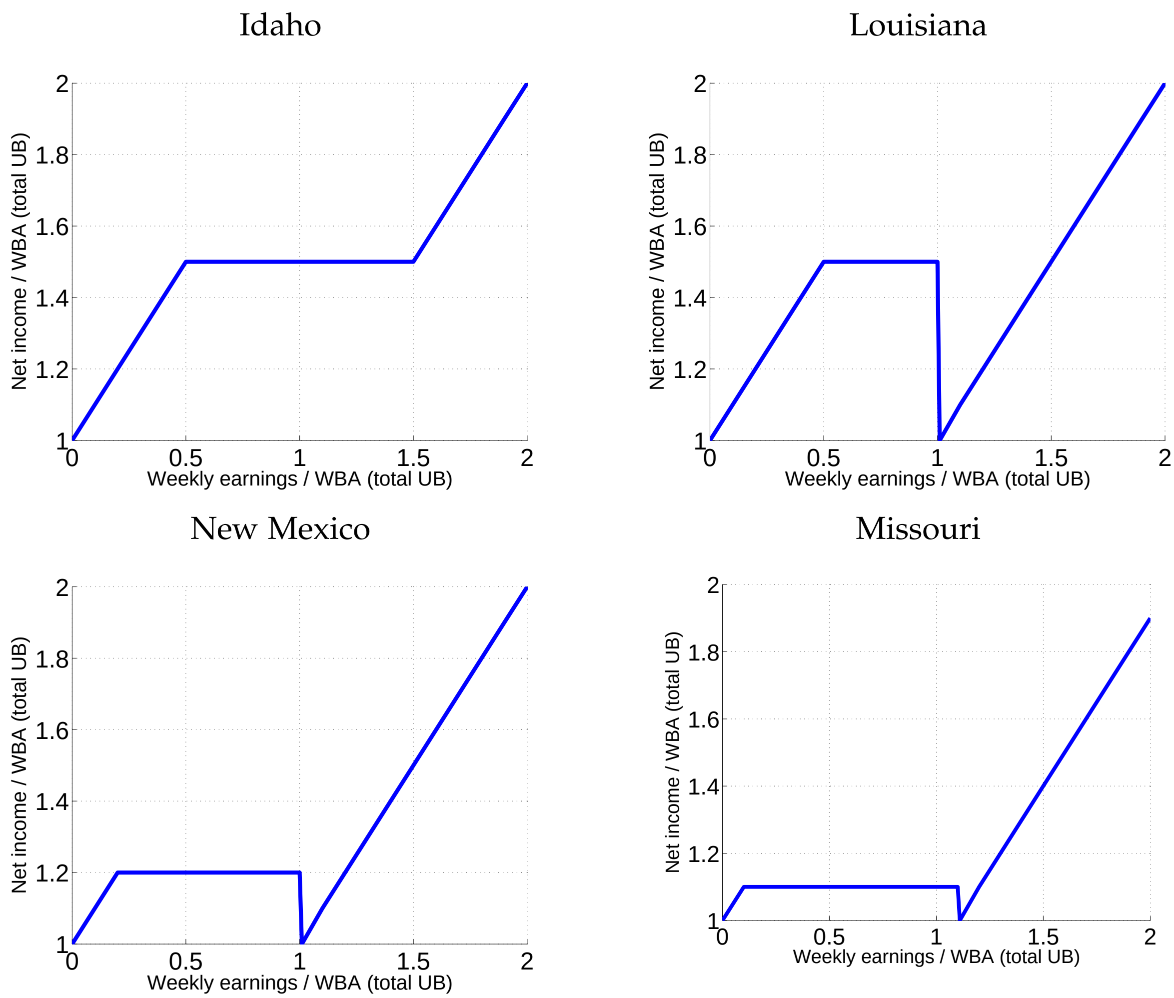
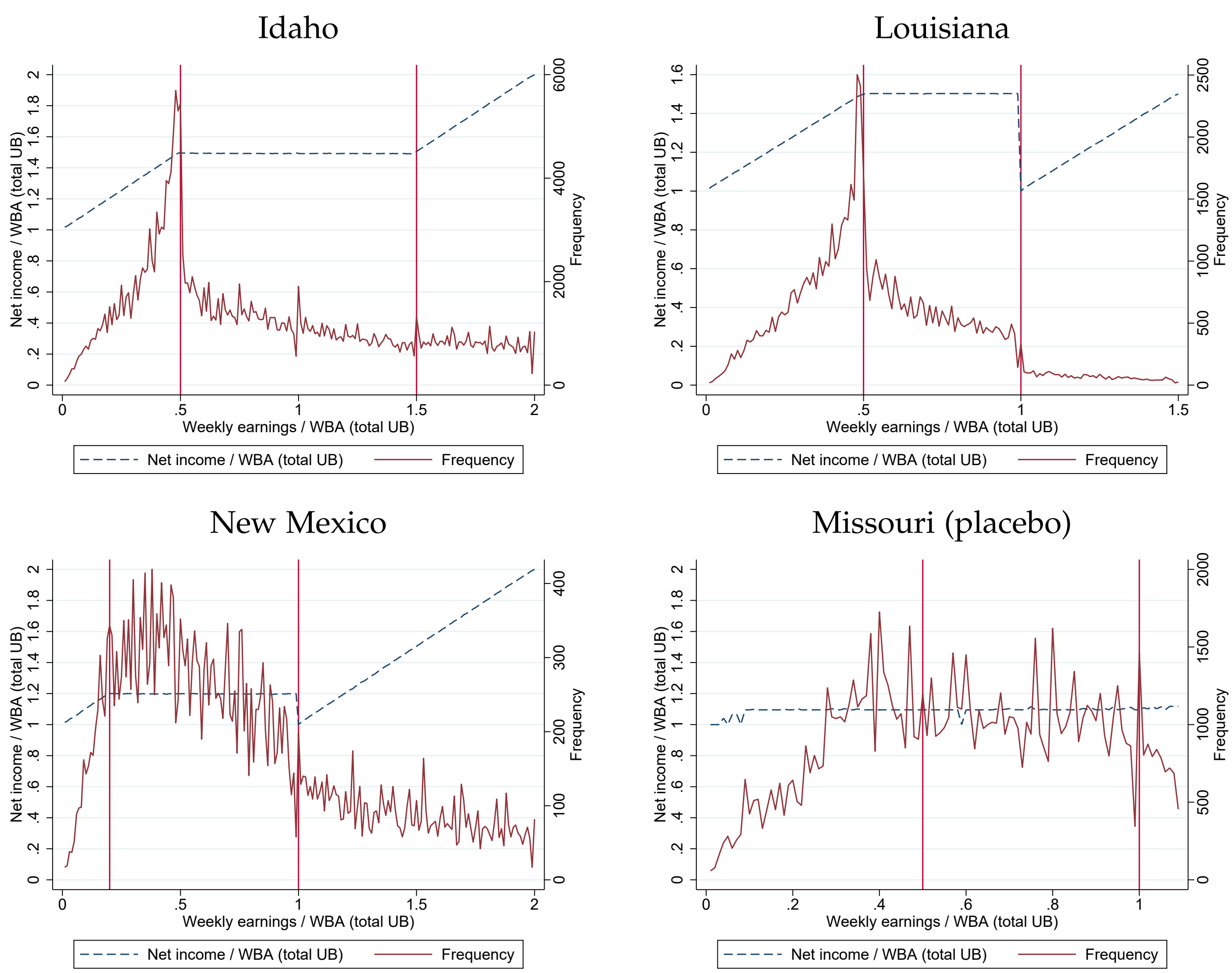
Figures

Figure 1: Partial-UI schedules from 1976 to 1984



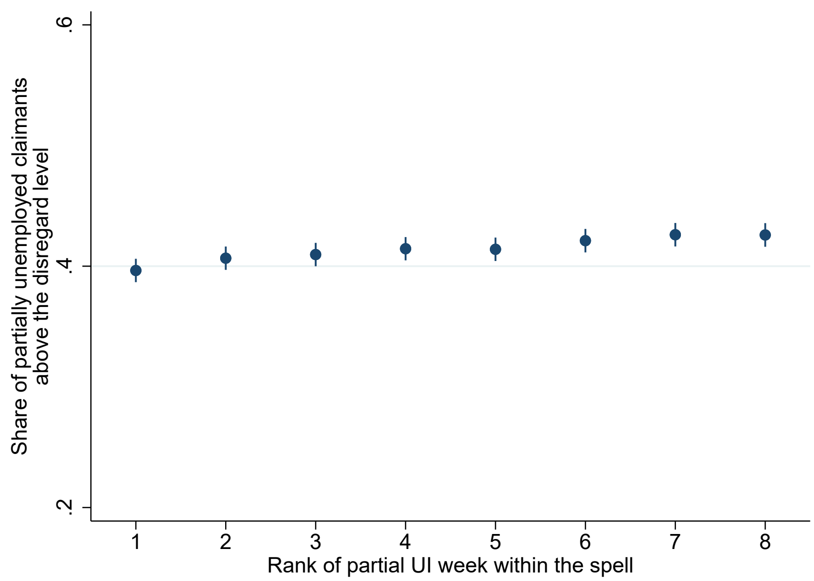
Source: U.S. Department of Labor, "Significant Provisions of State Unemployment Insurance Laws." Notes: This figure plots the theoretical schedules of partial unemployment insurance for the four U.S. states in the CWBH dataset. Each panel yields the net income (weekly labor earnings + unemployment benefit payments) of partially unemployed claimants as a function of their weekly labor earnings. Both net income and earnings are divided by the weekly benefit amount (UB paid in case of total unemployment). For Louisiana, I plot the schedule before April 1983. For Missouri, I consider a claimant whose WBA is $100.

Figure 2: Weekly earnings density and empirical schedule of partial UI.



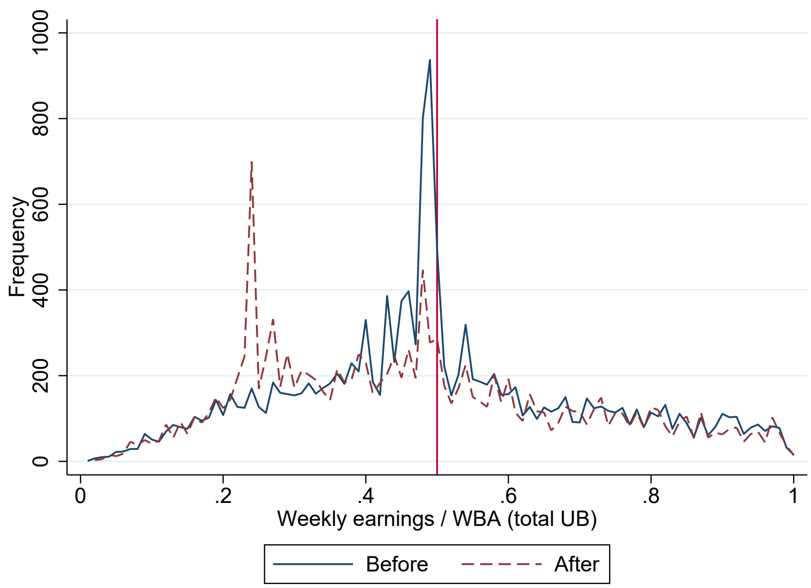
Source: CWBH. Notes: This figure plots - in solid red line - the distributions of weekly labor earnings of partially unemployed claimants divided by the weekly benefit amount (UB paid in case of total unemployment). Corresponding frequencies are on the right-hand Y-axis. Each panel corresponds to one of the four U.S. states in the CWBH dataset. The figure also plots - blue dashed line - the empirical schedule of partial unemployment insurance: net income (left-hand Y-axis) as a function of labor earnings. Red vertical lines show the kinks and notches of the partial UI schedule, except for Missouri (the placebo state). For Louisiana, I use data before April 1983.

Figure 3: Time evolution of the share of partially unemployed claimants with earnings above the disregard level



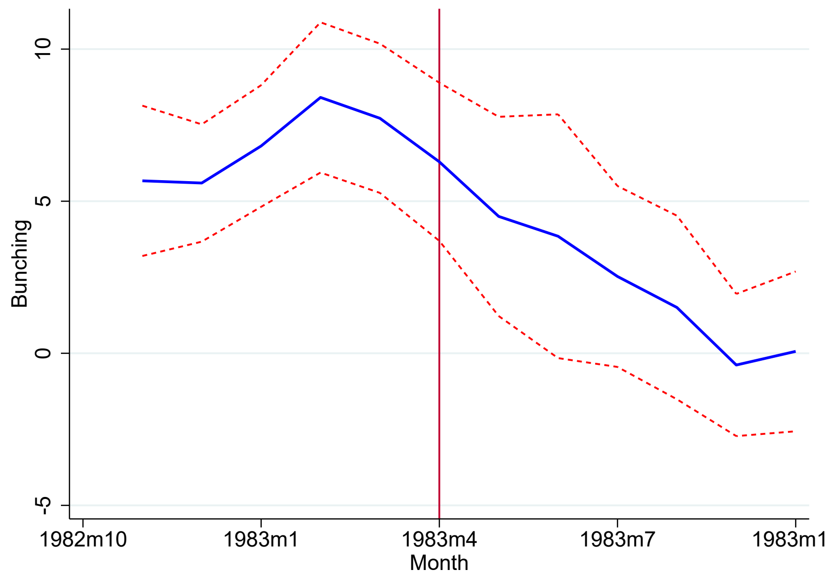
Source: for Idaho 1976-84 and Louisiana 1979-83Q1. Notes: I regress a dummy for having earnings above the disregard level on a set of dummies for the rank of the partial-UI week. The figure plots the coefficients of the rank dummies together with their 95% confidence interval as vertical lines. The sample is restricted to claimants with at least eight weeks under partial UI within the benefit year, and includes their partial UI weeks up to the 8^{th} week.

Figure 4: Weekly earnings density before and after the April 1983 reform in Louisiana



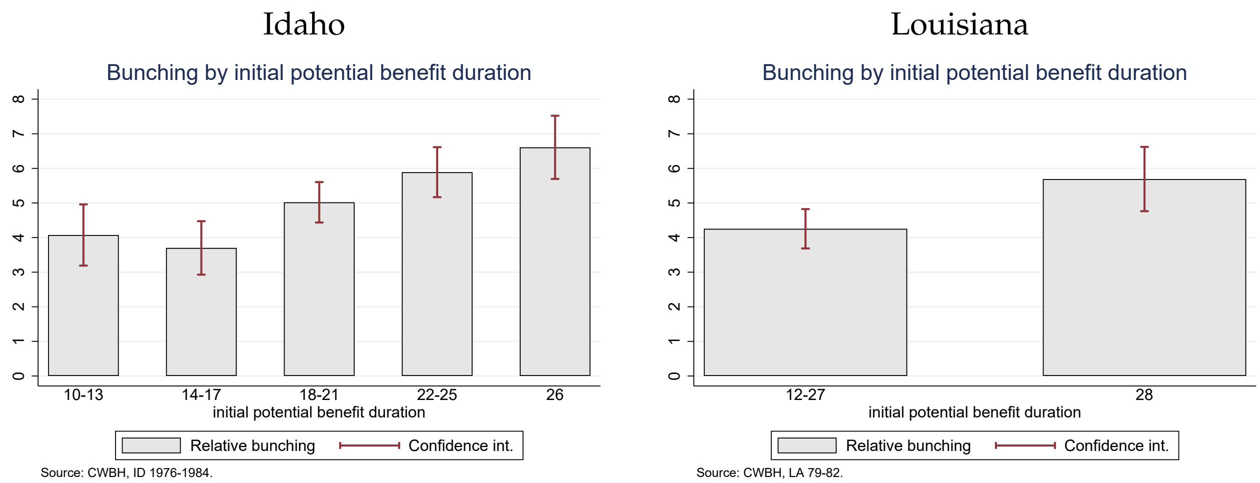
Source: CWBH, Louisiana October 1982 to September 1983. Notes: the figure plots the density of weekly labor earnings over the weekly benefit amount (WBA) under total unemployment, for partial UI claimants before the April 1983 reform (solid line in blue) and after the reform (dashed line in red). The sample is restricted to workers with weekly benefit amount (WBA) greater than 100$, for whom the reform changes the disregard level. The vertical line corresponds to the earnings disregard (kink) before the reform: 50% of WBA. After the reform, the disregard level amounts to 50$ for all workers, so that when expressed as a fraction of WBA it varies and bunching is less salient.

Figure 5: Bunching before and after the April 1983 reform in Louisiana



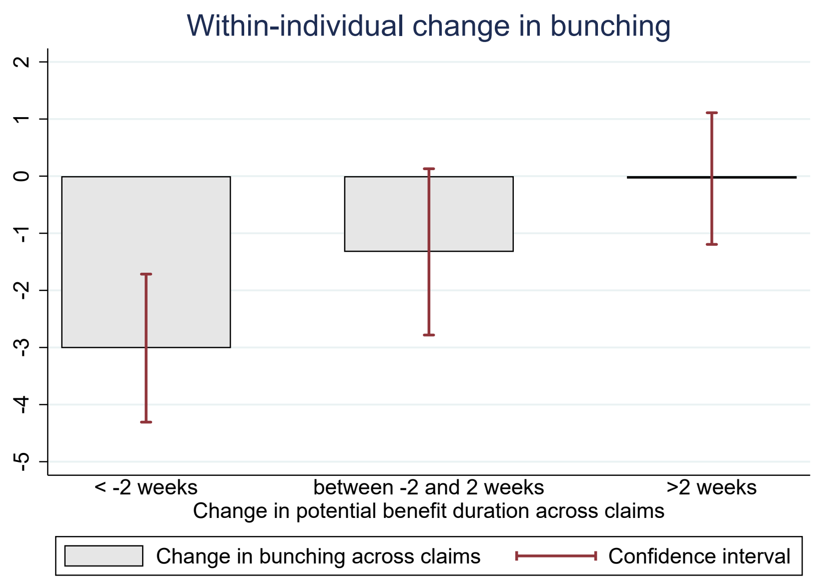
Source: CWBH, Louisiana October 1982 to September 1983. Notes: the figure plots the monthly bunching estimates at the earnings disregard (kink) before the reform (50% of WBA). The sample is restricted to workers with weekly benefit amount (WBA) greater than 100$, for whom the reform changes the disregard level. 95% confidence interval in dashed lines.

Figure 6: *Bunching by initial potential benefit duration*



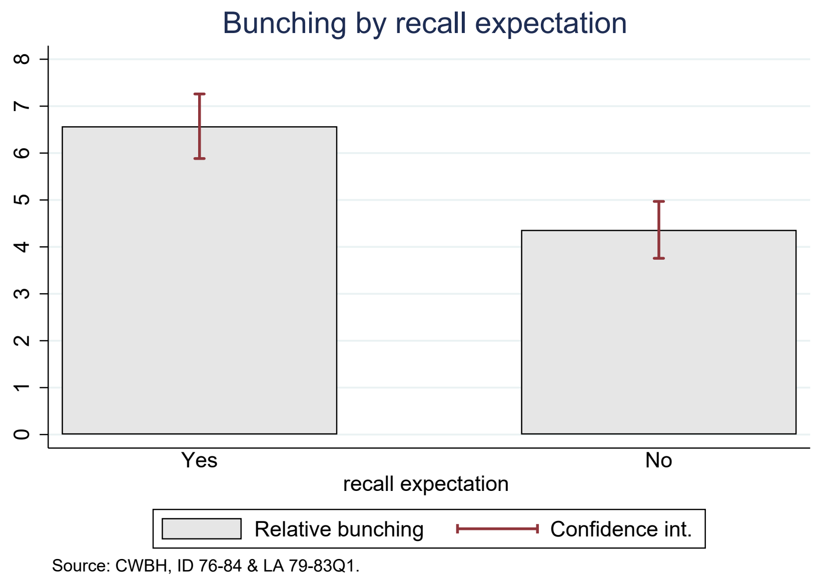
Source: CWBH for Idaho 1976-84 and Louisiana 1979-83Q1. Notes: The figures plot excess mass at the disregard amount (bunching at the kink) by potential benefit duration at the beginning of the claim (in weeks). Confidence interval at the 95% level in red.

Figure 7: *Change in bunching by change in potential benefit duration across claiming spells: within-worker design*



Source: CWBH for Idaho 1976-84 and Louisiana 1979-1984. Notes: The figure plots the across-spell change in excess mass at the disregard amount (bunching at the kink) for claimants experiencing a large negative shock in potential benefit duration across spells (left-hand bar), a small change in PBD (center bar) or a large positive shock in PBD (right-hand bar). Confidence interval at the 95% level in red.

Figure 8: *Bunching by recall expectation*



Source: CWBH for Idaho 1976-84 and Louisiana 1979-83Q1. Notes: The figure plots the excess mass at the disregard amount (bunching at the kink) for claimants expecting to be recalled to their previous employer (left-hand bar) and for claimants not expecting any recalls (right-hand bar). Confidence interval at the 95% level in red.

Tables

Table 1: Descriptive Statistics

	Idaho	Louisiana	New Mexico	Missouri
Male	.634	.703	.607	.504
Age	31.6	34.9	34.0	36.2
Education (years)	11.7	11.3	11.7	10.9
White	.946	.625	.460	.884
Pre-U Industries				
Construction	.146	.270	.124	.157
Manufacturing	.386	.261	.270	.397
Trade	.218	.115	.231	.142
Services	.122	.143	.196	.202
Pre-U Occupations				
Prof., tech. and managers	.058	.067	.104	.051
Clerical and sales	.149	.132	.196	.153
Structural work	.234	.265	.179	.074
Pre-U firm in private sector	.899	.964	.932	.951
Pre-U weekly wage (current $)	332	327	277	221
Weekly Benefit Amount (current $)	102	137	105	89
Replacement rate	.448	.468	.435	.483
Potential Benefit duration (weeks)	30.5	36.5	38.7	33.4
Actual Benefit duration (weeks)	17.5	19.7	16.7	16.1
Inflow (no)	40,792	25,320	12,163	26,907
Sample Years	76-84	79-84	80-84	78-84

Source: CWBH. Notes: Means are computed over the sample of partial UI claimants. Inflow reports the number of new claims with at least one week of partial unemployment.

Agriculture, Mining, Transportation, Finance, Insurance and Real Estate industries are not reported for the sake of space. Occupations correspond to the standard DOT (Dictionary of Occupation Titles). I only report the most common occupations and exclude service, agricultural, processing, machine trades and benchwork occupations from the table.

Wages and benefits amounts are in current dollars. Replacement rate is the ratio of weekly benefit amount over pre-unemployment wages. Potential Benefit duration is the maximum number of weekly benefit payments the worker is entitled to under total unemployment. It accounts for extended benefits programs, triggered in high-unemployment periods. Actual benefit duration is the number of weeks within the spell with actual benefit payments.

Table 2: Bunching, dynamic marginal tax rates and earnings elasticity estimates.

	Idaho	Louisiana	New Mexico	Missouri *placebo*
Excess bunching mass ($\mathcal{B}$)	5.334***	4.814***	1.247	-.7767**
	(.2417)	(.279)	(.7767)	(.3647)
Hazard rate $(1-p)$	.042	.033	.039	
Effective Marginal Tax Rate (τ)	.538	.576	.606	
Earnings elasticity to net-of-tax rate (e)	.187***	.129***	0.096	
	(.0099)	(.0073)	(.0594)	
Partial-UI weeks (no)	230,535	69,024	31,103	99,451
Sample Years	76-84	79-83	80-84	78-84

Source: CWBH. Notes: This table reports the estimates for excess bunching mass at the disregard level in the distribution of weekly earnings of partial UI claimants, for four US states (in columns). The earnings disregard level is at 50% of weekly benefit amounts (WBA) for Idaho and Louisiana (up to April 1983) and 20% for New Mexico. For Missouri, I consider a placebo earnings disregard at 50% WBA: this amounts to 45 current dollars on average, while the actual disregard level in Missouri is $10. For non-placebo states, the table also reports the average predicted hazard rate out of compensated unemployment $(1-p)$, which is an important building block of the dynamic effective marginal tax rate (τ) (see Equation 7). Quantities in the first three rows allow to estimate the earnings elasticity to the net-of-tax rate in the last row (see bunching relation in Equation 12). Standard errors are in parentheses below estimates. *** p<0.01, ** p<0.05, * p<0.1

Table 3: Bunching and potential benefit duration: IV strategy with UI extension programs (EB and FSC).

	Bunching ($\mathcal{B}$)		
Potential Benefit Duration	.072***	.075**	.036**
	(.013)	(.035)	(.017)
State FE	Y	Y	Y
Quarter FE		Y	
State quarterly unemployment rate			Y
# observations			
state X quarter	46	46	46
weekly claims	322,450	322,450	322,450

Source: CWBH, Idaho and Louisiana 1979-1984. Notes: This table reports the coefficient of potential benefit duration in a regression of excess bunching. I instrument potential benefit duration by the number of benefit extension programs available in state s in quarter t (2SLS estimation). The data of over 300,000 weekly earnings are collapsed at the state X quarter level into 23 observations for each state. Bootstrapped standard errors, accounting for the first-step estimation of excess bunching, are in parentheses below estimates. In Column (1), I control for state fixed effects. In Column (2), I add quarter dummmies from 1979Q1 to 1984Q3. Column (2) is my preferred design (short-period difference-in-difference). In Column (3), I control for state fixed effects, and the quarterly state unemployment rate. I also add dummies for Q1 to Q4 to control for seasonality. The mean outcome is around 5 (see bunching estimates in Table 2). *** p<0.01, ** p<0.05, * p<0.1

Online Appendix

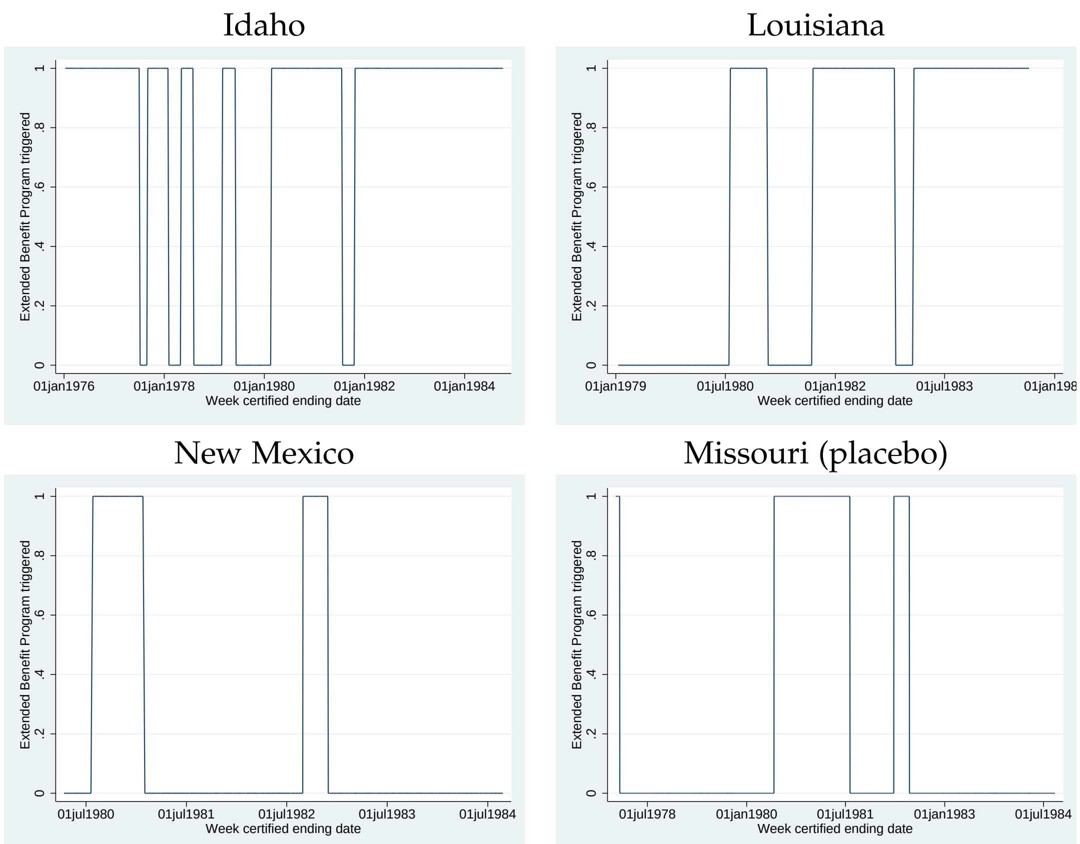
A Institutional background

Between the late 70s and early 80s, the unemployment insurance (UI) rules, in Idaho, Louisiana, New Mexico and Missouri, are as follows. First, UI claimants must meet a monetary eligibility requirement. They must have accumulated a sufficient amount of earnings during a one-year base period before job separation. Second, UI claimants must meet non-monetary eligibility requirements. They must not have quit their previous job, they must not have been fired for misconduct. They must search and be available for work.

When claimants meet the above requirements, states compute their weekly benefit amount (WBA). This would be their weekly unemployment benefit payment when they earn less than the partial UI disregard. The WBA is a fraction (between 1/20 and 1/26) of the high quarter wages (HQW), defined as the wages earned in the quarter of the base period (BP) with the highest earnings. The BP is the first four calendar quarters of the five completed quarters before job separation. The WBA is subject to a maximum and minimum benefit level. As maximum levels are quite low, a large fraction of claimants have their WBA capped. For example, in the first quarter of 1980, the maximum amount was $121 in Idaho. The above rule implies a decreasing gross replacement rate between 50% and 40%. States also compute a potential benefit duration (PBD). This is usually a fraction (between 2/5 and 3/5) of base period wages (BPW), subject to a minimum and maximum number of weeks. The maximum PBD is 26 weeks, except in Louisiana before 1983 where it is 28 weeks. The total entitlement is defined as the product of the WBA and of the PBD. It represents the total amount of unemployment benefits that the claimant can be paid over the benefit year (BY), i.e. the continuous one-year period starting at the first claim. Note that, after the end of the BY, no unemployment benefits can be paid from the corresponding claim, but the claimant can be eligible for a new claim. States observe a waiting period of one week at the beginning of the claim, during which no unemployment benefits are paid.

During periods of high unemployment, the potential duration of unemployment benefits is extended, either by the Federal-state extension benefit (EB) program, or the federal supplemental compensation (FSC) program. Those programs are triggered, when federal or state unemployment are over certain levels. In Figure A1, I plot the EB periods in each state. The EB program extended the initial entitlement period by 50% up to a total of 39 weeks when the state unemployment rate reached a certain trigger. The FSC program, in action from September 1982 to March 1985 in all four states considered, extended the entitlement period of individuals who had exhausted their regular and EB entitlement, by a rate ranging from 50% to 65% up to a maximum of weeks depending on the FSC phase and the U.S. state (see Grossman 1989 for more details on the FSC).

Figure A1: Extended Benefit Program.



Source: Trigger reports. Note: this does not account for Federal Supplemental Program (FSC) extensions.

There was one major change in UI rules in Louisiana in April 1983. The partial-UI disregards have been capped at $50. In addition, the maximal potential duration of usual benefits was reduced from 28 weeks to 26 weeks. Last, Appendix Table A1 summarizes parameters of the partial UI rules detailed in the main text.

Table A1: Partial-UI rules from 1976 to 1984

	Disregard	Maximum earnings
Idaho	$0.5 \times WBA$	$1.5 \times WBA$
Louisiana bef. Apr. 1983	$0.5 \times WBA$	WBA
Louisiana aft. Apr. 1983	$\min(0.5 \times WBA, \$50)$	WBA
New Mexico	$0.2 \times WBA$	WBA
Missouri	\$10	WBA+\$10

Source: U.S. Department of Labor, "Significant Provisions of State Unemployment Insurance Laws.".

Note: the table reports disregard and maximum levels. Taking into account inflation, \$10 (resp. \$50) in 1978 represent around \$37 (resp. \$185) in 2016.

B Excess bunching estimation

In this appendix, I detail how I estimate excess bunching. I follow the procedure of Chetty et al. (2011a). I fit a polynomial on the earnings density of partial-UI claimants, taking into account that there is bunching in a bandwidth around the disregard, and that the bunching mass comes from the earnings distribution above the disregard.

First, the earnings distribution is centered around the disregard amount. Let C_j be the count of individuals earning between j and $j+1$ dollars above the disregard level (when they earn below the disregard, j is negative), and let Z_j be the dollar amount earned by claimants in bin j ($Z_j = j$), centered around the disregard level. I estimate the following equation:

$$C_j \left(1 + \mathbb{1}[j > \overline{R}] \frac{\hat{B}_N}{\sum_{j>\overline{R}} C_j}\right) = \sum_{k=0}^{q} \beta_k (Z_j)^k + \sum_{i=-\underline{R}}^{\overline{R}} \gamma_i \mathbb{1}[Z_j = i] + \epsilon_j \tag{14}$$

where $\hat{B}_N = \sum_{i=-\underline{R}}^{\overline{R}} \hat{\gamma}_i$ is the excess mass taken off the earnings distribution above the disregard.[41] The order of the polynomial q and the width of the bunching window $(-\underline{R}, \overline{R})$ are not estimated, but set after visual inspection. Robustness checks of the estimation results with respect to those two parameters are presented below.

Equation (14) defines the counterfactual distribution (with no benefit reduction): $\hat{C}_j = \sum_{k=0}^{q} \hat{\beta}_k (Z_j)^k$. Then the estimator of excess bunching equals:

$$\hat{\mathcal{B}} = \frac{\hat{B}_N}{\sum_{j=-\underline{R}}^{\bar{R}} \hat{C}_j / (\underline{R} + \overline{R} + 1)}. \tag{15}$$

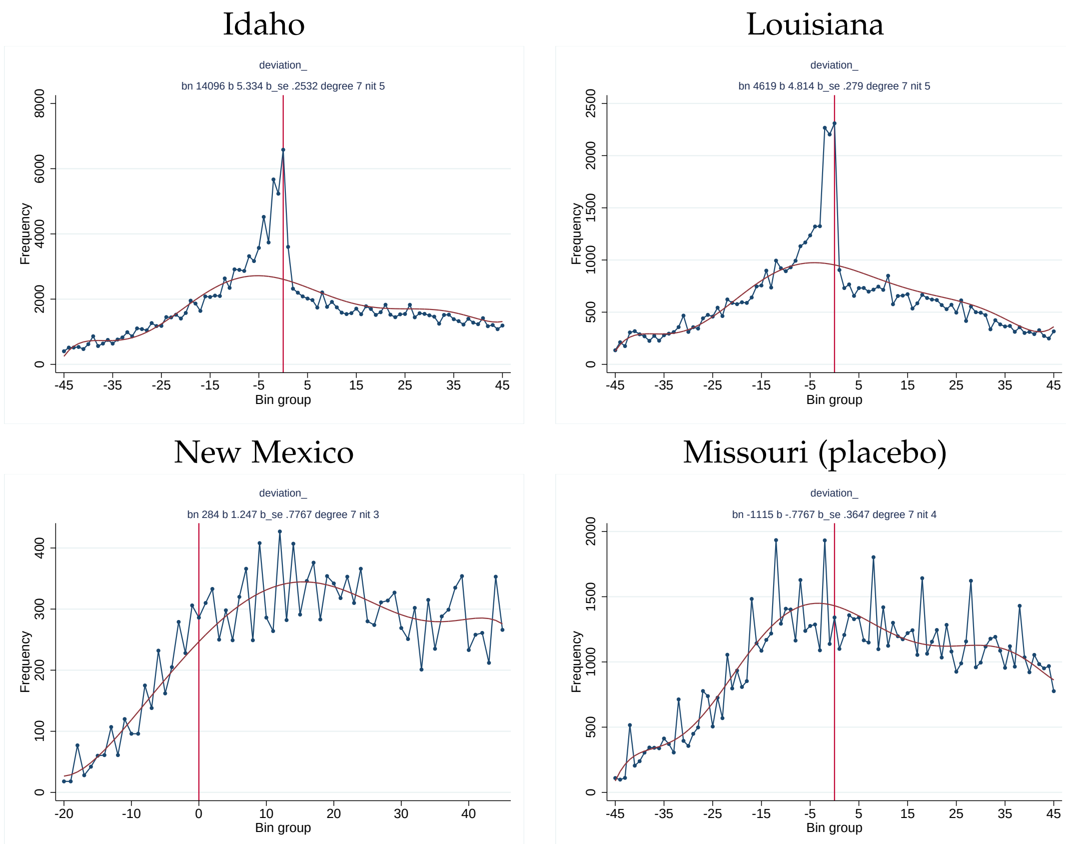
The recursive estimation is bootstrapped to obtain standard errors. The bootstrap procedure draws new error terms (ϵ_j) among the estimated distribution.

Appendix Figure B1 illustrates the estimation procedure for each state. It plots the partial claimants' earnings density in bins of one dollar centered around the disregard level, together with the counterfactual density estimated along the lines of Chetty et al. (2011a). In practice, the procedure fits a polynomial of degree 7. The bandwidth is such that $-\underline{R} = -5$ and $\bar{R} = 2$. Appendix Figure B1 confirms that the counterfactual density compares well to the actual data. Appendix Figure B1 reveals some periodicity in the earnings distribution in Missouri. Claimants report earnings that are multiples of ten dollars. This may bias bunching estimates, especially if there are heaps in the window where bunching is ex-

[41] Because $\hat{B}_N$ depends on $\hat{\gamma}_i$, I follow an iterative procedure to estimate the equation. At each step, $\hat{B}_N$ is computed with past estimates of $\hat{\gamma}$, and the procedure stops when a fixed point is obtained.

pected. I verify that the bunching estimate does not change if I modify the earnings density by smoothing the heaping points.

Figure B1: Centered weekly earnings density of partial-UI claimants.



Source: CWBH. Notes: Earnings are in dollars centered at the disregard level. Empirical earnings density in blue. Counterfactual density in red.

C Theoretical model

In this appendix, I derive in detail the solution of the claimants' program. Then I introduce risk-aversion and stepping-stone/crowding-out effects in the theoretical model and discuss identification in those cases.

C.1 Model Solution

I derive in detail the solution of the claimants' program:

$$U(B_t; n_i) = \max_{c_t, z_t} u(c_t, z_t; n_i) + \beta \left[p(z_t) U(B_{t+1}; n_i) + (1 - p(z_t)) W \right]$$

$$\text{such that}$$

$$\begin{cases} c_t & = z_t + \min(b, B_t) - T(z_t) \\ B_{t+1} & = B_t - \min(b, B_t) + T(z_t) \\ B_{t+1} & \geq 0. \end{cases}$$

By definition of the partial-UI schedule, I have that $T(z_t) \leq b$ when $B_t > b$ and $T(z_t) \leq B_t$ when $B_t < b$. As a consequence, the capital stock B_t depreciates or stays constant over time: $B_{t+1} \leq B_t$. I first discuss the existence of stationary solutions.

A stationary solution U_i for individual i with $B > b$ (resp. $b > B$) satisfies $T(z) = b$ (resp. $T(z) = B$). Then the program simplifies as:

$$U_i = \max_{c,z} u(c, z; n_i) + \beta \left[pU_i + (1 - p) W \right] \text{ such that } c = z.$$

The first order condition writes:

$$u_c(c, z; n_i) + u_z(c, z; n_i) = 0. \tag{16}$$

This determines the level of consumption, while the Bellman equation determines the value of claiming U_i:

$$U_i = \frac{u(c, z; n_i) + \beta(1 - p) W}{1 - \beta p}. \tag{17}$$

The two previous equations (16) and (17) show that the stationary value of unemployment U_i and the corresponding earnings z do not depend on the level of UB capital B.

Recall that, when $B > b$, the typical partial-UI schedule is such that there exists a unique $\bar{z}$ such that for any $z \geq \bar{z}$, $T(z) = b$ and the marginal tax rate is 100% just below $\bar{z}$.

Let me consider the marginal individual who would supply $\overline{z}$, she would benefit from deviating from the stationary path during one period by decreasing her labor supply by δz. Actually, her flow income is not affected, while she enjoys more leisure. A consequence of this manipulation is that her UB capital is depreciated. However her future utility is not affected as the value of stationary unemployment does not depend on UB capital. Then, this deviation necessarily increases her welfare and a stationary equilibrium does not exist for this worker with ability verifying $B > b$. Of course there may exist some very able individuals whose stationary z is well above $\overline{z}$. To rule out the existence of such individuals, it is sufficient to assume that there is a fixed flow cost to claim. Such a cost decreases the relative gain of stationary claiming.

The previous reasoning also applies when $B \in (0, b)$. Recall that, for any $B \in (0, b)$, the typical partial-UI schedule is such that there exists $\overline{z}(B) = B + z^*$ an exit point from partial UI. Let me consider as above the marginal claimant supplying $\overline{z}(B)$. The similar reasoning as above applies: the marginal claimant finds it beneficial to deviate from the stationary path and consume her UB capital. The previous argument does not apply to individuals with $z > \overline{z}(B)$. In the remainder, I implicitly restrict the analysis to individuals with preferences inconsistent with stationarity. An alternative solution could be to introduce a fixed flow cost to claim. This would make the group of job-seekers with ability consistent with stationary claiming arbitrarily small.

While claiming, UB capital is thus strictly decreasing over the spell. I define $\overline{t} < \infty$ the finite exhaustion date (first date when $B_t = 0$). The program becomes stationary only when job-seekers run out of benefits. I denote U_i the value of unemployment of job-seeker i when benefits are exhausted.

Let me now solve the program. When $B_t > b$, it simplifies as:

$$U(B_t; n_i) = \max_{z_t} u(z_t + b - T(z_t), z_t; n_i) + \beta \left[p.U(B_t - b + T(z_t); n_i) + (1-p)W \right].$$

When $B_t \in (0, b)$, it is given by:

$$U(B_t; n_i) = \max_{z_t} u(z_t + B_t - T(z_t), z_t; n_i) + \beta \left[p.U(T(z_t); n_i) + (1-p)W \right].$$

Both sub-programs share the same first order condition:

$$u_c(c_t, z_t; n_i) \left(1 - T'(z_t)\right) + \beta p(z_t) T'(z_t) U'_{t+1}(B_{t+1}; n_i) = -u_z(c_t, z_t; n_i). \tag{18}$$

Using the envelope theorem, I show that the marginal value of UB capital satisfies the following recursive equation:

$$U'(B_t; n_i) = \begin{cases} \beta p U'(B_{t+1}; n_i) & \text{when } b < B_t, \\ u_c(c_t, z_t; n_i) & \text{when } 0 < B_t < b. \end{cases}$$

For simplicity I assume that the individual only claims one period when $0 < B_t < b$. This can be rationalized by introducing a fixed flow cost of claiming. Then this period verifies $t = \bar{t} - 1$. Consequently, the third term of the marginal gain of labor earnings can be written as:

$$\beta p T'(z_t) U'(B_{t+1}; n_i) = T'(z_t) \beta^{\bar{t}-t-1} p^{\bar{t}-t-1} u_c(c_{\bar{t}-1}, z_{\bar{t}-1}; n_i) \tag{19}$$

where $p^{\bar{t}-t-1}$ is the probability to exhaust benefits conditional on claiming at date t.

Using Equation (30) and the utility definition, the FOC in Equation (18) can be simplified. The rest of the derivation is in the main text.

C.2 Risk-aversion

In this section, I consider the behavior of risk-averse job-seekers. I assume that the per-period utility of a claimant writes:

$$u(c, z; n_i) = \frac{c^{1-\sigma}}{1-\sigma} - \frac{n_i}{1+1/e} \left(\frac{z}{n_i} \right)^{1+1/e} \tag{20}$$

where σ is the coefficient of relative risk-aversion. The derivation of the solution path follows the same lines as in the main text and Appendix C.1. We focus on the case when $B_t > b$ as in the main text. Solving the job-seekers' program yields the following FOC:

$$u_c(c_t, z_t; n_i) \left(1 - T'(z_t)\right) + T'(z_t)(\beta p)^{\bar{t}-t-1} u_c(c_{\bar{t}-1}, z_{\bar{t}-1}; n_i) = -u_z(c_t, z_t; n_i). \tag{21}$$

Using the definition of utility of risk-averse job-seekers, the FOC simplifies as:

$$1 - T'(z_t) \left(1 - (\beta p)^{\bar{t}-t-1} \left(\frac{c_t}{c_{\bar{t}-1}} \right)^{\sigma} \right) = \left(\frac{z_t}{n} \right)^{1/e} (c_t)^{\sigma}. \tag{22}$$

Compared to Equation (6), the wedge between the static benefit-reduction rate and the dynamic marginal tax rate depends on the coefficient of risk-aversion and on the ratio of current consumption to consumption in the last week of claim.

While the behavior of risk-averse job-seekers is more complex, bunching still identifies the parameter e *if the coefficient of risk-aversion is separately identified*. The intuition follows. First, consider the individual who bunches from below. Her ability n^* satisfies the following FOC:

$$1 = \left(\frac{z^*}{n^*}\right)^{1/e} (b + z^*)^{\sigma}. \tag{23}$$

Second, consider the individual who bunches from above. Her ability $n^* + \delta n(t)$ satisfies the following FOC:

$$(\beta p)^{\bar{t}-t-1} \left(c_{\bar{t}-1}\right)^{-\sigma} = \left(\frac{z^*}{n^* + \delta n(t)}\right)^{1/e}. \tag{24}$$

As $c_{\bar{t}-1}$ is a function of the ability $n^* + \delta n(t)$ and of the other parameters of the model (β, p, σ, B_t, b, z^* and e), Equations (23) and (24) identify e when excess bunching is observed in the data. More precisely, we obtain the consumption in the last week of claim $c_{\bar{t}-1}$ using the FOC of the program when $B_{\bar{t}-1} < b$:

$$1 = \left(\frac{z_{\bar{t}-1}}{n^* + \delta n(t)}\right)^{1/e} \left(c_{\bar{t}-1}\right)^{\sigma}. \tag{25}$$

The budget constraint in the last week of claim writes: $c_{\bar{t}-1} = B_{\bar{t}-1} + z_{\bar{t}-1}$. Assuming that the remaining UB capital in the last week of claim is negligible, we have $c_{\bar{t}-1} = z_{\bar{t}-1}$. Then Equation (25) shows that $c_{\bar{t}-1}$ only depends on $n^* + \delta n(t)$, e and σ. Replacing the implicit expression of $c_{\bar{t}-1}$ in Equation (24), we obtain that excess bunching identifies e from Equations (23) and (24).

We now quantify the order of magnitude of the bias in the estimate of e when risk-aversion is neglected. We re-write Equations (23) and (24):

$$n^* = z^* \left(c_t\right)^{e\sigma}, \tag{26}$$

$$n^* + \delta n(t) = z^* (1 - \tau_t)^{-e} \left(c_{\bar{t}-1}\right)^{e\sigma}. \tag{27}$$

Taking the difference between these two equations and using first-order approximations, we obtain:

$$\frac{\delta n(t)}{z^*} = e \left(\tau_t + \sigma \frac{\Delta c}{c_t}\right)$$

where $\Delta c = c_{\bar{t}-1} - c_t$. Rearranging terms, we obtain the following identification formula:

$$e = \frac{\delta n(t)}{z^* \left(\tau_t + \sigma \frac{\Delta c}{c_t}\right)}. \tag{28}$$

Taking the ratio of the above expression and Equation (11), we obtain the ratio of elasticity estimates with or without risk-aversion: $1 - \sigma \frac{\Delta c}{c_t} \frac{1}{\tau_t}$. In the main text, I quantify the order of magnitude of this ratio.

C.3 Stepping-stone/crowding-out effects

In this section, I account for stepping-stone and/or crowding-out effects of low-earnings jobs. I assume that the probability to find a permanent job depends on the earnings level in the current low-wage job: $1 - p(z_t)$. The claimants' program then becomes:

$$U(B_t; n_i) = \max_{c_t, z_t} u(c_t, z_t; n_i) + \beta \left[p(z_t) U(B_{t+1}; n_i) + (1 - p(z_t)) W \right]$$

$$\text{such that}$$

$$\begin{cases} c_t & = z_t + \min(b, B_t) - T(z_t) \\ B_{t+1} & = B_t - \min(b, B_t) + T(z_t) \\ B_{t+1} & \geq 0. \end{cases}$$

I can show, as in the previous appendix C.1, that there exists a solution where the UB capital is strictly decreasing up to a finite exhaustion date $\bar{t} < \infty$. Considering the case when $B_t > b$, we obtain the following FOC:

$$\underbrace{u_c(c_t, z_t; n_i) \left(1 - T'(z_t)\right)}_{(I)} + \underbrace{\beta p(z_t) T'(z_t) U'(B_{t+1}; n_i)}_{(II)} - \underbrace{\beta p'(z_t) \left(W - U(B_{t+1}; n_i)\right)}_{(III)}$$
$$= -u_z(c_t, z_t; n_i). \qquad (29)$$

Compared to the FOC in the baseline model (Equation 6), a third term (III) appears on the left-hand side. When working while on claim increases the future probability to find a permanent job - stepping-stone effect ($p' < 0$) -, the job-seeker is induced to work more. She has the opposite reaction when working while on claim crowds out job search for permanent jobs - crowding-out effect ($p' > 0$).

The FOC makes clear that it is only the *marginal* stepping-stone/crowding-out effect that matters: $p'(z)$. I expect *marginal* effects to be smaller than *average* stepping-stone/crowding-out effects that compare outcomes of partial UI claimants to those of total UI claimants who do not work while on claim. The empirical literature - McCall (1996) in the U.S. and Kyyra (2010), Caliendo et al. (2016), Kyyra et al. (2013), Fremigacci and Terracol (2013) and Godoy and Roed (2016) in European countries - provides estimates of *average* effects that are small. It seems then reasonable to neglect stepping-stone/crowding-out effects when studying the intensive margin of labor supply for low-earnings jobs. For the sake of completeness, this appendix further discusses assumptions that are sufficient to obtain the bunching formula when one does not neglect *marginal* stepping-stone/crowding-out effects.

Following the same reasoning as in the previous appendix C.1, I obtain a simplified expression of the second term in Equation (29):

$$\beta p(z_t) T'(z_t) U'_{t+1}(B_{t+1}; n_i) = T'(z_t) \beta^{\bar{t}-t-1} \left(\Pi_{i=t}^{\bar{t}-2} p(z_i) \right) u_c(c_{\bar{t}-1}, z_{\bar{t}-1}; n_i) \tag{30}$$

where $\Pi_{i=t}^{\bar{t}-2} p(z_i)$ is the probability to exhaust benefits conditional on claiming at date t. Using Equation (30) and the utility definition, the FOC in Equation (29) simplifies to:

$$1 - T'(z_t)\tau_t(z_t) - \beta p'(z_t) \left(W - U(B_{t+1}; n_i) \right) = \left(\frac{z_t}{n_i} \right)^{1/e} \tag{31}$$

where the wedge τ_t now depends explicitly on z_t:

$$\tau_t(z_t) = 1 - \beta^{\bar{t}-t-1} \Pi_{j=t}^{\bar{t}-2} p(z_j). \tag{32}$$

We now state two assumptions that are sufficient to obtain identification of the earnings elasticity to the net-of-tax rate. First, the marginal effect of earnings on the permanent job finding probability $p'(z)$ is continuous. Second, the net gain of permanent jobs ($W - U(B_{t+1}; n_i)$) depends continuously on earnings z_t and depends on individual ability only through earnings. These assumptions imply that there exists a continuous function π_t such that $\pi_t(z_t) = \beta p'(z_t) \left(U(B_{t+1}; n_i) - W \right)$.

Consequently, the FOCs can be written as:

$$1 + \pi_t(z_t) = \left(\frac{z_t}{n_i} \right)^{1/e} \text{ when } z_t < z^*, \tag{33}$$

$$1 - \tau_t(z_t) + \pi_t(z_t) = \left(\frac{z_t}{n_i} \right)^{1/e} \text{ when } z_t > z^*. \tag{34}$$

This leads me to define a lower threshold n_t^* and an upper threshold $n_t^* + \delta n_t$, such that:

$$n_t^* = \frac{z^*}{(1 + \pi_t^-(z^*))^e}, \tag{35}$$

$$n_t^* + \delta n_t = \frac{z^*}{\left(1 - \tau_t(z^*) + \pi_t^+(z^*) \right)^e}. \tag{36}$$

where π^+ and π^- are respectively the upper and lower limits of π. Because π is assumed continuous, the marginal gains induced by stepping-stone/crowding-out effects cancel out of the identifying relation (as long as $\pi_t(z^*) << 1$). Then the elasticity verifies the same identification relation: $e = \frac{\mathcal{B}_t}{z^* \tau_t(z^*)}$.

D Hazard model

In this appendix, I report results of the estimation of the hazard model used to compute the probability to remain claiming the following week (p). I follow the baseline assumptions of the theoretical model and neglect any duration dependence (p does not depend on t). I estimate the following exponential hazard model where covariates enter proportionally. For individual i, the hazard model is: $h_i = h_0.\exp(\beta X_i)$. The hazard model is estimated on a subsample of claimants, according to the local nature of the bunching estimate. I am interested in the hazard rate of claimants, close to bunching. I thus restrict the estimation to claimants whose benefits are not reduced because of partial UI.

It is well-established that hazard rates out of UI registers feature spikes at benefit exhaustion date. I verified that I obtain such patterns in the data from the Continuous Work and Benefit History (CWBH) project, as Katz and Meyer (1990b) do. As I want to capture the probability to remain claiming for individuals who are still entitled to unemployment benefits, observations are censored before exhaustion spikes. I use the theoretical exhaustion date in Tier 1 when claimants are totally unemployed along the whole claim ($\bar{t}^{Utot}$), in order to censor observations.

My objective is to estimate claimants' expectation about their hazard rates. Rational forward-looking claimants would use all available information to form their expectations. Consequently, covariates X capturing individual heterogeneity include: gender, age (and its square), years of initial education (and its square), ethnicity, calendar year of first week of claim, potential benefit duration (in Tier 1), weekly benefit amount and recall expectation. For each covariate, a specific dummy is included to account for missing values. Table D1 reports the coefficient estimates of the hazard model for each state (in columns).

Table D1: Results of hazard model estimation

	Idaho	Louisiana	New Mexico	Missouri
Male	.078***	.288***	.080***	.175***
	(.021)	(.015)	(.015)	(.015)
Age	-.021***	-.006**	-.031***	-.006*
	(.004)	(.003)	(.004)	(.003)
Age (square)	.0001**	-.00005	.0003***	-.00002
	(.00005)	(.00003)	(.00004)	(.00004)
Education (years)	-.111***	-.061***	.012	-.036
	(.021)	(.009)	(.011)	(.062)
Education (square)	.006***	.004***	.0007	.004
	(.0009)	(.0004)	(.0005)	(.003)
Black	-.042	-.216***	-.172***	-.583***
	(.103)	(.013)	(.049)	(.020)
Hispanic	.296***	.097*	-.223***	-.207
	(.044)	(.053)	(.014)	(.153)
American Indian	-.164*	-.084	-.231***	-.211
	(.093)	(.122)	(.024)	(.378)
Potential benefit duration	.053***	.031***	.059***	.044***
	(.002)	(.002)	(.007)	(.002)
Weekly benefit amount	-.001***	-.003***	-.001***	-.005***
	(.0003)	(.0001)	(.0002)	(.0004)
No recall expectation	-.484***	-.249***	-.398***	-.600***
	(.025)	(.016)	(.013)	(.016)
Constant	-2.846***	-3.327***	-3.929***	-3.279***
	(.143)	(.081)	(.198)	(.358)
Years fixed effects	Yes	Yes	Yes	Yes
No. spells	25274	55519	37937	41663
Log-likelihood	-32412.47	-75213.41	-53243.84	-56269.63

Source: CWBH. Notes: The reference is a white female with recall expectation whose claim starts in the first year of the sample. *** $p<0.01$, ** $p<0.05$, * $p<0.1$

E Difference-in-difference

In April 1983, Louisiana changed UI rules. The change in partial UI affected both the stock of individuals registered in April 1983 and new inflows after that point in time.[42] The disregard level was reduced from $0.5 \times WBA$ to \$50 for all claimants whose WBA is more than \$100. This is the treatment group. For all claimants with a WBA below \$100, the disregard was not reduced and remained equal to $0.5 \times WBA$. This is the control group. I select claims around the policy shocks, from April 1982 to March 1984. The sample covers a full year before the policy change and another full year after the new rules were implemented.

I expect that, if bunching is actually related to the partial-UI schedule, the bunching location would switch from the old to the new threshold in the treatment group, and remain the same in the control group. If bunching is due to norms or policies unrelated to the partial-UI program, bunching (in the treatment group) should not be altered by the policy change.

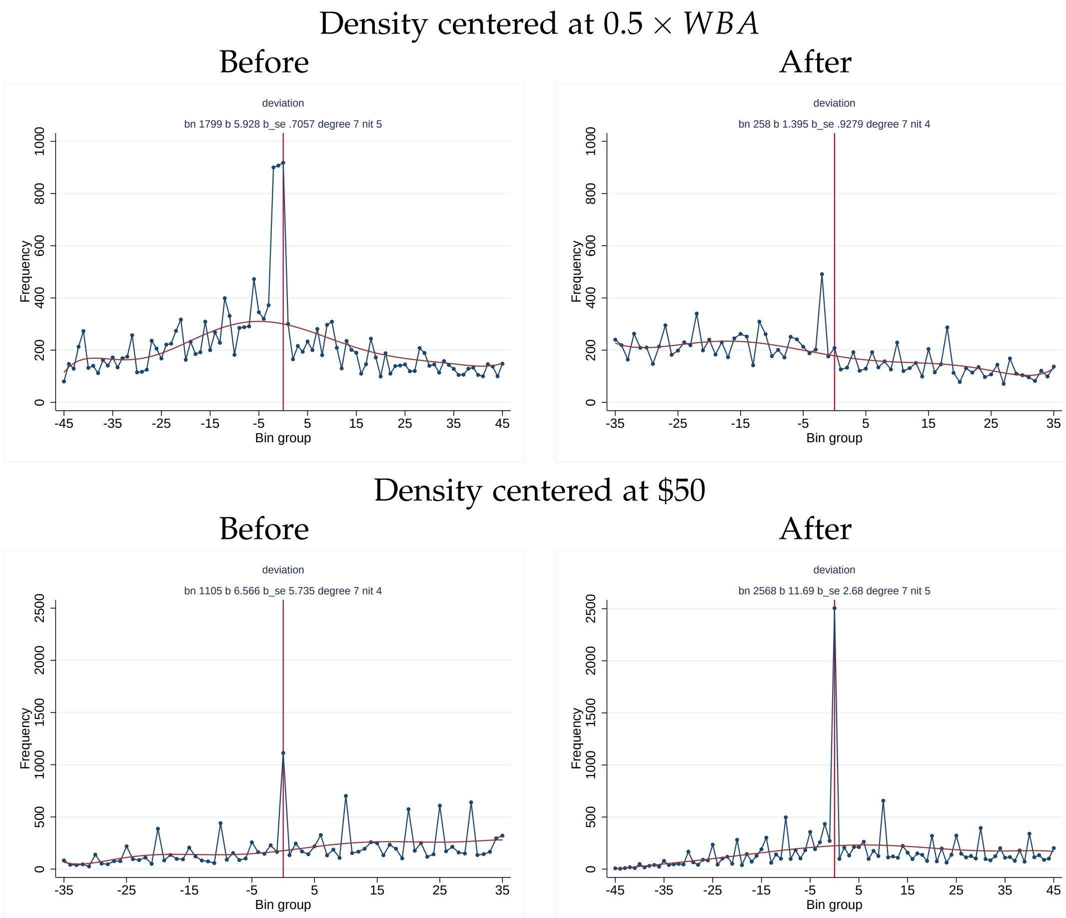
Figure E1 plots the earnings density of partial-UI claimants in the treatment group. In the upper panel, densities are centered at the pre-reform disregard ($0.5 \times WBA$). In the lower panel, they are centered at the post-reform disregard (\$50). Starting with the upper panel, bunching is considerably reduced from before the reform (left graph) to after the reform (right graph). Bunching estimate at the pre-reform disregard level is no longer statistically significant after the reform. The lower panel shows that claimants actually switch to the post-reform disregard after the reform. The mass of bunchers at \$50 doubles after April 1983. Note that there were actually some claimants at the \$50 threshold before the reform. This may be explained by norms unrelated to the partial-UI program. The important point here is that bunching increases after the reform. Note also that bunching is sharper when disregards are rounded amounts.

Figure E2, in which I repeat the same exercise for the control group, does not display any fundamental changes in the bunching pattern after the reform. Claimants in the control group continue to bunch at their relevant disregard amount ($0.5 \times WBA$). They do not switch to the post-reform disregard of the treatment group (\$50). The absence of bunching after the reform in the control group also suggests that bunching incentives mediated by the demand side of the labor market are weak in Louisiana. Suppose that firms actually internalize the partial-UI program and post wages at the disregard level. Because they cannot direct their search to claimants with certain disregard levels, it is likely that they would post the most common disregard (Chetty et al., 2011a). In Louisiana, the mode of

[42] There was also a reduction in the maximum number of entitlement weeks from 28 to 26 weeks. This could have affected the amount of bunching, but not its location.

the disregard distribution is $50 (the treatment group is twice as large as the control group). If the bunching incentives were mainly mediated by firms, I would expect to see bunching at $50 in the control group, which is not the case.

Figure E1: Centered weekly earnings density of partial-UI claimants in the treatment group.

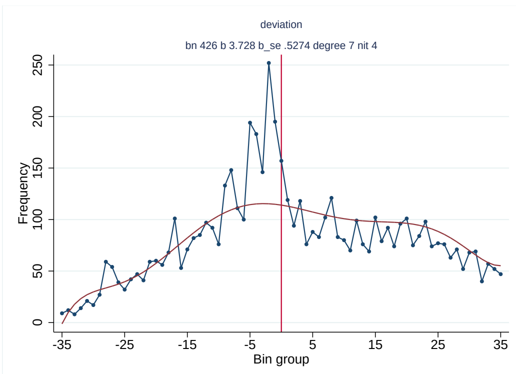


Source: CWBH. Notes: Earnings are in dollars. Empirical earnings density in blue. Counterfactual density in red.

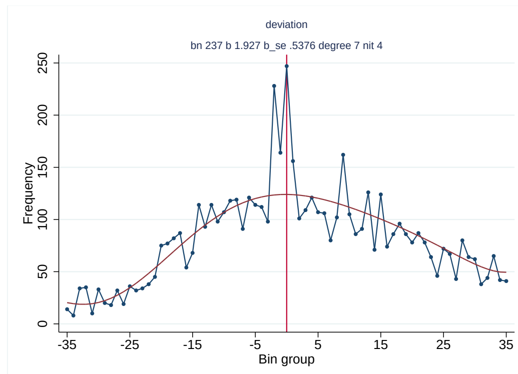
Figure E2: Centered weekly earnings density of partial-UI claimants in the control group.

Density centered at $0.5 \times WBA$

Before

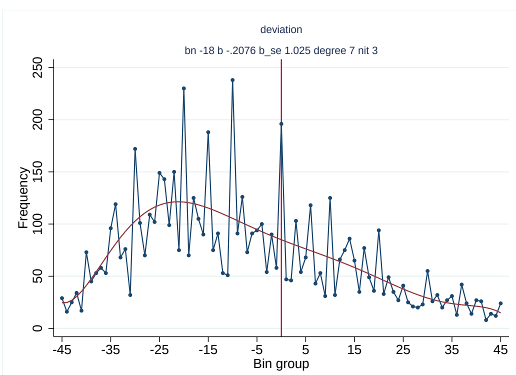


After

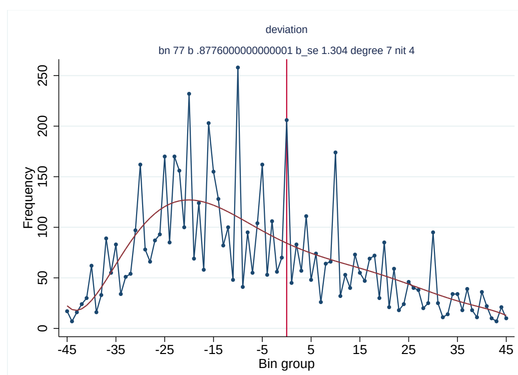


Density centered at \$50

Before



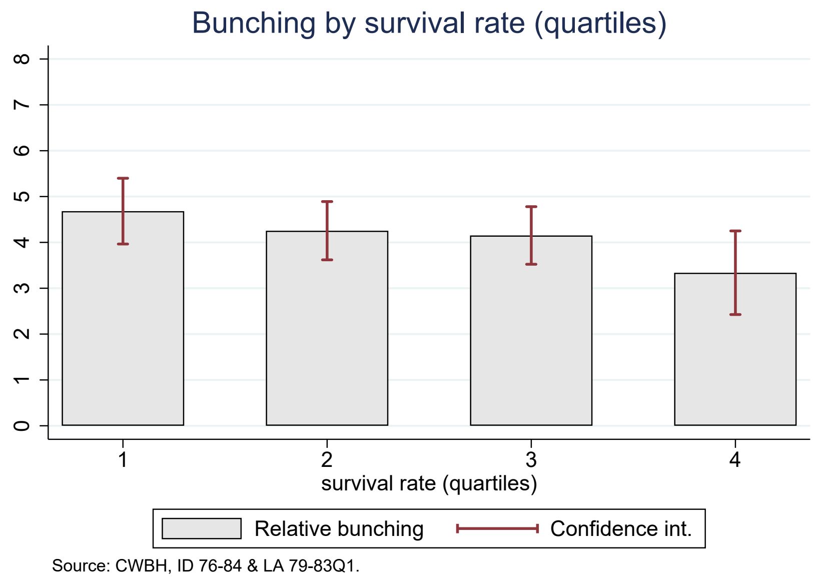
After



Source: CWBH. Notes: Earnings are in dollars. Empirical earnings density in blue. Counterfactual density in red.

F Supplementary Figures and Tables

Figure F1: Bunching by predicted survival rate.



Source: CWBH for Idaho 1976-84 and Louisiana 1979-83, Q1. Notes: Excess mass at the disregard amount (kink) by quartile of predicted survival rates. Confidence interval at the 95% level in red.

Table F1: Robustness of earnings elasticities to the net-of-tax rate varying estimation parameters

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Baseline	Lower bound			Upper bound		Polynomial degree		
Bandwidth	[-5,2]	[-15,2]	[-10,2]	[-3,2]	[-5,1]	[-5,3]	[-5,2]	[-5,2]	[-5,2]
Poly. Deg.	7	7	7	7	7	7	9	5	3
Idaho	0.187***	0.287***	0.264***	0.134***	0.188***	0.184***	0.167***	0.216***	0.276***
	(0.010)	(0.021)	(0.014)	(0.010)	(0.010)	(0.011)	(0.008)	(0.012)	(0.019)
Louisiana	0.129***	0.215***	0.168***	0.108***	0.142***	0.129***	0.129***	0.145***	0.187***
	(0.007)	(0.017)	(0.009)	(0.006)	(0.007)	(0.007)	(0.008)	(0.008)	(0.011)
New Mexico	0.096	.	0.197	0.100**	0.054	0.071	0.053	0.057	0.039
	(0.065)	.	(0.214)	(0.044)	(0.048)	(0.068)	(0.063)	(0.054)	(0.049)

Source: CWBH. Notes: This Table reports estimates of the earnings elasticity to the net-of-tax rate varying the estimation parameters. Column 1 recalls the results of the baseline estimation (in Table 2) for the three U.S. states: ID, LA and NM. In Columns 2 to 4, I increase the lower bound of the bunching window. In Columns 5 and 6, I increase the upper bound of the bunching window. In Columns 7 to 9, I decrease the degree of the polynomial fitting the density. Because the disregard level is around \$20 in NM, it does not make sense to consider a lower bound at -15, and the estimation results are not reported. Standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table F2: Earnings elasticity estimates without first-order approximation of the marginal tax rate.

	Idaho	Louisiana	New Mexico
Earnings elasticity to net-of-tax rate (e)	.130***	.092***	.063
	(.0065)	(.0055)	(.0457)

Source: CWBH. Notes : This Table reports estimates of earnings elasticity to the net-of-tax rate, computed with the exact identifying formula $e = -\mathcal{B}/z^* / \ln(1-\tau_t)$. Standard errors are in parentheses below estimates. *** p<0.01, ** p<0.05, * p<0.1