Optimal Partial Unemployment Insurance: Evidence from Bunching in the U.S.*

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Abstract

In this paper, I use kinks in the U.S. partial unemployment insurance schedule to study the response of claimants to the program. Partial unemployment insurance enables claimants to keep part of their unemployment benefits when they work in low-earnings jobs. When U.S. claimants earn over a state-specific threshold, termed the "disregard", their benefits are reduced at a 100% marginal tax rate above that amount. This reduction in current benefits leads to an increase in future benefits, with the result that forward-looking claimants are taxed according to a lower dynamic marginal tax rate. To account for these mechanisms, I develop a dynamic model of claimants, who work in part-time/temporary jobs while searching for permanent jobs. Using administrative data on weekly claims, I document substantial bunching of unemployment insurance claimants at the disregard level. I estimate that the earnings elasticity to the net-of-tax-rate (at the intensive margin) lies between 0.1 and 0.2. Using this estimate, simulations show that setting the benefit reduction rate at 80% is Pareto improving, as the current schedule induces claimants to inefficiently reduce their earnings.

Keywords: unemployment insurance, temporary/part-time work, tax bunching
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1 Introduction

The design of optimal unemployment insurance (UI) has been addressed by a large academic literature and is regularly at the center of the public debate.[1] Mostly, the literature emphasizes the issue of moral hazard associated with unemployment benefits. By reducing the net gain from employment relative to unemployment, UI reduces incentives for the unemployed to search for a job or to accept low-wage employment. To reduce such disincentives effects, many countries implement partial unemployment insurance rules. Partial UI enables claimants to keep their unemployment benefits (or a percentage of their benefits) while they work in low-earnings jobs – usually part-time or temporary work. In 2012, 12% of UI claimants in OECD countries work while on claim.[2] Despite its wide implementation, partial UI has not been fully incorporated in theoretical or empirical work on unemployment insurance. This research-policy gap is especially apparent in the U.S., where the most recent contribution is the empirical study by McCall (1996), while partial UI concerns almost 20% of claimants in some U.S. states. In this paper, I attempt to fill this gap by studying behavioral responses to the U.S. partial-UI program using administrative weekly UI data.

Partial unemployment insurance can be viewed as a form of in-work benefits. As such, it induces unemployed claimants to work in low-earnings jobs and thus affects the extensive margin of labor supply. This is confirmed by McCall (1996) who shows that part-time work is more prevalent in U.S states where partial UI is more generous. To the best of my knowledge, my paper makes a first contribution on the behavior of claimants at the intensive margin. I investigate, whether, conditional on working, UI claimants adjust earnings in response to the benefit reduction in the partial-UI schedule. Understanding the intensive margin is important to evaluate the efficiency of partial UI. A reduction of the benefits of partial-UI claimants, on behalf of the UI administration, could induce claimants to reduce their earnings. This resulting reduction in earnings could in turn increase the overall cost of partial UI. Using kinks in the partial UI schedule and the bunching of claimants at the corresponding earnings levels, I estimate that the earnings elasticity to the net-of-tax rate[3] lies between 0.1 and 0.2.[4] Using this estimate, simulations show that the (flat-rate) policy that minimizes the benefits paid to partially unemployed claimants, is to set the benefit-reduction rate at 80% (above the disregard).

[1] Baily (1978) is an early contribution on optimal unemployment insurance and has been followed by Chetty (2006) and Shimer and Werning (2007), among others.

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

[3] This is the elasticity of earnings to one minus the marginal tax rate.

[4] This is consistent with estimates found in previous micro empirical work using annual data (see the review of quasi-experimental estimates in Chetty (2012) or Chetty, Guren, Manoli and Weber (2011)).

Intertemporal aspects are key to understand the behavioral response of claimants to the partial-UI schedule. In most U.S. states, when claimants earn under the disregard level, the marginal benefit-reduction rate is zero. Then, for every dollar earned above the disregard, current benefits are reduced on a dollar-per-dollar basis: the static marginal benefit-reduction rate is 100%. However, the reduction in benefits because of the partial-UI program is not lost, it can be paid in a later week. The corresponding benefit transfers delay the potential benefit exhaustion date. As a consequence, forward-looking claimants make their labor supply decisions based on a dynamic marginal tax rate, which is lower than the static benefit-reduction rate. To account for these intertemporal effects, I develop a dynamic model of labor supply while on claim, where job-seekers also search for permanent jobs, which make them ineligible for partial UI. This model enables me to derive the analytical expression of the dynamic marginal tax rate, which depends not only on the discount factor and on the marginal benefit-reduction rate, but also on the claimant's expected survival probability in the UI registers. 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-transfer mechanism and her dynamic marginal tax rate is larger, closer to the static benefit-reduction rate. I therefore estimate a hazard model of exiting the UI registers, and calculate that claimants with rational expectations have on average a dynamic marginal tax rate of approximately 60%. My dynamic model shows that bunching at the (convex) kink of the partial-UI schedule identifies the earnings elasticity to the net-of-tax rate.

I compute bunching and elasticity 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.[5] The data set has the unique advantage of containing weekly unemployment benefits payments and weekly earnings of claimants. I find substantial bunching at the disregard level. In Idaho and Louisiana, the excess mass at the disregard is five times the population density that would work at this level 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 the dynamic marginal tax rate, rather than to the static 100% marginal tax rate.[6]

I perform two placebo exercises. First, disregard amounts are comparable in Idaho and Louisiana, around $50 (current dollars), however the cutoff is much lower in Missouri, around $10. I confirm that there is no bunching in Missouri at the disregard levels

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

[6] However this is not a definitive test of the dynamic aspects of my model, as adjustment costs/search frictions for small jobs could also explain why myopic claimants work for earnings above the disregard.

prevailing in Idaho and Louisiana. Second, there was a policy shock in Louisiana in April 1983, when the disregard amount was changed for claimants with high previous wages. The location of bunching for this treated group changed from the high previous disregard level to the low new disregard amount. The different placebo tests show that the bunching identified in the data is actually linked to the partial-UI schedule and is not an artifact of other labor legislations or social norms.

Bunching heterogeneity is consistent with the fact that claimants react to the benefit-transfer mechanism of the partial-UI program. Claimants with longer potential benefit duration bunch more in my data. As predicted by the dynamic model, they have lower incentives to use the partial-UI program to delay the benefits exhaustion date. I also verify that bunching estimates are larger for claimants with a low propensity to remain on the UI registers, such as claimants expecting to be recalled to their previous employer (Katz and Meyer, 1990*b*). Additionally, I study the evolution of bunching with unemployment duration, holding the sample of claimants constant (to avoid composition effects). The theoretical model predicts that bunching should decrease over the claim.[7] I find a moderate decrease in bunching estimates, close to the exhaustion date. But this decrease is not as steep as the model predicts. This could be explained by search frictions or adjustment costs.[8]

Finally, I consider the program of the UI administration that chooses the benefit-reduction rate to maximize social welfare while maintaining its costs (benefits payments) below a certain exogenous level. The optimal formula suggests that a constant benefit-reduction rate is a second-best solution. Costs could be further reduced by allowing rates to depend on the remaining benefit entitlement or the claiming duration. However, because constant benefit-reduction rates are easier to implement, I only perform simulations in this case. They suggest that the UI administration could reduce its costs (while increasing social welfare) by setting the benefit-reduction rate at 80%. Switching from a 100% benefit-reduction rate to this optimal level would reduce the UI costs by about 2%.

My paper contributes to the literature on partial UI. It is the first paper that studies specifically behavioral response at the intensive margin. In the U.S., McCall (1996) and the early contributions of Holen and Horowitz (1974) and of Kiefer and Neumann (1979) show that the partial unemployment rate is positively correlated with the state's disregard level and thus document the elasticity at the extensive margin.[9] All European studies on

[7] Claimants have a higher probability to receive transferred benefits when they are already at the end of their claim.

[8] I also find that, at the beginning of the claim, bunching estimates increase with time. This also suggests search frictions, or alternatively learning effects.

[9] Munts (1970) is an early descriptive contribution on partial UI in the U.S.

partial UI propose to estimate the effects of part-time jobs on the probability to find permanent jobs (Kyyrä, 2010; Caliendo et al., 2012; Kyyrä et al., 2013; Fremigacci and Terracol, 2013; Godoy and Røed, 2014). This literature asks whether partial-UI jobs act as stepping-stones to regular employment or whether they crowd out the time dedicated to job search. Using the timing of events approach, European studies point to mixed results depending on the socio-demographic group. I discuss how my identification strategy accommodates for stepping-stone or crowding-out effects of partial-UI jobs.

I also contribute to the literature on the estimation of intensive elasticities using kinks in tax (or benefit-reduction) schedule. I provide first empirical evidence of substantial bunching at kinks in benefits/tax schedule for UI claimants in the U.S. This complements the findings of Saez (2010) and Chetty et al. (2013) who show that U.S. *low-earnings self-employed* workers bunch at the kink of the EITC schedule. This also complements the results of Gelber et al. (2013) who find bunching among U.S. old-age wage-earners at the kink of the Social Security Annual Earnings Test (SSAET).[10] From a theoretical perspective, I extend the Saez (2010) formula by incorporating dynamic considerations when benefits or taxes can be transferred to the future. le Maire and Schjerning (2013) also consider dynamic aspects in income tax schedule, but they specifically model income shifting by the self-employed.[11] Brown (2013) adapts the Saez (2010) formula to a lifecycle model, where delaying retirement increases future annuities.

The paper is organized as follows. In Section 2, I present the U.S. partial unemployment insurance program. In Section 3, I develop a job-search model of a claimant working while on claim and derive the identification result. In Section 4, I detail the different steps of the estimation procedure and the data. In Section 5, I present my main estimates of bunching and the corresponding earned income elasticities to the net-of-tax rate, I also perform various placebo tests including a difference-in-difference analysis in Louisiana. In Section 6, I document the heterogeneity of bunching and its evolution along the claim. In Section 7, I perform a normative exercise. Finally, Section 8 concludes.

[10]Note that the population analyzed by Gelber et al. (2013) is much older than my sample, as that study concerns workers above the national retirement age.

[11]Gelber et al. (2013) also discuss intertemporal aspects of the U.S. Social Security Annual Earnings Test. As in my case, 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.

2 Institutional background

In the U.S., when unemployment insurance (UI) claimants work while on claim, they are eligible for partial unemployment benefits. The definition of eligibility varies across states, but in all those considered here, partial-UI claimants must not earn more than a maximum amount of labor income over a week.[12] This maximum amount is usually set as a fraction of the weekly benefit amount (WBA), which is the unemployment benefits (UB) payment when claimants do not work (i.e. total unemployment benefits). Partial-UI claimants are paid their weekly benefit amount when their weekly earnings are below the state-specific "disregard" threshold. When partial-UI claimants earn between the disregard and the maximum amount, their current benefits are reduced by their earnings minus the disregard. The static marginal benefit-reduction rate is then 100%. Table 1 displays the parameters of partial-UI rules for the four states analyzed in this paper. The most generous state is Idaho: it has the highest maximal wage threshold and its disregard is also high. The three other states have a maximal amount around the WBA (in practice less than 1.1 times the WBA). Missouri is the least generous state. Its disregard is stated in dollar values and, actually, does not exceed 10% of the WBA. In Louisiana, the disregard was reduced in April 1983. I will use this policy shock as a placebo exercise in a difference-in-difference analysis.

Figure 1 illustrates the partial-UI schedules in Idaho (ID), Lousiana (LA), New Mexico (NM) and Missouri (MO). I plot the weekly net income (earnings plus UB payments) against the weekly earnings while on claim. 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. The graphics clearly illustrate that the schedule is kinked at the disregard amount. From a static point of view, there are no incentives to work for a wage just above the disregard, as the net income is essentially a plateau above that level. The graphics also illustrate the notch at the maximal amount in Louisiana and New Mexico (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 threshold (Kleven and Waseem, 2013). Because of data limitations, I will not analyze the claimants' behavior around notches. The incentives to claim jump discontinuously at the notch value, so that individuals above the notch should leave the UI registers, and hence my data. As will become clear below, I take advantage of the absence of kinks in Missouri at the disregard level prevailing in Idaho and Louisiana ($0.5 \times WBA$) to perform a placebo exercise.

[12] All components of labor income are considered in the computation. The only exceptions are payments for jury service in New Mexico and wages from "service in the organized militia for training or authorized duty from benefit computation" in Missouri.

UI claimants have to pay income taxes on their labor earnings.[13] Before 1979, there were no federal taxes to be paid on unemployment benefits. Since 1979, unemployment benefits have been taxable for single tax filers with income over $20,000 and for married filers with income over $25,000. The thresholds were lowered in 1982 to respectively $12,000 and $18,000. The difference between income tax thresholds for labor earnings and for unemployment benefits may affect the relative gains of working while on claim. However, it is very unlikely that the weekly disregard level of partial UI corresponds to another kink or discontinuous change in the annual income tax schedule or welfare system. This ensures that my identification strategy below is robust to the existence of other incentives caused by the whole tax and benefit system.

I now turn to the dynamic aspects of the partial-UI schedule.[14] At the beginning of the claim, the UI administration computes a total entitlement amount (B_0), which depends on past earnings. The total entitlement can be thought of as a kind of UB capital that depreciates with UB payments. If claimants are totally unemployed all along their claim and receive each week their WBA, their benefits will lapse after B_0/WBA weeks, defined as their initial Potential Benefit Duration (PBD). The initial PBD typically varies between 10 and 26 weeks (see the Appendix for more details). 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, the UB capital depreciates at a slower pace. Working while on claim is thus a way to delay the benefit exhaustion date. Note, however, that the exhaustion date cannot be delayed forever, as any remaining UB capital is lost one year after the first claiming week, defined as the benefit year.[15]

Except for the above earning thresholds, there is no other specific eligibility condition for partial UI. Claimants must only meet the usual UI eligibility requirement (described in the Appendix). Partial-UI claimants are allowed to work for any employer, including their past employers; claimants who are temporarily laid off are also eligible for partial UI.[16] Also, individuals whose hours have been reduced at their current workplace are eligible for partial UI, so long as they can file a claim based on this reduction in hours worked. Note that claimants with reduced hours represent a small share of partial-UI claimants.[17]

[13] In practice, I expect that many UI claimants will be below the minimum income threshold of the income tax schedule.

[14] Partial-UI rules have not changed since the early 80s (see for example Chapter 3 of the 2013 DOLE booklet "Comparison of the State Unemployment Laws"). The dynamic aspects of the partial-UI schedule are also described in Kiefer and Neumann (1979) and McCall (1996).

[15] Also note that working while on claim is not sufficient to meet the eligibility requirement to register for a new claim, as the earnings in partial-UI jobs are too low.

[16] Indiana, which I do not study in this paper, is an exception: claimants are not eligible for partial UI if they work for their previous employers.

[17] In the CWBH data, I cannot distinguish between claimants taking up new jobs and claimants with

The partial-UI rules described above correspond to the default UI system (Tier I), but they are also valid when additional 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). First, the EB program, which was administrated by the state, extended the initial entitlement period by 50% up to a total of 39 weeks when the state unemployment rate reached a certain trigger. Second, the FSC program, in action from September 1982 to March 1985, 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. Extension programs induce uncertainty about claimants' entitlement period, which depends on the labor market conditions when regular benefits lapse.[18]

Eligibility to partial UI is based on reported earnings. There is thus scope for claimants manipulating their reports. However the UI administration takes action to limit false statements. The UI administration performs random audits of claimants' declarations.[19] If fraud is detected, it can be severely punished as a Class VI Felony.[20]

3 Theoretical model

In this section, I develop a dynamic model of working while on claim that incorporates the dynamic aspects of the partial-UI program. The model features reasonable assumptions from both the job-search literature and public finance literature on bunching: quasi-linear utility, frictions in the search for permanent jobs and liquidity constraints. 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. The identification of the earned income elasticity then follows a modified bunching formula (Saez, 2010).

3.1 Setup

I consider an infinitely lived individual i claiming benefits from period 0 on (with discount factor β). Following UI rules, periods are weeks in my model. Until she finds a permanent

reduced hours (Short Time Compensation), except in Louisiana since 1982. From 1982 to 1984, only 15.7% of partial-UI weeks in LA concerned claimants with reduced hours.

[18] Details about the trigger dates are in the Appendix.

[19] The UI administration currently cross-checks W-2 and new hires declarations of employers with claimants reported earnings.

[20] Criminal action may result in up to 2 years in prison and fines up to $150,000 for each false statement.

job, the job-seeker may work in a small job, defined as short-term, part-time work eligible for partial UI. Then she earns z_t in period 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.[21] In the baseline model, the job-seeker faces no frictions (or adjustment cost) to find a small job. I discuss this assumption below. The per period utility $u_i(c, z)$ of job-seeker i depends on consumption c and labor earnings in small jobs z. I also assume that the utility is quasi-linear (ruling out income effects) and that it depends on the individual talent n_i:

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

where e is my parameter of interest. e captures the earnings elasticity to the net-of-tax rate.

At each date $t > 0$, the job-seeker may find a permanent job with probability $(1 - p(z_t))$. 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. Claimants never decline permanent job offers. The probability to find a permanent job may depend on the amount of earnings on small jobs. This captures potential stepping-stone effects or job-search crowding-out effects of small jobs. Note that I model small and permanent jobs as totally separated markets. The market for small jobs is tight, there are no search frictions, however the utility derived from working in small jobs is low. The market for permanent jobs feature search frictions, but they yield a very high utility (which is assumed not to depend on talent in small jobs).

At the beginning of her claim, the job-seeker has a total benefit entitlement (or UB capital) equal to B_0 (not discounted). 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 is registered and 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 sufficiently large to pay b. If the job-seeker does not work at all along her unemployment spell, she receives benefits during $t_{exh}^{Utot} = \lceil B_0/b \rceil$ periods (t_{exh}^{Utot} corresponds to the date of exhaustion of total unemployment, i.e. the first date when UB capital is zero). When she takes up a small 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

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

unemployment registers, but she still looks for a permanent job and she may still work for a small job. The actual partial-UI schedule[22] $T(.)$ is defined as:

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

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 ($z^* + b$). As explained above, I will abstract from this second kink as data limitation prevents me to analyze behaviors around the maximal earnings amount.[23] This second kink does not affect my identification strategy that is local and around the first kink.

Let me define $U_t(B_t)$ the value of unemployment at time t when the UB capital is B_t. At each date, the job-seeker maximizes the following program:

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

such that

$$\begin{aligned} c_t &= z_t + b\mathbb{1}\left[B_t > b\right] + B_t\mathbb{1}\left[b > B_t > 0\right] - T(z_t)\mathbb{1}\left[B_t > 0\right] \\ B_{t+1} &= B_t - b\mathbb{1}\left[B_t > b\right] - B_t\mathbb{1}\left[b > B_t > 0\right] + T(z_t)\mathbb{1}\left[B_t > 0\right] \\ B_{t+1} &\geq 0 \end{aligned}$$

The first constraint of the program is a standard current budget constraint. I assume that agents cannot lend or borrow, as UI claimants are likely to be low-skilled workers who are credit-constrained. The only non-standard elements in this first constraint are the dummies variables that model the relation between current total UB payment (b) and the remaining UB entitlement (B_t). The second constraint captures the endogenous entitlement reduction (or UB capital depreciation). The UB capital is reduced by the UB payment $b - T(z_t)$, if the entitlement at the beginning of the period t is sufficiently large ($B_t > b$), and by $B_t - T(z_t)$, if it is not. The last constraint states that job-seekers cannot borrow UB entitlement from the UI administration.[24]

[22] In the definition of the partial-UI schedule, I assume 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 it focuses on earnings close to the kink.

[23] 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.

[24] For the sake of simplicity, I do not model the fact that any remaining entitlement at the end of

3.2 Model solution

I describe below solutions where the UB capital is strictly decreasing.[25] I thus define $t_{exh} < \infty$ the exhaustion date (first date when $B_t = 0$). Note that the exhaustion date is endogenous and depends on both the initial entitlement level and on the solution path of z_t. I describe the program solution when $B_t > b$, which is more relevant to the empirical analysis. For all z_t, such that $T(.)$ is differentiable at z_t, the first order condition is:

$$\underbrace{u_c(c_t, z_t)\left(1 - T'(z_t)\right)}_{(1)} - \underbrace{\beta p'(z_t)\left(W - U_{t+1}(B_{t+1})\right)}_{(2)} + \underbrace{\beta p(z_t) T'(z_t) U'_{t+1}(B_{t+1})}_{(3)} = -u_z(c_t, z_t) \tag{3}$$

where u_c is the marginal utility of consumption and u_z the marginal disutility of work. Equation 3 describes that the marginal gains of work (on the left-hand side) equal the marginal cost of effort (or disutility of work). The marginal gains have three components. The first term 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 represents the marginal gain induced by the stepping-stone effect of small jobs (when $p' < 0$). Alternatively it corresponds to the marginal cost induced by the crowding-out effect (when $p' > 0$). The third term is the marginal value of future UB capital. It depends on the discount factor, on the survival rate and on the marginal benefit-reduction rate.

Using the envelope theorem (at every future period), I compute the marginal value of UB capital, and the third term of Equation 3 simplifies as:[26]

$$\beta p(z_t) T'(z_t) U'_{t+1}(B_{t+1}) = T'(z_t) \beta^{t_{exh}-t-1} \left(\Pi_{j=t}^{t_{exh}-2} p(z_j)\right) u_c(c_{t_{exh}-1}, z_{t_{exh}-1}) \tag{4}$$

Note that $\Pi_{j=t}^{t_{exh}-2} p(z_j)$ is the probability to exhaust the whole benefit entitlement conditional on claiming at date t.

Using Equation 4 and the properties of the quasi-linear utility, Equation 3 simplifies as:

$$1 - T'(z_t)\tau_t - \beta p'(z_t)\left(W - U_{t+1}(B_{t+1})\right) = \left(\frac{z_t}{n}\right)^{1/e} \tag{5}$$

the benefit year is lost, as in the data, almost all job-seekers find permanent jobs or exhaust their UB entitlement before that date. Note that I also assume that the value of a permanent job does not depend on current entitlement. The UI rules enable the former claimants who have left the UI registers to work in a new job, to claim past remaining entitlement in the same benefit year if they lose their new job. However, in the data, this happens in very few occasions, so I do not model this possibility.

[25] I show, in the Appendix, that such a focus is relevant when studying the behavior of claimants around the disregard point.

[26] See the Appendix for computation details.

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

$$\tau_t = 1 - \beta^{t_{exh}-t-1} \Pi_{j=t}^{t_{exh}-2} p(z_j). \tag{6}$$

Compared to the usual bunching identification, Equation 5 has two supplementary elements: the marginal gain of labor earnings associated with the benefit transfer and the marginal gain induced by stepping-stone/crowding-out effects. In the baseline model, I assume away stepping-stone/crowding-out effects. The marginal gain induced by such effects is likely to be small, as it is scaled by the marginal effect of one extra dollar of earnings on the probability to find a permanent job. Previous empirical studies in European countries or in the U.S. find small effects of partial-UI jobs on permanent employment. They typically estimate an average effect compared to the counterfactual of no partial UI: $\int p(z)g(z)dz - p(0)$. This means that the marginal effect $p'(z)$ is likely to be even smaller. Alternatively (when $p'()$ cannot be considered small), I propose below theoretical conditions, under which kinks still allow to estimate the earnings elasticity to the net-of-tax rate.

Under the assumption of no stepping-stone/crowding-out effects, the wedge simplifies as $\tau_t = 1 - (\beta p)^{t_{exh}-t-1}$ and the FOC in Equation 5 further simplifies as:

$$1 - T'(z_t)\tau_t = \left(\frac{z_t}{n}\right)^{1/e} \tag{7}$$

Note that, if there was no benefit reduction at all ($T(z) = 0$ for all z), all individuals would supply $z_t = n$. The talent n_i of individual i can thus be interpreted as her potential earnings in small jobs. The actual partial-UI schedule ($T(.)$) 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 (corner solution).

To describe the bunching behavior, I define a first threshold at talent $n^* = z^*$ (recall that z^* is the disregard level). The FOC implies that all individuals with $n < n^*$ earn $z_t = n$. I define another threshold of talent $n^* + \delta n(t)$, such that all individuals with talent strictly over $n^* + \delta n(t)$ earn strictly more than the disregard z^*. Such individuals have their current benefits reduced and they earn $z_t = n(1 - \tau_t)^e$ (as $T'(z_t) = 1$). Using the FOC, the upper threshold then verifies:

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

Equation 8 illustrates that the upper threshold depends on the time period. More fundamentally, it shares the dependence structure of the dynamic marginal tax rate. 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.

Note that a priori the dynamic marginal tax rate τ_t may depend on talent n, as it depends on t_{exh}, which derives from the program solution. However, for individuals just above $n^* + \delta n(t)$, their exhaustion date is only delayed by one period and I can reasonably set $t_{exh} = t_{exh}^{Utot}$.[27]

To summarize, the earnings density function $g_t(z)$ at period t verifies:[28]

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

where $f_t(n)$ is the talent density of claimants at period t.

3.3 Identification strategy

The bunching mass at the disregard level $g_t(z^*)$ can be approximated by $f_t(n^*)\delta n(t)$. Then $\delta n(t)$ is identified by the data, given that $f_t(n^*)$ corresponds to the left limit of the earnings density at the disregard level. In other words, I observe the excess bunching at period t, denoted $\mathcal{B}_t$:

$$\mathcal{B}_t = \frac{1}{f_t(n^*)} \int_{n^*}^{n^*+\delta n(t)} f_t(n)dn \simeq \delta n(t) \tag{10}$$

Using Equation 8, I obtain identification of the earnings elasticity to the net-of-tax rate e. A first-order approximation of Equation 8 yields the following expression for the elasticity:[29]

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

The main difference between the static bunching formula of Saez (2010) and the above expression is the definition of the marginal tax rate. In my setting, the dynamic marginal tax rate depends on the discount factor and the probability to exhaust the initial benefit

[27] 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.

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

[29] 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, we estimate below that on average $\tau_t = 0.6$.

entitlement.[30]

The bunching formula 11 makes clear that the bunching behavior only depends on the remaining benefit duration before exhaustion. This simplifies the aggregation of earnings distributions over time. Let me define $\mathcal{B} = \frac{1}{f(z^*)} \int_t \int_{n^*}^{n^*+\delta n(t)} f(n) dn dG(t)$ where $G(t)$ is the cumulative distribution of remaining UB durations over the claims. Using Equations 10 and 11, I obtain the aggregate bunching formula:

$$e = \frac{\mathcal{B}}{z^* \mathbb{E}[\tau(t)]} \tag{12}$$

where $\mathbb{E}[\tau(t)] = \int_t \tau(t) dG(t)$ is the marginal tax rate that new claimants expect. It is also the population average on all claiming weeks.

3.4 Heterogeneity in bunching

A direct consequence of the bunching formula 11 is that bunching decreases as the claiming duration increases. As the job-seeker approaches the exhaustion date, she is more likely to benefit from the UB transfers and the dynamic marginal tax rate decreases over the spell.[31] This yields the following Proposition:

Proposition 1 (Bunching over the spell) *Bunching decreases over the spell.*

I next consider the predictions of the model when job-seekers differ in their propensity to find a permanent job (heterogeneity in p). Intuitively, 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.[32] I then have the following comparative statics result:

Proposition 2 (Bunching across p-strata) *At any given period t, excess bunching decreases with the probability to remain claiming p.*

In the Appendix, I further discuss the implications of heterogeneity in the permanent job finding rate. I show that, in this context, Proposition 1 is not only valid within a p-strata, but also on average. I also introduce heterogeneity in the structural elasticity (e).

[30] Note that one key condition for identification is that claimants are not myopic. Myopic individuals have no incentives to work above the disregard level. Then bunching is not informative about the earnings elasticity to the net-of-tax rate.

[31] Bunching is time-dependent through the dynamic marginal tax rate, which decreases over time: $\frac{d\tau_t}{dt} = \log(\beta p) \times (\beta p)^{t_{exh}-t-1} < 0$. Note that changes in bunching are larger as the exhaustion date gets closer.

[32] More formally, the dynamic marginal tax rate decreases with p. The exact derivation is given by: $\frac{d\tau_t(p)}{dp} = -(t_{exh} - t) \times (\beta p)^{t_{exh}-t}/p < 0$

I then show that the model still predicts that bunching decreases over the spell within a p-strata. However, bunching does not necessarily decrease with the permanent job finding rate, as the comparison across p-strata may be confounded by different average structural elasticities (of claimants with talent at the lower threshold).

3.5 Stepping-stone/crowding-out effects

I now consider that the probability to find a permanent job depends on earnings in small jobs. When working while on claim increases the future probability to find a permanent job, 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. When there is a kink at the disregard level in the partial-UI schedule, the earnings elasticity to the net-of-tax rate is still identified under two specific assumptions. First, the marginal effect of earnings on the permanent job finding probability $p'(z)$ is continuous. Second, the net gain of permanent jobs $(W - U_{t+1}(B_{t+1}))$ depends continuously on earnings z_t and depends on individual talent only through earnings. Details of the proof are in the Appendix. The second assumption is strong. However, as already discussed, it seems very likely that marginal stepping-stone/crowding-out effects are negligible ($p' << 1$).

3.6 Adjustment costs and matching frictions for low-earnings jobs

Claimants may not be able to find part-time/temporary jobs with earnings that exactly match their desired optimal labor supply. This may be due to firms' constraints in their productive processes and resulting schedules. Alternatively, this may be due to search frictions in the market for small jobs – it takes time for claimants to acquire information about vacancies that fit their labor supply desires. The model can be extended to account for such frictions. For example, I can assume that there is a fixed cost ϕ to adjust from total unemployment to work in small jobs. Totally unemployed claimants only accept jobs that deliver a net gain exceeding the fixed cost, i.e. jobs around their optimal earnings. Such extensions show that optimization frictions typically smooth bunching. Bunching then no longer identifies the structural elasticity e. However it is still informative about the behavioral costs of the partial-UI program.

Frictions also alter the evolution of bunching over the spell. First, they may lead to an increase in bunching during the first claiming weeks. The intuition is as follows. Totally unemployed claimants take the first job that delivers a positive net utility gain. As long as they work in that first job, they will only switch to a new job if it increases

their utility, or equivalently if it is closer to their optimal labor supply. Second, frictions may smooth the decrease in bunching later in the claim (see Proposition 1). Consider a claimant whose current job is close to her notional earnings, i.e. her current optimal earnings if there were no frictions. Notional earnings increase over the spell for claimants working above the disregard level. Claimants facing frictions will increase their earnings if the associated gains exceed the fixed adjustment cost. When claimants are close to their notional earnings, inertia is likely to be strong, as switching gains are typically second-order (the derivative of the utility function with respect to the earnings level is zero at the notional earnings). This is also the case for bunchers, so that debunching is likely to be slow.

4 Estimation and Data

4.1 Estimation

The estimation procedure has two main steps. First, I estimate the excess bunching $\mathcal{B}$, i.e. the numerator of Equation 12. Second, I estimate the expected dynamic marginal tax rate in the denominator of Equation 12.

4.1.1 Excess bunching estimation

Following the procedure of Chetty, Friedman, Olsen and Pistaferri (2011), 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 me define C_j the count of individuals earning between j and $j+1$ dollars above the disregard level (when they earn below the disregard, j is negative). I define Z_j the dollar amount earned by claimants in bin j ($Z_j = j$), centered around the disregard level. I estimate the following equation:

$$C_j \left(1 + \mathbb{1}[j > \overline{R}] \frac{\hat{B}_N}{\sum_{j>\overline{R}} C_j} \right) = \sum_{k=0}^{q} \beta_k (Z_j)^k + \sum_{i=-\underline{R}}^{\overline{R}} \gamma_i \mathbb{1}[Z_j = i] + \epsilon_j \tag{13}$$

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.[33] The order of the polynomial q and the width of the bunching window $(-\underline{R}, \overline{R})$

[33] Because $\hat{B}_N$ depend 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.

are not estimated, but set after visual inspection. I will check below the robustness of the estimation results with respect to those two parameters.

Equation 13 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:

$$\hat{\mathcal{B}} = \frac{\hat{B}_N}{\sum_{j=-\underline{R}}^{\bar{R}} \hat{C}_j / (\underline{R} + \overline{R} + 1)} \tag{14}$$

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

4.1.2 Dynamic marginal tax rate

Let me recall that the dynamic marginal tax rate for a claimant working just above the disregard in period t, is defined as:

$$\tau_t = 1 - (\beta p)^{t_{exh}^{Utot} - t} \tag{15}$$

I first calibrate the weekly discount factor β so that the annual interest rate equals 4%. The calibrated discount factor is very close to one. I then compute for each individual the potential benefit duration under total unemployment: t_{exh}^{Utot}. Thirdly, 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 $h_i = h \exp(\beta X_i)$. The details of the hazard model are reported in the Appendix. I then compute the predicted survival rate $\hat{p}_i$ for each individual. Note that, by using predicted rates, claimants are assumed to have rational expectations about their compensated unemployment duration.[34]

Finally, I obtain an estimate of the denominator of Equation 12 by averaging, over all individuals and weeks, the predicted dynamic marginal tax rate and by multiplying this average by the average of observed disregards. The standard errors of the elasticity estimate are obtained by the delta method.

4.2 Data

I use individual panel data from the Continuous Wage and Benefit History project (CWBH). The project collected weekly claims for a random subsample of UI claimants in the U.S.,

[34]Unobserved heterogeneity in the survival rate would bias the elasticity estimate, if it is correlated with unobserved heterogeneity in the elasticity.

and the resulting data has the unique advantage of including the weekly earnings that claimants report to the UI administration and the consecutive UB payments. I can thus characterize whether claimants are partially unemployed. The major drawback of the data set, that it covers the late 70s and early 80s, is mitigated by the fact that the partial-UI schedules have been hardly reformed since then. The data cover four U.S. states – Idaho, Louisiana, Missouri and New Mexico – and include all relevant information about the claim: weekly benefit amount, total entitlement, and highest quarter earnings. Socio-demographics characteristics are also available (gender, age, education, ethnicity, past firm status and industry, past occupation). In addition, the data set includes survey information about recall expectations for a subsample of claimants.

Table 2 reports descriptive statistics of UI claimants by state. The weekly benefit amount is around \$100 (current dollars) and the average replacement rate is between 40% and 50%. The potential benefit duration (PBD) is actually greater than 26 weeks (the maximal PBD in Tier 1) as the early 80s was a period of high unemployment, and UI extensions were triggered. The average claiming duration is around four months (excluding the waiting week). The last line of Table 2 reports the share of claimed weeks with positive reported earnings. The share of partial unemployment weeks amounts to 17.6% in Idaho where the partial-UI schedule is very generous: claimants can be partially unemployed up to 1.5 times their WBA. In Louisiana and Missouri, respectively 6.1% and 8.1% of claimed weeks concern partially unemployed claimants. The corresponding share in New Mexico is low at 2.5%, reflecting partly the fact that the disregard amount is quite low in this state.

Table 3 reports descriptive statistics for claimants in Louisiana from 1980 to 1982. It compares claimants with at least one week of partial unemployment during the benefit year in Column 1 to claimants always on total unemployment in Column 2. Column 3 reports the p-value of the test of equality between the two first columns. Partial-UI claimants are older, more educated and more frequently whites. Their previous employer is more frequently a firm from the private sector and operates in construction and manufacturing industries. Partial-UI claimants are less frequently in upper occupations, such as professionals, technicians or managers. Their pre-unemployment wage and their entitlement duration are higher.

5 Main results

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

Figure 2 displays the weekly earnings density reported by UI claimants together with the empirical partial-UI schedule for the four different U.S. states. In line with the partial-UI rules in Idaho, Louisiana and New Mexico, I normalize the weekly earnings by the weekly benefit amount (unemployment benefits in case of total unemployment). The empirical schedules, 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 clearly display bunching at the level of the disregard (50% of the weekly benefit amount). On the upper right Louisiana 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 obvious (lower left panel). The lower right panel illustrates a placebo test. In Missouri, the level of disregard is so low at a mere \$10 that the schedule is totally flat at the $0.5 \times WBA$ threshold. There is indeed no bunching at this placebo level. Thus the bunching observed in Idaho or Louisiana is unlikely to be an artifact of other labor legislations or norms, or hour constraints according to which claimants take some part-time jobs that provide roughly one fourth of their previous wages (given that the average replacement rate is around 50%).

Another important feature of the earnings distribution in Figure 1 is the substantial fraction of claimants working for earnings above the disregard level. This observation is consistent with claimants reacting to the dynamic marginal tax rate, rather than to the static 100% benefit-reduction rate. Indeed, myopic claimants have no incentive to work above the disregard level. However this observation is not a definitive test of the dynamic aspects of my model, as matching frictions for small jobs could also explain why myopic claimants work for earnings above the disregard.[35]

To conduct the bunching estimation, I consider earnings in absolute levels and I center the earnings density at the disregard level. Note that because the disregard is specified as a fraction of weekly benefit amounts, individuals with different weekly benefit amounts are subject to different disregard levels. The resulting centered earnings densities are plotted in Figure 3. Observations are grouped in bins of 1 dollar width. Figure 3 confirms the

[35]Of course, claimants, who are not aware of the partial UI schedule, would also work above the disregard.

bunching patterns observed in Figure 2. It also reveals some periodicity in the earnings distribution in Missouri. Claimants report earnings that are multiples of ten dollars.[36] Each panel also displays the counterfactual density in red, which is estimated along the lines of Chetty, Friedman, Olsen and Pistaferri (2011).[37]

Table 4 reports the results of the estimation for each state (in columns). 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, excess bunching is much lower and not statistically significant. In Missouri, there seems to be a missing mass of claimants at the threshold level. As a consequence, the placebo test confirms that in the absence of a kink there is no excess mass at the threshold level. The periodicity in the earnings density in Missouri may bias bunching estimates, especially if there are heaps in the window where bunching is expected. I verify that the bunching estimate does not change if I modify the earnings density by smoothing the heaping points.

To compute the dynamic marginal tax rate, I need to estimate the expected exhaustion probability in the sample. I estimate an exponential model of the hazard rate out of the unemployment registers. I include in this model various characteristics of claimants (gender, age, education and ethnicity), claim characteristics (WBA, PBD, recall expectations) and year fixed effects. Detailed estimation results are reported in the Appendix. The weekly hazard rates vary between 3% and 4% across states. Predicted hazard rates are then used to compute the expected UB exhaustion probability at each date, which takes into account the remaining number of entitled weeks. The average dynamic marginal tax rate is around 54% in Idaho and Louisiana; it is larger in New Mexico, where it amounts to 60%.

Finally, I use the identification relation to compute the earned income elasticities to the net-of-tax rate. I obtain statistically significant elasticities in Idaho and Louisiana, respectively 0.19 and 0.13. The elasticity in Missouri 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 window and the polynomial degree in the estimation procedure.[38] When I do not use first-order approximation of the logarithm of the net-of-tax rate, elasticity estimates are slightly lower, but still around 0.1.[39] My results are broadly consistent with the estimates of the intensive labor supply elasticity found in previous micro empirical

[36]The periodicity appears only in Missouri because it has a large mass of individuals at the maximal WBA. Then the disregard is the same for a large fraction of claimants.

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

[38]Robustness results are reported in the Supplementary Table 7 in the Appendix

[39]Robustness results are reported in the Supplementary Table 8 in the Appendix.

work (see the review of quasi-experimental estimates in Chetty (2012) or Chetty, Guren, Manoli and Weber (2011)). Note that I cannot be certain that this elasticity is purely driven by behavioral responses from the supply side of the labor market. In the next section, I present another placebo test that indicates that bunching is actually related to the partial-UI schedule.

I can compare my elasticities to estimates obtained specifically with bunching estimators. Most recent papers (Saez, 2010; Chetty et al., 2013; Chetty, Friedman, Olsen and Pistaferri, 2011; le Maire and Schjerning, 2013; Bastani and Selin, 2014) find bunching at the kinks of the annual income tax schedule among the self-employed, but very little bunching among wage-earners (especially among those at the bottom of the wage distribution, as in my case). My sample comprises former wage-earners – indeed, this is an eligibility condition for claiming. While I cannot rule out that some of them are actually self-employed while on claim, it is very likely that former wage-earners also work as wage-earners while on claim. This comparison suggests that my elasticity estimate is higher than most results from the bunching literature, with the notable exception of Gelber et al. (2013). Indeed, those authors 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.

5.2 Difference-in-difference in Louisiana

Louisiana's rules regarding unemployment insurance changed in April 1983. The change in partial UI affected both the stock of individuals registered in April 1983 and new inflows after that point in time.[40] 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.

[40]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.

Figure 4 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 5, 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 really direct their search to claimants with certain disregard levels, it is likely that they would post the most common disregard (Chetty, Friedman, Olsen and Pistaferri, 2011). In Louisiana, the mode of 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.

6 Bunching heterogeneity

In this section, I test predictions of the theoretical model regarding heterogeneity in bunching. To maximize statistical power, I jointly analyze partial UI in Idaho and Louisiana (before the reform in April 1983). Indeed, both states share the same disregard level ($0.5 \times WBA$).

6.1 Potential benefit duration

My theoretical model predicts that when total UB entitlement is more generous, bunching increases. The intuition is as follows. Let me consider a baseline claimant (X) with an initial UB entitlement B_0 yielding t_{exh}^{Utot} weeks of unemployment benefits. Her earnings

along the claim write $(z_0, z_1, ...)$. Now let me consider an identical claimant (Y) with UB capital enhanced by b. Given that the model depends on the past only through the UB capital, claimant (Y) behaves as claimant (X) with some period lags. Thus claimant (Y) has a higher marginal tax rate during her first claiming weeks. As a consequence, claimant (Y) is more likely to bunch. This prediction is a corollary of Proposition 1. As bunching decreases with time until the exhaustion date, individuals with longer potential benefit duration are more likely to bunch. In Idaho and Louisiana, the potential benefit duration in tier I varies from 10 to 28 weeks, in relation to past work history. Figure 6 shows that bunching is significantly greater when claimants have longer potential benefit durations. 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 Katz and Meyer (1990*a*) for an early contribution or Lalive et al. (2006) for evidence based on regression discontinuity design). Higher survival rates tend to decrease bunching (see Proposition 2), so that Figure 6 underestimates the positive relation between bunching and potential benefit duration.

6.2 Recall expectations

Job-seekers who expect to be recalled to their previous employer have different job-search behaviors than non-expecting claimants. Katz and Meyer (1990*b*) show that their unemployment duration is shorter, i.e. they have a lower probability to remain claimants. According to Proposition 2, claimants expecting recalls would then bunch more than non-expecting claimants. Indeed, expecting claimants have a lower probability to benefit from partial-UI transfers, as they may be recalled by their previous employer even before their calendar exhaustion date (while totally unemployed). In the joint Idaho-Louisiana sample, 65% of partial-UI claimants expect to be recalled to some previous employers.[41] Figure 7 shows that expecting claimants bunch significantly more. The bunching mass at the disregard level is 50% larger. I next propose a more systematic test of Proposition 2.

6.3 Survival rate

Proposition 2 states that bunching decreases with the expected survival rate (or increases with the permanent job finding rate). Figure 8 compares bunching across the quartiles of the predicted survival rate distribution. In both states, bunching tends to decrease from the second to the fourth quartile, confirming Proposition 2. Though the differences are not statistically significant (at the 5% level), their magnitude is important. From

[41] Recall expectations were obtained through a specific survey. As a result, recall expectations are missing for almost 50% of claimants.

the second to the fourth quartile, bunching is reduced by one third. The evidence is less clear at the bottom of the survival rate distribution: bunching increases from the first to the second quartile. As above, the comparison across quartiles may be confounded by other factors. Namely, individuals with high survival rate are expected to have on average less remaining benefits (under the assumption that the correlation between the initial potential benefit duration and the survival rate is negligible).

6.4 Bunching over the claim

As described in the theoretical section, two dimensions affect the dynamic marginal tax rate and thus the amount of bunching: weekly survival rates and time to benefit exhaustion. The previous heterogeneity analysis compared bunching along one dimension without holding constant the second dimension. I now take advantage of the panel structure of the data to control for heterogeneity in weekly survival rates. I first select claimants who have worked at least eight weeks while on claim. This cutoff corresponds to the third quartile of the distribution of the number of partial-UI weeks while on claim.[42] I then explore the evolution of bunching over those first eight weeks on partial UI. I further select the sample by excluding claimants expecting to be recalled to their previous employer. Predictions of the evolution of bunching over the claim are less clear for this excluded sample, as they may have a definite recall date.[43] Figure 9 shows that there is a steep increase in bunching between the first and second week of partial UI. This can be explained by learning effects (or alternatively by frictions). Figure 9 also shows that bunching estimates do not evolve after the second week of partial UI. I do not find any decreasing pattern in bunching over the spell. While the sample is held fixed across the different bars of Figure 9, each subgroup pulls together claimants with different horizon until exhaustion (subgroups are defined by the rank of the partial-UI week within the claim). Consequently, changes in horizon vary across subgroups, which may blur decreasing patterns. Moreover, the sample analyzed in Figure 9 may be composed of claimants at the beginning of their claim with a long horizon. This results in small changes in dynamic marginal tax rate over time, which makes it difficult to detect changes in bunching behavior.[44] Indeed, increasing the horizon by one week has a stronger effect on the dynamic marginal tax rate close to the exhaustion date then early in the claim. I thus propose another heterogeneity analysis that focuses on claimants later on the spell.

[42] See Supplementary Figure 13 in the Appendix.

[43] When claimants have a definite recall date, they have no uncertainty about finding a permanent job and their behavior is different from the predictions of the theoretical model. For this group of claimants, the dynamic marginal tax rate does not evolve over the spell (given that the discount factor is one).

[44] Supplementary Figure 14 in the Appendix show that the change in dynamic marginal tax rate between two consecutive weeks is around -2 percentage points.

I compute, for each claiming week, the remaining UB entitlement and the corresponding time to benefit exhaustion (under the assumption of total unemployment). The corresponding time to benefit exhaustion can be described as the current potential benefit duration. In the first week of claim, the current potential benefit duration is equal to the initial potential benefit duration (see section 6.1 above). In Figure 10, I select claimants who worked both the month before their exhaustion date and earlier in the claim (when they had at least five months of benefits left in their UB capital). When claimants are far from the exhaustion date, bunching tends to increase while the current potential benefit duration decreases. When there is only one month of benefits left, claimants tend to de-bunch. Confidence intervals (at the 95% level) around bunching estimates are as wide as 2. With bunching estimates around 5, this makes the comparisons not very powerful. Consequently, while the broad picture drawn by parameter estimates is consistent with the theoretical model amended with learning effects, I cannot formally reject that claimants' bunching behavior is flat over the spell.

7 Optimal benefit-reduction rate

This section is devoted to normative considerations. Based on the theoretical model and on the estimate of the earned income elasticity, I propose modifications to the partial-UI schedule to maximize claimants' welfare subject to a UI administration revenue constraint. I consider the following formal program:

$$SWF = \max_{T(.)} \int_n \omega_n U_0(B_0, n, b, T(.)) dF(n)$$
$$\text{such that} \int_n C(B_0, n, b, T) dF(n) < \underline{C}$$

where ω_n are Pareto weights. Let me recall that $U_0(B_0, n, b, T)$ is the expected utility of a new claimant with talent n and total initial entitlement B_0, when the UI administration chooses the level of unemployment benefits b (in case of total unemployment) and the partial-UI schedule $T(.)$. U_0 is computed according to the baseline model (without stepping-stone/crowding-out effects). Note that I assume no heterogeneity in the initial entitlement B_0. Let me denote $C(B_0, n, b, T)$ the expected benefits paid by the UI administration to a new claimant (n, B_0). The UI administration budget constraint states that the aggregate expected benefit payments cannot exceed an exogenous upper threshold $\underline{C}$.

In the above program, the UI administration maximizes claimants' welfare modifying the partial-UI schedule, holding the benefit level constant. Indeed, my focus here is on the optimal partial-UI schedule. I do not consider how the UI administration sets the

benefit level b, and I solve for the optimal partial schedule conditional on b. Note that the program above differs from the standard optimal UI program in several dimensions. First, I consider, as in the baseline model, that utility is quasi-linear. I am not interested in the usual welfare effect due to consumption smoothing. Second, the UI administration does not internalize the increase in its revenue due to the taxes levied on permanent jobs. Indeed, in the baseline model without stepping-stone/crowding-out effects, working while on claim has no effect on the probability to find a permanent job. Third, there is no unobserved job-search effort driving the UI efficiency cost. However, there is adverse selection associated to the claimants' labor supply in low-earnings jobs.

To design the optimal partial-UI program, I vary the (static) benefit-reduction rate (T).[45] I thus define the modified partial-UI schedule:

$$T(z) = \begin{cases} 0 & \text{if } z < z^* \\ T(z - z^*) & \text{if } z \in (z^*, z^* + b) \\ b & \text{if } z > z^* + b \end{cases} \tag{16}$$

In the current schedule, the static marginal benefit-reduction rate is 100% ($T = 1$). The expected cost for the administration of a UI claim with initial entitlement B_0 depends on talent and equals:

$$C(B_0, n, b, T) = \begin{cases} \Sigma_{t=0}^{t_{exh}-1}(\beta p)^t b & \text{when } n < n^* \\ \Sigma_{t=0}^{t_{debunch}-1}(\beta p)^t b + \Sigma_{t=t_{debunch}}^{t_{exh}-1}(\beta p)^t (b - T(z - z^*)) & \text{when } n > n^* \end{cases} \tag{17}$$

where I denote $t_{debunch}$ the debunching date, i.e. the first date when claimants work strictly above the disregard level.[46] As in the derivation of the baseline model, I neglect the costs associated with claimants with high talents (who exit the partial-UI program before benefits lapse). I also assume that the UI administration shares the same discount factor as claimants and the same expectation about the permanent job finding probability (p).

I consider the effect of a small increase in the benefit-reduction rate $dT > 0$. This manipulation does not affect benefits paid to claimants with low talent. Let me consider an individual with talent n such that $t_{debunch} < t_{exh}$. First, the manipulation generates a

[45] Alternatively, I could have modified the disregard level or the maximal earnings (eligible for partial UI). However, to study those modifications, I would need to predict the earnings elasticity away from the actual disregard. This is left for future work.

[46] The debunching date can be zero if claimants do not even bunch in the first week of claim.

mechanical decrease in benefits paid:

$$dM = \Sigma_{t=t_{debunch}}^{t_{exh}-1}(\beta p)^t dT(z_t - z^*) \tag{18}$$

Second, the decrease in benefits paid creates a social welfare loss. According to the usual argument (envelope theorem), the social welfare loss is only due to the decrease in net income (note also that utility is quasi-linear). The corresponding social welfare loss (in terms of UI administration funds) equals:

$$dW = -\omega_n/\lambda dM \tag{19}$$

where λ is the Lagrange multiplier associated to the UI administration budget constraint and ω_n is the Pareto weight of claimants with talent n. Third, the manipulation triggers a decrease in total earnings due to behavioral response of the claimant. The corresponding decrease in UI administration revenue is:[47]

$$\begin{aligned} dB &= \Sigma_{t=t_{debunch}}^{t_{exh}-1}(\beta p)^t T z_t \frac{d\ln z_t}{d\ln(1-T\tau_t)} d\ln(1-T\tau_t) &(20)\\ &= -\Sigma_{t=t_{debunch}}^{t_{exh}-1}(\beta p)^t z_t e \frac{T\tau_t}{1-T\tau_t} dT &(21)\end{aligned}$$

The overall effect of the manipulation on claimant with talent n equals:

$$\Sigma_{t=t_{debunch}}^{t_{exh}-1}(\beta p)^t (z_t - z^*)\left(1 - \omega/\lambda - a_t e \frac{T\tau_t}{1-T\tau_t}\right) dT \tag{22}$$

where a_t is defined as $a_t = \frac{z_t}{z_t - z^*}$. The above formula abstracts from changes in $(t_{debunch}, t_{exh})$, which are second order. To obtain the overall effect of the manipulation, I aggregate the expression 22 over all claimants with earnings above the disregard (at least in one period). The corresponding set of claimants, denoted $\mathcal{Q}$, represents a fraction q of the weekly claims. For the sake of simplicity, I assume that the Pareto weights ω are constant in the population above the disregard and that the discount factor is equal to one. I also assume, as in the baseline model, that there is no heterogeneity in the elasticity (e), nor in the permanent-job finding probability $(1-p)$. The overall effect then equals:

$$dSWF = \left[(\overline{z} - z^*)(1 - \omega/\lambda) - e\left(\int_{n\in\mathcal{Q}} \Sigma_{t=t_{debunch}(n)}^{t_{exh}(n)-1} p^t z_t(n) \frac{T\tau_t}{1-T\tau_t} dF(n)/q\right)\right] q dT \tag{23}$$

where $\overline{z}$ is the average earnings in the population of claimants with benefit reduction.

[47] I adjust the FOC of the baseline model to account for T: $1 - T\tau_t = \left(\frac{z_t}{n}\right)^{1/e}$. I obtain that the behavioral response is: $\frac{d\ln z_t}{d(1-T\tau_t)} = e$.

The welfare-maximizing schedule is obtained when Equation 23 is set to zero. In general, there does not exist a constant static benefit-reduction rate T, solution to Equation 23. It seems that the optimal static benefit-reduction rates depend on horizon $t_{exh} - t - 1$ (through τ_t). Solving analytically for the optimal time-varying benefit-reduction rate is left for future research.

To draw normative considerations, I perform numerical simulations of the program with constant benefit-reduction rates. I consider a representative claimant earning above the disregard and I search for the cost-minimizing benefit-reduction rate (this corresponds to the welfare-maximizing rate when Pareto weights are zero for partial-UI claimants). I set the different elements of Equation 17 at their average values observed (or estimated) in Idaho: $p = 0.96$ and $e = 0.2$. I consider a claimant entitled to 26 weeks of total UB with weekly benefit amount equal to \$100 (and discount factor equal to one). Her talent is normalized to 100 (this corresponds to her potential earnings in part-time/temporary jobs in a world without benefit reduction, $T = 0$). In the actual schedule ($T = 1$), the claimant earns \$73 in the first week of her spell. Her earnings increase over the spell, up to \$99 in the exhaustion week (40 weeks after the first claim in this example). The left panel in Figure 11 shows the expected cost born by the UI administration (C) for different static benefit-reduction rates. The expected cost in the actual schedule is around \$1,376. This cost can be decreased by \$25 (almost 2%) when the static benefit-reduction rate is set down to 80%. The right panel in Figure 11 shows that claimants' welfare could be improved by further decreasing the benefit-reduction rate. Whether the UI administration is willing to do so depends on the social marginal welfare weight ω/λ. For example, decreasing the benefit-reduction rate from 80% to 50% decreases expected UI revenue by \$50, and increases claimants' welfare by \$125. This would be optimal if the marginal welfare weight is greater than 2.

The above exercise calls for several comments. First, the optimal schedule depends on the wedges τ_t which in turn depend on the permanent job finding probability and the discount factor. This means that the optimal schedule may depend on aggregate labor market conditions, which could make its implementation difficult in practice. Second, the existence of stepping-stone/crowding-out effects influence the optimal schedule. Suppose that earnings increase the probability of exiting the UI registers; then the UI administration has an extra motive to reduce benefit-reduction rates. The above exercise does not take this into account.

8 Conclusion

In this paper, I study claimants' behavioral response to the rules of partial unemployment insurance in the U.S. I observe that claimants bunch at the kinks of the partial-UI schedule, and to interpret this bunching mass, I build a dynamic model of working while on claim. The model highlights that forward-looking claimants react to the dynamic marginal tax rate of the partial-UI schedule, which is lower than the static benefit-reduction rate. The wedge between the two rates is primarily due to benefit transfers. Claimants internalize that the reduction in current benefits leads to an increase in future benefits. I extend the approach of Saez (2010) to account for this mechanism and I estimate that the earnings elasticity to the net-of-tax rate lies between 0.1 and 0.2. Simulations of the model, calibrated with these estimates, suggest that decreasing the static benefit-reduction rate from 100% to 80% would decrease overall benefits paid to partial claimants by the UI administration and still improve the claimants' welfare.

Empirical evidence on bunching heterogeneity is broadly consistent with the predictions of the dynamic model. The theoretical model predicts (i) that claimants with higher survival probability should bunch less and (ii) that claimants should bunch less and less over time within their claim. While I find empirical support for the first prediction, the decrease in bunching over the spell is not as steep as the dynamic model predicts. Moreover, I find an increase in bunching at the beginning of the spell. This could be the result of adjustment costs/frictions in the search for low-earnings jobs or the consequences of claimants progressively learning the partial-UI schedule. Understanding the respective roles of those two mechanisms is a promising direction for future work.

My normative exercise provides insights on the efficiency of the partial-UI schedule and guidelines to policymakers. It focuses on modifications of the static benefit-reduction rate, a parameter of great interest. In France, the benefit-reduction rate was extensively discussed during the 2014 UI reform. Policymakers could also be interested in the effect of other parameters of the schedule, such as the disregard level. My work could be extended to study such modifications. It could also be extended, in a more theoretical direction, to derive the optimal flexible partial-UI schedule.

Finally, 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. For example in the U.S., a reduction in current pension benefits increases future benefits (SSAET). Gathering evidence on such schemes would help us understand more broadly how individuals make intertemporal decisions under uncertainty.

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9 Tables and Figures

Table 1: 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."

Table 2: Descriptive Statistics

	Idaho	Louisiana	New Mexico	Missouri
Years	76-84	79-84	80-84	78-84
Inflow (nb)	91,162	95,675	62,030	78,065
Pre-U weekly wage (current dollars)	337	316	265	247
Weekly benefit amount (WBA)	96	131	94	88
Replacement rate	.425	.471	.405	.455
Potential duration (weeks)	28	35.8	34.7	32.7
Actual UB duration	15.3	18.1	15.0	15.6
Share of partial-UI weeks	.176	.067	.025	.085

Source: CWBH. Notes: Means are computed over the sample of claimants, except the share of partial UI computed over all claiming weeks.

Table 3: Selection into partial UI (Louisiana 1980-82)

	At least one week of partial UI	No partial UI	Diff- test p-value
Male	0.70	0.71	0.38
Age	34.90	33.98	0.00
Education (years)	11.39	11.26	0.00
White	0.62	0.60	0.00
Private firm	0.97	0.95	0.00
Industries			
Construction	0.32	0.31	0.03
Manufacturing	0.23	0.19	0.00
Trade	0.10	0.14	0.00
Services	0.14	0.17	0.00
Occupations			
Prof., tech. and managers	0.07	0.09	0.00
Clerical and sales	0.12	0.15	0.00
Structural work	0.31	0.30	0.03
Pre-U weekly wage (current dollars)	352.18	315.34	0.00
Weekly benefit amount (WBA)	144.47	130.58	0.00
Replacement rate	0.46	0.47	0.00
Potential duration (weeks)	38.99	37.28	0.00
Actual UB duration	24.06	20.18	0.00
Inflow (nb)	13,174	37,139	

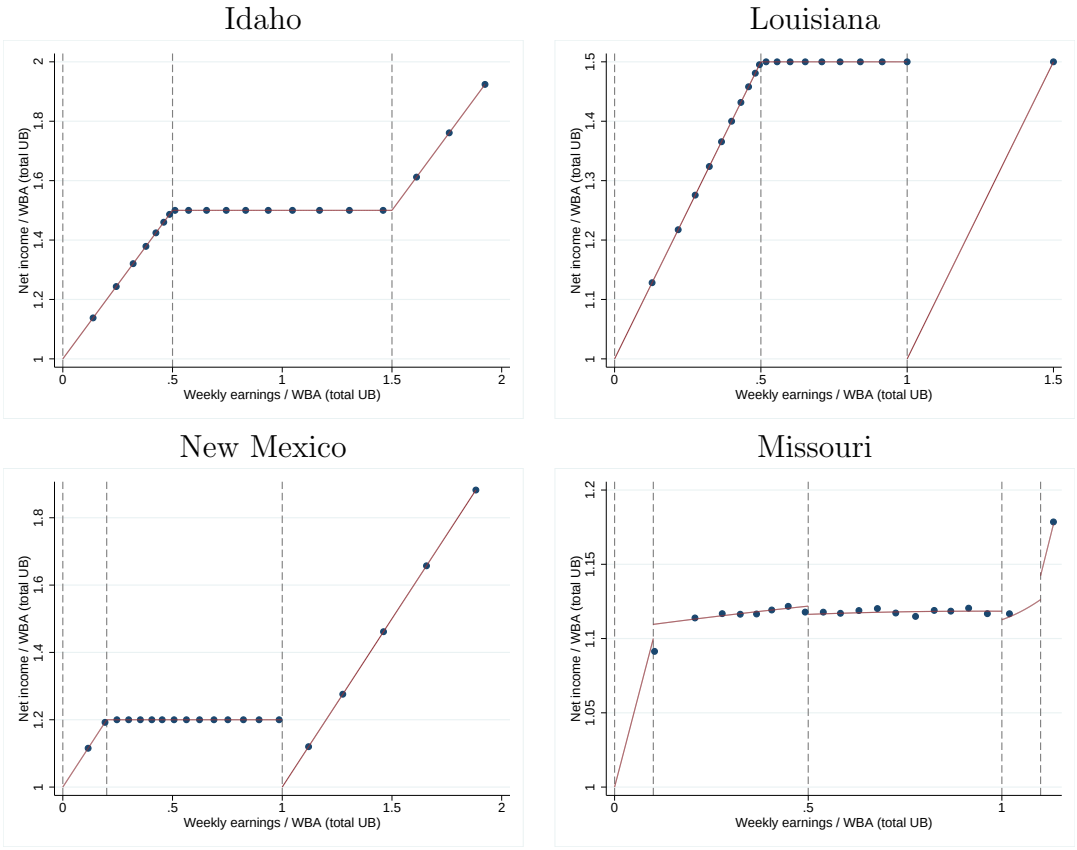
Source: CWBH. Notes: Agriculture, Mining, Transportation and the FIRE industries (Finance, Insurance and Real Estate) are not reported for the sake of space (around 20% of the sample). Occupations corresponds to the standard DOT (Dictionary of Occupation Titles). I only report the most common occupations and exclude service, agricultural, processing, machine trades and benchwork occupations from the table.

Table 4: Bunching, dynamic marginal tax rates and earnings elasticity estimates.

	Idaho	Louisiana	New Mexico	Missouri
Years	76-84	79-83	80-84	78-84
Partial-UI weeks (nb)	230,535	77,602	31,103	91,147
Disregard/kink level (z^*)	\$53	\$64	\$21	\$45
z^* as a fraction of WBA	0.5	0.5	0.2	0.5
				placebo
Excess mass ($\mathcal{B}$)	5.334	4.814	1.247	-.7820
	(.2417)	(.279)	(.8757)	(.3395)
Hazard rate ($1-p$)	.042	.033	.039	
Implicit MTR (τ)	.538	.554	.606	
Earnings elasticity	.187	.134	0.096	
to net-of-tax rate (e)	(.0099)	(.0073)	(.0594)	

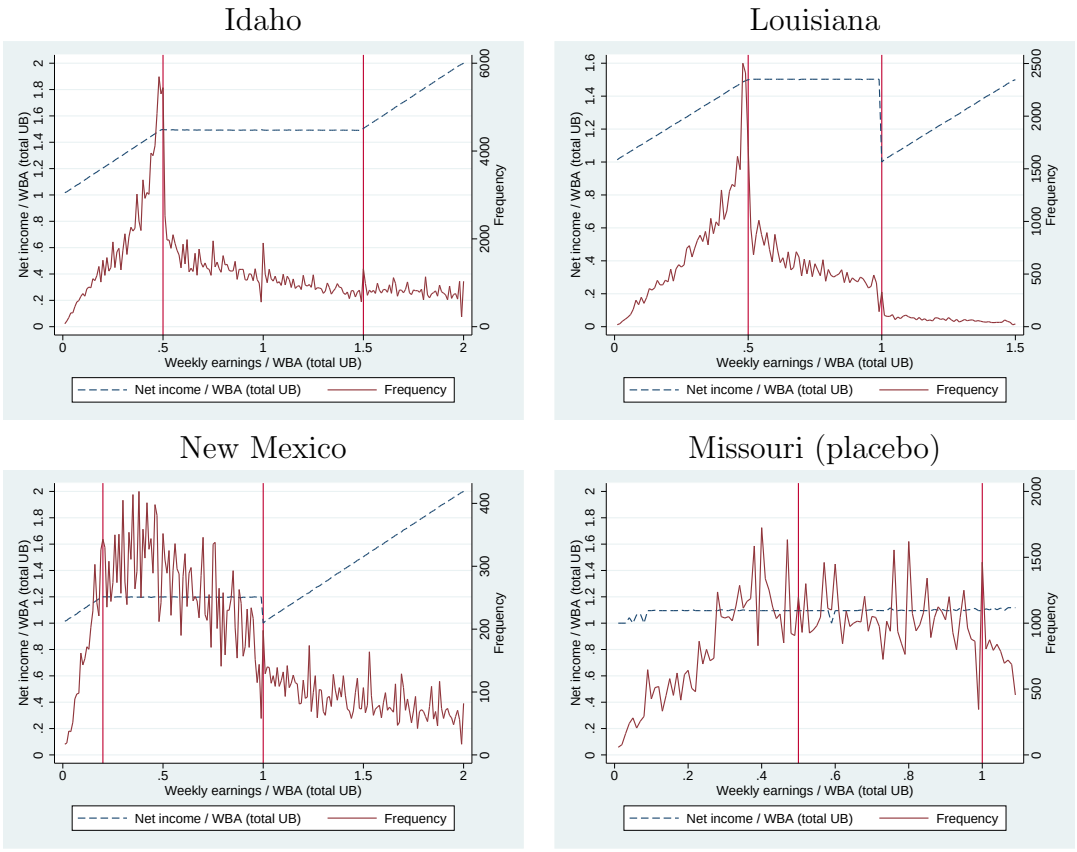
Source: CWBH. Notes: Standard errors are in parentheses below estimates.

Figure 1: Partial-UI schedules from 1976 to 1984



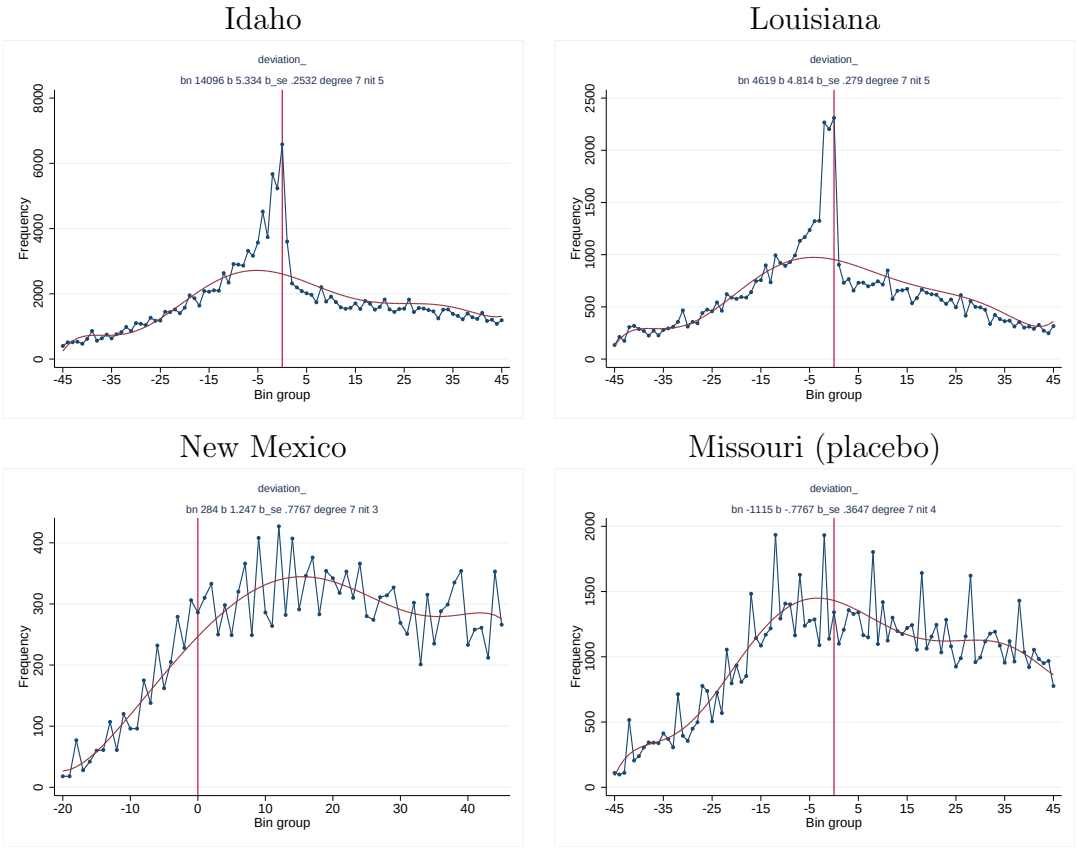
Source: U.S. Department of Labor, "Significant Provisions of State Unemployment Insurance Laws." Notes: X-axis corresponds to weekly earnings divided by the weekly benefit amount (UB paid in case of total unemployment). Y-axis corresponds to the net income (earnings + UB payments) divided by the WBA. The different panels show the theoretical schedules. The graph is noisier for Missouri, because weekly earnings and net income are divided by the WBA, whereas the theoretical schedule is set in dollar values. Red vertical lines show the kinks and notches of the partial UI schedule, except the line at 0.5 in Missouri (the placebo test).

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



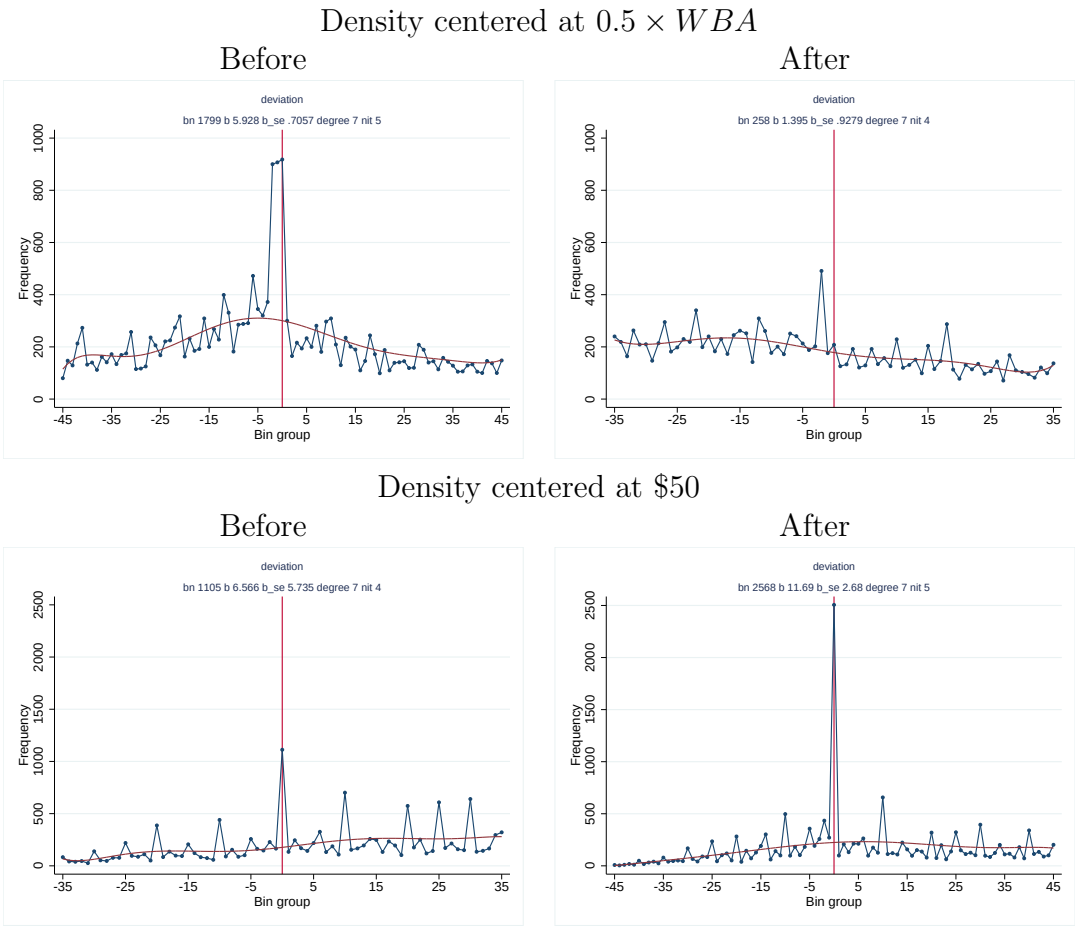
Source: CWBH. Notes: X-axis corresponds to weekly earnings divided by the weekly benefit amount (UB paid in case of total unemployment). Earnings density is plotted in red. Corresponding frequencies are on the right Y-axis. The partial-UI schedule corresponds to the blue dashed line. It plots the net income divided by the WBA. Red vertical lines show the kinks and notches of the partial UI schedule, except for Missouri (the placebo state).

Figure 3: Centered weekly earnings density of partial-UI claimants.



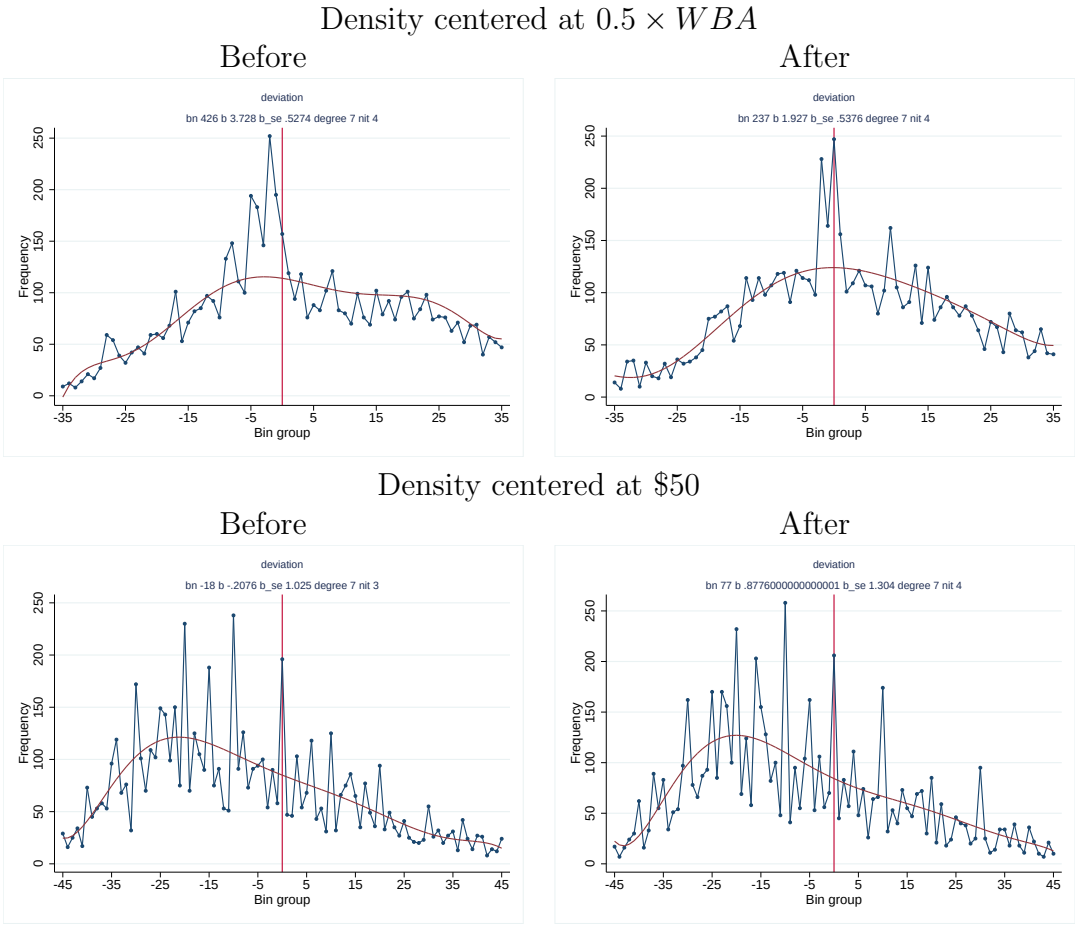
Source: CWBH. Notes: Earnings are in dollars centered at the disregard level. Empirical earnings density in blue. Counterfactual density in red.

Figure 4: Centered weekly earnings density of partial-UI claimants in the treatment group.



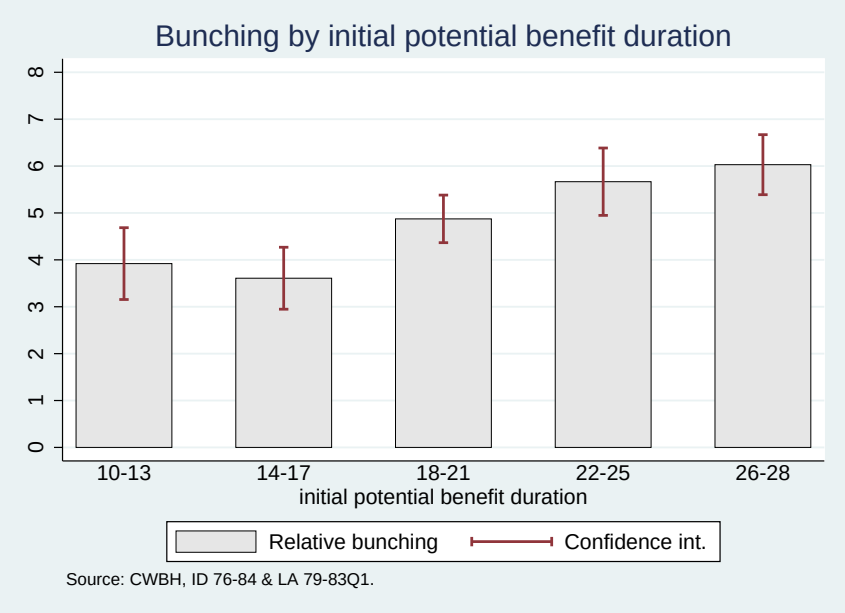
Source: CWBH. Notes: Earnings are in dollars. Empirical earnings density in blue. Counterfactual density in red.

Figure 5: Centered weekly earnings density of partial-UI claimants in the control group.



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

Figure 6: *Bunching by initial potential benefit duration*



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

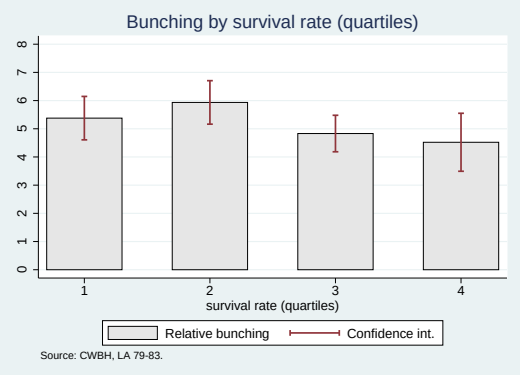
Figure 7: *Bunching by recall expectation*



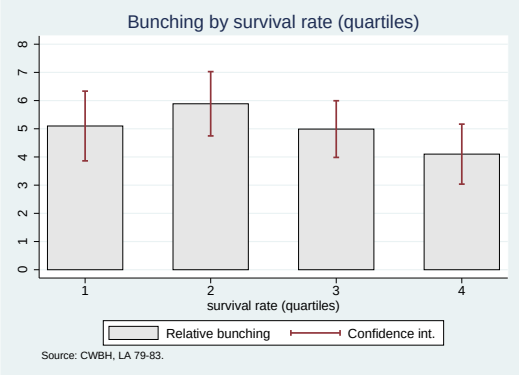
Source: CWBH for Idaho 1976-84 and Louisiana 1979-83, Q1. Notes: Excess mass at the disregard amount (kink) for claimants expecting to be recalled to their previous employer (left bar) and for claimants not expecting any recalls (right bar). Confidence interval at the 95% level in red.

Figure 8: Bunching by survival rate.

Idaho

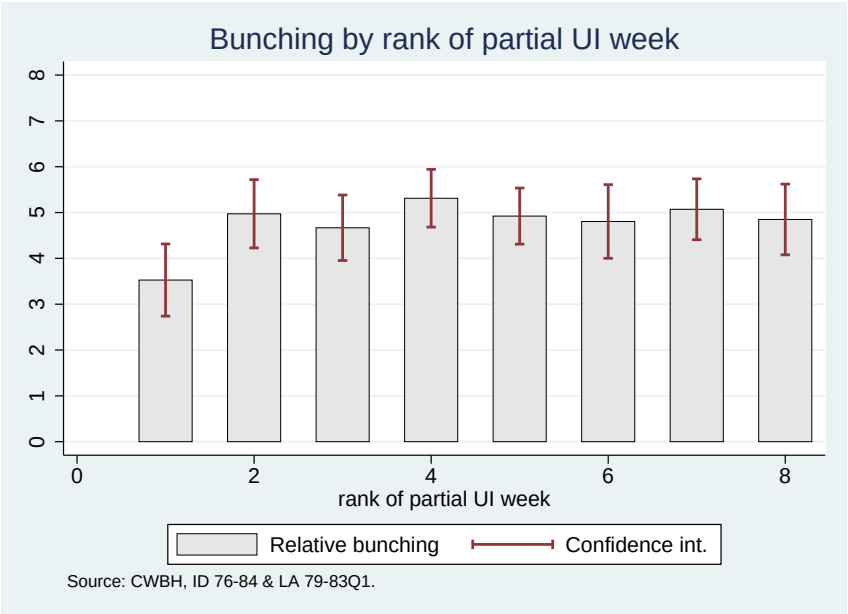


Louisiana



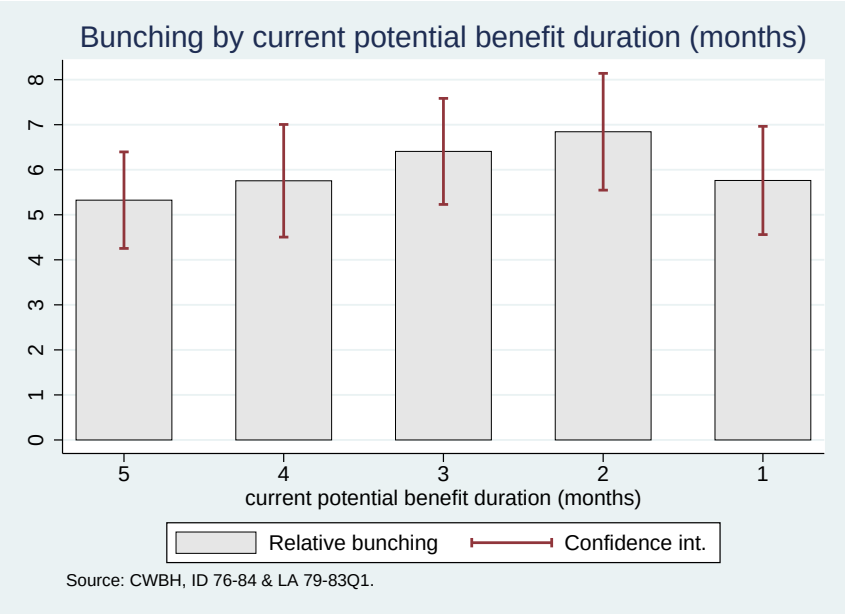
Source: CWBH for Idaho 1976-84 and Louisiana 1979-83, Q1. Notes: Excess mass at the disregard amount (kink) by quartile of predicted survival rates. Idaho in the left panel, Louisiana in the right panel. Confidence interval at the 95% level in red.

Figure 9: *Bunching by rank of the partial-UI week*



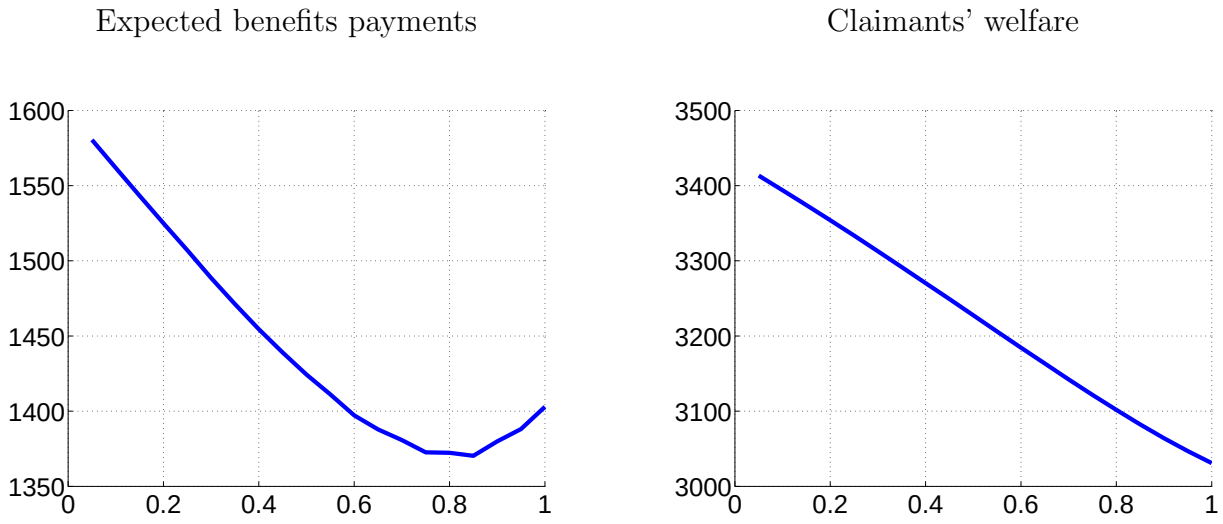
Source: CWBH for Idaho 1976-84 and Louisiana 1979-83, Q1. Notes: For every individual, I rank, within her claim, their weeks with positive earnings (partial UI). I select claimants with at least eight weeks with positive earnings. I estimate the excess mass at the disregard amount (kink) for each of those first eight weeks of partial UI. Confidence interval at the 95% level in red.

Figure 10: *Bunching by current potential benefit duration*



Source: CWBH for Idaho 1976-84 and Louisiana 1979-83, Q1. Notes: For every individual and period, I compute her current potential benefit duration. I estimate the excess mass at the disregard amount (kink) at different levels of current PBD. Confidence interval at the 95% level in red.

Figure 11: *Expected benefits payments and claimants' welfare by static benefit-reduction rates*



Note: X-axis corresponds to different static benefit-reduction rates (from 10% to 130%). Y-axis corresponds to the expected total amount of benefits paid to the claimant over her claim (left panel) and to the expected utility of the claimant over the benefit year.

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 nonmonetary 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.

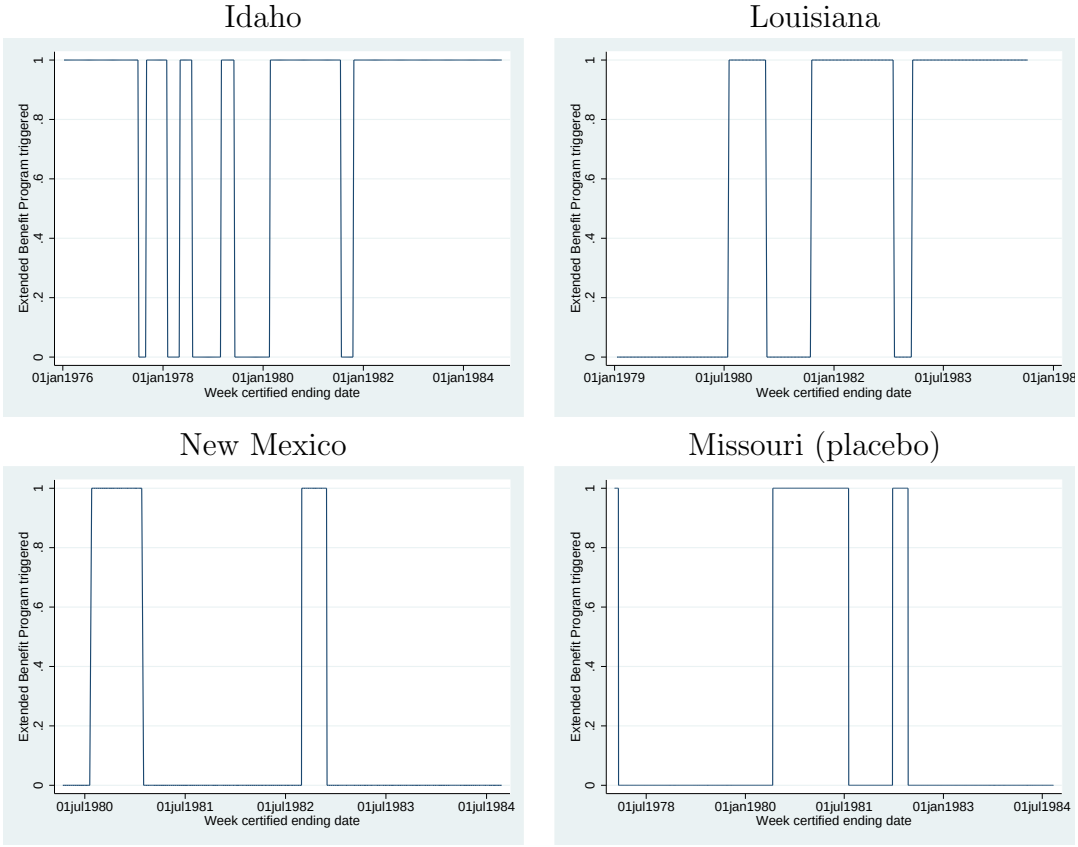
When claimants meet the above requirements, states compute their weekly benefit amount (WBA). This would be their weekly unemployment benefit payment when they earn less than the partial UI disregard. The WBA is a fraction (between 1/20 and 1/26) of the high quarter wages (HQW), defined as the wages 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 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. The FSC program was active from September 1982 up to March 1985 in all four states considered. It consisted of four distinct phases (see Grossman 1989 for more details on the FSC). In Figure 12, I plot the EB periods in each state.

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.

Figure 12: Extended Benefit Program.



Source: Trigger reports.

B Model Solution

In this Appendix, I derive in detail the solution of the claimants' program:

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

such that

$$\begin{aligned} c_t &= z_t + b\mathbb{1}\left[B_t > b\right] + B_t\mathbb{1}\left[b > B_t > 0\right] - T(z_t)\mathbb{1}\left[B_t > 0\right] \\ B_{t+1} &= B_t - b\mathbb{1}\left[B_t > b\right] - B_t\mathbb{1}\left[b > B_t > 0\right] + T(z_t)\mathbb{1}\left[B_t > 0\right] \\ B_{t+1} &\geq 0 \end{aligned}$$

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 with $B > b$ (resp. $b > B$) satisfies $T(z) = b$ (resp. $T(z) = B$). Then the program simplifies as:

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

Then the first order condition is:

$$u_c(c, z) + u_z(c, z) + \beta p'(z) (U - W) = 0 \tag{24}$$

This determines the level of consumption together with the definition of U from the Bellman equation:

$$U = \frac{u(c, z) + \beta(1 - p(z))W}{1 - \beta p(z)} \tag{25}$$

Note that the two previous equations 24 and 25 show that the stationary value of unemployment U and the corresponding earnings z do not depend on the level of UB capital B. However they depend on the talent n_i of the individual.

Recall that, when $B > b$, the typical partial-UI schedule is such that there exists a unique $\overline{z}$ such that for any $z \geq \overline{z}$, $T(z) = b$ and the marginal tax rate is 100% just below $\overline{z}$. Let me consider the marginal individual whose talent is consistent with supplying $\overline{z}$, 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 talent with $B > b$. Of course there may exist some very talented 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 to 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 with talent consistent with stationary claiming arbitrarily small.

While claiming, UB capital is thus strictly decreasing over the spell. I define $t_{exh} < \infty$ the exhaustion date (first date when $B_t = 0$). The program becomes stationary only when job-seekers run out of benefits. I denote U the value of unemployment when benefits are exhausted: $U \equiv U_t(0)$ for all $t \geq t_{exh}$.

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

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

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

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

Both sub-programs share the same first order condition:

$$u_c(c_t, z_t) \left(1 - T'(z_t)\right) + \beta p'(z_t) \left(U_{t+1}(B_{t+1}) - W\right) + \beta p(z_t) T'(z_t) U'_{t+1}(B_{t+1}) = -u_z(c_t, z_t) \tag{26}$$

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

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

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 $t = t_{exh} - 1$. Consequently, the third term of the marginal gain of labor earnings can be written as:

$$\beta p(z_t) T'(z_t) U'_{t+1}(B_{t+1}) = T'(z_t) \beta^{t_{exh} - t - 1} \left(\Pi_{i=t}^{t_{exh}-2} p(z_i) \right) u_c(c_{t_{exh}-1}, z_{t_{exh}-1}) \tag{27}$$

where $\Pi_{i=t}^{t_{exh}-2} p(z_i)$ is the probability to exhaust benefits conditional on claiming at date t.

Using Equation 27 and the utility definition, the FOC in Equation 26 can be simplified. The rest of the derivation is in the main text.

B.1 Heterogeneity

Heterogeneity in the permanent job finding rate. In this extension, job-seekers are characterized by their talent n_i and their probability to remain claimants p_i, distributed according to a joint distribution $f_t(n,p)$ at period t. The excess bunching is then a weighted average of excess bunching in each p-strata: $\mathcal{B}_t \simeq \int_p f_t(p|n=z^*)\delta n(\tau(t,p))dp$. Consequently, the identification formula 11 is only slightly modified:

$$e = \frac{\mathcal{B}_t}{z^* \mathbb{E}_t[\tau(t,p)|n=z^*]} \tag{28}$$

The relevant dynamic marginal tax rate is then averaged over the population still claiming at period t. When job-seekers differ by their permanent job finding rate, there is dynamic selection over the spell. The fraction of job-seekers with high p increases over the spell. This reinforces the decrease in the marginal tax rate across periods. As a consequence, Proposition 1 is not only valid within a p-strata, but also on average.

Heterogeneity in the structural elasticity. I can also extend the model by allowing for different individual elasticities e_i. At time t, talents and elasticities are distributed according to a joint distribution $f_t(n,e)$. The excess bunching at time t then equals: $\mathcal{B}_t \simeq \int_e f_t(e|n=z^*)\delta n(\tau_t,e)de \simeq z^* \mathbb{E}_t[e\tau_t|n=z^*]$. I obtain the following identification formula:

$$\mathbb{E}_t[e|n=z^*] = \frac{\mathcal{B}_t}{z^*\tau_t} \tag{29}$$

Equation 29 shows that bunching at kinks identifies a specific local average of structural elasticities. Note that Proposition 1 is still relevant in this context, as $\mathbb{E}_t[e|n=z^*]$ does not depend on t.

Heterogeneity in the permanent job finding rate and in the structural elasticity I define $f_t(n,p,e)$ the joint distribution of talents, permanent job finding probabilities and elasticities, at period t. The excess bunching within a p-strata is:

$$\begin{aligned}
\mathcal{B}_t(p) &= \int_e \delta n(\tau(t,p),e) f_t(e|p,n=z^*)de \\
&= E_t[\delta n(\tau(t,p),e)|n=z^*,p] \\
&= z^*\tau(t,p)E_t[e|n=z^*,p]
\end{aligned}$$

When there are heterogeneity in all three dimensions, Proposition 1 still holds within a p-strata, because $E_t[e|n=z^*,p] = E_{t+1}[e|n=z^*,p]$. However Proposition 2 is not verified. Heterogeneity across p can be confounded by heterogeneity in elasticities (e), as there may exist p_1 and p_2 such that $E_t[e|n=z^*,p_1] \neq E_t[e|n=z^*,p_2]$.

B.2 Stepping-stone/crowding-out effects

I follow the same reasoning as in the main text. The assumptions specified in the main text imply that there exists a continuous function π_t such that $\pi_t(z_t) = \beta p'(z_t)\left(U_{t+1}(B_{t+1}) - W\right)$. I explicitly denote the dependency of τ to the partial-UI earnings:

$$\tau_t(z_t) = 1 - \beta^{t_{exh}-t-1}\Pi_{i=t}^{t_{exh}-2}p(z_i)$$

Consequently, the FOCs can be written as:

$$1 + \pi_t(z_t) = \left(\frac{z_t}{n}\right)^{1/e} \text{ when } z_t < z^* \tag{30}$$

$$1 - \tau_t(z_t) + \pi_t(z_t) = \left(\frac{z_t}{n}\right)^{1/e} \text{ when } z_t > z^* \tag{31}$$

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

$$n_t^* = \frac{z^*}{(1 + \pi_t^-(z^*))^e} \tag{32}$$

$$n_t^* + \delta n_t = \frac{z^*}{\left(1 - \tau_t(z^*) + \pi_t^+(z^*)\right)^e} \tag{33}$$

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: $e = \frac{\mathcal{B}_t}{z^*\tau_t(z^*)}$

C 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: $h_i = h.\exp(\beta X_i)$. 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 (1990*b*) 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 (t_{exh}^{Utot}), in order to censor observations.

My objective is to capture 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 5 reports the coefficient estimates of the hazard model for each state (in columns).

Table 5: 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
Nb. 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.

D Supplementary Tables and Figures

Table 6: Descriptive Statistics

	Idaho	Louisiana	New Mexico	Missouri
Years	76-84	79-84	80-84	78-84
Inflow (nb)	91,162	95,675	62,030	78,065
Male	.673	.708	.670	.606
Age	31.1	34.5	33.6	34.7
Education (years)	11.8	11.3	11.6	11.2
White	.946	.622	.423	83.8
Private firm	.889	.952	.912	.939
Industries				
Construction	.148	.308	.224	.157
Manufacturing	.330	.178	.131	.397
Trade	.218	.135	.199	.142
Services	.139	.168	.199	.202
Occupations				
Prof., tech. and managers	.076	.095	.121	.072
Clerical and sales	.145	.153	.192	.162
Structural work	.221	.312	.278	.147

Source: CWBH. Notes : Agriculture, Mining, Transportation, Finance, Insurance and Real Estate industries are not reported for the sake of space (around 20% of the sample). Occupations corresponds 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.

Table 7: Robustness of earnings elasticities to the net-of-tax rate varying estimation parameters

	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	0.134	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)

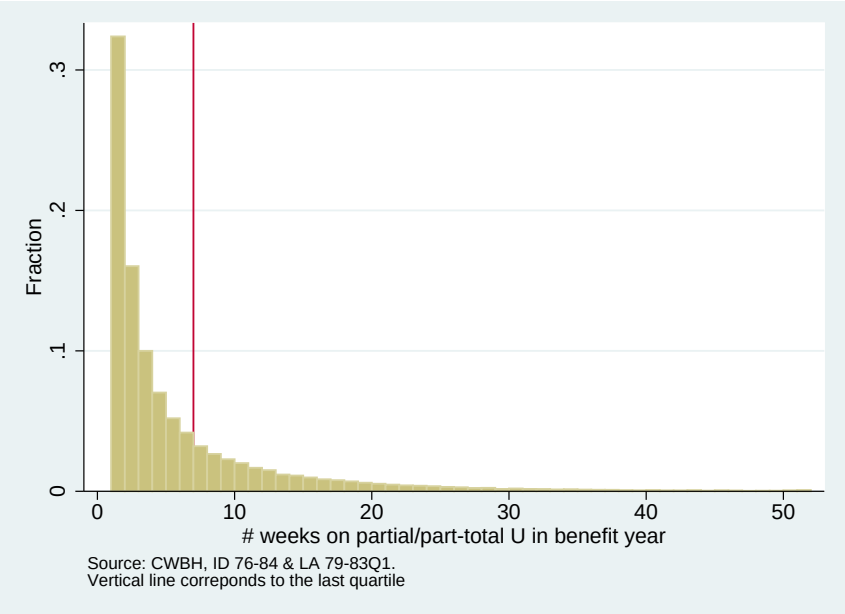
Source: CWBH. Notes: This Table reports estimates of the earnings elasticity to the net-of-tax rate varying the estimation parameters. Column 1 recalls the results of the baseline estimation (in Table 4) for the three U.S. states: ID, LA and NM. In Columns 2 to 4, I increase the lower bound of the bunching 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. Standard errors are in parentheses. 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.

Table 8: Earnings elasticity estimates without first-order approximation of the marginal tax rate.

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

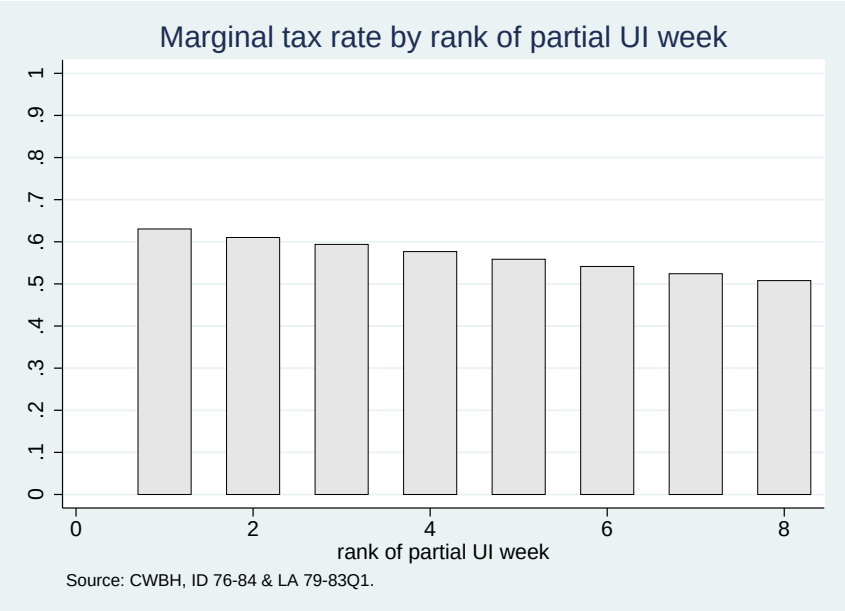
Source: CWBH. Notes : This Table reports estimates of earnings elasticity to the net-of-tax rate, computed with the exact identifying formula $e = -\mathcal{B}/z^* / \ln(1-\tau_t)$. Standard errors are in parentheses below estimates.

Figure 13: *Distribution of the number of partial-UI weeks over the claim*



Source: CWBH ID 76-84 & LA 79-83Q1. Notes: For each individual claim, I compute the total number of weeks with positive earnings (partial UI).

Figure 14: *Dynamic marginal tax rate by rank of the partial-UI week*



Source: CWBH for Idaho 1976-84 and Lousiana 1979-83, Q1. Notes: For every individual, I rank, within her claim, their weeks with positive earnings (partial UI). I select claimants with at least eight weeks with positive earnings. I estimate the average dynamic marginal tax rate for each of those first eight weeks of partial UI.