# Supplemental Appendix

## Taxes Today, Benefits Tomorrow

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The online appendix is divided in seven sections from A to G. Section A provides a detailed description of the Unemployment Insurance rules in the U.S., specifically for claimants who work while on claim. Section B describes the estimation process for excess bunching. Section C provides the full solution of the main theoretical model, and some model extensions (risk-averse workers, borrowing/saving margins, and stepping-stone/crowding-out effects). Section D estimates the hazard model out of unemployment, used to compute rational expectations of survival. Section E performs a difference-in-difference exercise, as a placebo test for the bunching source. Section F presents supplementary test for heterogeneity in bunching across groups with different job finding predictions. Section G displays supplementary Figures and Tables.

# A Institutional background

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

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

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

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

Table A1: Partial-UI rules from 1976 to 1984

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

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

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

# B Excess bunching estimation

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

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

$$C_j \left(1 + \mathbb{1}[j > \overline{R}] \frac{\hat{B}_N}{\sum_{j>\overline{R}} C_j}\right) = \sum_{k=0}^{q} \beta_k (Z_j)^k + \sum_{i=-\underline{R}}^{\overline{R}} \gamma_i \mathbb{1}[Z_j = i] + \epsilon_j \tag{15}$$

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.[31] The order of the polynomial $q$ and the width of the bunching window $(-\underline{R}, \overline{R})$ are not estimated, but set after visual inspection. Robustness checks of the estimation results with respect to those two parameters are presented below.

Equation (15) 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:

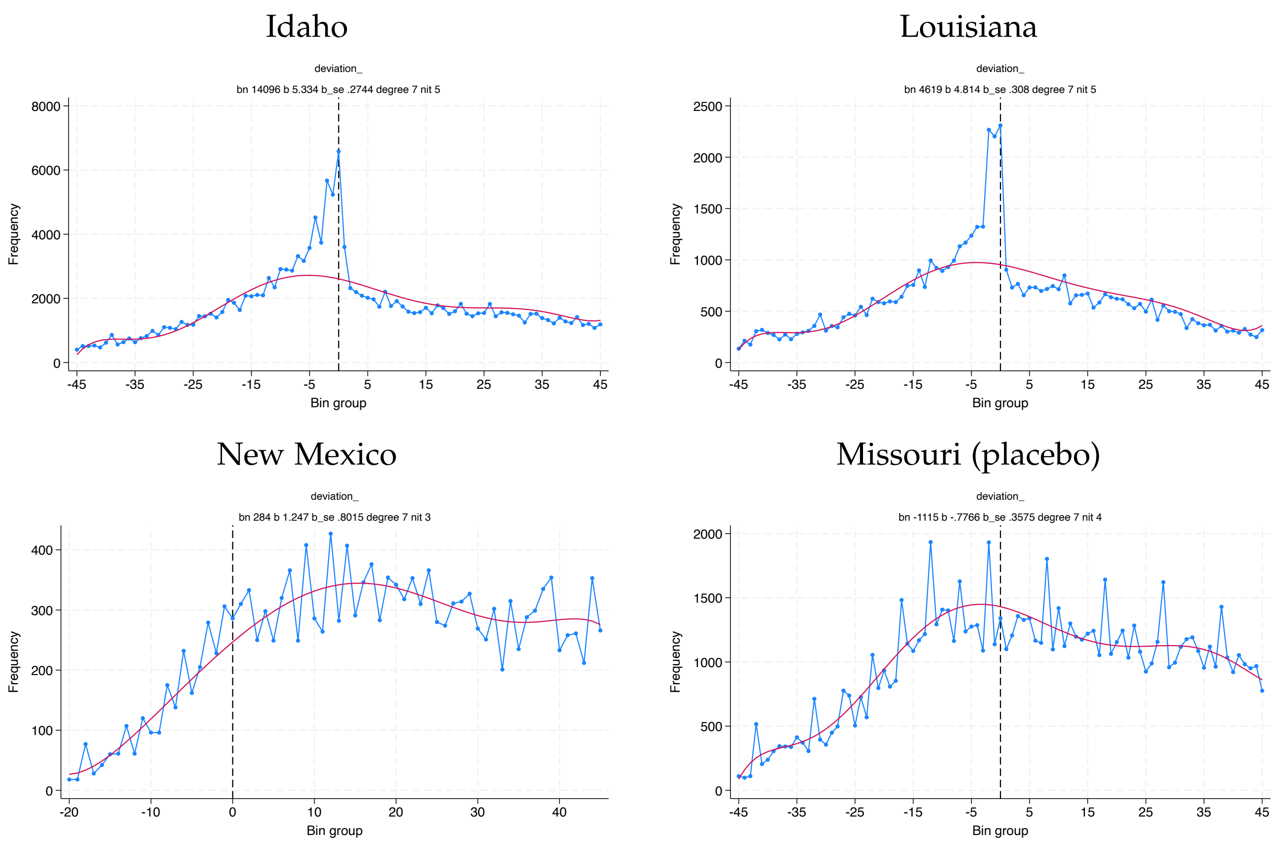
$$\hat{\mathcal{B}} = \frac{\hat{B}_N}{\sum_{j=-\underline{R}}^{\overline{R}} \hat{C}_j / (\underline{R} + \overline{R} + 1)}. \tag{16}$$

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

Appendix Figure B1 illustrates the estimation procedure for each state. It plots the partial claimants' earnings density in bins of one dollar centered around the disregard level, together with the counterfactual density estimated along the lines of Chetty et al. (2011*b*). In practice, the procedure fits a polynomial of degree 7. The bandwidth is such that $-\underline{R} = -5$ and $\bar{R} = 2$. Appendix Figure B1 confirms that the counterfactual density compares well to the actual data. Appendix Figure B1 reveals some periodicity in the earnings distribution in Missouri. Claimants report earnings that are multiples of ten dollars. This may bias bunching estimates, especially if there are heaps in the window where bunching is expected. I verify that the bunching estimate does not change if I modify the earnings density by smoothing the heaping points.

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

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




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

# C Theoretical model

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

## C.1 Model Solution

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

$$U(B_t; n_i) = \max_{c_t, z_t} u(c_t, z_t; n_i) + \beta \left[ pU(B_{t+1}; n_i) + (1-p)W \right]$$
$$\text{such that}$$
$$\begin{cases} c_t & = z_t + \min(b, B_t) - T(z_t) \\ B_{t+1} & = B_t - \min(b, B_t) + T(z_t) \\ B_{t+1} & \geq 0. \end{cases}$$

By definition of the partial-UI schedule, I have that $T(z_t) \leq b$ when $B_t > b$ and $T(z_t) \leq B_t$ when $B_t < b$. As a consequence, the capital stock $B_t$ depreciates or stays constant over time: $B_{t+1} \leq B_t$. The main model insights belong to the case of UB capital decreasing over time. Before solving the model under this case, I explain why stationary solutions with constant UB capital can be ruled out. My reasoning is by contradiction. I characterize the stationary solution and explain why deviations increase workers' welfare.

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

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

The first order condition writes:

$$u_c(c, z; n_i) + u_z(c, z; n_i) = 0. \tag{17}$$

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

$$U_i = \frac{u(c, z; n_i) + \beta(1-p)W}{1 - \beta p}. \tag{18}$$

The two previous equations (17) and (18) show that the stationary value of unemployment $U_i$ and the corresponding earnings $z$ do *not* depend on the level of UB capital $B$. This is an important property of stationary solutions.

Recall that, when $B > b$, the typical partial-UI schedule is such that there exists a unique maximum earnings $\overline{z}$ such that for any $z \geq \overline{z}$, $T(z) = b$ and the marginal tax rate is 100% just below $\overline{z}$. Workers with stationary consumption have earnings above $\overline{z}$. Let me consider the marginal individual who would supply $\overline{z}$, she would benefit from deviating from the stationary path during one period by decreasing her labor supply by $\delta z$. Actually, her flow income is not affected, while she enjoys more leisure. A consequence of this manipulation is that her UB capital is depreciated. However her future utility is not affected as the value of stationary unemployment does not depend on UB capital. Then, this deviation necessarily increases her welfare and a stationary equilibrium does not exist for this worker.

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

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

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

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

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

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

Both sub-programs share the same first order condition:

$$u_c(c_t, z_t; n_i) \left(1 - T'(z_t)\right) + \beta p T'(z_t) U'(B_{t+1}; n_i) = -u_z(c_t, z_t; n_i). \tag{19}$$

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

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

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

$$\beta p T'(z_t) U'(B_{t+1}; n_i) = T'(z_t) \beta^{\bar{t}-t-1} p^{\bar{t}-t-1} u_c(c_{\bar{t}-1}, z_{\bar{t}-1}; n_i) \tag{20}$$

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

Using Equation (39) and the utility definition, the FOC in Equation (19) can be simplified. The rest of the derivation is in the main text.

## C.2 Risk-aversion

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

$$u(c, z; n_i) = \frac{c^{1-\sigma}}{1-\sigma} - \frac{n_i}{1+1/e} \left(\frac{z}{n_i}\right)^{1+1/e} \tag{21}$$

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

$$u_c(c_t, z_t; n_i) \left(1 - T'(z_t)\right) + T'(z_t)(\beta p)^{\bar{t}-t-1} u_c(c_{\bar{t}-1}, z_{\bar{t}-1}; n_i) = -u_z(c_t, z_t; n_i). \tag{22}$$

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

$$1 - T'(z_t) \left(1 - (\beta p)^{\bar{t}-t-1} \left(\frac{c_t}{c_{\bar{t}-1}}\right)^{\sigma}\right) = \left(\frac{z_t}{n}\right)^{1/e} (c_t)^{\sigma}. \tag{23}$$

Compared to Equation (34), the wedge between the static benefit-reduction rate and the dynamic marginal tax rate depends on the coefficient of risk-aversion and on the ratio of current consumption to consumption in the last week of claim.

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

$$1 = \left(\frac{z^*}{n^*}\right)^{1/e} (b + z^*)^{\sigma}. \tag{24}$$

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

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

As $c_{\bar{t}-1}$ is a function of the ability $n^* + \delta n(t)$ and of the other parameters of the model ($\beta$, $p$, $\sigma$, $B_t$, $b$, $z^*$ and $e$), Equations (24) and (25) identify $e$ when excess bunching is observed in the data. More precisely, we obtain the consumption in the last week of claim $c_{\bar{t}-1}$ using the FOC of the program when $B_{\bar{t}-1} < b$:

$$1 = \left(\frac{z_{\bar{t}-1}}{n^* + \delta n(t)}\right)^{1/e} \left(c_{\bar{t}-1}\right)^{\sigma}. \tag{26}$$

The budget constraint in the last week of claim writes: $c_{\bar{t}-1} = B_{\bar{t}-1} + z_{\bar{t}-1}$. Assuming that the remaining UB capital in the last week of claim is negligible, we have $c_{\bar{t}-1} = z_{\bar{t}-1}$. Then Equation (26) shows that $c_{\bar{t}-1}$ only depends on $n^* + \delta n(t)$, $e$ and $\sigma$. Replacing the implicit expression of $c_{\bar{t}-1}$ in Equation (25), we obtain that excess bunching identifies $e$ from Equations (24) and (25).

We now quantify the order of magnitude of the bias in the estimate of $e$ when risk-aversion is neglected. We re-write Equations (24) and (25):

$$n^* = z^* \left(c_t\right)^{e\sigma}, \tag{27}$$

$$n^* + \delta n(t) = z^*(1 - \tau_t)^{-e} \left(c_{\bar{t}-1}\right)^{e\sigma}. \tag{28}$$

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

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

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

$$e = \frac{\delta n(t)}{z^*\left(\tau_t + \sigma\frac{\Delta c}{c_t}\right)}. \tag{29}$$

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

## C.3 Borrowing-saving

In this section, I no longer assume that job-seekers are hand-to-mouth. They have access to safe assets with return rate $r$. This introduces a new state variable in the job-seeker's program: asset capital $A_t$. Asset $A$ can be transferred across state, so that the value of permanent jobs also depends on $A_t$. The job-seeker program is modified as follows:

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

such that

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

where the modified budget constraint allows for saving and borrowing through $A$. The last constraint prevents the job-seeker from borrowing infinite amounts.
The solution is characterized by two FOCs (when constraint inequalities do not bind). For the first FOC, I differentiate wrt $z_t$, and for the second FOC wrt $A_{t+1}$:

$$u_c(c_t, z_t; n_i) \left(1 - T'(z_t)\right) + \beta p T'(z_t) U_B(B_{t+1}, A_{t+1}; n_i) = -u_z(c_t, z_t; n_i) \quad (30)$$
$$-u_c(c_t, z_t; n_i) + \beta \left[ pU_A(B_{t+1}, A_{t+1}; n_i) + (1-p)W_A(A_{t+1}) \right] = 0 \quad (31)$$

where $U_B$ and $U_A$ denote the differentials of the intertemporal value $U$ wrt to its first and second state variable. Using the envelope theorem when $B_t > b$, the FOCs write:

$$U_B(B_t, A_t; n_i) = \beta p U_B(B_{t+1}, A_{t+1}; n_i), \quad (32)$$
$$U_A(B_t, A_t; n_i) = (1+r) u_c(c_t, z_t; n_i). \quad (33)$$

When $B_t < b$, I have the same terminal condition as in the baseline model: $U_B = u_c(c_{\bar{t}}, z_{\bar{t}}; n_i)$. For the sake of simplicity, I do not write down the full program when the worker works in permanent jobs. I assume that the marginal value of wealth in this state relates to the marginal utility of consumption $W_A(A_{t+1}) = (1+r)\bar{u}_c$.

Before solving further, I note that under risk neutrality, the borrowing-saving margin is irrelevant for bunching. The marginal utility of consumption is equal to one and it does not depend on consumption level. The FOC (30) simplifies as before and this pins down the path of low-wage earnings $z_t$ independently of the wealth sequence $A_t$:

$$1 - T'(z_t)\tau_t = \left( \frac{z_t}{n_i} \right)^{1/e} \quad (34)$$

The more interesting case is that of risk-aversion. I assume that job-seekers are risk-averse as in Section C.2 with per-period utility:

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

As in Section C.2, solving the job-seekers' program forward yields the following FOC wrt $z_t$:

$$u_c(c_t, z_t; n_i)\left(1 - T'(z_t)\right) + T'(z_t)(\beta p)^{\bar{t}-t-1} u_c(c_{\bar{t}-1}, z_{\bar{t}-1}; n_i) = -u_z(c_t, z_t; n_i). \tag{35}$$

The previous expression can be re-written as:

$$1 - T'(z_t)\left(1 - (\beta p)^{\bar{t}-t-1}\frac{u_c(c_{\bar{t}-1}, z_{\bar{t}-1}; n_i)}{u_c(c_t, z_t; n_i)}\right) = \left(\frac{z_t}{n}\right)^{1/e}\frac{1}{u_c(c_t, z_t; n_i)}. \tag{36}$$

We note as in Section C.2 that the tax wedge depends on the ratio of marginal utility across periods. When intertemporal transfers are allowed on financial markets, the ratio of marginal utility is pinned down by the Euler equation, expressed in the FOC wrt $A_t$ in Equation (31). Substituting relation (32) and the expression of $W_A$, Equation (31) writes:

$$u_c(c_t, z_t; n_i) = \beta\left[p(1+r)u_c(c_{t+1}, z_{t+1}; n_i) + (1-p)(1+r)\bar{u}_c\right]$$

Iterating the previous expression at future dates allows to express current marginal utility as an expression of marginal utility at date $\bar{t}-1$:

$$u_c(c_t, z_t; n_i) = \left(\beta p(1+r)\right)^{\bar{t}-1-t} u_c(c_{\bar{t}-1}, z_{\bar{t}-1}; n_i) + (1-p)\sum_{k=1}^{\bar{t}-1-t}\left(\beta(1+r)\right)^k p^{k-1}\bar{u}_c$$

I further assume that discount factor and interest rates are such that: $\beta(1+r) = 1$. I use the analytical formula for the sum of first terms of a geometric sequence and obtain a compact expression of the current marginal utility:

$$u_c(c_t, z_t; n_i) = p^{\bar{t}-1-t} u_c(c_{\bar{t}-1}, z_{\bar{t}-1}; n_i) + \left(1 - p^{\bar{t}-1-t}\right)\bar{u}_c$$

It allows to express the ratio of marginal utility as:

$$\frac{u_c(c_{\bar{t}-1}, z_{\bar{t}-1}; n_i)}{u_c(c_t, z_t; n_i)} = \frac{1 - \frac{\bar{u}_c}{u_c(c_t, z_t; n_i)}}{p^{\bar{t}-1-t}} + \frac{\bar{u}_c}{u_c(c_t, z_t; n_i)} \tag{37}$$

From the above expression, it is not clear how the ratio compares to one. Intuitively, the ability to borrow and save smooths marginal utility across periods. This would bring the ratio of marginal utility closer to one and bunching behavior of risk-averse workers would be closer to that of risk-neutral workers than to that of hand-to-mouth risk-averse workers.

The bias quantification exercise performed at the end of Section C.2 is still valid when borrowing-saving is allowed. It relies on the same FOC (wrt $z_t$) that is left unchanged when introducing borrowing and saving. The bias quantification essentially relies on supplementary data on the curvature of the utility function and on consumption drops over the spell. Only the interpretation of the consumption drop changes when borrowing and saving are allowed. In the borrowing-saving case, it follows Formula (37).

However, without consumption data, it seems more difficult to exactly identify the earnings elasticity $e$ using the bunching formula derived from the FOC (30) taken for marginal bunchers. The identification proof used for hand-to-mouth risk-averse workers does not hold any longer (even when one assumes identification of the utility curvature). Compared to the hand-to-mouth case, earnings levels are not sufficient to identify consumption, which the proof leverages in Equation (24) and (26) of Section C.2. In the borrowing-saving case, savings/borrowings introduce an unobserved wedge between consumption and earnings. Moreover, the Euler equation is unlikely to solve the identification problem as the marginal utility while on claim depends on the marginal utility in permanent jobs, which is neither observed.

## C.4 Stepping-stone/crowding-out effects

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

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

such that

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

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

$$\underbrace{u_c(c_t, z_t; n_i) \left(1 - T'(z_t)\right)}_{(I)} + \underbrace{\beta p(z_t) T'(z_t) U'(B_{t+1}; n_i)}_{(II)} - \underbrace{\beta p'(z_t) \left(W - U(B_{t+1}; n_i)\right)}_{(III)} = -u_z(c_t, z_t; n_i). \tag{38}$$

Compared to the FOC in the baseline model (Equation 34), a third term (III) appears on the left-hand side. When working while on claim increases the future probability to find a permanent job - stepping-stone effect ($p' < 0$) -, the job-seeker is induced to work more. She has the opposite reaction when working while on claim crowds out job search for permanent jobs - crowding-out effect ($p' > 0$).

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

Following the same reasoning as in the previous appendix C.1, I obtain a simplified expression of the second term in Equation (38):

$$\beta p(z_t) T'(z_t) U'_{t+1}(B_{t+1}; n_i) = T'(z_t) \beta^{\bar{t}-t-1} \left( \Pi_{i=t}^{\bar{t}-2} p(z_i) \right) u_c(c_{\bar{t}-1}, z_{\bar{t}-1}; n_i) \tag{39}$$

where $\Pi_{i=t}^{\bar{t}-2} p(z_i)$ is the probability to exhaust benefits conditional on claiming at date $t$. Using Equation (39) and the utility definition, the FOC in Equation (38) simplifies to:

$$1 - T'(z_t)\tau_t(z_t) - \beta p'(z_t) \left( W - U(B_{t+1}; n_i) \right) = \left( \frac{z_t}{n_i} \right)^{1/e} \tag{40}$$

where the wedge $\tau_t$ now depends explicitly on $z_t$:

$$\tau_t(z_t) = 1 - \beta^{\bar{t}-t-1} \Pi_{j=t}^{\bar{t}-2} p(z_j). \tag{41}$$

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

Consequently, the FOCs can be written as:

$$1 + \pi_t(z_t) = \left(\frac{z_t}{n_i}\right)^{1/e} \text{ when } z_t < z^*, \tag{42}$$

$$1 - \tau_t(z_t) + \pi_t(z_t) = \left(\frac{z_t}{n_i}\right)^{1/e} \text{ when } z_t > z^*. \tag{43}$$

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

$$n_t^* = \frac{z^*}{(1 + \pi_t^-(z^*))^e}, \tag{44}$$

$$n_t^* + \delta n_t = \frac{z^*}{\left(1 - \tau_t(z^*) + \pi_t^+(z^*)\right)^e}. \tag{45}$$

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

# D Hazard model

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

It is well-established that hazard rates out of UI registers feature spikes at benefit exhaustion date. I verified that I obtain such patterns in the data from the Continuous Work and Benefit History (CWBH) project, as Katz and Meyer (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 ($\bar{t}^{Utot}$), in order to censor observations.

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

Table D1: Results of hazard model estimation

| | Idaho | Louisiana | New Mexico | Missouri |
|---|---|---|---|---|
| Male | .078*** | .288*** | .080*** | .175*** |
| | (.021) | (.015) | (.015) | (.015) |
| Age | -.021*** | -.006** | -.031*** | -.006* |
| | (.004) | (.003) | (.004) | (.003) |
| Age (square) | .0001** | -.00005 | .0003*** | -.00002 |
| | (.00005) | (.00003) | (.00004) | (.00004) |
| Education (years) | -.111*** | -.061*** | .012 | -.036 |
| | (.021) | (.009) | (.011) | (.062) |
| Education (square) | .006*** | .004*** | .0007 | .004 |
| | (.0009) | (.0004) | (.0005) | (.003) |
| Black | -.042 | -.216*** | -.172*** | -.583*** |
| | (.103) | (.013) | (.049) | (.020) |
| Hispanic | .296*** | .097* | -.223*** | -.207 |
| | (.044) | (.053) | (.014) | (.153) |
| American Indian | -.164* | -.084 | -.231*** | -.211 |
| | (.093) | (.122) | (.024) | (.378) |
| Asian | .127 | -.139* | -.264*** | .230 |
| | (.113) | (.075) | (.091) | (.218) |
| 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) |
| Years fixed effects | Yes | Yes | Yes | Yes |
| No. spells | 25274 | 55519 | 37937 | 41663 |
| Log-likelihood | -32412.47 | -75213.41 | -53243.84 | -56269.63 |

Source: CWBH. Notes: The reference is a white female with recall expectation whose claim starts in the first year of the sample. The estimation includes a constant and missing categories for ethnicity and recall expectation that we do not report in the table. *** $p<0.01$, ** $p<0.05$, * $p<0.1$

# E Difference-in-difference

In April 1983, Louisiana changed UI rules. The change in partial UI affected both the stock of individuals registered as unemployed in April 1983 and new inflows after that point in time.[32] 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.

**Placebo test** 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.

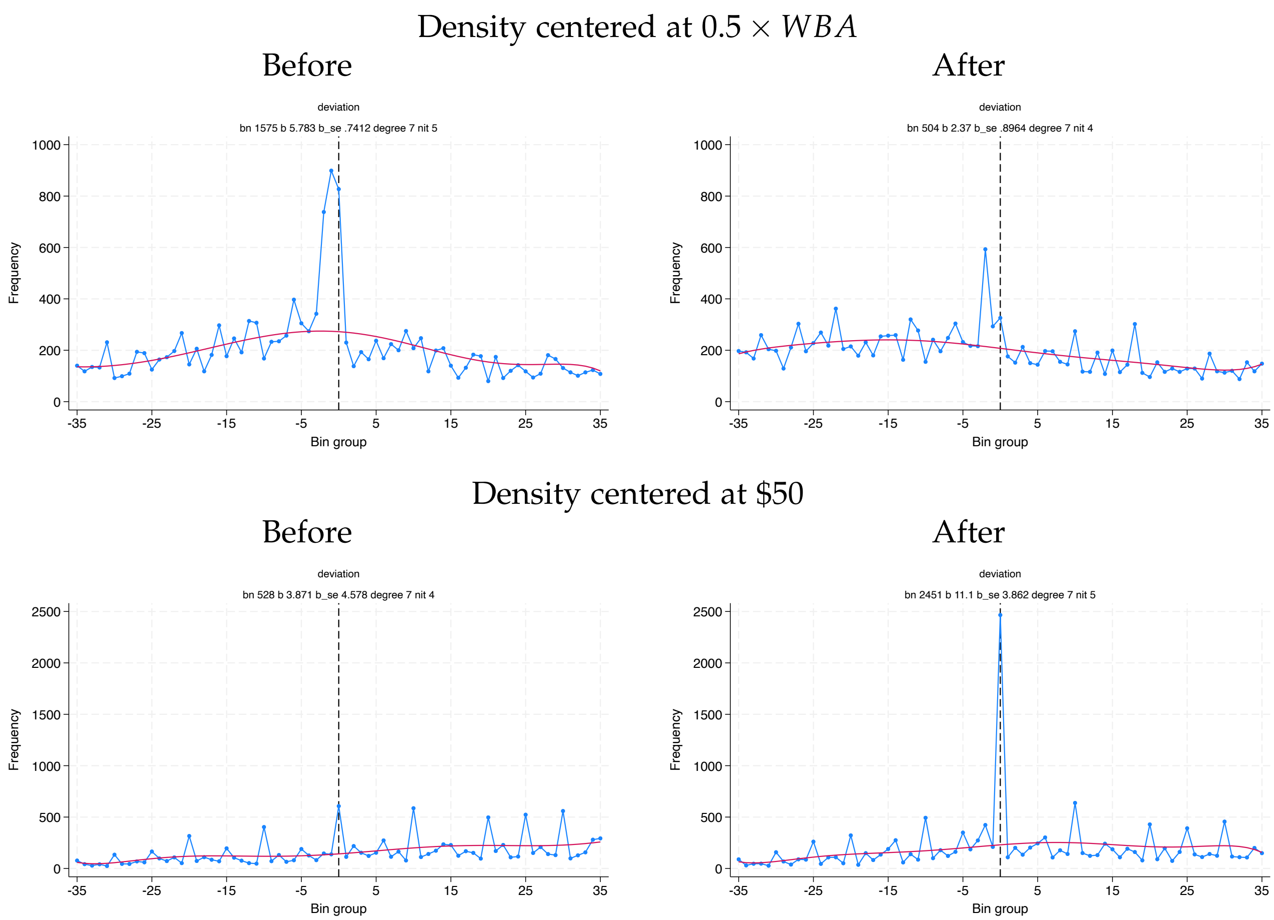
Appendix Figure E1 plots the earnings density of partial-UI claimants in the treatment group. In the upper panel, densities are centered at the pre-reform disregard ($0.5 \times WBA$). In the lower panel, they are centered at the post-reform disregard (\$50). Starting with the upper panel, bunching is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

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

[32] 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.

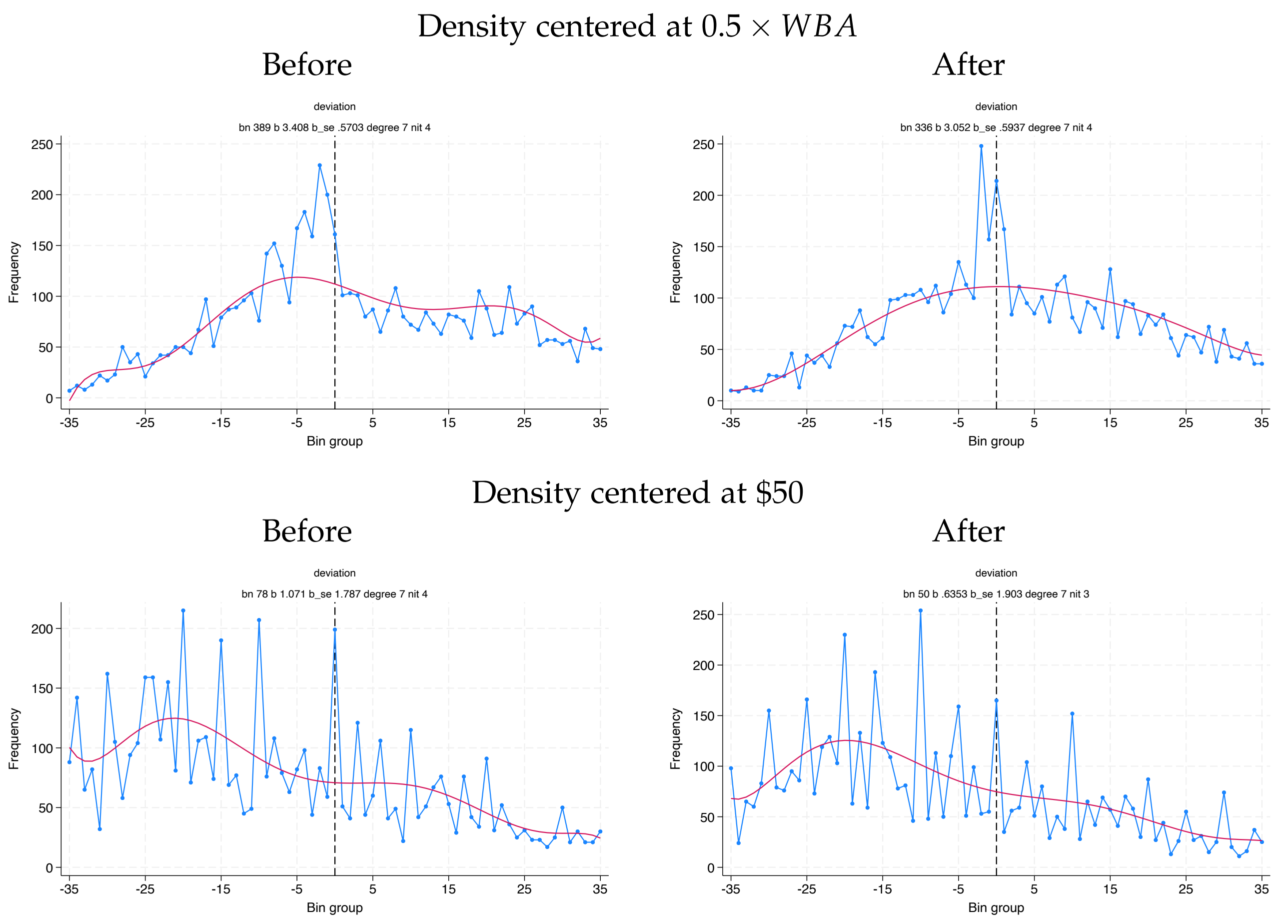
bunching at $50 in the control group, which is not the case.

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




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

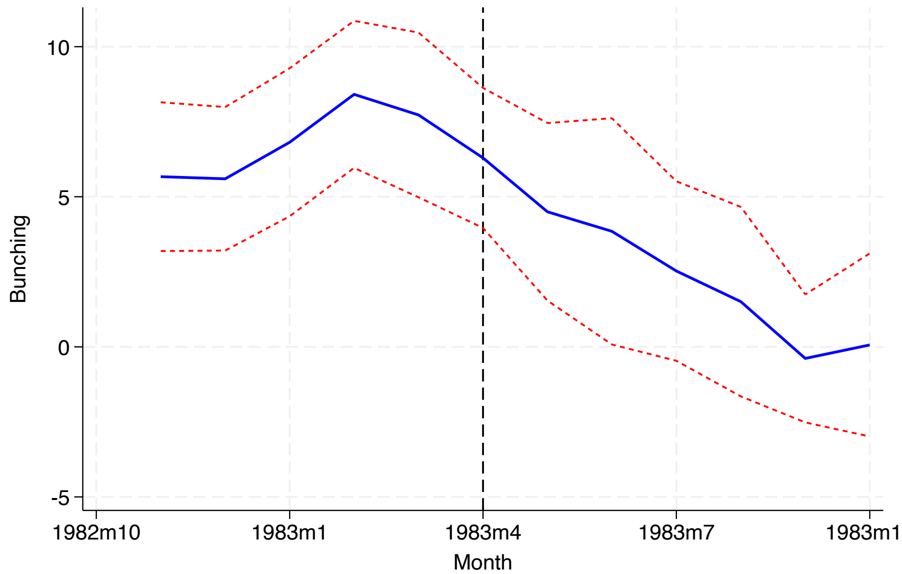
Figure E2: Centered weekly earnings density of partial-UI claimants in the control group.




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

**Speed of adjustment** Appendix Figure E3 plots the monthly evolution of the bunching estimates at the pre-reform disregard for the treated group. The pre-reform bunching pattern disappears within a few months after the reform. This suggests that adjustment costs in the market for low-wage jobs are small.

Figure E3: Bunching before and after the April 1983 reform in Louisiana

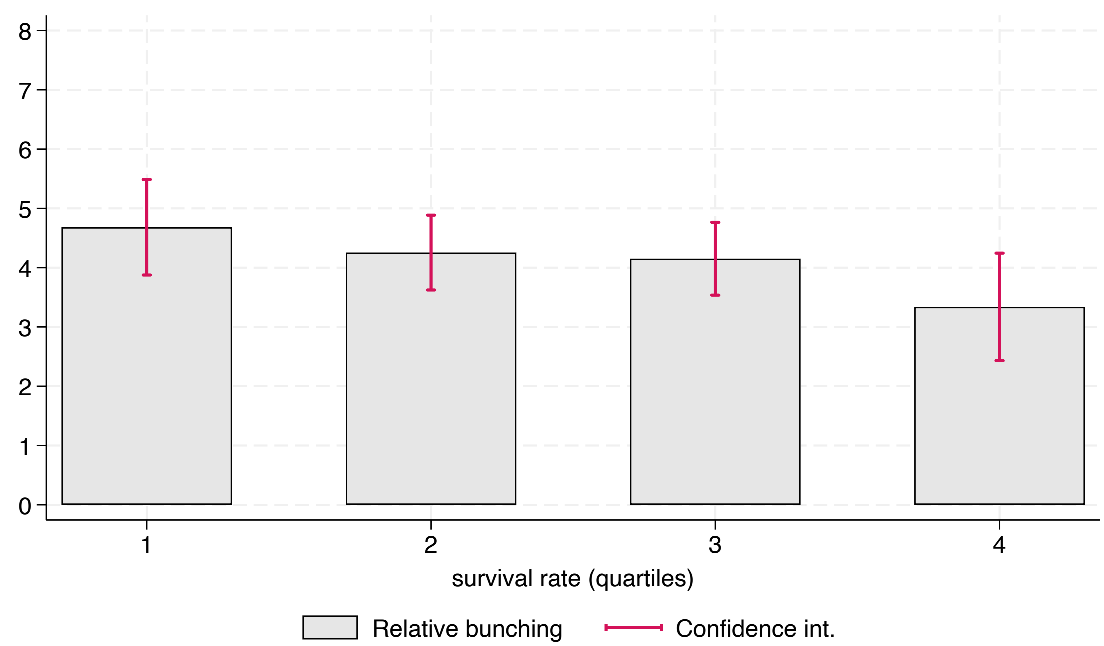



Source: CWBH, Louisiana October 1982 to September 1983. Notes: the figure plots the monthly bunching estimates at the earnings disregard (kink) before the reform (50% of WBA). The sample is restricted to workers with weekly benefit amount (WBA) greater than 100$, for whom the reform changes the disregard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

# F Supplementary bunching heterogeneity test

Appendix Figure F1 provides further evidence that supports Proposition 1. Under the assumption of rational expectations for claimants' job finding rates, I predict for each worker her expected survival rate (see Section V.A for more details). I then compare bunching across the quartiles of the predicted survival rate distribution. Overall, bunching tends to decrease from the first to the fourth quartile, confirming Proposition 1. The relation between bunching and predicted survival rate is significantly negative: the test of zero slope is rejected with p-value 0.03. The empirical evidence is consistent with Proposition 1.
Note that biased beliefs in job finding rates do not compromise the heterogeneity test in Appendix Figure F1 to the extent that biased beliefs preserve the rank of workers from rational-expectations estimates. However, it may lead to an attenuation bias when estimating the bunching elasticity to the expected survival rate, as workers with high realized job finding rates tend to be over-pessimistic and workers with low realized job finding rates over-optimistic (see Mueller, Spinnewijn and Topa, 2021).

Figure F1: Bunching by predicted survival rate.

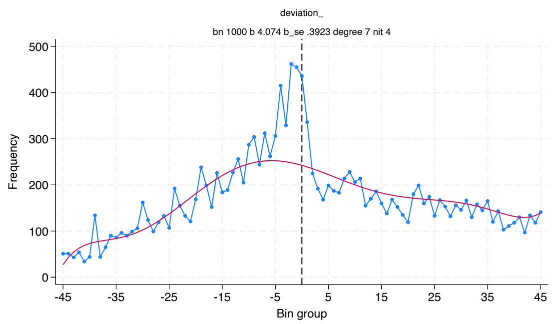



Source: CWBH for Idaho 1976-84 and Louisiana 1979-83, Q1. Notes: Excess mass at the disregard amount (kink) by quartile of predicted survival rates. Confidence interval at the 95% level in red.

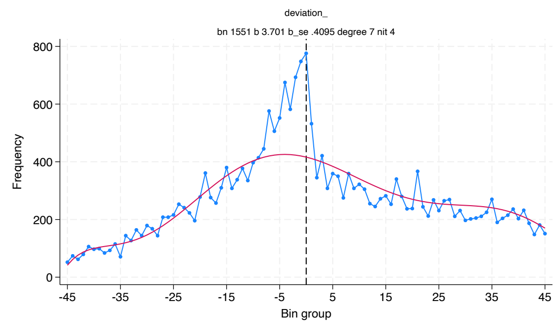
# G Supplementary Figures and Tables

Figure G1: Bunching by initial potential benefit duration in Idaho: Centered weekly earnings density of partial-UI claimants.

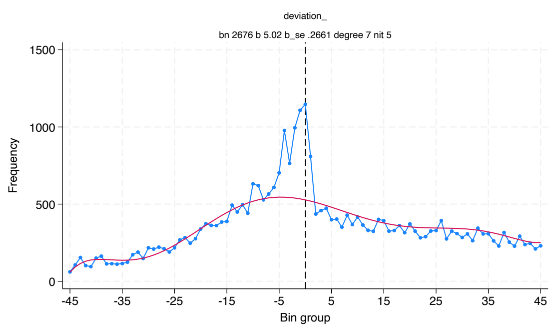
Panel A: 10-13 PBD weeks



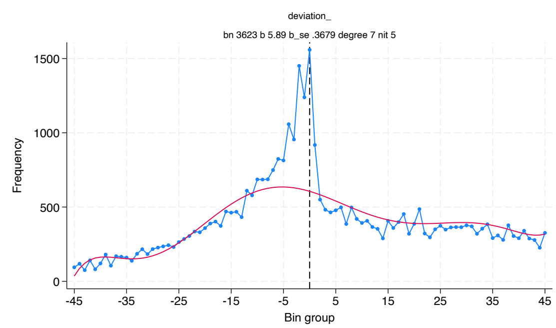

Panel B: 14-17 PBD weeks



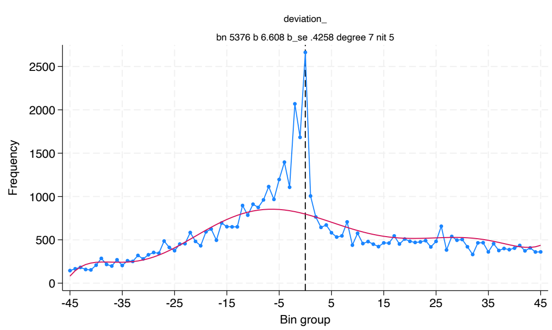

Panel C: 18-21 PBD weeks




Panel D: 22-25 PBD weeks




Panel E: 26 PBD weeks

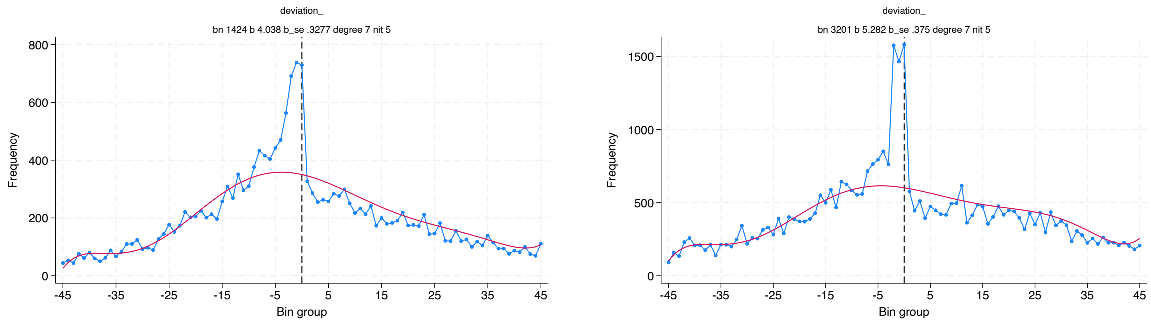



Source: CWBH. Notes: This figure plots centered weekly earnings density of partial-UI claimants, underlying the bunching estimates in Figure 3. Earnings are in dollars centered at the disregard level. Empirical earnings density in blue. Counterfactual density in red.

Figure G2: Bunching by initial potential benefit duration in Louisiana: Centered weekly earnings density of partial-UI claimants.

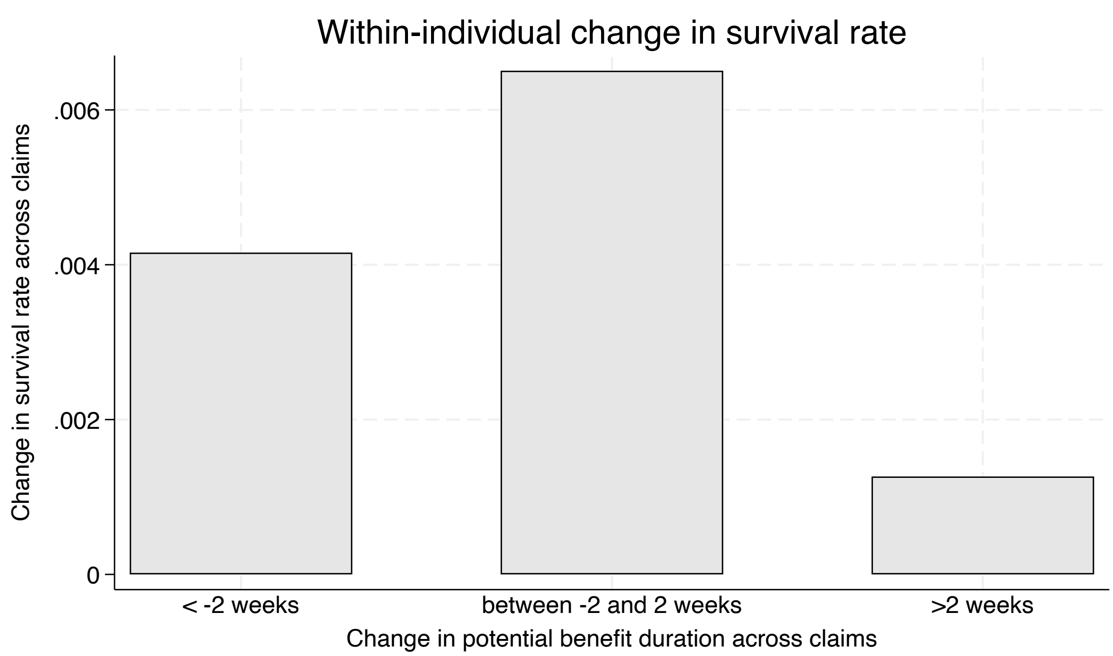
Panel A: 12-27 PBD weeks

Panel B: 28 PBD weeks




Source: CWBH. Notes: This figure plots centered weekly earnings density of partial-UI claimants, underlying the bunching estimates in Figure 3. Earnings are in dollars centered at the disregard level. Empirical earnings density in blue. Counterfactual density in red.

Figure G3: Change in expected weekly survival rate by change in potential benefit duration across claiming spells: within-worker design




Source: CWBH for Idaho 1976-84 and Louisiana 1979-1984. Notes: The figure plots the across-spell change in expected survival for claimants experiencing a large negative shock in potential benefit duration across spells (left-hand bar), a small change in PBD (center bar) or a large positive shock in PBD (right-hand bar). The average weekly survival rate is around 0.96.

Table G1: Robustness of earnings elasticities to the net-of-tax rate varying estimation parameters

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (11) | (12) |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | 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] | [-5,2] | [-5,2] | [-5,2] |
| Poly. Deg. | 7 | 7 | 7 | 7 | 7 | 7 | 9 | 8 | 6 | 5 | 4 | 3 |
| *Idaho* | | | | | | | | | | | | |
| Elasticity | .187 | .287 | .264 | 0.134 | 0.188 | 0.184 | 0.167 | 0.175 | 0.196 | 0.218 | 0.228 | 0.275 |
| s.e. | .00907 | .0194 | .0144 | 0.010 | 0.010 | 0.010 | 0.010 | 0.010 | 0.011 | 0.011 | 0.014 | 0.019 |
| *Louisiana* | | | | | | | | | | | | |
| Elasticity | .129 | .207 | .161 | 0.104 | 0.136 | 0.124 | 0.124 | 0.128 | 0.136 | 0.144 | 0.148 | 0.182 |
| s.e. | .0071 | .018 | .011 | 0.007 | 0.007 | 0.008 | 0.008 | 0.009 | 0.008 | 0.009 | 0.008 | 0.010 |
| *New Mexico* | | | | | | | | | | | | |
| Elasticity | .0962 | . | .197 | 0.100 | 0.054 | 0.071 | 0.053 | 0.102 | 0.099 | 0.081 | 0.094 | 0.054 |
| s.e. | .0646 | . | .17 | 0.040 | 0.050 | 0.065 | 0.080 | 0.070 | 0.066 | 0.053 | 0.047 | 0.056 |

Source: CWBH. Notes: This Table reports estimates of the earnings elasticity to the net-of-tax rate varying the estimation parameters. Column 1 recalls the results of the baseline estimation (in Table 2) for the three U.S. states: ID, LA and NM. In Columns 2 to 4, I increase the lower bound of the bunching window. In Columns 5 and 6, I increase the upper bound of the bunching window. In Columns 7 to 12, I decrease the degree of the polynomial fitting the density. Because the disregard level is around $20 in NM, it does not make sense to consider a lower bound at −15, and the estimation results are not reported.

Table G2: Earnings elasticity estimates without first-order approximation of the marginal tax rate.

| | Idaho | Louisiana | New Mexico |
|---|---|---|---|
| Elasticity | .13 | .0867 | .0626 |
| s.e. | .00694 | .00459 | .0404 |
| Obs. | 230535 | 69024 | 31103 |

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

Table G3: Bunching, dynamic marginal tax rates and earnings elasticity estimates with biased beliefs correction.

| | Idaho | Louisiana | New Mexico | Missouri *placebo* |
|---|---|---|---|---|
| Excess bunching mass ($\mathcal{B}$) | 5.33 | 4.81 | 1.25 | -.777 |
| | ( .315 ) | ( .299 ) | ( .811 ) | ( .351) |
| Hazard rate ($1-p$) | .0505 | .04 | .0472 | . |
| Effective Marginal Tax Rate ($\tau$) | .588 | .631 | .661 | . |
| Earnings elasticity | .171 | .118 | .0882 | . |
| to net-of-tax rate ($e$) | ( .0101 ) | ( .00734 ) | ( .0574 ) | . |
| Partial-UI weeks (no) | 230,535 | 69,024 | 31,103 | 99,451 |
| Sample Years | 76-84 | 79-83 | 80-84 | 78-84 |

Source: CWBH. Notes: This table replicates Table 2 of the main text, updating the expected hazard rates to account for biased beliefs. Standard errors are in parentheses below estimates.