The Welfare Effects of Conditioning the Potential Duration of UI Benefits on Prior Work History

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Abstract

Most unemployment insurance (UI) systems condition the potential duration of benefits on claimants' work history, yet little evidence exists on whether doing so is socially optimal. We study this question using 25 regression discontinuities in Canada's UI system, where benefit duration increases discretely when prior hours worked crosses various thresholds. Using administrative records covering the universe of UI claims from 1997 to 2018, we estimate the fiscal effects of extending potential duration across the work-history distribution. On average, an additional week lengthens unemployment by 0.30 weeks and costs the government $96. However, this average effect masks significant heterogeneity: the mechanical cost falls steeply with prior hours, while the behavioral cost increases modestly. Consequently, the fiscal externality of extending UI rises with labor market attachment. Embedding these estimates in a sufficient statistics framework, we find that the fiscal externality of a uniform extension is 76% larger than that of an extension targeted at claimants with short histories. Moreover, aggregate welfare can be raised at zero fiscal cost by redistributing benefits from high to low hour claimants.

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I Introduction

A fundamental challenge in the design of social safety net programs is determining the duration of benefit eligibility. In the context of unemployment insurance (UI), most OECD countries condition the potential duration of benefits on claimants' labor market activity prior to job loss. For instance, the majority of states in the U.S. scale the length of benefits according to claimants' earnings in the year before unemployment (Department of Labor, 2022). In contrast, seven states offer the same duration of benefits to all claimants, regardless of their prior work history.[1] These policy differences imply that the same UI claimant can receive up to double the number of weeks of benefits in a state with a uniform benefit schedule relative to a state that indexes the potential duration of benefits to prior labor market attachment. Despite the large impact that differences in the benefit schedule may have on UI claimants and the government's budget, there exists scarce empirical research on how alternative benefit schedules affect workers' behavior and program outcomes.

Our paper studies the welfare implications of conditioning the potential duration of UI benefits on claimants' past work history.[2] Prior research on the effects of increasing the potential duration of unemployment benefits have focused primarily on changes to the maximum potential duration offered by the UI system.[3] However, existing evidence suggests that around half of UI claimants do not even qualify for the maximum weeks of benefits (Card et al., 2015a; Landais, 2015; Lee et al., 2021). Due to a lack of policy variation targeting UI claimants with short work histories, little is known about the costs and benefits of extending the duration of benefits for these types of workers. As a result, despite the prevalence of UI systems that condition the potential duration of unemployment benefits on claimants' past work history, existing estimates are insufficient for determining the implications of this common practice.

To evaluate the welfare consequences of linking the potential duration of benefits to prior work history, we develop a sufficient statistics model that captures the conditions under which it would be welfare-improving to shift toward a more uniform benefit schedule. Extending the

[1] These states are Connecticut, Hawaii, Illinois, Maryland, New Hampshire, New York, and West Virginia.

[2] By potential duration, we mean the duration of benefits to which an individual claimant is entitled, but may not necessarily exhaust (i.e. if their unemployment spell ends first).

[3] For example, studies using age-discontinuities in Europe restrict their data to only individuals who qualify for the maximum potential duration (Schmieder et al., 2012), studies using regression-kink designs by definition rely on the response of claimants at the margin of receiving the maximum weeks of benefits (Landais, 2015), and even difference-in-difference studies rely on region specific changes to the maximum potential duration (Farber et al., 2015).

standard Baily-Chetty framework to incorporate worker heterogeneity, our model highlights two key forces in determining the optimal benefit schedule (Baily, 1978; Chetty, 2009). On the one hand, individuals with longer employment spells likely have higher incomes and savings, suggesting a lower marginal utility from government transfers. This would be an argument for giving claimants with longer employment histories *fewer* weeks of UI benefits. On the other hand, it is empirically unclear how the moral hazard cost varies across claimants with different work histories. If claimants with stronger labor market attachment also generate larger fiscal externalities, then the two forces push in the same direction and the model unambiguously calls for flattening the benefit schedule. Thus, fundamental to the decision of how much to condition the potential duration of benefits on claimants' past employment spells is the distribution of moral hazard costs across workers with different work histories.

We empirically estimate the key parameters of the model using a series of regression discontinuities unique to the Canadian unemployment insurance system. UI claimants in Canada receive an additional week of benefits for every 70 hours they worked in the year before unemployment, up until they reach 1400 hours, after which they receive one more week of potential duration for every 35 hours of work, giving us a total of 25 discontinuities. Leveraging administrative data covering all UI claimants in Canada from 1997 to 2018, we use the discontinuities to estimate the fiscal externality of marginally increasing job seekers' potential duration of benefits along the hours distribution.

We report four sets of results from our analysis. First, pooling the discontinuities together, we find that on average, an additional week of potential duration increases claimants' unemployment duration by 0.3 weeks and raises the cost to the government by $96 per claim. To validate our empirical design, we find neither discontinuities in covariates nor in the density of hours at the policy thresholds, suggesting that our estimates reflect the causal effect of the increase in potential duration and not selection bias due to unobservables that are correlated with hours. The magnitude of the effect on the weeks of unemployment is comparable to that of previous studies (Schmieder et al., 2016). However, the average effect masks substantial heterogeneity across claimants. Understanding this heterogeneity is central to evaluating how UI generosity should be allocated across workers with different employment histories.

Second, we unpack this heterogeneity by showing that the fiscal impacts of increasing the potential duration vary significantly across claimants with different work histories. The total cost of an additional week of potential duration decreases from a maximum of $217 per claim for individuals with short work histories to $65 per claim for individuals who accumulated many hours prior to unemployment. Decomposing the total cost, we find that the heterogeneity is driven primarily by differences in the mechanical effect of increasing UI generosity, rather than the behavioral effect. Since claimants with short work histories qualify for fewer weeks of unemployment insurance, a greater share of them exhaust their benefits, so we find that increasing the potential duration for this group has a larger direct mechanical effect despite weekly benefit amounts being smaller on average. In contrast, our estimates of the behavioral cost increase with respect to workers' prior labor market attachment.

Third, we show that the heterogeneity in fiscal impacts translates into meaningful differences in the welfare effects of extending UI benefits. In the context of the model, the deadweight loss of raising potential duration is captured by the fiscal externality, computed as the ratio of the behavioral and mechanical costs. Our estimates suggest that the fiscal externality is \$0.06 to \$0.11 larger for claimants who worked 100 more hours in the year prior to unemployment. For comparison, the mean fiscal externality in our sample is \$0.38. Although the costs of expanding UI generosity increase with workers' past hours, we find suggestive evidence that the value of insurance decreases. Claimants with weaker labor market attachment tend to have lower wages, receive less weekly UI benefits, and live in regions with higher unemployment rates. As a result, low-hour claimants likely face greater liquidity constraints and possess a higher marginal utility for UI benefits.

Fourth, we combine these sufficient statistics in our welfare framework to show that it would be welfare improving to redistribute UI benefits from claimants with strong labor market attachment to those with less. We find that the fiscal externality of a uniform increase in potential duration is 76% larger than a policy that increases the potential duration proportional to the distance from the top of the hours schedule. Moreover, our model finds that even if the social planner is unable to raise taxes to expand UI benefits, they can increase social welfare while keeping costs constant by reducing benefits for those with the most weeks of insurance and redistributing these benefits to those with the shortest work history. We also show that this tax-neutral pivot point is independent of claimants' valuation of insurance, and can thus be computed from our estimates of the fiscal effects without additional assumptions about the distribution of claimants' preferences.

Our results contribute to three areas of research in the unemployment insurance literature. First, we contribute to studies on the response of job seekers to increases in the generosity of unemployment insurance benefits (for a review, see Le Barbanchon et al., 2024). This literature has relied primarily on policy variation to the maximum potential duration and, as a result, focuses on the behavior of workers with significant attachment to the labor force (Lalive et al., 2006; Card et al., 2007b; Landais, 2015; Schmieder et al., 2016; Johnston and Mas, 2018; Karahan et al., 2025).[4] The few papers that examine marginally attached

[4] Aside from changes to the potential duration, a large literature also studies changes to the benefit

workers focus on the eligibility margin of benefits (Leung and O'leary, 2020; Moore and McQuillan, 2025; Jackson et al., 2025). In comparison, we provide the first evidence on how the behavioral and fiscal effects of extending UI benefits vary across the entire distribution of prior work histories.

Second, we contribute to a normative literature that measures the welfare effects of extending unemployment insurance benefits (for a review, see Schmieder and von Wachter, 2016). Whereas existing work has focused on the optimal *level* of UI generosity, we study the optimal *allocation* of benefits across workers (Gruber, 1997; Chetty, 2008; Landais, 2015; Ganong and Noel, 2019; Landais and Spinnewijn, 2021; Cohen and Ganong, 2026). Our findings indicate that the fiscal externality of extending UI benefits is higher for workers with stronger labor market attachment. These heterogeneous effects imply that the welfare effect of extending UI benefits depends not only on the magnitude of the policy, but also on its distribution.

Third, we contribute to the literature on the optimality of heterogeneous UI benefits. This literature has studied whether the generosity of UI benefits should vary over the business cycle (Schmieder et al., 2012; Kroft and Notowidigdo, 2016; Bell et al., 2024; Ganong et al., 2024), by age (Michelacci and Ruffo, 2015), during the unemployment spell (Kolsrud et al., 2018; Lindner and Reizer, 2020), or with the presence of informal labor markets (Gerard and Gonzaga, 2021). Hopenhayn and Nicolini (2009) examines from a theoretical perspective whether eligibility for unemployment benefits should depend on workers' employment history. We add to this literature by empirically studying whether the potential duration of benefits should also vary by claimants' prior labor market activity. Our results indicate that the social planner could improve aggregate welfare by moving towards a more uniform benefit schedule across workers.

The remainder of this paper is organized as follows. In Section II, we explain our data and the factors that determine the potential duration of benefits for UI claimants in Canada. Section III outlines our sufficient statistics model and our empirical strategy for identifying the key parameters of the model. Section IV reports our main results on average and heterogeneous effects of increasing the potential duration of benefits. We conclude in Section V by summarizing our findings and areas for future research.

amount (Card et al., 2015b,c; Lee et al., 2021; Ganong et al., 2024; McQuillan and Moore, 2025).

II Institutional Details and Data

In this section, we describe the unemployment insurance system in Canada and the data we use for our analysis.

II.a Canada's Unemployment Insurance System

Unemployed individuals in Canada are entitled to receive between 14 and 45 weeks of benefits, depending on the unemployment rate in the region they reside and their "insurable hours", defined as the number of hours they worked in the year before the start date of their claim. Table 1 shows the weeks of potential duration that a claimant would be eligible for given their insurable hours and place of residence. There are three primary features to the benefits schedule. First, the minimum number of insurable hours required to be eligible for any benefits decreases with each percent increase in the regional unemployment rate.[5] Second, the minimum and maximum potential durations also become more generous as the regional unemployment rate increases. Third, within region, the potential duration of benefits increases by 1 week for every 70 hours of work until 1,400 insurable hours, and then by 1 week for every 35 hours of work up to a maximum of 45 weeks. In short, the Canadian UI system offers more weeks of potential benefits to individuals living in high unemployment regions and to claimants with long work histories.

Key to our research design is that the potential duration of benefits is a step function that varies discontinuously with claimants' insurable hours.[6] These discontinuities allow us to identify the causal effect of giving claimants an additional week of potential duration along the entire hours distribution by comparing the outcomes of workers just above and below each cutoff. Comparing these local average causal effects at different hours thresholds allows us to then test how the fiscal externality of extending the duration of benefits varies across workers with different levels of prior labor market attachment. For our interpretation to be valid, other parameters of the UI system must evolve continuously as a function of insurable hours. We note three features of the Canadian UI system that may interact with our identifying assumption.

First, while claimants' weekly benefit amount also depends on their work history and region of residence, they do not vary discontinuously by hours of work. To calculate the

[5] For the purposes of UI benefits, Canada is divided into 62 economic regions. The unemployment rate within each region is updated monthly as a three month moving average.

[6] Refer to Appendix B for a discussion of how our institutional framework differs from previous settings studied in the literature.

amount of benefits that a claimant receives each week, the government first determines the individual's "best weeks" over the past year, defined as the weeks in which the claimant had their highest income. The weekly benefit amount is then calculated as 55% times the average income in the best weeks. To make the program more generous for claimants in weak labor markets, the formula uses the 22 best weeks in regions with lower than 6% unemployment rate, but decreases the number of weeks used by 1 for each percent increase in the regional unemployment rate. Given that the weekly benefit amount does not increase discontinuously by insurable hours, it does not affect our analysis on the effect of increasing individuals' potential duration.

Second, regions differ in the range of insurable hours for which there is a discontinuity. For example, regions with unemployment rates under 6% have discontinuities for insurable hours between 700 and 1820, whereas regions with over 16% unemployment rates have discontinuities between 420 and 1,330 hours. Although this means that different RD estimates may be identified using claimants from different regions, it does not fundamentally alter the welfare interpretation. Similar to most studies in the literature, our estimates capture the effect of increasing the potential duration of benefits while *assuming that all other features of the UI system remain constant.* Thus, given the structure of the rest of the UI system, we answer whether it is welfare improving for a social planner to relax the degree to which the potential duration of benefits depend on past work hours.

Third, we note that besides the regular benefits schedule, Canada has also implemented multiple pilot programs and temporary policies during the period of our study. For example, in 2016, Canada extended UI benefits to a maximum of 50 weeks in 12 regions that experienced a rather large increase in the unemployment rate over the past year. To the extent that these policies preserve the discontinuities required by our analysis, they do not affect our empirical strategy. The only rule that interacts with our regression discontinuity design is a requirement that "new entrants and re-entrants" must have at least 910 insurable hours to claim any amount of benefits. This policy implies that any discontinuity in outcomes at 910 hours reflects not only differences in potential duration of benefits, but also selection into eligibility for claimants who have low attachment to the labor force. To account for this selection effect, our preferred specifications will drop the 910 cutoff from the analysis.

II.b Administrative Claimants Data

Our analysis leverages over 20 years of unemployment insurance records for all of Canada from 1997 to 2018. These administrative microdata come from the Employment Insurance Status Vector (EISV) provided by the Canadian Research Data Centre Network. Included in the data are three types of information vital to our analysis. First, the data contains information used to compute claimants' benefits such as their work history and region of residence. Second, the data includes the weeks of benefits and the weekly benefit amount to which claimants are entitled. Third, the data also reports the outcomes of each claim, namely the number of weeks claimed and the total dollar amount claimed. Together, the data allows us to use work history data to generate quasi-random variation in potential duration, which we then use to identify the effect on unemployment duration and fiscal costs.

We make three restrictions to the data to avoid bunching and selection bias at specific hours. First, as described in section II.a, "new entrants and re-entrants" to the labor market have different eligibility thresholds than regular claimants so we drop all such individuals from our analysis. Second, we also drop industry codes associated with police, military, and teachers as these occupations tend to bunch insurable hours at specific values. In particular, full-time members of the Canadian Forces or a police force are credited 35 insurable hours per week regardless of their actual hours. Teachers tend to bunch at exactly 1,505 insurable hours because they work 35 hours per week for 43 weeks per year. Third, we drop insurable hours that are multiples of 5 to account for salaried workers. Since employers are not required to monitor the hours of salaried workers, the Canadian UI system allows employers and employees to agree upon an insurable hour, often resulting in bunching masses at round numbers.[7] Together, the aim of these restrictions is to ensure that the unobserved characteristics of claimants are continuous along the distribution of hours. We show in Section C that dropping round numbers and other point masses in the hours distribution yields a well-defined sub-population for which the regression discontinuity design can be implemented.

Figure 1 compares the distribution of insurable hours before and after our sample restriction. Panel (a) shows clear bunching masses along the distribution as well as a large discontinuity at 910 insurable hours. However, the majority of these spikes in the distribution do not occur at the cutoffs, suggesting that they are driven by other factors unrelated to the benefits schedule. Panel (b) shows that our sample restrictions eliminate the bunching masses, but not the discontinuity at 910 hours. We suspect that there are other potential pilot programs not captured in the data that target the 910 threshold. Given that the identifying assumptions for a regression discontinuity clearly fail at the 910 cutoff, we eliminate this threshold from our analysis. Besides the discontinuity at 910, panel (b) also finds small discontinuities at and below the 700 threshold. These jumps in the number of claimants simply reflect variation in the minimum insurable hours required to qualify for benefits across

[7] If an agreement is not made, then the government calculates a claimant's insurable hours as the minimum of their earnings divided by the minimum wage, 7 hours per day, or 35 hours per week.

regions. To avoid this margin of selection affecting our estimates, we also restrict each cutoff to regions where we can observe claimants on both sides of the discontinuity.[8] After these adjustments, the restricted distribution is continuous, suggestive that for this subset of claimants, there is no manipulation of insurable hours. For the remainder of our analysis, we restrict the sample to this subsample.

To understand the representativeness of our analysis sample, Table 2 reports summary statistics comparing our sample to the universe of claimants. The complete data set contains approximately 33 million claims filed over a 20 year period. The average claim is for an individual with 1,380 insurable hours, with implied weekly earnings of $640 during their best weeks before unemployment, and lives in a region with an unemployment rate of 9.16%. This translates to the average claim receiving 34 weeks of potential duration and a weekly benefit amount of $355. Although the average claimant qualifies for 34 weeks of benefits, they only collect 20 weeks, cumulating in $6,495 benefits paid per claim. Our restrictions to the data results in claimants with fewer insurable hours and less benefits on all dimensions. In total, these restrictions drop 67% of the sample but still leave over 10 million claims.

III Empirical Strategy

III.a Sufficient Statistics Model

In this section, we develop a sufficient statistics model of the central parameters needed for welfare evaluation of competing UI policies. We then use the model to answer two questions. First, under what conditions is it welfare enhancing to pivot the potential duration schedule rather than uniformly increasing the length of UI benefits for all claimants? Second, under what conditions can the social planner improve aggregate welfare by redistributing UI benefits without increasing taxes? Our framework builds on the model developed by Lee et al. (2021), extending it to allow for heterogeneity along a dimension of labor market attachment and therefore consider the possibility of redistribution across UI claimants.

Consider a social planner who chooses the potential duration of UI benefits P and a tax

[8]For example, claimants need at least 700 hours to qualify for UI in regions with unemployment rates less than or equal to 6%. In those regions, we only observe claimants to the right of the 700 cutoff, so we drop these regions from the analysis. Thus, the discontinuity at 700 is only identified from regions with above 6% unemployment rate.

parameter τ to maximize aggregate social welfare while maintaining a balanced budget:

$$\max_{P,\tau} W \text{ s.t. } E[B_i - T_i] = 0$$
$$\text{where } W := E\Big[G\big(V_i(B_i, T_i, Y_i)\big)\Big] \tag{1}$$

Individual i's utility $V_i(\cdot)$ depends on the amount of benefits received $B\big(Y, P\big)$ and taxes paid $T(Y, \tau)$, which in turn depend on their behavior Y such as job-search intensity. Utility also depends on Y directly. The function $G(\cdot)$ allows for the social planner to assign different weights across individuals. In what follows, we suppress the index for individual i.

Using this model, we consider the welfare implications of three different policy interventions. First, suppose the government increases the potential duration of UI benefits for all claimants. Taking the total derivative of the social welfare function with respect to P gives:

$$\frac{1}{\gamma}\frac{dW}{dP} = \underbrace{\frac{E\Big[G' \cdot V_1 \cdot \frac{\partial B}{\partial P}\Big] - \gamma E\big[\frac{\partial B}{\partial P}\big]}{\gamma}}_{\text{Benefit of Insurance}} - \underbrace{E\Big[\frac{\partial B}{\partial Y} \cdot \frac{dY}{dP} - \frac{\partial T}{\partial Y} \cdot \frac{dY}{dP}\Big]}_{\text{Behavioral Cost}} \tag{2}$$

where $\gamma = E\big[G' \cdot V_2 \cdot \frac{\partial T/\partial \tau}{dE[B-T]/d\tau}\big]$ represents the disutility of raising taxes to offset the cost of providing additional weeks of UI benefits. The "benefit of insurance" is therefore the increase in utility from more benefits relative to the cost of higher taxes. The "behavioral cost" represents the moral hazard cost from changes in individuals' behavior Y. By the envelope theorem, marginal changes in behavior have no first order impact on individuals' utility, but penalizes the government's budget constraint leading to a deadweight loss. For example, the behavioral cost captures the fiscal burden of reduced search effort in response to more generous UI benefits.

Equation (2) shows that to evaluate the welfare effect of increasing the potential duration of UI benefits, the social planner needs to know the value of insurance and the behavioral cost. Given that the value of insurance depends on utility parameters that we do not observe in the data, many studies focus on estimating the behavioral cost. In order to make our estimate comparable to previous policies studied in the literature, we follow the common practice of scaling the behavioral cost by the mechanical cost $E\big[\frac{\partial B}{\partial P}\big]$ of increasing the potential duration of benefits (Schmieder and von Wachter, 2017):

$$\frac{BC_P}{MC_P} = \frac{E\Big[\frac{\partial B}{\partial Y} \cdot \frac{dY}{dP} - \frac{\partial T}{\partial Y} \cdot \frac{dY}{dP}\Big]}{E\big[\frac{\partial B}{\partial P}\big]} \tag{3}$$

The mechanical cost represents the amount that benefits would increase by if the government increases P and individuals' behavior remains constant. The ratio of the behavioral and mechanical effects, called the fiscal externality, captures the deadweight loss per dollar of mechanical transfer.

In our second policy experiment, suppose the social planner is deciding whether to make the potential duration of benefits more uniform across workers with different work histories. In other words, should the social planner make the angle θ in Panel (a) of Figure 2 less steep? A challenge in estimating the welfare impacts of making UI benefits more uniform is that there is seldom policy variation in the slope of potential duration P with respect to past work history z. However, suppose we can estimate the marginal welfare effect, $\frac{dW}{dP}(z)$, at every point z. In that case, the expected change in social welfare from a pivot in the benefit schedule can be represented as

$$\frac{dW}{d\theta} = \int \frac{dW}{dP(z,\theta)} \frac{dP(z,\theta)}{d\theta} f(z) dz \tag{4}$$

where $\frac{dP(z,\theta)}{d\theta}$ is the amount by which potential benefits at work history z would change as result of a pivot in the benefits schedule, and $f(z)$ is the share of individuals with work history z. In other words, the effect of pivoting the potential duration schedule (i.e. $\frac{dW}{d\theta}$) is equivalent to a weighted sum of the effect of raising the potential duration of benefits (i.e. $\frac{dW}{dP}$) across the domain of work histories. The weights are higher for individuals with smaller z's, since $\frac{dP(z,\theta)}{d\theta}$ is larger there.

By iterating expectations over Z, the scaled fiscal externality of pivoting the potential duration schedule is

$$\frac{BC_\theta}{MC_\theta} = \frac{\int BC_{P_z} \frac{dP(z,\theta)}{d\theta} f(z) dz}{\int MC_{P_z} \frac{dP(z,\theta)}{d\theta} f(z) dz} \tag{5}$$

where $BC_{P_z} = E\left[\frac{\partial B}{\partial Y}\frac{dY}{dP} - \frac{\partial T}{\partial Y}\frac{dY}{dP} | Z = z\right]$ and $MC_{P_z} = E\left[\frac{\partial B}{\partial P} | Z = z\right]$ are the behavioral and mechanical cost of a marginal increase in $P(z)$, respectively.[9]

Besides being able to calculate the fiscal externality of pivoting claimants' benefit schedule, the distribution of BC_{P_z} and MC_{P_z} also enables policymakers to determine whether a uniform or redistributive change in the potential duration of benefits is more efficient. For instance, suppose the social planner has already decided to increase the potential duration of benefits. If we find that $\frac{BC_{P_z}}{MC_{P_z}}$ is increasing in z, then it is less costly to pivot the benefit

[9]Summing the behavioral cost across z's assumes that the disutility of taxation γ is independent of claimants' prior history z. This is a reasonable assumption given that taxes for unemployment insurance benefits are often levied on only employed workers, not the unemployed individuals for whom z affects their potential duration and behavior.

schedule than to uniformly increase all claimants' potential duration. Following usual assumptions about diminishing marginal utility of income, if we further assume that the value of insurance is greater for individuals with little work history, then the model implies that pivoting claimants' benefits schedule would increase welfare more than a uniform increase in generosity. Intuitively, these conditions imply that the social benefit of insurance is higher for low z individuals, but the deadweight loss is lower. We summarize this scenario in the following statement.

Theorem 1. *Let $P_\theta(z) \equiv \frac{\partial P(z,\theta)}{\partial \theta}$ be the change in potential duration at work history z from a marginal pivot, and suppose the pivot is a net expansion in benefits, $\int MC_{P_z}\, P_\theta(z) f(z)\, dz > 0$. Define*

$$\phi(z) = \frac{E\left[G' V_1 \frac{\partial B}{\partial P} \mid z\right]}{\gamma\, E\left[\frac{\partial B}{\partial P} \mid z\right]}, \qquad e(z) = \frac{BC_{P_z}}{MC_{P_z}}, \qquad \nu(z) = \phi(z) - e(z).$$

Suppose (i) $P_\theta(z)$ is decreasing in z so that the pivot gives more benefits for claimants with less work history; (ii) the per-dollar value of insurance $\phi(z)$ is decreasing in z; (iii) the local fiscal externality $e(z)$ is nondecreasing in z; and (iv) the disutility of taxation, $E\left[G' \cdot V_2 \cdot \frac{\partial T/\partial \tau}{dE[B-T]/d\tau} \middle| Z = z\right]$, is the same for all z and hence equal to γ (see footnote 9). Then pivoting the potential duration schedule raises social welfare per dollar of mechanical transfer strictly more than a uniform increase:

$$\frac{dW/d\theta}{MC_\theta} > \frac{dW/dP}{MC_P} \qquad \textit{and} \qquad \frac{BC_\theta}{MC_\theta} < \frac{BC_P}{MC_P}.$$

Proof. See Appendix D.a. ■

For our final policy experiment, suppose the social planner has no way of raising tax revenues. In this case, if they want to increase benefits for one group of individuals, they must decrease benefits for another. Panel (b) of Figure 2 provides a visual interpretation of this policy. Following the same assumptions as Theorem 1, we show that there exists a unique point z^* at which rotating the schedule of benefits would increase aggregate welfare while also being tax neutral. Intuitively, if the value of insurance decreases and the fiscal externality increases with z, then it is relatively welfare enhancing to transfer benefits from individuals with long work histories to those with short work histories. The social planner can redistribute these benefits without raising taxes by selecting the appropriate pivot point, z^*. We show in the following theorem that z^* is simply the cost-weighted mean work history. Since z^* depends only on the costs of increasing potential duration, and not claimants' unobserved valuations, we are able to compute it in the data.

Theorem 2. *Suppose z is continuously distributed on $[z_{\min}, z_{\max}]$ with density f, and:*

(i) the total fiscal cost per week $c(z) \equiv MC(z) + BC(z) > 0$;

(ii) the per-dollar value of insurance $\phi(z) = \dfrac{E[G'V_1 \frac{\partial B}{\partial P} \mid z]}{\gamma\, MC(z)}$ is nonnegative and nonincreasing in z;

(iii) the local fiscal externality $e(z) = \dfrac{BC(z)}{MC(z)}$ is nondecreasing in z.

(iv) the disutility of taxation, $E\big[G' \cdot V_2 \cdot \frac{\partial T/\partial \tau}{dE[B-T]/d\tau}\big|Z = z\big]$, is the same for all z and hence equal to γ (see footnote 9).

Then the work history

$$z^* = \frac{\int z\,\big[MC(z) + BC(z)\big]\,f(z)\,dz}{\int \big[MC(z) + BC(z)\big]\,f(z)\,dz} \tag{6}$$

is the unique point at which rotating the potential-duration schedule about $\big(z^*, P(z^*)\big)$ — raising P for $z < z^*$ and lowering it for $z > z^*$ — is

1. ***Tax neutral:*** $\displaystyle\int \big[MC(z) + BC(z)\big]\frac{dP(z,\theta)}{d\theta} f(z)\,dz = 0$*; and*

2. ***Welfare increasing:*** $\dfrac{dW}{d\theta} \geq 0$*, with strict inequality if ϕ is strictly decreasing or e strictly increasing on a neighborhood of z^*.*

Proof. See Appendix D.b. ■

In summary, the sufficient statistics framework developed in this section provides conditions under which the social planner can improve social welfare by pivoting the potential duration schedule or by redistributing benefits. The key parameters needed to evaluate the welfare implications of these policies are the mechanical cost of increasing potential duration MC_{P_z}, the behavioral cost BC_{P_z} and the distribution of work histories $f(z)$.

III.b Regression Discontinuity

In this section, we describe how we estimate the necessary parameters of our model using the regression discontinuities available in the Canadian UI system. Our method proceeds in four steps following the framework developed by Lee et al. (2021).

First, we define the outcome variable $\mathcal{S}_i = B_i - T_i$ as the total amount of benefits claimed by individual i minus the amount of taxes paid. We directly observe the amount of benefits B_i in the data. Since we do not observe the taxes paid in the data, our analysis will focus on the fiscal externality resulting from a change in the amount of benefits paid. We believe this to be an innocuous assumption given that previous studies have found that lost tax revenue plays a minor role in the fiscal externality (Lee et al., 2021). Intuitively, taxes are levied on the entire working population to support the UI system, so the share of costs borne by the claimants themselves is fairly small. As such, the main costs of claimants delaying entry into the labor market due to an extension of UI benefits is the direct cost of the benefits rather than foregone taxes.

Second, we compute a regression discontinuity at each cutoff using $\mathcal{S}_i$ as the outcome. The RD estimate at a cutoff z_0 identifies the local average treatment effect among those for whom P increases at z_0:

$$\begin{aligned} RD_{z_0} &= \frac{\lim_{Z\to z_0^+} E[\mathcal{S}|Z=z] - \lim_{Z\to z_0^-} E[\mathcal{S}|Z=z]}{\lim_{Z\to z_0^+} E[P|Z=z] - \lim_{Z\to z_0^-} E[P|Z=z]} \\ &= E\Big[\mathcal{S}(P+1) - \mathcal{S}(P)|Z=z_0, P(z_0^+) > P(z_0^-)\Big] \\ &\approx E\Big[\frac{d\mathcal{S}}{dP}|Z=z_0\Big] \end{aligned}$$

where $P(z_0^+) > P(z_0^-)$ denotes "compliers" for whom P increases by one unit at the cutoff. The third line treats this discrete difference of one week as a derivative, and notes that for most z_0, all regions have a discontinuity such that $P(z_0^+) = P(z_0^-) + 1$ for all claimants. This allows us to interpret the treatment effect as if the RDD were sharp (though we still use the fuzzy estimand to allow for exceptions).

The RD estimate is therefore the net effect of increasing the potential duration of benefits on the government's budget. This can be further decomposed into the mechanical effect and behavioral effect by expanding the total derivative:

$$E\Big[\frac{d\mathcal{S}}{dP}|Z=z_0\Big] = E\Big[\frac{\partial B}{\partial P}|Z=z_0\Big] + E\Big[\frac{\partial B}{\partial Y}\frac{\partial Y}{\partial P} - \frac{\partial T}{\partial Y}\frac{\partial Y}{\partial P}|Z=z_0\Big] \tag{7}$$

Third, we compute the mechanical effect. By definition, the mechanical effect is simply the cost of increasing the potential duration by 1 week if claimants do not change their behavior. We assume that individual i receives another week of benefits, equivalent to their "weekly benefit amount" (wba) if the potential duration increases and they do not change their behavior, so the mechanical effect is simply $E\Big[wba_i \cdot \mathbb{1}(i \text{ exhausts benefits})|z=z_0\Big] =$

$E\left[\tilde{B}_i|z=z_0\right]$, where we define the variable:

$$\tilde{B}_i = \begin{cases} 0, & \text{if individual } i \text{ does not exhaust benefits} \\ wba_i, & \text{if using all their benefits} \end{cases} \tag{8}$$

Fourth, equipped with the total cost and mechanical costs from the previous steps, we compute all the statistics in which we are interested. Following equation (7), the behavioral cost is simply the difference between the total and mechanical effect. Moreover, we can scale the behavioral cost by the mechanical effect to estimate the fiscal externality, where the standard errors are estimated via the Delta method. In order to easily estimate the joint asymptotic variance of the RD estimate along with the mechanical effect $E\left[\tilde{B}_i|Z=z_0\right]$, we estimate the mechanical effect through a second RDD, also centered at $v=z-z_0$ in which $\tilde{B}_i$ is treated as the outcome variable, and stack both regressions. The mechanical cost is the constant term in this second regression.

To estimate the fiscal externality of a uniform increase in benefits (i.e. equation (3)), we partition the data so that each cutoff appears once in each subsample, and then we stack the samples together to estimate our RD estimate for an "average" individual in the data (Cattaneo et al., 2016). Then to estimate the fiscal externality of a pivot in the potential duration schedule, we compute the RD estimate at each cutoff separately and aggregate them together using equation (5), where the density $f(z)$ is simply the share of observations in the partition with the cutoff.[10] We assume the pivot makes the duration schedule perfectly uniform across work histories when defining $\frac{dP(z,\theta)}{d\theta}$ in equation (5). Finally, to test whether the conditions for theorems (1)-(2) hold, we regress the local fiscal externalities on the running variable z, the number of insurable hours. In Appendix section E, we describe how to conduct inference on the slope of the fiscal externality with respect to z.

IV Results

This section presents our estimates of the fiscal externality of extending the potential duration of unemployment benefits and traces how that externality varies across the distribution of prior work histories. Our identification strategy exploits the fact that the potential duration of benefits is a step function of insurable hours, so that two otherwise similar claimants who fall on opposite sides of a threshold are entitled to a different number of weeks of

[10]For example, if there are 3 cutoffs at hours 700, 770, and 840, then $f(770)$ is the share of observations with insurable hours between 735 and 805.

benefits.

Figure 3 illustrates this variation and previews the descriptive patterns that motivate our welfare analysis. Panel (a) shows the source of our variation: potential duration rises in a series of discrete steps, jumping by one week at each threshold. However, to the left of 700 insurable hours, the relationship between hours and benefits appears to run "backward," with lower-hour claimants receiving *more* weeks of benefits. Panel (b) shows that this backward-relationship is an artifact of regional selection into eligibility rather than a causal effect. A claimant with very few insurable hours can qualify for benefits only in a high-unemployment region, and such regions are precisely the ones that grant the most weeks of potential duration. As explained in Section II.b, we account for this selection effect by estimating the RDD at each threshold using only region-years in which claimants are observed on both sides of the cutoff.

In Panels (c) and (d), we find evidence that increasing the potential duration of benefits leads to longer periods of unemployment and larger costs to the government's budget. Focusing on the cutoffs above 700 insurable hours, there are discontinuous jumps in the number of weeks claimed and the total amount of benefits received to the right of each threshold. This effect is especially pronounced in the 910 to 1,400 range, but becomes less clear at the shorter intervals in the right of the distribution. To rigorously quantify the magnitude of the discontinuities, we next turn to our regression discontinuity analysis.

IV.a Uniform Increase in Potential Duration

We begin by estimating the average effect of a uniform one-week increase in potential duration. Following the procedure in Section III.b, we center insurable hours at each interior threshold and stack the resulting samples into a single pooled regression, so that the estimated discontinuity reflects the response of an average claimant in our data.

Figure 4 presents results of the stacked design. Panel (a) confirms the first stage effect showing that crossing a threshold raises potential duration by approximately one week. Panel (b) shows how this in turn leads to an increase in the number of weeks of benefits claimed by about 0.3 weeks. Thus, not all claimants use the additional week of UI benefit to which they are eligible. Consistent with the increase in weeks of benefits claimed, panel (c) finds that the total amount of UI benefits paid over the course of a claim increased by approximately $100 for workers to the right of the cutoff.

To validate that the discontinuity in benefits claimed is not subject to selection bias, we test whether the observable characteristics of claimants are continuous across the threshold. Panel (d) plots an index of covariates as a function of insurable hours, where the index is a linear combination of baseline characteristics computed from an OLS regression of total benefits claimed on gender, age, age-squared, and firm and industry fixed effects. Unlike the actual amount of benefits claimed, we find no discontinuity in the covariate index. Similarly, Appendix Figure A.1 finds no discontinuity in workers' weekly earnings prior to unemployment, except at the 910 cutoff that we have already omitted from our regressions. Consequently, Appendix Figure A.2 shows that there is no discontinuity in weekly benefit amounts at the policy thresholds. Taken together, the continuity of claimants' predetermined characteristics supports our identifying assumption that workers are not sorting around the cutoffs.

Recall from the sufficient statistics model that our target parameters are measured relative to the change in potential duration, so we need to take the ratio of the discontinuities to compute our desired statistics. We estimate the following IV RDD regression by 2SLS:

$$y_{ij} = \beta_1 D_{ij} + \beta_2 p(V_{ij}) + \beta_3 p(V_{ij}) Z_{ij} + \varepsilon_{ij} \tag{9}$$

where y_{ij} is an outcome of claim i in stack j, such as the amount of UI benefits received. A claim can appear in two stacks to estimate the discontinuity to its left and to its right. D_{ij} is the number of weeks of potential duration, $p(V_{ij})$ is a polynomial with respect to the number of insurable hours, V_i, recentered such that $V_i = 0$ at a policy threshold. We allow the slopes of the polynomial trend to differ on either side of the threshold by interacting the polynomial with the indicator variable $Z_{ij} = 1[V_{ij} \geq 0]$. Our object of interest is β_1, which we identify by instrumenting D_{ij} with Z_{ij}.

Our preferred specification makes the following parameter decisions. First, we use a polynomial order of degree 4 with a square kernel. In practice, a local linear regression achieves similar results. Second, we use a bandwidth of 35 for cutoffs that are separated from the nearest other cutoff by 35 hours, and a bandwidth of 70 when the nearest cutoff is 70 hours away. Standard errors are clustered by claim, so that the copies are not treated as independent.

Table 3 reports the estimates. The first three columns present reduced-form coefficients on the above-threshold indicator Z, which measure the jump in each outcome at the cutoff. Crossing a threshold raises potential duration by 1.044 weeks and increases the number of weeks of benefits claimed by 0.314 weeks. Dividing the reduced-form response of weeks claimed by the first stage implies that each additional week of potential duration lengthens the realized benefit spell by roughly 0.30 weeks, a pass-through comparable to the duration responses estimated in the literature. For comparison, the review by Schmieder and von Wachter (2016) finds a mean effect of 0.26 across 13 studies, with a range of 0.05 to 0.65. Examining the fiscal impacts of crossing the threshold, column (3) finds that claimants to the right of the cutoff receive \$100.70 more UI benefits over their entire claim. Since the policy only increases potential duration by 1 week at each threshold, the IV estimate in column (4) is very similar to the reduced form estimate, raising fiscal costs by \$96.49.[11]

IV.b Heterogeneity by Work History

Next, we show that the pooled RD estimates mask significant heterogeneity in response across the distribution of insurable hours. In addition, we estimate the sufficient statistics described in section III.a that allow us to determine whether a social planner can increase social welfare by reallocating potential duration between claimants with different work histories.

Figure 5 plots the four objects of interest as a function of insurable hours, together with pointwise confidence bands. Panel (a) reports the total effect on benefits at each threshold. We find that the cost of increasing the potential duration of benefits decreases with insurable hours through most of the hours distribution. Appendix Figure A.3 finds a similar pattern when plotting the effect on the number of weeks of benefits claimed. At the left tail of the distribution, an additional week of potential duration can cost up to \$217 per claim, whereas the cost decreases to less than \$100 per claim for workers near the top of the distribution. However, at the very right tail, the negative relationship between total costs and insurable hours reverses trend, with the right most discontinuity exhibiting the largest increase in costs.

The increased cost to the government is not necessarily a deadweight loss. This increase in benefits reflects both a mechanical effect whereby workers who would have otherwise exhausted their benefits receive another week, and a behavioral effect whereby some claimants reduce their search effort and remain unemployed longer. As described in the sufficient statistics model, the impact of the increased fiscal costs on social welfare depends only on the share of the costs that is driven by a change in claimants' behavior.

Panel (b) plots the mechanical effect of raising the potential duration of benefits by one week. We compute this following the procedure described in Section III.b, where the mechanical effect is the constant term in a regression discontinuity regression with the outcome variable being the weekly benefit amount for workers who exhausted their benefits (see equation (8)). This component of the total cost declines steadily with insurable hours, falling from roughly \$155 at the lowest thresholds to about \$62 at the highest. The decline reflects two offsetting forces documented in Appendix Figure A.5: as insurable hours rise, the share of claimants who exhaust their benefits falls while the weekly benefit amount rises. Since the

[11] We also find in unreported regressions that our fuzzy RDD results do not vary appreciably with the inclusion of linear trends that vary by stack, or the inclusion of region-by-year fixed effects.

mechanical cost is the product of these two terms (i.e. $MC(z) = \mathbb{E}[\text{wba} \cdot \mathbf{1}(\text{exhaust}) \mid z]$) the net decline indicates that the fall in the exhaustion rate dominates. Intuitively, claimants with short work histories qualify for fewer weeks of benefits and are therefore far more likely to be at the exhaustion margin, so extending their potential duration mechanically pays out benefits to a larger share of them.

Panel (c) plots the behavioral effect, obtained by differencing the mechanical effect from the total effect. In contrast to the total and mechanical effects, the behavioral cost displays a much flatter, and slightly positive, relationship with respect to claimants' past work history. Thus, the downward trend in the total effect on the government's budget is driven nearly entirely by the mechanical effect. Finally, panel (d) plots the fiscal externality, the ratio of the behavioral to the mechanical effect. This is the key parameter needed to assess whether a social planner should increase potential duration for one subpopulation or another. Since the mechanical cost falls with insurable hours while the behavioral cost is relatively flat, the fiscal externality rises with work history.

Table 4 formally tests the relationship between the fiscal externality and claimants' insurable hours. Column (1) reports a simple OLS regression of the fiscal externality estimate on insurable hours, where standard errors are computed following the approach described in Appendix E. The slope of 0.00113 implies that going from the first threshold at 490 to the last threshold at 1785 would increase the fiscal externality by \$1.46 per dollar of mechanical transfer. In comparison, the mean fiscal externality across the distribution is only 0.38.

In columns (2) to (6), we show our results are robust to alternative specifications and sample restrictions. Column (2) weighs each threshold-specific estimate by the number of claimants in the bandwidth used to compute the fiscal externality. Column (3) restricts the bandwidth used to calculate each regression discontinuity so that none of the bandwidths overlap between adjacent thresholds. In both cases, we find similar results to our baseline specification.

We repeat our analysis in columns (4) to (6) after dropping three outliers estimates at 490, 700, and 1785. In particular, Figure 5 finds that the magnitude of the discontinuities at the very left and right tails of the distribution (i.e. 490 and 1785) are significantly larger than the other estimates. Moreover, our placebo test in Appendix Figure A.4 finds no evidence of manipulation around any of the thresholds except at 700 and the last threshold. Although dropping the outlier estimates shrink the magnitude of the trends, we nevertheless continue to find a strong positive correlation between the fiscal externality and claimants' insurable hours.

IV.c Comparing Welfare Effects Between Different Policies

In this section, we use the heterogeneity in fiscal externalities to compare the welfare implications of two policies: a uniform increase in potential duration for all claimants, and a pivot that raises potential duration for claimants with short work histories while leaving high-hour claimants unchanged.

The heterogeneity in response across the insurable hours distribution may reflect multiple factors. First, by definition, the population at each threshold differs in their past labor market engagement. Second, claimants at each threshold may differ along other dimensions too such as age, gender, and race. Third, the generosity of unemployment benefits also differs along the distribution. For example, the meta-analysis by Cohen and Ganong (2026) finds that jurisdictions with less generosity like Florida tend to have smaller unemployment elasticities than generous jurisdictions like France. However, a key insight of our sufficient statistics model is that the welfare implications of increasing UI generosity does not depend on the underlying source of the heterogeneity in responses.

Under our sufficient statistics framework, comparing the welfare effects of the two policies depends only on the heterogeneity in valuation for insurance and the heterogeneity in fiscal externalities. Although the value of insurance is unobservable in the data, Appendix Figure A.1 shows that job seekers with a short work history had lower wages when they worked. Under standard assumptions of diminishing marginal utility with respect to benefits, the descriptive evidence suggests that claimants at the left of the hours distribution likely have a higher marginal valuation of insurance.[12] At the same time, the empirical results in section IV.b find that those at the left tail also have lower fiscal externalities. Taken together, Theorem 1 predicts that aggregate welfare increases more from a pivot in the benefit schedule than from a uniform increase in potential duration.

Theorem 1 also implies that the fiscal externality from a pivot is lower than from a comparable uniform increase in potential duration. We test this prediction directly in Table 5, where we report the total cost, mechanical cost, behavioral cost, and fiscal externality of each reform. To calculate these statistics, we aggregate the threshold-specific estimates following equations (4) and (5), where we set $\frac{dP(z,\theta)}{d\theta} = 1$ for a uniform increase and $\frac{dP(z,\theta)}{d\theta} = \frac{z_{max}-z}{\sum_z f(z)(z_{max}-z)}$ for a pivot.[13] Thus, both policies increase average potential duration by one

[12] For example, Gruber (1997) and Cohen and Ganong (2026) assume the value of consumption-smoothing decreases linearly with the replacement rate.

[13] The total cost of a uniform increase in potential duration using this method does not perfectly match our estimates using the stacked regression in Table 3 because this method allows for the slope of the running variable to vary for each threshold.

week, but the pivot puts more weight on claimants farther from the maximum threshold, z_{max}.

Consistent with the theory, we find that the fiscal externality of a uniform increase in potential duration is 76% larger than a pivot in the benefit schedule. The first column Table 5 finds that the total cost of the pivot is actually larger than the uniform benefit increase. However, this difference is primarily driven by the mechanical effect since claimants with shorter work histories are more likely to exhaust their benefits. The mechanical cost has no deadweight loss since it is equivalent to a lump sum transfer. After removing the mechanical effect, we find that the uniform increase in potential duration has a fiscal externality of \$0.425, whereas the fiscal externality of a pivot is only \$0.241. For comparison, a review of the literature found that the fiscal externality of increasing potential duration ranges between \$0.11 and \$2.13, with a median of \$0.60 (Schmieder and von Wachter, 2016).

The comparison between a uniform and redistributive increase in potential duration is informative about which policy has a more positive effect on welfare. This is useful if policymakers already intend to expand the unemployment insurance system. However, our results do not claim that either policy is socially optimal. For example, even if pivoting the benefit schedule increases welfare more than a uniform increase, both policies may negatively affect aggregate welfare. Although our results are not informative about whether a social planner should expand and shrink the UI system, the model does provide a tax-neutral welfare-increasing policy.

Theorem 2 shows that if the fiscal externality increases with insurable hours and the value of insurance decreases, then there is a pivot point at which the social planner can rotate the benefit schedule and increase welfare without changing net costs. We calculate this point using equation (6) and report the estimate in the bottom row of Table 5. This statistic is simply the cost-weighted average insurable hours. We find that policymakers can increase welfare by rotating the potential duration schedule at 1,234 hours, giving more benefits to all claimants to the left of that cutoff and taking away benefits from claimants to the right.

V Conclusion

This paper studies whether the potential duration of unemployment insurance benefits should depend on workers' prior labor market attachment. Although many unemployment insurance systems allocate more weeks of benefits to workers with longer employment histories, there has been little empirical evidence on whether this common design improves social welfare.

We develop a sufficient-statistics framework showing that the optimal relationship between benefit duration and prior work history depends on two objects: how the value of insurance and the fiscal externality of extending benefits vary across workers. We estimate the relevant fiscal effects using 26 regression discontinuities in the Canadian unemployment insurance system, where potential benefit duration increases discretely with prior hours worked. While an additional week of potential duration increases unemployment duration by roughly 0.3 weeks on average, we find substantial heterogeneity across claimants. In particular, the mechanical cost declines sharply with work history whereas fiscal externalities increase with prior labor market attachment.

These heterogeneous responses have important welfare implications. Under standard assumptions that workers with weaker labor market attachment place a higher value on insurance, our estimates imply that a more redistributive expansion in benefit duration dominates a uniform increase in potential duration. Moreover, we show that a social planner can increase social welfare without increasing program costs by reallocating weeks of potential duration from workers with the longest employment histories to those with shorter work histories.

More broadly, our findings suggest that the design of social insurance depends not only on the overall generosity of benefits, but also on how generosity is allocated across recipients. While much of the literature has focused on determining the optimal level of unemployment insurance, our results highlight the importance of targeting benefits toward workers for whom the social value of insurance is highest and the efficiency costs are lowest. Documenting the characteristics that most accurately predict claimants' response to expansions in UI benefits would be a fruitful avenue for future research.

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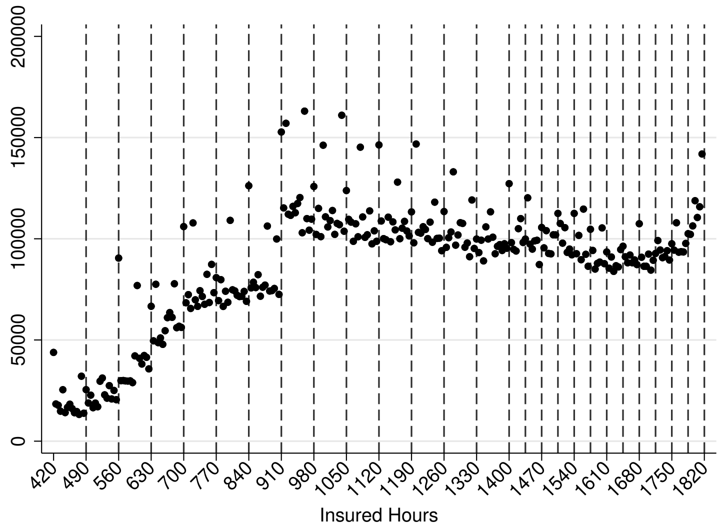
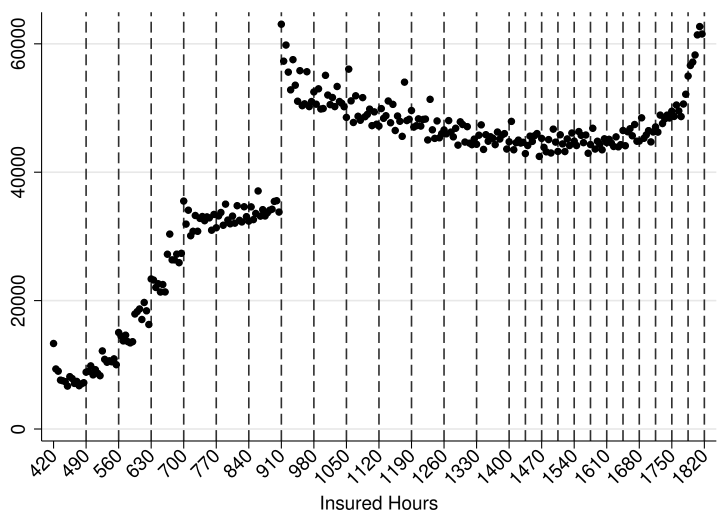


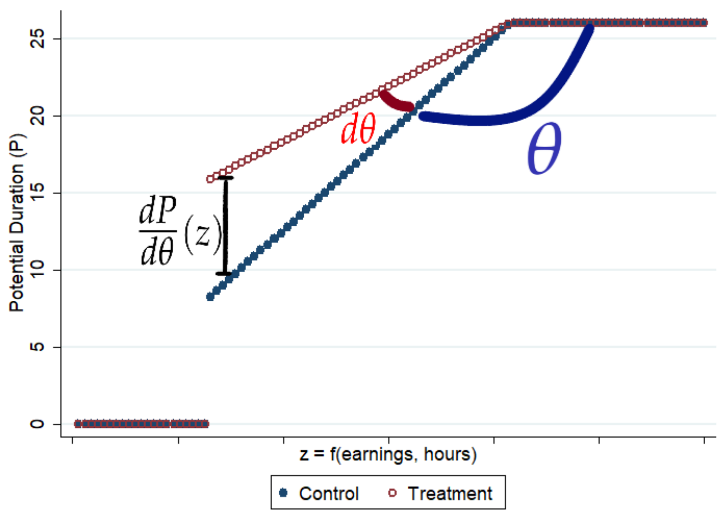
(a) All Claimants



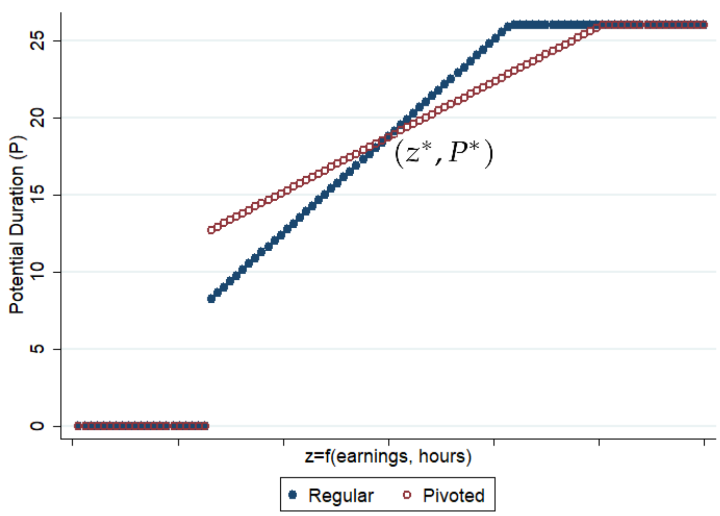
(b) Restricted Sample

Figure 1: Distribution of Insurable Hours

Notes. Panel (a) shows the distribution of insurable hours among all claimants in the data. Panel (b) drops claimants with insurable hours at multiples of 5, teachers/police/military, and new entrants and re-entrants.



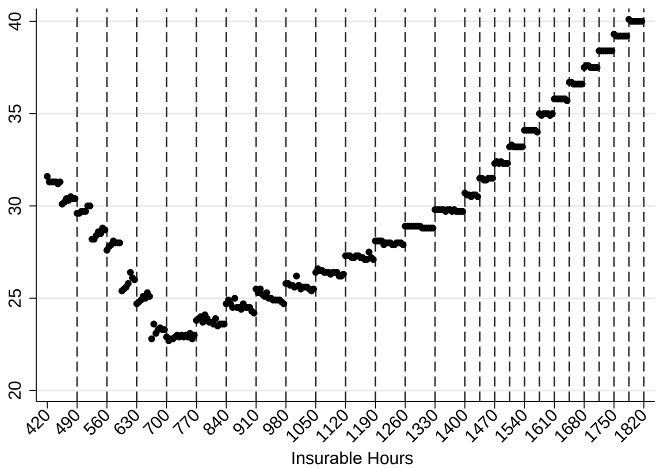
(a) Flattening the PD Schedule



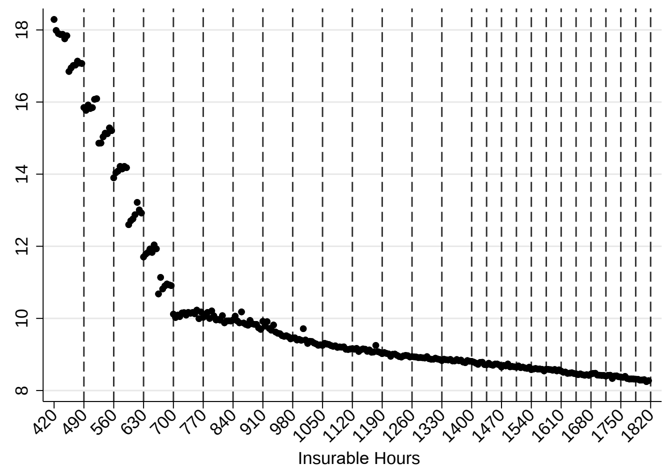
(b) Redistributing UI Benefits

Figure 2: Conceptual Policy Experiments

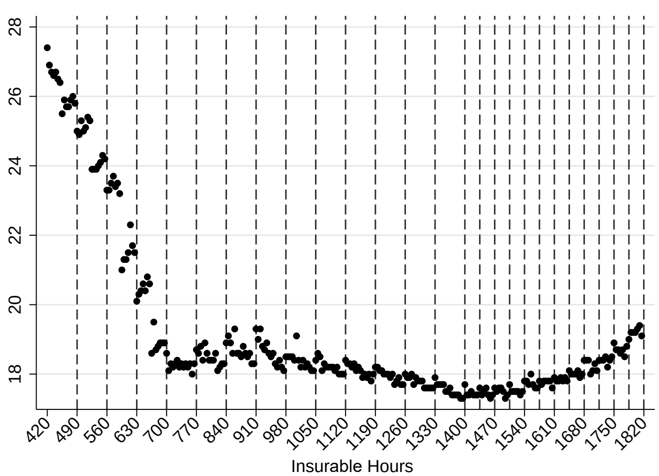
Notes. Panel (a) shows a hypothetical policy to flatten the potential duration schedule as a function of claimants' past work history. Panel (b) shows a hypothetical policy where the social planner redistributes weeks of UI benefits from claimants with long work histories to those with short ones.



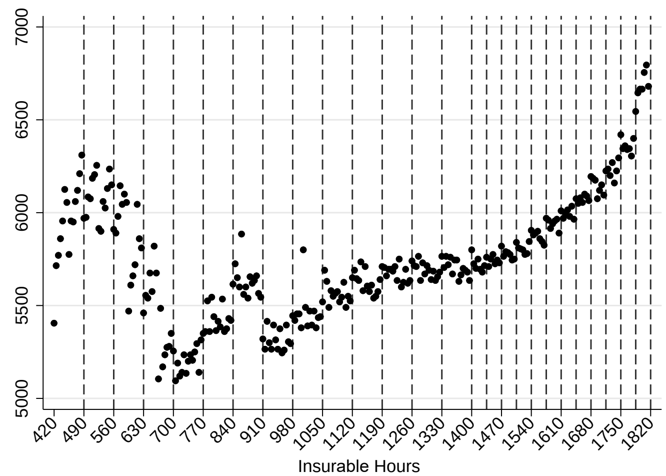
(a) Potential Duration of Benefits



(b) Regional Unemployment Rate



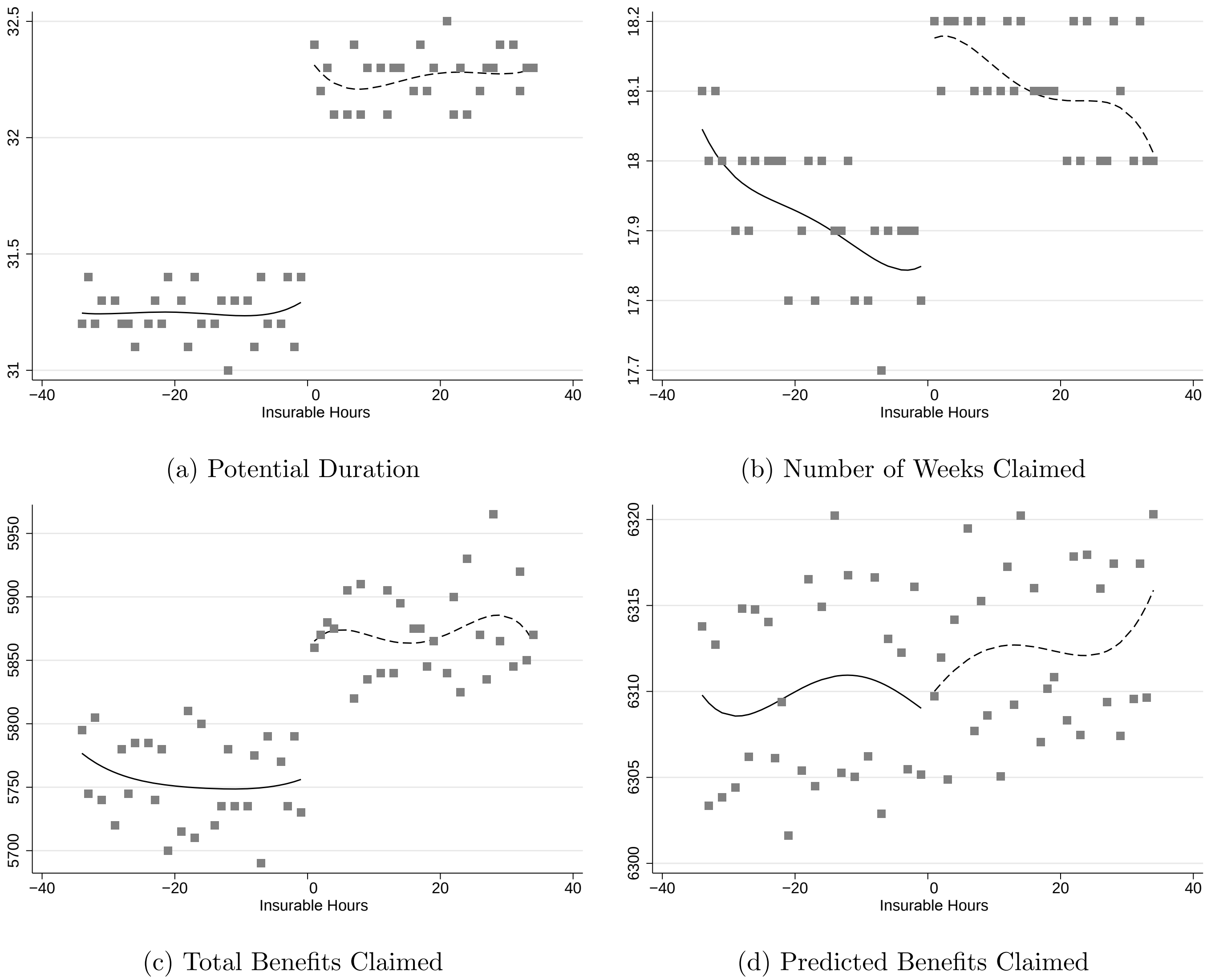
(c) Number of Weeks Claimed



(d) Total Amount of Benefits Claimed

Figure 3: Claimant Characteristics Conditional on Insurable Hours

Notes. This figure shows the potential duration, regional unemployment rate, the number of weeks claimed, and the amount of benefits claimed averaged across all claims for each insurable hour.



(a) Potential Duration

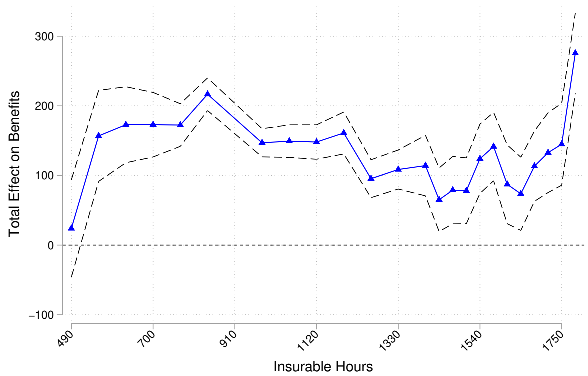
(b) Number of Weeks Claimed

(c) Total Benefits Claimed

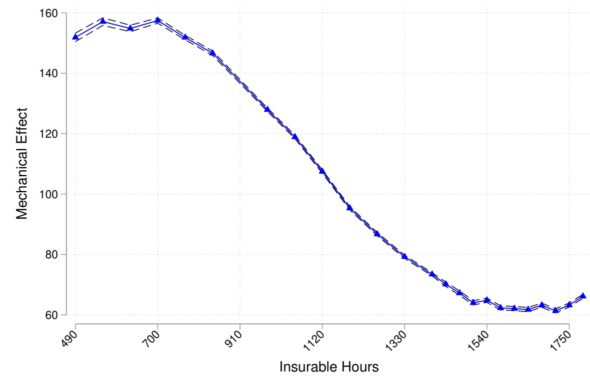
(d) Predicted Benefits Claimed

Figure 4: Stacked Regression Discontinuity

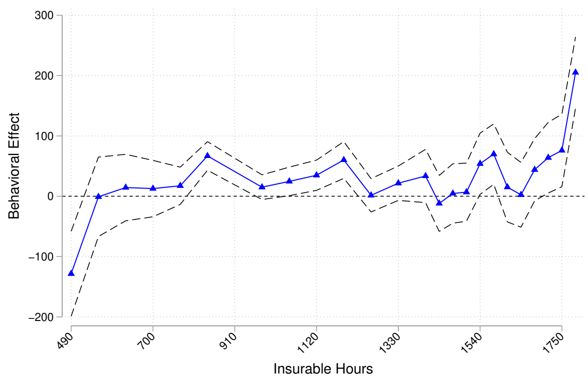
Notes. This figure plots four outcome variables as a function of insurable hours from a stacked sample that centers insurable hours of each stack at a discontinuity in the potential duration schedule. Panel (a) plots the average potential duration as a function of the recentered insurable hours. Panels (b) and (c) plot the average number of weeks of UI claimed and amount of benefits claimed over the unemployment spell, respectively. Panel (d) plots a placebo test using predicted benefits claimed over the spell, where the predicted value is computed from an OLS regression of benefits claimed on gender, age, age-squared, with employer and industry fixed effects.



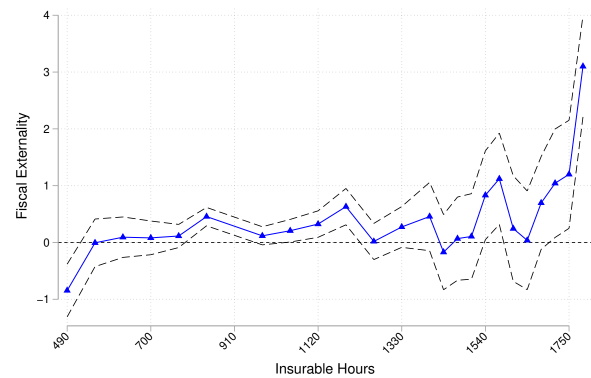
(a) Total Effect



(b) Mechanical Effect



(c) Behavioral Effect



(d) Fiscal Externality

Figure 5: Heterogeneity in Mechanical and Behavioral Effects

Notes. This figure compares the effects of a one week increase in potential duration along the distribution of insurable hours. Panel (a) plots the total effect on the government's budget, estimated from a regression discontinuity design. Panel (b) plots the mechanical cost assuming that behavior remains constant. Panel (c) plots the behavioral effect, computed as the difference between the total and mechanical effect. Panel (d) plots the ratio of the behavioral and mechanical effects.

Table 1: Number of weeks of regular UI benefits payable by regional rate of unemployment

Insurable Hours	**Regional Rate of Unemployment**											
	≤ 6%	**6.1-7%**	**7.1-8%**	**8.1-9%**	**9.1-10%**	**10.1-11%**	**11.1-12%**	**12.1-13%**	**13.1-14%**	**14.1-15%**	**15.1-16%**	**> 16%**
420 to 454	0	0	0	0	0	0	0	0	26	28	30	32
455 to 489	0	0	0	0	0	0	0	24	26	28	30	32
490 to 524	0	0	0	0	0	0	23	25	27	29	31	33
525 to 559	0	0	0	0	0	21	23	25	27	29	31	33
560 to 594	0	0	0	0	20	22	24	26	28	30	32	34
595 to 629	0	0	0	18	20	22	24	26	28	30	32	34
630 to 664	0	0	17	19	21	23	25	27	29	31	33	35
665 to 699	0	15	17	19	21	23	25	27	29	31	33	35
700 to 734	14	16	18	20	22	24	26	28	30	32	34	36
735 to 769	14	16	18	20	22	24	26	28	30	32	34	36
770 to 804	15	17	19	21	23	25	27	29	31	33	35	37
805 to 839	15	17	19	21	23	25	27	29	31	33	35	37
840 to 874	16	18	20	22	24	26	28	30	32	34	36	38
875 to 909	16	18	20	22	24	26	28	30	32	34	36	38
910 to 944	17	19	21	23	25	27	29	31	33	35	37	39
945 to 979	17	19	21	23	25	27	29	31	33	35	37	39
980 to 1,014	18	20	22	24	26	28	30	32	34	36	38	40
1,015 to 1,049	18	20	22	24	26	28	30	32	34	36	38	40
1,050 to 1,084	19	21	23	25	27	29	31	33	35	37	39	41
1,085 to 1,119	19	21	23	25	27	29	31	33	35	37	39	41
1,120 to 1,154	20	22	24	26	28	30	32	34	36	38	40	42
1,155 to 1,189	20	22	24	26	28	30	32	34	36	38	40	42
1,190 to 1,224	21	23	25	27	29	31	33	35	37	39	41	43
1,225 to 1,259	21	23	25	27	29	31	33	35	37	39	41	43
1,260 to 1,294	22	24	26	28	30	32	34	36	38	40	42	44

Continued on next page

Table 1 – continued from previous page

Hours	≤ 6%	6.1-7%	7.1-8%	8.1-9%	9.1-10%	10.1-11%	11.1-12%	12.1-13%	13.1-14%	14.1-15%	15.1-16%	> 16%
1,295 to 1,329	22	24	26	28	30	32	34	36	38	40	42	44
1,330 to 1,364	23	25	27	29	31	33	35	37	39	41	43	45
1,365 to 1,399	23	25	27	29	31	33	35	37	39	41	43	45
1,400 to 1,434	24	26	28	30	32	34	36	38	40	42	44	45
1,435 to 1,469	25	27	29	31	33	35	37	39	41	43	45	45
1,470 to 1,504	26	28	30	32	34	36	38	40	42	44	45	45
1,505 to 1,539	27	29	31	33	35	37	39	41	43	45	45	45
1,540 to 1,574	28	30	32	34	36	38	40	42	44	45	45	45
1,575 to 1,609	29	31	33	35	37	39	41	43	45	45	45	45
1,610 to 1,644	30	32	34	36	38	40	42	44	45	45	45	45
1,645 to 1,679	31	33	35	37	39	41	43	45	45	45	45	45
1,680 to 1,714	32	34	36	38	40	42	44	45	45	45	45	45
1,715 to 1,749	33	35	37	39	41	43	45	45	45	45	45	45
1,750 to 1,784	34	36	38	40	42	44	45	45	45	45	45	45
1,785 to 1,819	35	37	39	41	43	45	45	45	45	45	45	45
1,820 and more	36	38	40	42	44	45	45	45	45	45	45	45

Table 2: Summary Statistics

	Full Sample	Final Sample
Insurable Hours	1,380	1,231
	(408.67)	(354.10)
Weekly Earnings	640	600
	(212.73)	(230.39)
Unemployment Rate	9.16	9.37
	(4.02)	(4.16)
Weekly Benefit Amount	355	330
	(117.09)	(126.76)
Potential Duration (Weeks)	33.9	29.7
	(9.79)	(8.23)
Weeks of Benefits Received	19.8	18.4
	(13.46)	(10.99)
Share Exhaust Benefits	0.30	0.32
	(0.46)	(0.47)
Total Benefits Paid	6,495	5,735
	(5,310.42)	(4,381.42)
Observations	33,677,140	10,952,970

Note: This table reports the average characteristics of the universe of claims and the sample that drops claimants with insurable hours at multiples of 5, teachers/policy/military, and new entrants and re-entrants. The first three rows report the factors that determine the amount of benefits available to the claimant: insurable hours, average weekly earnings, and the unemployment rate at the time of filing. The next two rows report the amount of benefits claimants receive per week and the potential weeks of benefits. The last three rows report the duration of benefits that claimants actually take-up, the share who exhaust their benefits, and the total amount of benefits received over the duration of their claim. Standard deviations are reported in brackets.

Table 3: Regression Discontinuity Estimates

	(1) Potential Duration	(2) Weeks Received	(3) Total Benefits	(4) Total Benefits
Z	1.044***	0.314***	100.7***	
	(0.0253)	(0.0359)	(14.26)	
D				96.49***
				(13.29)
Polynomial controls	Yes	Yes	Yes	Yes
Observations	11,071,975	11,071,975	11,071,975	11,071,975

Note: This table reports the regression discontinuity estimates from equation (9). Columns (1)-(3) show reduced form estimates of the effect of crossing a policy threshold. Column (4) shows the IV estimate of the effect of raising the potential duration by 1 week, where potential duration is instumented by an indicator for being to the right of a threshold. All regressions control for a polynomial of degree 4 in the number of insurable hours that varies above and below the discontinuities. Standard errors in parentheses are clustered by UI claim. $^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$.

Table 4: Trend in Fiscal Externality by Insurable Hours

	All Thresholds			Drop Outliers		
	(1)	(2)	(3)	(4)	(5)	(6)
Hours	0.00113***	0.00114***	0.00096***	0.00065***	0.00066***	0.00094***
	(0.00016)	(0.00017)	(0.00024)	(0.00019)	(0.00019)	(0.00028)
Non-overlapping bandwidth			Yes			Yes
Proportional weighting		Yes	Yes		Yes	Yes

Note: This table reports the slope from a regression of the per-threshold fiscal externality estimate on insurable hours. Columns (1)-(3) include all thresholds except 910. Columns (4)-(6) additionally drops thresholds 490, 700 and 1785. Columns (1) and (4) assigns equal weight to each threshold and allows each claim to be used in estimating the discontinuity to its left and its right thresholds. Columns (2) and (5) weighs each threshold according to the number of claimants within the bandwidth. Columns (3) an (6) halves the bandwidth to prevent sample overlap between adjacent thresholds. $^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$.

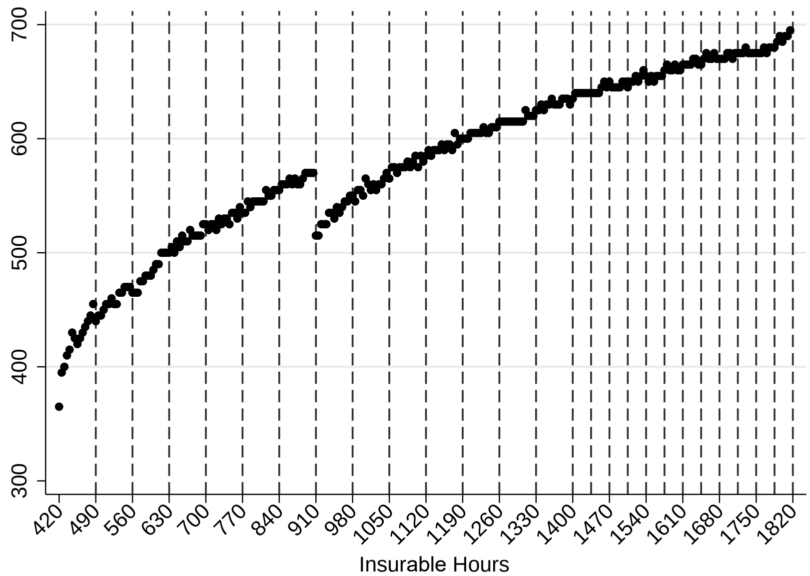
Table 5: Welfare Statistics: Uniform Increase and Pivot in Potential Duration

	Total Cost	Mechanical Cost	Behavioral Cost	Fiscal Externality
Panel A: Uniform Increase				
	136.29***	95.62***	40.67***	0.425***
	(4.15)	(0.07)	(4.15)	(0.043)
Panel B: Pivot				
	143.57***	115.71***	27.86***	0.241***
	(4.19)	(0.09)	(4.19)	(0.036)
Tax-neutral pivot point z^*		1,234 insured hours		

Note: Panel A reports the density-weighted average mechanical cost, behavioral cost, total cost, and fiscal externality of a uniform one-week increase in potential duration across all thresholds. Panel B reports the corresponding statistics for a pivot in the potential duration schedule, normalized to the same average increase in potential duration as the uniform increase. Dollar amounts are in Canadian dollars per claimant. The fiscal externality is the ratio of behavioral to mechanical cost. The tax-neutral pivot point z^* is the mean insured hours weighted by the total cost at each discontinuity. This is the unique pivot point at which rotating the schedule raises welfare without increasing taxes (Theorem 2). Standard errors are calculated using the delta method, assuming independence across thresholds. $^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$.

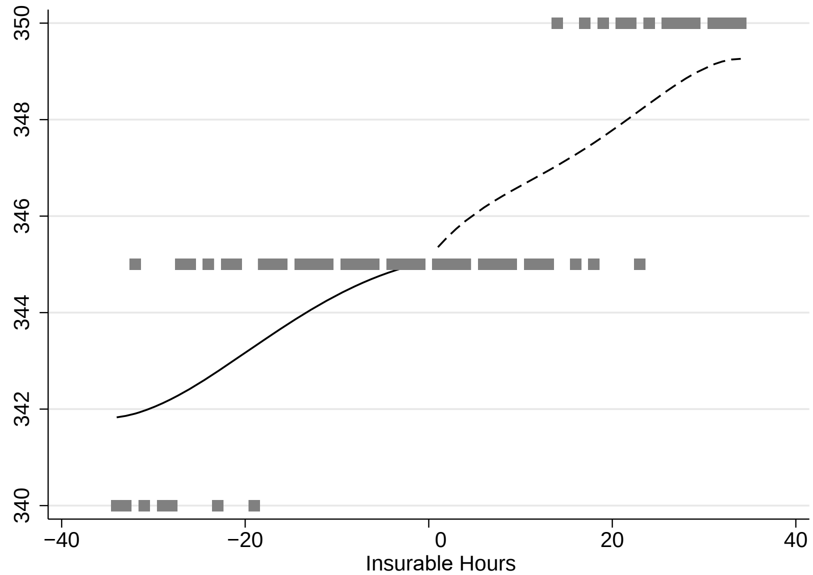
Appendix A. Additional figures and tables

Appendix Figure A.1: Weekly Earnings in Best Weeks



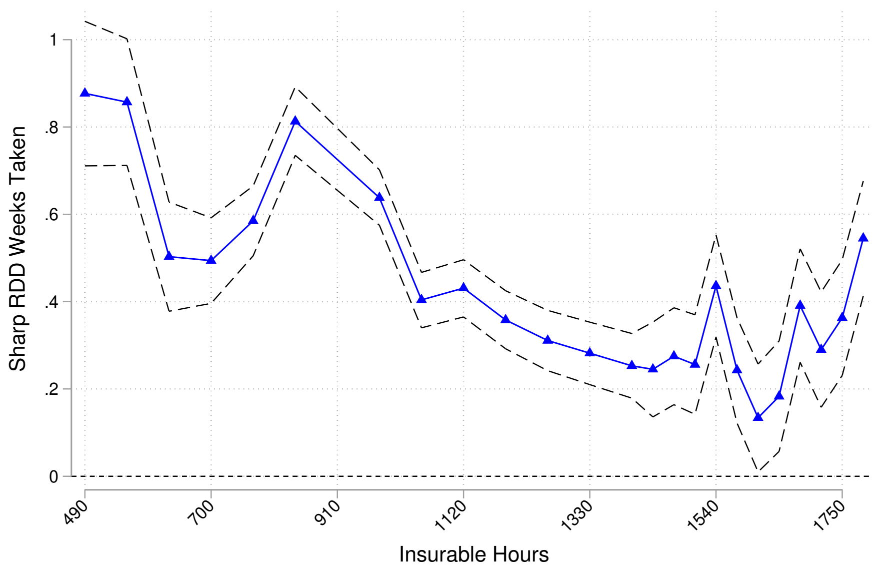
Note: This figure plots claimants' average weekly earnings in the year before unemployment as a function of their insurable hours.

Appendix Figure A.2: Pooled RD Plot: Weekly Benefit Amount



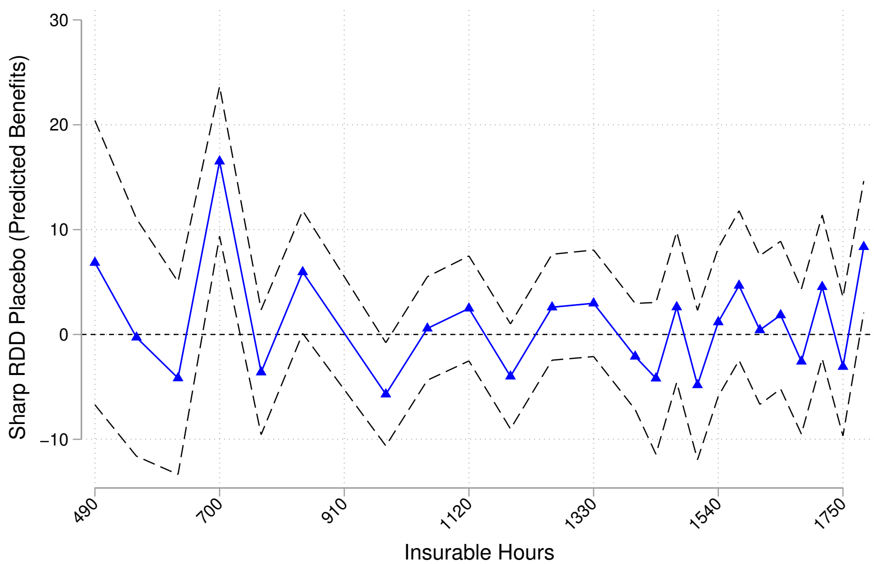
Note: This figure plots claimants' weekly benefit amount as a function of their insurable hours from a stacked sample that re-centers hours of each stack at a discontinuity in the potential duration schedule. Note: means by hour are rounded to the nearest multiple of $5 (based on data release rules), however the regression line on either side of the cutoff is computed from the unrounded data.

Appendix Figure A.3: Heterogeneous Effect on Number of Weeks of Benefits Claimed

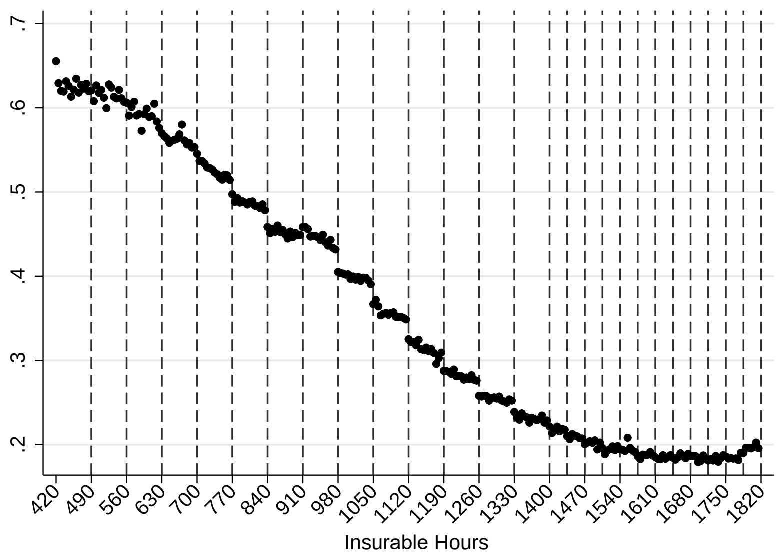


Note: This figure plots RDD estimates of the effect of a one week increase in potential duration on the number of weeks of UI benefits claimed.

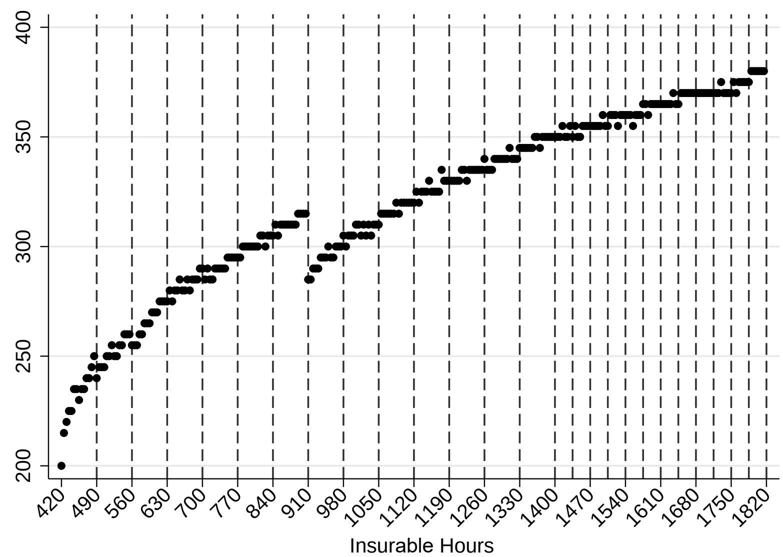
Appendix Figure A.4: Heterogeneous Effect on Predicted Benefits Claimed



Note: This figure plots RDD estimates of the effect of a one week increase in potential duration on a covariate index, computed as the predicted value from an OLS regression of benefits claimed on gender, age, age-squared, with employer and industry fixed effects.



(a) Share of Claims Exhausted



(b) Weekly Benefit Amount

Appendix Figure A.5: Heterogeneity in Exhaustion Rate and Weekly Benefit Amount

Note: Panel (a) plots the share of claims that exhaust all weeks of benefits as a function of their insurable hours. Panel (b) plots the average weekly benefit amount as a function of insurable hours.

Appendix B. Comparison to Other UI Systems

The unique feature of our institutional setting is that we are able to examine how the causal impact of increasing the potential duration of benefits differs across workers with different work histories. This rich heterogeneity enables us to test whether it is more cost effective to increase benefits for claimants with low or high labor market attachment.

To illustrate the advantage of our methodology, Figure B.1 compares common empirical designs used in the literature. Panel (a) shows a general benefits schedule that is used across many states in the US. Similar to Canada, the United States generally offers workers more weeks of potential duration if they were more active in the labor market prior to unemployment, as measured by a function of their hours and earnings. Unlike Canada, most US systems use a continuous function rather than a step function in their determination of potential duration, so workers can claim less than a full week's benefits. Previous studies have leveraged this continuous feature to identify the effects of increasing potential duration using a regression kink design (Landais, 2015).[14] By construction, this empirical strategy estimates a local average treatment effect specifically for workers at the margin of qualifying for the maximum weeks of benefits, and is less informative about workers with less labor force attachment.

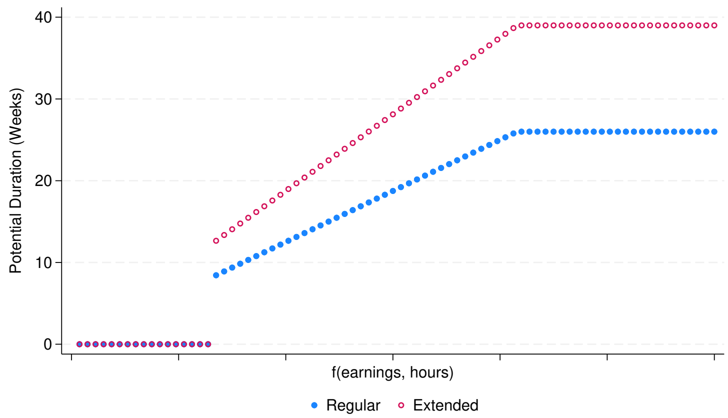
In addition to a kink feature, many states also increase their potential duration when the unemployment rate exceeds a certain threshold. Although there is some variation across states, most define these "extended benefits" as an additional 50% of claimants' regular weeks of benefits (e.g. *Texas Labor Code, Chapter 209: Extended Benefits*, 2025). As a result, estimates of the effects of the extended benefits program in the United States capture a weighted average effect of raising the potential duration of benefits across workers with different work histories (Farber et al., 2015; Farber and Valletta, 2015; Chodorow-Reich et al., 2019).

[14] Card et al. (2015b) and Lee et al. (2021) likewise use a RKD to estimate the effect of increasing the weekly benefit amount.

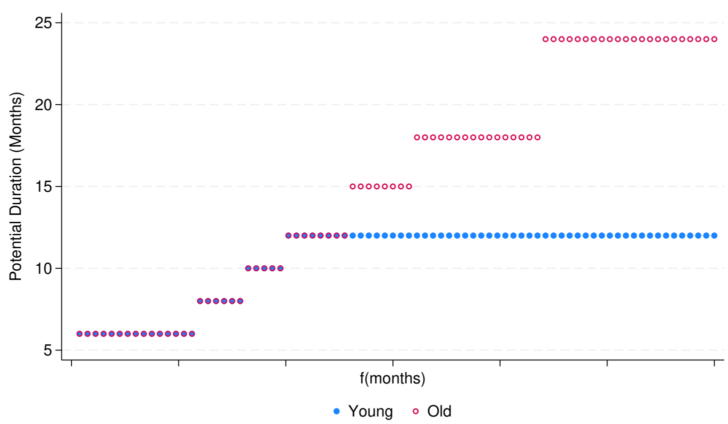
Panel (b) shows a typical potential duration schedule for countries in Europe. Benefits tend to be far more generous in Europe than either the U.S. or Canada, with potential duration being measured in months rather than weeks. However, similar to Canada, the potential duration of benefits often follows a step function conditional on claimants' tenure at their previous job. Previous studies have used these discontinuities to estimate the effect of increasing potential duration on job search in Austria (Card et al., 2007a) and France (Le Barbanchon, 2016). However, given the scarce number of discontinuities, especially among prime working age adults, these papers do not focus on the heterogeneity in estimates.

Aside from the discontinuity by job tenure, many European systems also have a discontinuity by age. Prior work has used the sharp discontinuity in maximum potential duration at age thresholds to identify the effects of increasing the length of UI benefits (Lalive, 2008; Schmieder et al., 2012, 2016). However, these age-discontinuities only affect the *maximum* potential duration, not the entire schedule. As a result, papers using this type of variation have intentionally restricted their analysis to individuals who qualify for the maximum. A similar restriction is also imposed by studies using sudden cuts to the maximum potential duration in the United States (Johnston and Mas, 2018).

In summary, most UI systems define the potential duration of benefits as a function of workers' labor market attachment. However, there is scarce empirical evidence for whether such a strategy is socially optimal as existing studies lack sufficient variation along the distribution of claimants' work history. By estimating treatment effects along the hours distribution, we can simulate the fiscal impact of flattening the benefits schedule.



(a) U.S. Regular vs. Extended Benefits



(b) Europe Discontinuity by Age and Tenure

Appendix Figure B.1: Types of Potential Benefit Schedules

Notes. Panel (a) shows a stylized example of potential duration as a function of claimants' past work history, separately for periods of regular and extended benefits. Panel (b) shows a stylized example of potential duration as a function of claimants' work history, separately by young and old workers.

Appendix C. Regression Discontinuity on Subsample

In this section, we show that dropping insurable hours at multiples of 5 or other point masses in the hours distribution yields a well-defined population upon which to conduct our RDD analysis. Suppose that claims i are classified by two latent categories: *genuine* (indicated by $G_i = 1$) and *non-genuine* (indicated by $G_i = 0$). For claims with $G_i = 1$, we observe the true number of (insurable) hours X_i^* in the insured hours variable X_i, and for claims with $G_i = 0$, we observe a rounded version of X_i^* or an arbitrary number indicating "typical" hours. For example, $G_i = 0$ includes claimants whose hours are approximated to a round number because the employer does not track their hours.

We want to allow for the possibility that G_i is correlated with potential outcomes $Y_i(w)$, where w is weeks of UI. For example, salaried workers or the armed forces may draw different amounts of benefits than hourly workers. We assume that $F_{X|G=0}$ consists of a set of point-masses only, (e.g. at round numbers), in some set $\mathcal{X} = supp\{F_{X|G=0}\}$. Meanwhile $F_{X^*|G=1}$ is continuously distributed. This means that the set $\mathcal{X}$ is identified, by looking for the point masses in F_X.

Then suppose we drop these point-mass values of X from the sample. The resulting distribution then yields the conditional distribution for the $G_i = 1$ individuals:

$$F_{X|X\notin\mathcal{X}}(x) = \frac{P(X_i \leq x, X_i \notin \mathcal{X}, G_i = 1)}{P(G_i = 1)} = \frac{P(X_i \leq x, G_i = 1)}{P(G_i = 1)} = F_{X^*|G=1}(x)$$

since $G_i = 1(X \notin \mathcal{X})$, with probability one, then using that X_i is continuously distributed given $G_i = 1$.

Meanwhile, for any $x \notin \mathcal{X}$ and $V_i \in \{Y_i, W_i\}$ (either benefit draw Y_i or maximum benefit weeks W_i):

$$\mathbb{E}[V_i|X_i = x, X_i \notin \mathcal{X}] = \mathbb{E}[V_i|X_i = x, G_i = 1]$$

From this, it follows that the fuzzy RDD estimand on the "no-spikes" sample recovers the fuzzy RDD estimand on the latent "genuine" subsample ($G_i = 1$):

$$\frac{\mathbb{E}[Y_i|X_i = t^+, X_i \notin \mathcal{X}] - \mathbb{E}[Y_i|X_i = t^-, X_i \notin \mathcal{X}]}{\mathbb{E}[W_i|X_i = t^+, X_i \notin \mathcal{X}] - \mathbb{E}[W_i|X_i = t^-, X_i \notin \mathcal{X}]} = \frac{\mathbb{E}[Y_i|X_i = t^+, G_i = 1] - \mathbb{E}[Y_i|X_i = t^-, G_i = 1]}{\mathbb{E}[W_i|X_i = t^+, G_i = 1] - \mathbb{E}[W_i|X_i = t^-, G_i = 1]}$$

where t is a threshold between two maximum weeks values. This yields a conditional LATE (conditional on $G_i = 1$) provided that $\mathbb{E}[Y_i(w)|X_i = x, G_i = 1]$ and complier shares are continuous at $x = t$.

Appendix D. Proofs for Sufficient Statistics Model

D.a Theorem 1

Let $P_\theta(z) \equiv \frac{\partial P(z,\theta)}{\partial \theta}$ be the change in potential duration at work history z from a marginal pivot, and suppose the pivot is a net expansion in benefits, $\int MC_{P_z}\, P_\theta(z) f(z)\, dz > 0$. Define

$$\phi(z) = \frac{E\big[G' V_1 \frac{\partial B}{\partial P} \mid z\big]}{\gamma\, E\big[\frac{\partial B}{\partial P} \mid z\big]}, \qquad e(z) = \frac{BC_{P_z}}{MC_{P_z}}, \qquad \nu(z) = \phi(z) - e(z).$$

Suppose (i) $P_\theta(z)$ is decreasing in z so that the pivot gives more benefits for claimants with less work history; (ii) the per-dollar value of insurance $\phi(z)$ is decreasing in z; (iii) the local fiscal externality $e(z)$ is nondecreasing in z, and (iv) the disutility of taxation, $E\big[G' \cdot V_2 \cdot \frac{\partial T/\partial \tau}{dE[B-T]/d\tau} \big| Z = z\big]$, is the same for all z and hence equal to γ. Then pivoting the potential duration schedule raises social welfare per dollar of mechanical transfer strictly more than a uniform increase:

$$\frac{dW/d\theta}{MC_\theta} > \frac{dW/dP}{MC_P}, \qquad \text{and} \qquad \frac{BC_\theta}{MC_\theta} < \frac{BC_P}{MC_P}.$$

Proof. Conditioning equation (2) on z and writing $MC_{P_z} = E[\frac{\partial B}{\partial P} \mid z]$ gives the change in welfare from increasing potential duration at a particular z:

$$\frac{1}{\gamma} \frac{dW}{dP}(z) = MC_{P_z}\big(\nu(z) - 1\big) \tag{10}$$

By equation (4), the welfare per dollar of mechanical transfer from pivoting the potential duration schedule is the weighted average of equation (10):

$$\begin{aligned}
\frac{1}{\gamma MC_\theta}\frac{dW}{d\theta} &= \int \frac{1}{\gamma MC_\theta}\frac{dW}{dP(z,\theta)}\frac{dP(z,\theta)}{d\theta}f(z)dz \\
&= \int \frac{1}{MC_\theta}MC_{P_z}\big(\nu(z)-1\big)\frac{dP(z,\theta)}{d\theta}f(z)dz \quad \text{by eq. (10)} \\
&= \int \big(\nu(z)-1\big)\frac{MC_{P_z}P_\theta(z)f(z)}{\int MC_{P_{z'}}P_\theta(z')f(z')dz'}dz
\end{aligned}$$

Analogously, the welfare per dollar of mechanical transfer from a uniform increase in potential duration is

$$\frac{1}{\gamma MC_P}\frac{dW}{dP} = \int \big(\nu(z)-1\big)\frac{MC_{P_z}f(z)}{\int MC_{P_{z'}}f(z')dz'}dz$$

In other words, the welfare effect of either reform is simply a weighted average of $\nu(z)-1$, the welfare effect at a given z. The weights for a pivot $w_\theta(z)$ and for a uniform increase $w_P(z)$ share the following relationship:

$$\begin{aligned}
w_\theta(z) =& \frac{MC_{P_z}P_\theta(z)f(z)}{\int MC_{P_{z'}}P_\theta(z')f(z')dz'} \\
&= \frac{MC_{P_z}P_\theta(z)f(z)}{\int MC_{P_{z'}}f(z')dz'}\frac{\int MC_{P_{z'}}f(z')dz'}{\int MC_{P_{z'}}P_\theta(z')f(z')dz'} \\
&= w_P(z)P_\theta(z)\frac{1}{E_{w_P}(P_\theta)}
\end{aligned}$$

where for a function $g(z)$ we let $E_{w_P}(g) := \int MC_{P_{z'}}g(z')f(z')dz' / \int MC_{P_{z'}}f(z')dz'$.

The difference in the welfare effects can then be represented by:

$$
\begin{aligned}
\frac{dW/d\theta}{MC_\theta} - \frac{dW/dP}{MC_P} &= E_{w_\theta}(\nu - 1) - E_{w_P}(\nu - 1) \\
&= E_{w_\theta}(\nu) - E_{w_P}(\nu) \\
&= \frac{E_{w_P}(\nu P_\theta)}{E_{w_P}(P_\theta)} - E_{w_P}(\nu) \\
&= \frac{\mathrm{Cov}_{w_P}(\nu, P_\theta) + E_{w_P}(\nu) E_{w_P}(P_\theta)}{E_{w_P}(P_\theta)} - E_{w_P}(\nu) \\
&= \frac{\mathrm{Cov}_{w_P}(\nu, P_\theta)}{E_{w_P}(P_\theta)} \\
&> 0
\end{aligned}
$$

where the last inequality follows from the assumptions of the theorem. By (ii)–(iii), ν is decreasing in z; by (i), $P_\theta(z)$ is also decreasing in z; the integrand $\mathrm{Cov}_{w_P}(\nu, P_\theta)$ is thus positive. Since the reform is expanding UI generosity, the denominator $E_{w_P}[P_\theta]$ is also positive. By a similar argument, $\frac{BC_\theta}{MC_\theta} - \frac{BC_P}{MC_P} = \mathrm{Cov}_{w_P}(e, P_\theta)/E_{w_P}[P_\theta] \leq 0$ by (i) and (iii). ■

D.b Theorem 2

Suppose z is continuously distributed on $[z_{\min}, z_{\max}]$ with density f, and:

(i) the total fiscal cost per week $c(z) \equiv MC(z) + BC(z) > 0$;

(ii) the per-dollar value of insurance $\phi(z) = \dfrac{E[G' V_1 \frac{\partial B}{\partial P} \mid z]}{\gamma\, MC(z)}$ is nonnegative and nonincreasing in z;

(iii) the local fiscal externality $e(z) = \dfrac{BC(z)}{MC(z)}$ is nondecreasing in z.

(iv) the disutility of taxation, $\gamma = E\big[G' \cdot V_2 \cdot \frac{\partial T/\partial \tau}{dE[B-T]/d\tau}\big]$, is independent of z.

Then the work history

$$
z^* = \frac{\int z \left[MC(z) + BC(z)\right] f(z)\, dz}{\int \left[MC(z) + BC(z)\right] f(z)\, dz}
$$

is the unique point at which rotating the potential-duration schedule about $\big(z^*, P(z^*)\big)$ — raising P for $z < z^*$ and lowering it for $z > z^*$ — is

1. **Tax neutral:** $\displaystyle\int \big[MC(z) + BC(z)\big]\frac{dP(z,\theta)}{d\theta} f(z)\, dz = 0$; and

2. **Welfare increasing:** $\dfrac{dW}{d\theta} \geq 0$, strictly if ϕ is strictly decreasing or e strictly increasing on a neighborhood of z^*.

Proof. Index pivots by their center through the rotation $\frac{\partial P}{\partial \theta} = z^* - z$, which is strictly decreasing in z, positive for $z < z^*$ and negative for $z > z^*$.

Tax neutrality. With $c(z) = MC(z) + BC(z)$, the budget effect of a rotation at z^* is $\mathcal{B}(z^*) = \int c(z)(z^* - z) f(z)\, dz = z^* \int cf - \int z\, cf$. Solving for $\mathcal{B}(z^*) = 0$ implies

$$z^* = \frac{\int z \big[MC(z) + BC(z)\big] f(z)\, dz}{\int \big[MC(z) + BC(z)\big] f(z)\, dz}$$

Welfare increasing. Conditioning (2) on z and factoring out $MC(z) = E[\partial B/\partial P \mid z]$ gives $\frac{1}{\gamma}\frac{dW}{dP}(z) = MC(z)\phi(z) - c(z)$, where $c = MC + BC$. Substituting into (4) with and using tax neutrality $\int c(z)(z^* - z) f\, dz = 0$, the cost term drops out:

$$\frac{1}{\gamma}\frac{dW}{d\theta} = \int MC(z)\, \phi(z)\, (z^* - z)\, f(z)\, dz \ =:\ J.$$

Writing $\phi(z) = [\phi(z) - \phi(z^*)] + \phi(z^*)$,

$$J = \underbrace{\int \big[\phi(z) - \phi(z^*)\big](z^* - z) MC(z) f(z)\, dz}_{T_1} + \phi(z^*) \underbrace{\int (z^* - z) MC(z) f(z)\, dz}_{T_2}.$$

Let $d\nu(z) = MC(z) f(z)\, dz$ (a positive measure, since $MC > 0$), with total mass $M_\nu = \int d\nu$ and mean $\bar{z}_\nu = M_\nu^{-1} \int z\, d\nu$. With this notation:

$$J = \underbrace{\int \big[\phi(z) - \phi(z^*)\big](z^* - z)\, d\nu}_{T_1} + \phi(z^*) \underbrace{\int (z^* - z)\, d\nu}_{T_2}.$$

Term T_1. Because ϕ is nonincreasing by assumption (ii), $\phi(z) - \phi(z^*)$ and $z^* - z$ have the same sign at every z, so the integrand is nonnegative and $T_1 \geq 0$.

Term T_2. Since $T_2 = M_\nu\,(z^* - \bar{z}_\nu)$ and $\phi(z^*) \geq 0$, it suffices that $z^* \geq \bar{z}_\nu$. Using the definition of z^* and $c(z) = (1 + e(z))\,MC(z)$,

$$
\begin{aligned}
z^* - \bar{z}_\nu &= \frac{\int z\,c(z)\,f(z)dz}{\int c(z)\,f(z)dz} - E_\nu(z) \qquad \text{by definition} \\
&= \frac{\int z(1+e(z))\,d\nu}{\int (1+e(z))\,d\nu} - E_\nu(z) \\
&= \frac{E_\nu(z(1+e)) - E_\nu(z)E_\nu(1+e)}{E_\nu(1+e)} \\
&= \frac{\mathrm{Cov}_\nu(z,e)}{E_\nu(1+e)} \\
&\geq \; 0
\end{aligned}
$$

where $\mathrm{Cov}_\nu(z,e) \geq 0$ because e is nondecreasing in z and ν is a positive measure. Therefore $J \geq 0$ and $\frac{dW}{d\theta} = \gamma J \geq 0$. The inequality is strict whenever ϕ is strictly decreasing on a neighborhood of z^* (making $T_1 > 0$), or e is strictly increasing and $\phi(z^*) > 0$ (making $T_2 > 0$). ■

Appendix E. Inference for the Trend in FE

In this section, we describe how we assess the statistical significance of a regression of the fiscal externality on z. Let $\widehat{FE}(z_0)$ be our estimate of the fiscal externality at a given threshold z_0. Given a large sample approximation, the distribution of $\widehat{FE}(z_0)$ is approximately $N(FE(z_0), V(z_0))$, where we have a standard error $se(z_0)$ that provides a consistent estimator of $V(z_0)$ at each z_0.

Now consider the slope from a weighted OLS fit to the $\widehat{FE}(z_0)$, with weights $w(z_0)$:

$$\hat{\beta} = \frac{\sum_{z_0} w(z_0) \cdot (z_0 - \bar{z}) \cdot \widehat{FE}(z_0)}{\sum_{z_0} w(z_0) \cdot (z_0 - \bar{z})^2}$$

where $\bar{z} := (\sum_{z_0} w(z_0) \cdot z_0)/(\sum_{z_0} w(z_0))$ is the weighted average value of z_0. In our empirical results, we consider two weighting schemes: $w(z_0) = 1/N_z$ where $N_z = 26$ is the number of thresholds, and $w(z_0) = n(z_0)/\sum_{z'} n(z')$ that weights in proportion to the number $n(z_0)$ of claims z_0 within a bandwidth of z_0.

We consider inference for the parameter

$$\beta := \frac{\sum_{z_0} w(z_0) \cdot (z_0 - \bar{z}) \cdot FE(z_0)}{\sum_{z_0} w(z_0) \cdot (z_0 - \bar{z})^2}$$

conditional on the sample distribution of insurable hours (i.e. the $w(z_0)$ are non-random, in either weighting scheme).

The limiting distribution of the $\sqrt{n}(\widehat{FE}(z) - FE(z))$ stacked as a vector over the z are jointly normal and, given that $Cov(\widehat{FE}(z), \widehat{FE}(z')) = 0$ for any $z \neq z'$ (no observations are repeated in the window around two adjacent thresholds), independent. As a consequence:

$$\hat{\beta} - \beta = \sum_{z} \frac{w(z) \cdot (z - \bar{z})}{\sum_{z'} w(z') \cdot (z' - \bar{z})^2} \cdot (\widehat{FE}(z) - FE(z))$$

is distributed approximately as $N(0, V)$, where $V = \sum_z \left\{ \frac{w(z) \cdot (z - \bar{z})}{\sum_{z'} w(z') \cdot (z' - \bar{z})^2} \right\}^2 \cdot V(z)$, and thus

$(\hat{\beta} - \beta)/se(\hat{\beta}) \sim N(0,1)$ approximately, where

$$se(\hat{\beta}) = \sqrt{\sum_z \left\{ \frac{w(z) \cdot (z - \bar{z})}{\sum_{z'} w(z') \cdot (z' - \bar{z})^2} \right\}^2 \cdot se(z_0)^2}$$

When the samples used to compute $\widehat{FE}(z)$ and $\widehat{FE}(z')$ for two adjacent thresholds $z > z'$ overlap, we argue that standard error above will tend to be conservative. In particular, the primary ingredient in $\widehat{FE}(z)$ is the regression discontinuity of benefits received $\mathcal{S}_i$ with respect to hours Z_i. If an individual with a particularly high value of $\mathcal{S}_i$ is sampled, and their Z_i puts them between z and z', then this will lead $\widehat{FE}(z)$ to become higher (this individual is on the right for the RDD around z) and $\widehat{FE}(z')$ to become *lower* (this individual is on the left for the RDD around z'). Thus we expect $Cov(\widehat{FE}(z), \widehat{FE}(z')) \leq 0$, differing from zero only when z and z' are adjacent. Since the sign of $z - \bar{z}$ and $z' - \bar{z}$ will be the same (except for the single z,z' pair that crosses $\bar{z}$), this means the expression above for V is missing negative covariance terms, and is hence no smaller than the true asymptotic variance V_{true} of $\hat{\beta}$.

$$V_{true} = V + \sum_{\substack{z,z' \\ \text{adjacent}}} \underbrace{\left\{ \frac{w(z) \cdot (z - \bar{z})}{\sum_{z''} w(z'') \cdot (z'' - \bar{z})^2} \right\} \cdot \left\{ \frac{w(z') \cdot (z' - \bar{z})}{\sum_{z''} w(z'') \cdot (z'' - \bar{z})^2} \right\}}_{\geq 0 \text{ for all but one pair}} \cdot \underbrace{\text{Cov}(\widehat{FE}(z), \widehat{FE}(z'))}_{\leq 0}$$