Unemployment Insurance after Incarceration*

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Abstract

Unemployment insurance may have different effects for workers facing severe barriers to employment. We use Polish administrative unemployment insurance records on formerly incarcerated workers to provide quasi-experimental evidence on this population, leveraging a policy discontinuity in county unemployment rates that extends potential benefit duration from 6 to 12 months. Doubling potential duration raises benefit receipt by 2.3 months but unemployment duration by only 1.3 months. Despite the longer unemployment spells, it raises transitions to employment by 4.5 percentage points, while having no detectable effect on reincarceration. The fiscal externality for formerly incarcerated workers is well below that for the Polish workforce as a whole. We also examine heterogeneity across eligibility pathways. Behavioral responses appear more muted among workers who qualified through in-prison work, with an implied fiscal externality roughly half that of workers who qualified through pre-prison employment. Extending benefits to workers with limited labor-market opportunities may therefore generate relatively modest behavioral and fiscal costs, while potentially supporting transitions to employment without increasing reincarceration.

JEL Classification Codes: *H55, J65, K42*

Keywords: *Unemployment insurance*, *potential benefit duration*, *formerly incarcerated workers*, *prison work*, *regression discontinuity*

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1 Introduction

Unemployment insurance is a major source of income support for individuals facing severe structural barriers to employment. In particular, workers leaving prison form a population of considerable interest due to their unique reintegration challenges. This marginalized workforce faces restricted access to job opportunities due to employer discrimination, social stigma, and human capital depreciation. Limited labor-market opportunities may constrain behavioral responses to more generous benefits, while also making successful reentry more difficult and increasing the risk of returning to crime. Understanding how extended benefits affect labor-market and criminal-justice outcomes for this population is therefore an important economic and policy question, yet empirical evidence remains limited.

If unemployment among formerly incarcerated workers primarily reflects limited employment opportunities, more generous benefits may ease liquidity constraints with relatively modest effects on search behavior, while supporting transitions to employment (Apel and Sweeten, 2010; Pager, 2008).[1] If, instead, their search behavior responds to benefit generosity similarly to that of the broader workforce (Chetty, 2008; Cohen and Ganong, 2026; Schmieder and von Wachter, 2017; Schmieder et al., 2012, 2016), longer entitlement may substantially prolong non-employment, potentially hindering successful reentry and increasing the risk of reoffending. Answering this question requires detailed administrative data on formerly incarcerated benefit recipients together with exogenous variation in benefit generosity. To our knowledge, no quasi-experimental evidence of this kind exists.

In this paper, we study how extending unemployment insurance affects labor market and criminal outcomes among formerly incarcerated workers, exploiting a policy discontinuity in Poland whereby individuals residing in counties with unemployment rates above 150% of the national average are eligible for 12 rather than 6 months of unemployment benefits. We apply a regression discontinuity (RD) design to administrative records on benefit spells of formerly incarcerated workers between 2009 and 2018. The setting offers a sharp threshold, while the administrative records allow us to identify formerly incarcerated benefit recipients and observe

[1] These barriers matter for criminal participation as well as for employment. Since Becker (1968) and Ehrlich (1973), a large empirical literature documents the structural determinants of offending, including employment, detection, and the returns to crime (Braakmann et al., 2024; Draca et al., 2019; Gould et al., 2002; Hjalmarsson et al., 2024; Mastrobuoni, 2020; Di Tella and Schargrodsky, 2013), which is reviewed in Draca and Machin (2015), Chalfin and McCrary (2017), Doleac (2019) and Apel (2022).

their benefit and unemployment spells. A further feature of the Polish system, which we exploit below, is that entitlement can be earned either through employment preceding incarceration or through work performed while in prison. We refer to workers qualifying by these eligibility pathways as pre-prison work qualifiers and in-prison work qualifiers, respectively.

We find that doubling potential benefit duration (PBD) raises actual benefit receipt by 2.3 months but unemployment duration by only 1.3 months. Despite the longer unemployment spells, longer entitlement also raises the probability of exiting to employment by 4.5 percentage points, while leaving re-imprisonment unchanged. The implied elasticity of unemployment duration with respect to PBD is 0.14 and substantially lower than the corresponding elasticity of 0.29 reported for the Polish workforce as a whole (Jessen et al., 2025). Because our estimate and the corresponding estimate for the broader Polish workforce exploit the same policy threshold, they provide a particularly informative comparison of the effects of longer benefit entitlement across populations, suggesting that the behavioral response and associated moral hazard are relatively modest among formerly incarcerated workers.

We also find suggestive evidence of heterogeneity across eligibility pathways. The increase in benefit receipt is similar across the two groups, while the estimated increases in both unemployment duration and transitions to employment are smaller among in-prison work qualifiers than among pre-prison work qualifiers. Although the differences between the subgroup estimates are not statistically significant, this pattern is consistent with more limited labor-market opportunities among in-prison work qualifiers, who tend to have longer prison stays. Facing greater barriers to employment, these workers appear less likely to extend their unemployment spells, but also less likely to transition into employment, in response to longer benefit entitlement, suggesting less scope for labor-market adjustment on either margin. For criminal-justice outcomes, effects on reincarceration are small and statistically insignificant for both groups.

We further quantify the fiscal consequences of longer benefit entitlement using the ratio of behavioral to mechanical costs (BC/MC) (Schmieder and Von Wachter, 2016). The ratio is 1.6 for formerly incarcerated benefit recipients overall, substantially below the corresponding estimate of 2.5 for the broader Polish workforce reported by Jessen et al. (2025). Across eligibility pathways, the ratio is 1.6 for in-prison work qualifiers and 2.8 for pre-prison work qualifiers, reinforcing the finding that behavioral responses are particularly muted among in-prison qualifiers. Assuming a 10% consumption drop, as in Jessen et al. (2025), the resulting MVPFs range from 0.4 to 0.7

for formerly incarcerated benefit recipients, with higher MVPFs among in-prison than pre-prison work qualifiers, compared with 0.3 to 0.5 for the broader Polish workforce. Thus, the welfare calculations point to lower fiscal externalities and correspondingly higher MVPFs among formerly incarcerated workers, particularly among those who qualified through in-prison work.

Our study connects two largely separate literatures. The first examines social insurance, transitional income support and recidivism, mostly for newly released prisoners in the United States. Yang (2017) shows that restoring public assistance eligibility for convicted drug offenders reduces recidivism, while Tuttle (2019) and Deshpande and Mueller-Smith (2022) show that withdrawing comparable support raises reincarceration and criminal charges, and Mueller-Smith et al. (2023) document spillovers onto the children of banned adults. Gollu and Zapryanova (2022), Aslim et al. (2022) and Palmer et al. (2019) report parallel results for health insurance and emergency financial assistance, and Ludwig and Schnepel (2025) provide a recent synthesis.[2] Related work examines how prison programs and other interventions affect subsequent employment and recidivism. Bhuller et al. (2020) show that incarceration in Norway lowers reoffending and raises employment among individuals who were not working beforehand, effects they attribute in part to in-prison programs, while Raphael (2011) surveys the labor-market barriers faced by released prisoners in the United States and interventions designed to reduce them.[3]

The second literature estimates the disincentive costs of unemployment insurance (UI) for the general workforce (Chetty, 2008; Cohen and Ganong, 2026; Schmieder and von Wachter, 2017; Schmieder et al., 2012, 2016), including for Poland (Jessen et al., 2025, 2026).[4] We connect these literatures by providing quasi-experimental evidence on how unemployment insurance affects formerly incarcerated workers, a population facing particularly limited labor-market opportunities. Our findings suggest that the behavioral and fiscal effects of UI may differ substantially

[2]Early contributions include Berk et al. (1980), Mallar and Thornton (1978) and Rauma and Berk (1987). On the timing of benefit payments see Stam et al. (2024).

[3]On prison work programs see Maguire et al. (1988), Saylor and Gaes (1996), Wilson et al. (2000), Bushway et al. (2003), Hopper (2013) and Cox (2016) for the U.S., Simon (1999) for Britain, Alós et al. (2015) for Spain and Baeza and Grau (2017) for Chile; Nur and Nguyen (2023) review this evidence, while Zanella (2025), Mastrobuoni and Terlizzese (2022), Schnepel (2018), Siwach (2018) and Agan and Makowsky (2023) provide recent quasi-experimental estimates, although not every job works (Ramakers et al., 2017). On prison conditions and re-offending see Chen and Shapiro (2007), Cook et al. (2015) and Mueller-Smith (2015) for the U.S. and Drago et al. (2011) for Italy. Apel (2025) reviews social spending and recidivism more broadly.

[4]See Cohen and Ganong (2026) for a meta-analysis of dis-employment effects across contexts, Nekoei and Weber (2017) and Moore and McQuillan (2025) on whether benefit extensions improve job quality, and Krueger and Meyer (2002), Meyer (2002) and Lopes (2022) for earlier reviews.

across populations, with especially muted responses among workers facing greater barriers to employment.

The paper proceeds as follows. Section 2 describes the institutional setting and the data, Section 3 sets out the empirical strategy, Section 4 reports the results, Section 5 presents the welfare calculations, and Section 6 concludes. Additional exhibits are in the online Appendix.

2 Institutional Setting and Data

Formerly incarcerated workers are subject to special eligibility conditions if they register within 30 days of release from prison (Figure 1, panel (a)). To qualify for unemployment benefits, they must have accumulated at least one year of eligible employment over a look-back period that combines their entire incarceration spell, regardless of its length, with the 18 months preceding imprisonment. Employment during these two periods is counted jointly toward the one-year requirement.[5] Both employment during imprisonment and employment in the 18 months preceding imprisonment count toward the one-year requirement, and employment across the two periods is aggregated. Not every type of employment counts: the individual must have earned at least the minimum wage and have had social contributions paid into the Labor Fund, which finances unemployment benefits. At registration the employment office checks with day-level precision whether these conditions have been met, and only then grants the right to benefits.

Although Polish law stipulates that every inmate should undertake employment, framed both as a legal obligation and as part of rehabilitation, its implementation diverges markedly from these provisions.[6] In practice, prison work is not universally accessible. Qualitative evidence suggests that the opportunity to work is perceived as a privilege or reward: inmates must demonstrate good behavior and formally apply for a position, and initial assignments typically involve unpaid labor within internal prison services, with the possibility of promotion to paid employment (Felczak, 2025). Access to paid employment therefore occurs only when behavioral, institutional and market-related conditions align (Felczak and Wilamowska, 2025).

The PBD is determined at the county level and depends on how the local unemployment rate compares with the national average. The rate on 30 June of year $t-1$ determines the regime for

[5] Under standard conditions, an individual must have worked for at least one year within the 18 months prior to registering with the employment office, and must have been dismissed from their most recent job as well as from any other position ending within the six months preceding registration.

[6] Articles 67 and 121 of the Executive Penal Code Act of 6 June 1997.

year t. These figures are published by Statistics Poland at the end of September each year, so the applicable regime is known in advance.[7] From February 2009 until May 2025, the relative unemployment rate threshold allowing county authorities to extend benefit payments from 6 to 12 months was set at 150%.[8] Panel (b) of Figure 1 shows how counties were distributed across regimes during 2009–2018. The basic benefit amount is indexed annually to inflation and applies uniformly to all recipients.[9]

We use administrative data from the universe of public employment offices in Poland, covering formerly incarcerated workers who registered as unemployed between February 2009 and December 2018.[10] Focusing on this period ensures a uniform institutional environment with respect to potential benefit duration. Moreover, to maintain institutional homogeneity, we restrict the analysis to individuals younger than 50 years of age. The data are collected at registration and include incarceration histories, including previous spells of imprisonment. The data cover a substantial share of released prisoners. Former inmates have institutional incentives to register upon release and do so within nine days on average.[11] A comparison with annual statistics published by the Polish Prison Service suggests that the public employment office data cover approximately 20.7% of the prison population recorded at the end of a given calendar year.

The unit of observation is an unemployment spell; however, repeated spells are uncommon. Only 83 individuals in the full sample and 23 individuals in the estimation sample contribute more than one spell, representing less than 1% of the respective samples. Given this near-complete correspondence between spells and individuals, we refer to individuals throughout the paper for ease of exposition, although all analyses are conducted at the unemployment-spell level. Deregis-

[7] Jessen et al. (2026) document strategic behavior and temporal substitution by workers at the turn of the year. This is not a concern in our setting, since formerly incarcerated workers must register within 30 days of release to qualify under the more favorable eligibility conditions.

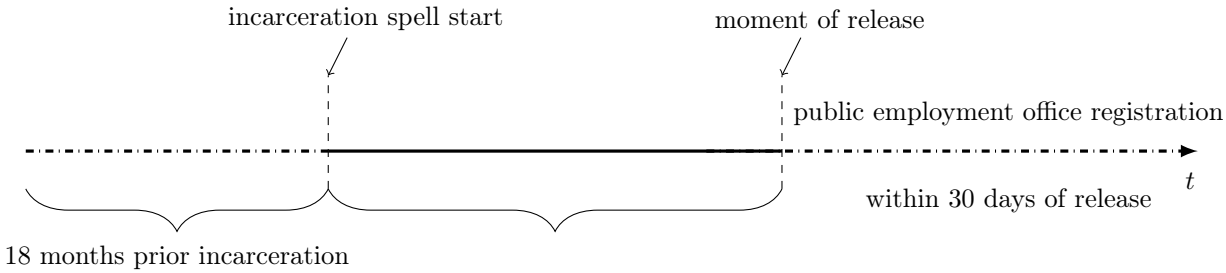
[8] The statutory threshold applies to counties whose unemployment rate exceeds 150% of the national average, and equation (1) follows this convention. No county-year in our sample sits exactly at 150. The threshold had previously been set at 125%, and at 100% before June 2004.

[9] In its current form the benefit amounted to PLN 1,662 (approximately EUR 385) per month during the first 90 days of receipt between 1 June 2024 and 31 May 2025, falling by approximately 21.5% thereafter. Nominal benefit levels were lower over our sample period. OECD indicators classify the Polish system as of 2025 as providing a net replacement rate of approximately 59% for a single individual without children earning 67% of the average wage prior to unemployment (OECD, 2025). Benefit levels and entitlement duration also vary with age and employment history, for example individuals aged 50 and above with sufficiently long employment records are eligible for more generous and longer-lasting benefits.

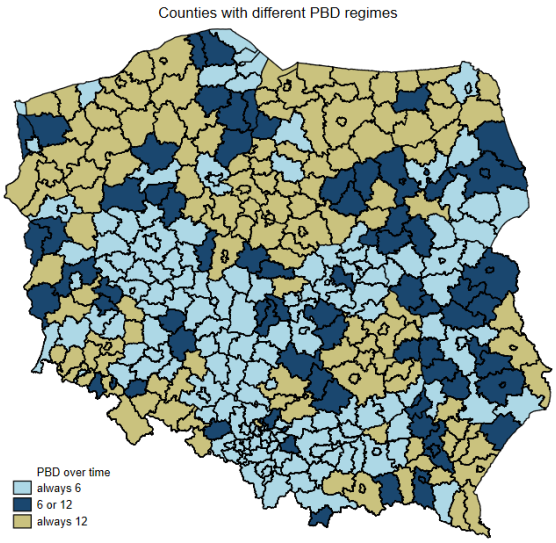
[10] The data end on 31 July 2021, affecting 0.18% of spells which are ongoing and therefore right-censored.

[11] These incentives include access to public health insurance and eligibility for means-tested social assistance programs. Unemployment benefit claimants constitute approximately 7.9% of formerly incarcerated individuals who register with the Public Employment Office, while the remaining 92.1% are not eligible for unemployment benefits.

Figure 1: Eligibility conditions and PBD regimes in Poland



(a) UI eligibility criteria for formerly incarcerated workers registering within 30 days of release



(b) Counties with different PBD regimes

Notes: Panel (a) illustrates the time frame for meeting eligibility conditions when a worker registers within 30 days of release. Panel (b) shows how regimes were distributed across counties over 2009–2018.

tration can occur through several channels. One is exit to employment, defined as deregistration following the start of employment or self-employment. We identify reincarceration either when an individual is deregistered due to pre-trial detention or the serving of a custodial sentence, or when a subsequent registration indicates that the individual has returned after another spell of imprisonment. Other reasons for deregistration include failure to confirm availability for work and voluntary deregistration. Employment and reincarceration are not mutually exclusive outcomes, as an individual may leave the register upon finding a job and subsequently experience another incarceration spell. The full sample comprises 9,004 spells, while the RD estimation sample comprises 3,026 spells within the bandwidth of 35, including 917 pre-prison work qualifiers and 1,621 in-prison work qualifiers.

Table 1 reports descriptive statistics for the full sample and separately by eligibility pathway, excluding from each subgroup those who would also qualify via the other route. The average individual is a low-educated male aged around 34, who, prior to registering as unemployed, had spent approximately 3.5 years in prison and approached the public employment office after being incarcerated about 1.7 times. On average, ex-inmates collect unemployment benefits for 5.2 months and reside in counties where the statutory potential benefit duration is equal to six months. Slightly more than half of formerly incarcerated individuals exhaust the full potential benefit duration (either 6 or 12 months), while their average unemployment spell lasts approximately 9.2 months. Around 43.8% ex-inmates find employment, while the average reincarceration rate is 18.4%.[12]

The in-prison work group is larger than the pre-prison work group. Both groups have similar average ages and shares of women. Those who qualified through in-prison employment were less educated (11.8% versus 28.7% with education above the vocational level) and, on average, spent 3.3 additional years in prison (4.9 versus 1.6 years), which likely increased their opportunity to meet the UI eligibility requirement of at least one year of in-prison employment. Individuals in the in-prison work group collected unemployment benefits for 0.9 months longer (5.6 versus 4.7 months) and were 16.1 percentage points more likely to exhaust the full potential benefit duration (57.2% versus 41.1%). They also remained unemployed for 2.3 months longer (10.0 versus 7.7 months) and were 23.5 percentage points less likely to find employment (34.3% versus 57.8%). The share who were reincarcerated was 11.8 percentage points higher (22.8% versus 11.0%).

[12]Official Ministry of Justice statistics show five-year recidivism rates among individuals sentenced to imprisonment of approximately 25–26% (Ministry of Justice, 2017).

Despite these differences, the two groups resided in counties with similar relative unemployment rates and hence faced similar average potential benefit durations.

Table 1: Summary statistics of unemployment benefit spells by eligibility group

	Eligibility group		
Variable	All benefit recipients	Pre-prison work qualifiers	In-prison work qualifiers
No. of observations	9,004	2,955	4,673
Age in years	34.040	33.125	34.766
	(6.998)	(6.860)	(6.962)
Female (0/1)	0.031	0.032	0.033
	(0.173)	(0.176)	(0.178)
Education above vocational (0/1)	0.184	0.287	0.118
	(0.387)	(0.452)	(0.323)
Prison duration in years	3.468	1.603	4.884
	(3.219)	(1.689)	(3.465)
Incarceration ordinal number	1.676	1.271	1.865
	(1.572)	(0.683)	(1.549)
Benefit duration in months	5.234	4.658	5.560
	(3.076)	(3.050)	(3.028)
Benefits exhausted (0/1)	0.513	0.411	0.572
	(0.500)	(0.492)	(0.495)
Unemployment duration in months	9.188	7.658	10.019
	(10.171)	(8.685)	(10.424)
Exit to employment (0/1)	0.438	0.578	0.343
	(0.496)	(0.494)	(0.475)
Reincarcerated (0/1)	0.184	0.110	0.228
	(0.387)	(0.313)	(0.419)
County unemployment rate (% of national avg.)	112.705	110.145	112.472
	(52.312)	(52.048)	(52.024)

Notes: The table reports sample means, with standard deviations in parentheses, over 2009–2018 and based on the start date of the unemployment spell. Exit to employment denotes deregistration due to employment; reincarcerated denotes workers who either deregistered due to pre-trial detention or the serving of a custodial sentence, or experienced a subsequent incarceration spell observed in the administrative records. Each eligibility pathway group excludes workers who also qualified via the other route. All variables are measured at the beginning of the unemployment spell unless stated otherwise.

3 Empirical Modeling

The institutional setting creates a sharp discontinuity that allows us to identify the causal impact of a longer potential benefit duration on labor market and criminal outcomes among formerly incarcerated workers. A key feature of the assignment mechanism is that counties with nearly identical unemployment rates, for example 150.2% and 149.8% of the national average, are assigned to different PBD regimes.

To implement the regression discontinuity (RD) design, we estimate the following linear model by OLS:

$$y_{ict} = \beta_0 + f(ur_{ct}) + \beta_1 \cdot I(ur_{ct} > 150\%) + \beta_2 \cdot X_{ict} + county_c + year_t + \epsilon_{ict}, \tag{1}$$

where y_{ict} is the outcome of interest for worker i in county c in year t, namely benefit duration, unemployment duration, exit to employment, or likelihood of reincarceration. The running variable ur_{ct} is the relative unemployment rate in county c in year t. The indicator $I(ur_{ct} > 150\%)$ equals one for workers registering at an employment office where the county unemployment rate exceeds 150% of the national average, and zero otherwise. The coefficient of interest β_1 captures the local average treatment effect of longer potential benefit duration. We control for individual characteristics X_{ict}, as well as county and year fixed effects. We estimate a local linear regression using a triangular kernel, with $f(ur_{ct})$ linear in ur_{ct}, so that $f(\cdot) = \delta_0 \cdot ur_{ct} + \delta_1 \cdot ur_{ct} \cdot I(ur_{ct} > 150\%)$. Bandwidths are selected using the procedure of Calonico et al. (2020), estimated separately for each outcome and sample and yielding a grand average of 34.30 (Table B.2); we therefore adopt a common bandwidth of 35 in the baseline specification.[13] Standard errors are clustered at the county level.

Because the running variable varies at the county-year level and counties can cross the threshold over time, identification comes from both cross-county comparisons near the cutoff and within-county changes in regime. Within the bandwidth of 35, the estimation sample contains 3,026 spells drawn from 202 counties and 973 county-years, of which 71 counties are observed on both sides of the threshold during the sample period (in 31% of cases they switched PBD regimes more than once). County fixed effects absorb time-invariant local labor-market condi-

[13] Table B.1 reports descriptive statistics for observations within the selected bandwidth, separately for workers eligible for 6 and 12 months of PBD.

tions and year fixed effects absorb national shocks, so β_1 is identified from deviations in the relative unemployment rate around 150% net of these components.

Formerly incarcerated workers have very limited scope to manipulate their place of registration, since the relevant public employment office is determined by the county of registered residence. Under the regulations in force throughout our sample period, workers could register only at the office corresponding to their permanent address, which is set through an administrative procedure, so former inmates typically registered at the office linked to their pre-incarceration address.[14] Compared with the general population of unemployed workers studied by Jessen et al. (2026) and Jessen et al. (2025), formerly incarcerated workers also face additional restrictions on the timing of registration: they must register within 30 days of release for the entire period of incarceration and the preceding 18 months to count toward eligibility. Consistent with this, a local-linear density test following Cattaneo et al. (2018) finds no evidence of a discontinuity in the density of the running variable at the threshold ($t = 0.458$), based on county-clustered bootstrap standard errors. Jessen et al. (2026) similarly find no density discontinuity in the full sample of unemployed workers, despite the potentially greater scope for endogenous responses to the threshold among workers facing fewer restrictions on registration.

We also examine whether the composition of formerly incarcerated workers registering with public employment offices changes at the threshold. Although release and registration occur after the county unemployment rate determining benefit duration is set, we test directly for discontinuities in observable characteristics around the cutoff. As shown in Table B.3, there are no statistically significant discontinuities in predetermined characteristics, providing no evidence of observable selection into the sample.

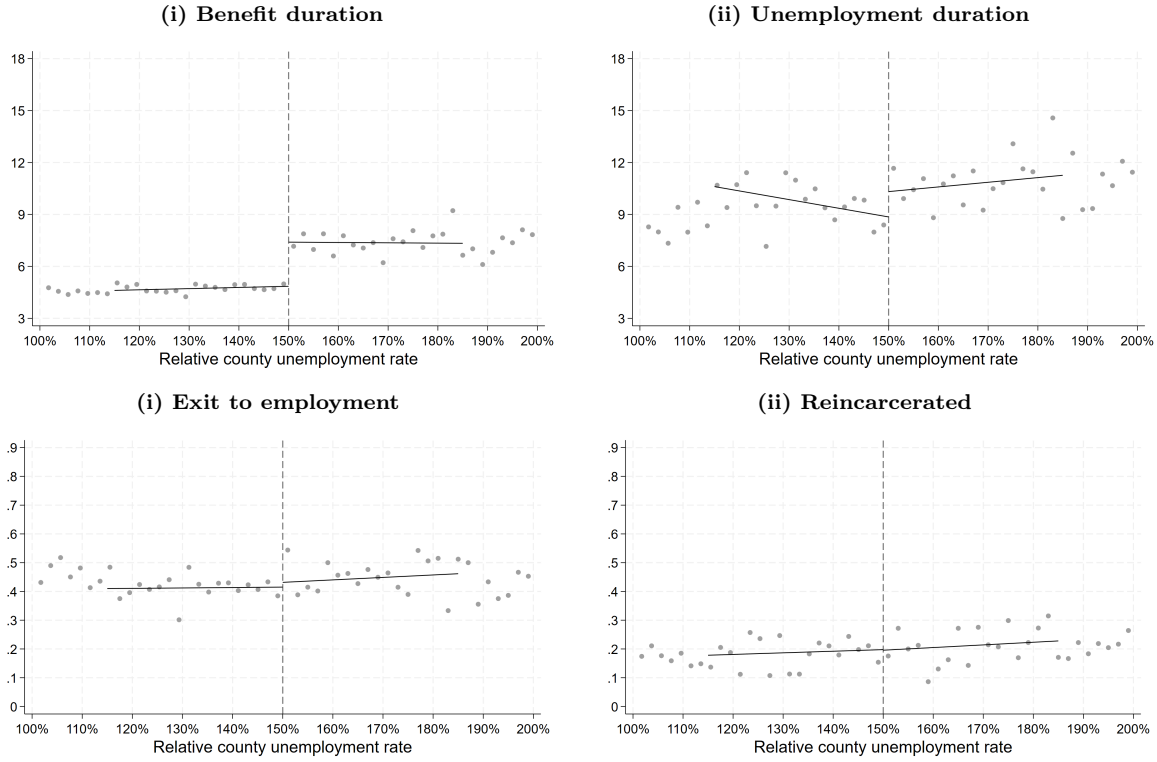
4 Results

Figure 2 plots the outcome variables against relative unemployment rates, together with the RD regression fits from equation 1, excluding control variables and county and year fixed effects. Figure A.1 presents the corresponding plots separately for individuals qualifying through pre-prison and in-prison work. The upper panels show discontinuous increases at the 150% relative

[14] From June 2025, registration is allowed at the office corresponding to a worker's current place of residence. This change, and any scope for manipulating the registration location that it creates, postdates our sample period.

unemployment threshold of approximately 2.5 months in benefit duration and 1.5 months in unemployment duration. By contrast, the lower panels show only a small discontinuity in the likelihood of exiting to employment and virtually no discontinuity in the likelihood of reincarceration.

Figure 2: RD plots around the PBD threshold for all benefit recipients



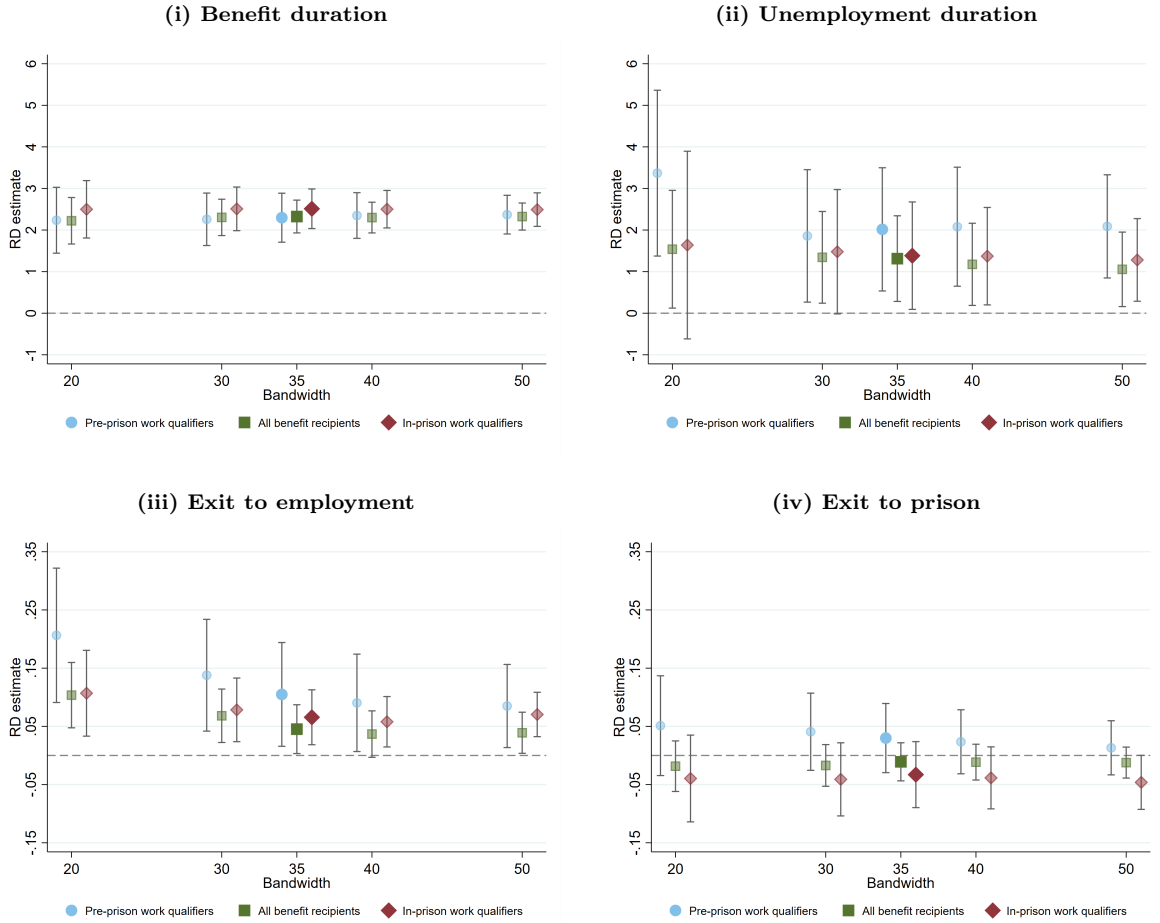
Notes: The figure presents RD plots around the PBD threshold for all benefit recipients over 2009–2018, using a bandwidth of 35. Points represent average outcomes within 2-percentage-point bins, and solid lines denote local linear fits estimated separately on either side of the threshold. Corresponding estimates are reported in Table B.4.

Figure 3 plots the estimates of β_1 across bandwidths, while Table B.4 reports the RD estimates for our preferred bandwidth of 35. The estimates are highly stable across bandwidths. In the upper panels, increasing potential benefit duration from 6 to 12 months raises actual benefit receipt by 2.3 months, indicating that recipients use approximately 38% of the additional six-month entitlement. Relative to the mean benefit duration of 4.8 months below the threshold, this represents an increase of 48%. The corresponding increase in unemployment duration is 1.3 months, or 14% relative to the below-threshold mean of 9.7 months. The smaller effect on unemployment duration suggests that part of the additional benefit receipt accrues to individuals who would have remained unemployed even without the extension. Since PBD doubles, the implied elasticity of unemployment duration with respect to PBD is 0.14, compared with 0.29 for the

Polish workforce as a whole (Jessen et al., 2025). Thus, previously incarcerated benefit recipients exhibit a behavioral response only about half as large as that of the population receiving unemployment benefits, consistent with the limited employment opportunities they face.

In the lower panels, we find a small but positive effect of approximately 5 percentage points on the likelihood of exiting unemployment for employment, suggesting that longer PBD may facilitate employment transitions despite extending unemployment duration. Importantly, although raising PBD from 6 to 12 months increases unemployment duration, it does not increase the likelihood of reincarceration. The estimated effects on reincarceration are small, negative, and statistically insignificant across bandwidths.

Figure 3: Estimated RD effects across alternative bandwidths by eligibility group



Notes: The figure presents individual-level RD estimates from equation (1) across alternative bandwidths over 2009–2018. The horizontal axis reports bandwidths ranging from 20 to 50, and vertical bars denote 95% confidence intervals. Controls and fixed effects are as in Table B.4, which also reports the baseline estimates at a bandwidth of 35.

Figure 3 also plots the RD estimates separately for pre-prison and in-prison work qualifiers. Both groups show similar increases in benefit receipt, of approximately 2.3–2.5 months.

The effects on unemployment duration exhibit greater heterogeneity: the estimated increase is 2.0 months among pre-prison qualifiers, approximately 50% larger than the 1.4-month increase among in-prison qualifiers. The estimated effect on the likelihood of entering employment is also larger among pre-prison qualifiers, at approximately 11 percentage points, compared with 7 percentage points among in-prison qualifiers. As in-prison work qualifiers have longer prison stays and likely face more restricted labor-market opportunities, they may have less scope to respond behaviorally to longer benefit entitlement. The estimated effects on reincarceration are close to zero for both groups, although the estimate is positive for pre-prison qualifiers and negative for in-prison qualifiers. Nevertheless, none of the differences between the subgroup estimates is statistically significant, reflecting the smaller sample sizes and the substantial overlap in their 95% confidence intervals. The evidence of heterogeneous responses should therefore be viewed as suggestive rather than conclusive. Overall, extending unemployment benefits to individuals with limited labor-market opportunities may generate meaningful benefits at modest efficiency costs, given their relatively muted behavioral responses. These costs may be particularly limited among the most disadvantaged recipients.

We next re-estimate the baseline specification at the county level, which is the level at which the policy is implemented. Figure A.2 presents the corresponding RD estimates, while Table B.5 reports the baseline estimates obtained using the common bandwidth of 35. County-level estimates mirror the individual-level findings, with lower precision given the smaller number of observations. At the county level the effect on unemployment duration for in-prison work qualifiers falls to 0.51 months and is no longer distinguishable from zero, while the corresponding estimate for pre-prison work qualifiers rises to 2.62 months and remains significant at the 1% level. The two levels of aggregation therefore bracket the in-prison response between roughly 0.5 and 1.4 months, and widen rather than narrow the estimated gap between eligibility pathways.

5 Welfare Analysis

Unemployment benefits generate behavioral responses among claimants, as improved outside options may reduce job-search intensity and prolong benefit receipt or unemployment duration. A broad literature evaluates the moral hazard and welfare effects of transfers by comparing their mechanical fiscal cost with the behavioral responses they induce (Finkelstein and Hendren, 2020;

Hendren and Sprung-Keyser, 2020). In the unemployment insurance context, this framework is used to analyze the fiscal externalities and disincentive effects of benefit generosity, often summarized by BC/MC, the ratio of behavioral cost to mechanical cost (Cohen and Ganong, 2026; Schmieder and von Wachter, 2017; Schmieder et al., 2012, 2016).

In our computations, we draw on Jessen et al. (2025), who build on the framework developed by Chetty (2008) and Schmieder et al. (2016) to provide comprehensive welfare calculations for the unemployment insurance system in Poland. We adopt the simplified computational approach proposed by Schmieder et al. (2016):

$$\frac{BC}{MC} = -\frac{dD}{dP} \cdot \frac{1}{S_P} \cdot \left(\xi + \frac{\tau}{b}\right),$$

where $\frac{dD}{dP}$ is the effect of longer PBD on unemployment duration D. Assuming a constant hazard of exit from unemployment, s, we have $D = \frac{1}{s}$. The exhaustion rate is given by

$$S_P = e^{-P \cdot s} = e^{-\frac{P}{D}},$$

and

$$\xi = 1 - (1 + P \cdot s)e^{-P \cdot s} = 1 - \left(1 + \frac{P}{D}\right) e^{-\frac{P}{D}}.$$

We compute the BC/MC ratio using the estimated marginal effects together with descriptive statistics. Specifically, we set $P = 6$ and define D as the mean unemployment duration in counties with $PBD = 6$, which is 9.71, 7.77 and 10.36 months for all benefit recipients, pre-prison work qualifiers and in-prison work qualifiers respectively. We obtain $\frac{dD}{dP}$ by dividing the estimates in column (2) of Table B.4 by 6, which yields the effect of a one-month increase in PBD on unemployment duration. We set $\frac{\tau}{b} = 3.72$ following Jessen et al. (2025), where τ denotes the tax rate and b the unemployment benefit level.

The BC/MC ratio measures the fiscal externality and provides a basis for welfare analysis. The marginal value of public funds (MVPF) captures the social value generated by an additional unit of public spending and thus summarizes the welfare effect of the policy (Finkelstein and Hendren, 2020; Hendren and Sprung-Keyser, 2020; Jessen et al., 2025). Under constant relative risk aversion (CRRA) utility, the MVPF of a one-unit increase can be written as:

$$MVPF = \frac{\text{Willingness to pay}}{\text{Net cost to government}} = \frac{\text{Willingness to pay}}{\text{Policy expenditure} + \text{Fiscal externality}} = \frac{\left(\frac{c_u}{c_e}\right)^{-\gamma}}{1 + \frac{BC}{MC}}, \quad (2)$$

where $\frac{c_u}{c_e}$ denotes the ratio of consumption under unemployment (c_u) to consumption under employment (c_e), and therefore reflects the consumption drop associated with unemployment. Evidence from the United States, where UI replacement rates are broadly similar to those in Poland (OECD, 2025), suggests an unemployment-related consumption drop of about 10% under typical benefit generosity, that is, $\frac{c_u}{c_e} = 0.9$ (Chetty and Finkelstein, 2013). We consider values of γ equal to 1, 2 and 5 (see Jessen et al. (2025) for a comprehensive discussion).

Table 2 reports the results. For all benefit recipients the BC/MC ratio is 1.6, implying that a mechanical transfer of an additional zloty (or €1, for international comparability) induces a further behavioral cost of 1.6 zloty (or €1.6). The estimates differ across eligibility pathways: the ratio is 1.6 for in-prison work qualifiers and 2.8 for pre-prison work qualifiers.

Table 2: BC/MC and MVPF of the longer PBD

	BC/MC	MVPF			Observations
CRRA coefficient:		1	2	5	
Consumption loss:		10%	10%	10%	
	(1)	(2)	(3)	(4)	(5)
Panel A: all benefit recipients	1.56	0.43	0.48	0.66	3,026
Panel B: pre-prison work qualifiers	2.84	0.29	0.32	0.44	917
Panel C: in-prison work qualifiers	1.58	0.43	0.48	0.66	1,621

Notes: Authors' calculations. BC/MC set $P = 6$ and take D as the mean unemployment duration below the threshold (9.71, 7.77 and 10.36 months respectively); dD/dP divides column (2) of Table B.4 by 6; $\tau/b = 3.72$. MVPF from equation (2) with $c_u/c_e = 0.9$.

This pattern is intuitive. Workers who qualified through in-prison work generally spent longer in prison and face particularly poor labor market prospects, which likely reduces the scope for behavioral adjustment in response to longer potential benefit duration. The behavioral response in this group is therefore present but attenuated, both relative to pre-prison work qualifiers and relative to the general Polish workforce. For comparison, Jessen et al. (2025) report BC/MC ratios ranging from about 2.5 to 3.6 for Poland as a whole, while Schmieder et al. (2016) report a median estimate of 1.8 across the literature. The in-prison figure of 1.6 lies below both benchmarks; the pre-prison figure of 2.8 lies close to overall Polish workforce result and above the international median.

Columns (2) to (4) of Table 2 report the corresponding MVPFs under alternative values of the CRRA coefficient. All estimates lie below one, implying that an extension of potential benefit duration would not be welfare-improving within a standard welfare framework, for either group and at every value of risk aversion we consider. Under the assumption of a consumption drop of approximately 10%, for formerly incarcerated workers overall, an additional euro of public spending generates a marginal social benefit of 0.4 to 0.7 euros. Estimates for ex-prisoners who qualified through in-prison work are nearly identical, reflecting the similarity of behavioral responses in this group and in the full sample. In contrast, the MVPF is substantially lower for ex-prisoners who qualified through pre-prison work, consistent with their stronger behavioral responses to changes in benefit duration. Therefore, while the benefit extension does not pay for itself, its fiscal externality is substantially smaller among formerly incarcerated workers, especially those who earned entitlement through in-prison employment, resulting in MVPFs of approximately 0.4–0.7, compared with 0.3–0.5 for the broader Polish workforce (Jessen et al., 2025).

These estimates should nevertheless be interpreted as lower bounds, since our calculations do not explicitly incorporate the social costs of crime or the costs borne by the criminal justice system. Furthermore, although we do not find evidence that longer potential benefit duration affects reincarceration rates in either direction, more generous unemployment insurance benefits may still reduce reliance on illegal activity among formerly incarcerated individuals, as many offenses remain undetected or do not result in imprisonment.[15] Consequently, a comprehensive assessment of the social welfare effect of providing unemployment benefits to this group should go beyond fiscal externalities and direct program costs to also account for the broader costs of operating the correctional system.

6 Discussion and Conclusion

We investigate whether more generous unemployment benefits for formerly incarcerated workers facilitate successful reintegration or exacerbate moral hazard. Doubling potential benefit duration from 6 to 12 months raises actual benefit receipt by roughly 2.3–2.5 months, a result that is

[15] As reincarceration reflects only a subset of criminal activity, namely offenses that are detected and lead to a custodial sentence, the absence of an effect on reincarceration should not be interpreted as evidence that benefit generosity has no effect on crime more broadly.

remarkably similar across all benefit recipients and both eligibility pathways. Longer entitlement also raises the unemployment duration by 1.3–2.0 months. However, the likelihood of transition to employment appears to increase, while the likelihood of reincarceration remains unchanged.

The extension therefore does not appear to discourage successful labor-market reentry, despite the increase in unemployment duration. The behavioral response we estimate is attenuated relative to that documented by Jessen et al. (2025) for the general Polish unemployment benefit population, which in turn is broadly comparable to estimates in the wider literature (Cohen and Ganong, 2026; Schmieder et al., 2016). It is not, however, absent. Moral hazard associated with extending potential benefit duration can operate through reduced job-search effort during the initial eligibility period and through longer observed spells following the extension itself. As Schmieder and von Wachter (2017) emphasize, these jointly constitute the behavioral cost of income transfers delivered through longer entitlement. Our estimates may capture this combined behavioral response, which is substantially smaller among formerly incarcerated workers than in the general Polish unemployment benefit population.

We find that the behavioral response appears to vary with the pathway through which benefit eligibility was established. The implied fiscal externality is approximately half as large for workers who qualified through in-prison work as for those who qualified through pre-prison employment. The weak behavioral response among in-prison work qualifiers may reflect two complementary mechanisms. A compositional channel is likely to play a role: individuals who qualify through work performed during imprisonment differ from pre-prison work qualifiers in observable characteristics, including lower education and longer prison spells, which may reflect more limited labor-market opportunities and less scope for behavioral responses to longer benefit entitlement. A transformational channel may also operate, since participation in prison work programs can strengthen labor market attachment, improve employability and increase motivation to pursue legal work after release (Duwe and Henry-Nickie, 2021; Nur and Nguyen, 2023). In the Polish context eligibility is acquired through paid prison work, which is often granted as a reward following a period of unpaid work (Felczak, 2025; Felczak and Wilamowska, 2025), so prison employment may reflect both stronger engagement during incarceration and better preparation for reentry.

A related question is whether more generous benefits attract additional claimants into the unemployment insurance system. This issue has received increasing attention in the unemployment

insurance literature, where endogenous inflows are viewed as a potential channel through which benefit generosity affects both program costs and observed labor-market outcomes (Van Doornik et al., 2023; Jessen et al., 2025; Lachowska et al., 2025; McQuillan and Moore, 2025). In our setting, inflow responses would be particularly informative because they would primarily capture a take-up effect. Individuals are released from prison before registration and therefore cannot adjust employment, separation, or labor-supply decisions in anticipation of benefit eligibility. As a result, any discontinuity in inflows at the eligibility threshold would primarily reflect changes in the propensity to register and claim benefits.

Examining inflows would also provide a useful complement to the covariate-balance evidence reported in Table B.3, allowing us to bound the potential magnitude of any compositional change in the observed sample. In practice, however, the available data do not provide sufficient statistical power for such an exercise. Inflow rates can only be constructed at the county-year level, and the number of formerly incarcerated benefit recipients entering the system is very small. The average inflow equals 3.1 individuals per county-year in the full sample. At this level of aggregation, even small random changes in the number of entrants translate into large percentage fluctuations, making the resulting estimates highly noisy and difficult to interpret. Reassuringly, the evidence that is available points to little scope for compositional changes around the threshold. We find no discontinuities in predetermined individual characteristics (Table B.3), nor do we find any statistically significant effects from analogous balance tests based on county-level inflow averages. Although future work with larger samples could directly examine inflow responses and benefit take-up among formerly incarcerated workers, our data provide no evidence that changes in the composition of claimants drive the estimated effects.

While the evidence provides little support for compositional changes around the threshold, several features of our empirical setting should nevertheless be considered when interpreting the results. First, our sample comprises formerly incarcerated workers who register with an employment office and qualify for benefits, a group with prior formal employment, so the estimates should not be read as effects for all released prisoners. Second, our welfare framework is designed around consumption smoothing and fiscal externalities, and it abstracts from the broader social costs associated with criminal activity. Because our reincarceration margin captures detected offenses resulting in a custodial sentence, it provides a specific metric of institutional return rather than a comprehensive measure of all criminal behavior. Third, despite the stable order-

ing of behavioral responses across eligibility pathways, which holds across bandwidths, levels of aggregation, and the implied fiscal externalities, the difference between our point estimates lacks the statistical precision to be formally distinguished.

Our findings therefore show that extending unemployment insurance for formerly incarcerated workers increases both benefit receipt and unemployment duration, but does not appear to discourage transitions to employment or increase reincarceration. More broadly, the results suggest that the behavioral and fiscal costs of unemployment insurance may differ substantially across populations. Among workers facing severe labor-market barriers, particularly those who qualify through in-prison work, longer benefit entitlement generates considerably smaller behavioral responses and fiscal externalities than estimates for the general workforce would suggest. These findings highlight the importance of accounting for recipients' labor-market opportunities when assessing the costs and benefits of unemployment insurance.

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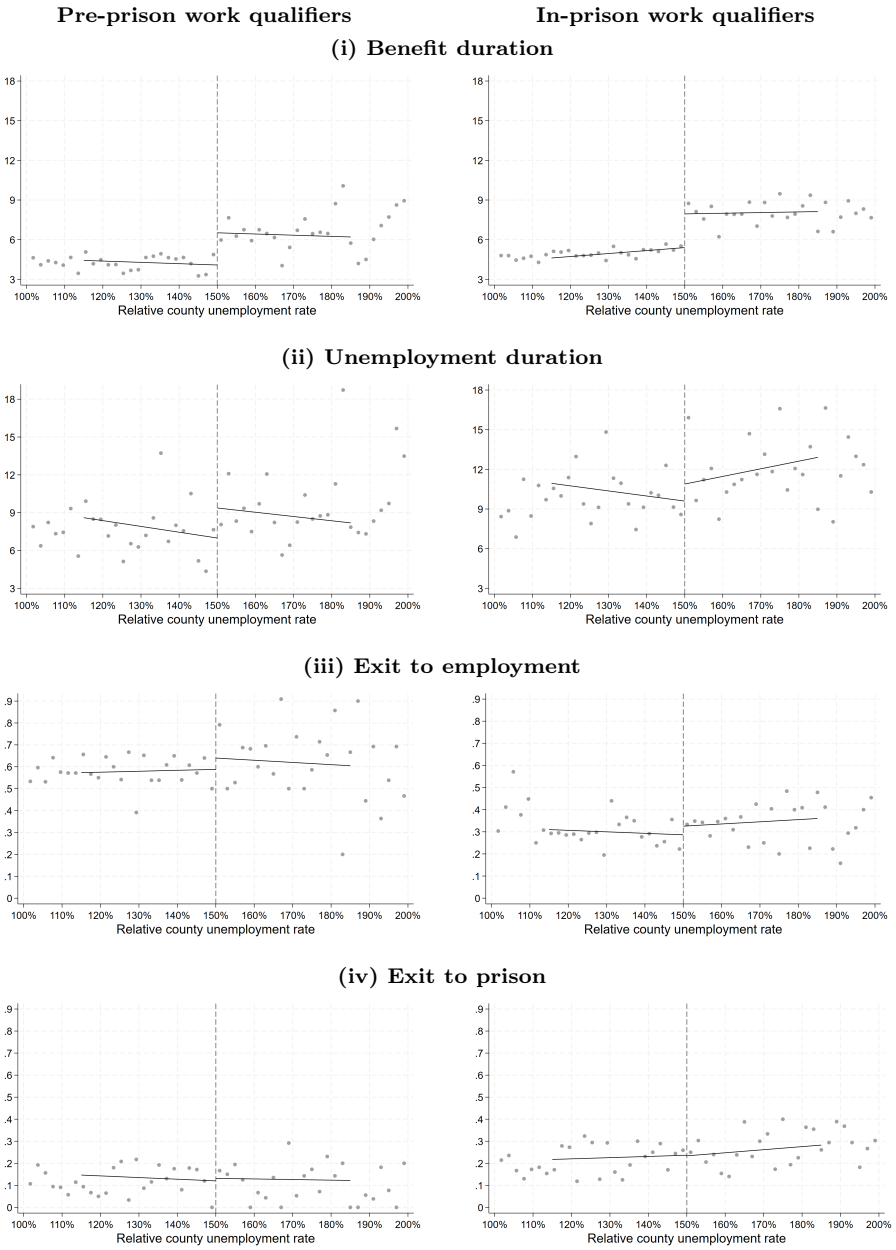
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APPENDIX (FOR ONLINE PUBLICATION)

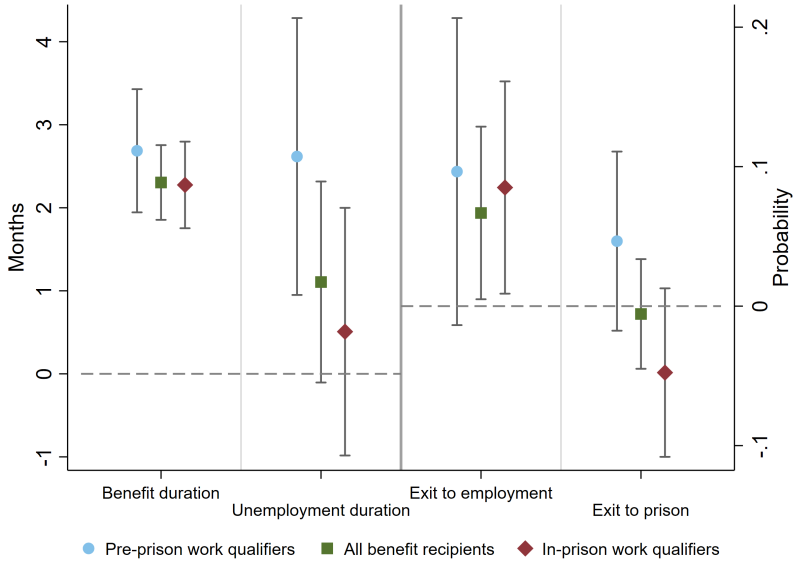
Figures and Tables

Figure A.1: RD plots around the PBD threshold by eligibility group



Notes: The figure presents RD plots around the PBD threshold for pre-prison work qualifiers in the left column and in-prison work qualifiers in the right column, over 2009–2018 and using a bandwidth of 35. Points represent average outcomes within 2-percentage-point bins, and solid lines denote local linear fits estimated separately on either side of the threshold. Corresponding estimates are reported in Table B.4.

Figure A.2: County-level RD estimates by eligibility group



Notes: The figure presents county-level RD estimates from equation (1) over 2009–2018, using a bandwidth of 35. Benefit duration and unemployment duration are plotted against the left axis and measured in months, while exit to employment and exit to prison are plotted against the right axis and measured in percentage points. Vertical bars denote 95% confidence intervals. Controls and fixed effects are as in Table B.5, which also reports the baseline estimates.

Table B.1: Summary statistics of unemployment benefit spells by eligibility group and potential benefit duration (PBD), bandwidth = 35

	All benefit recipients		Pre-prison work qualifiers		In-prison work qualifiers	
Variable / PBD	6 months	12 months	6 months	12 months	6 months	12 months
No. of observations	1,707	1,319	536	381	907	714
Age in years	33.842	33.754	33.312	33.181	34.443	34.195
	(7.064)	(7.049)	(6.938)	(7.086)	(7.125)	(6.955)
Female (0/1)	0.023	0.030	0.024	0.024	0.026	0.032
	(0.151)	(0.172)	(0.154)	(0.152)	(0.161)	(0.177)
Education above vocational (0/1)	0.189	0.151	0.308	0.228	0.108	0.098
	(0.392)	(0.358)	(0.462)	(0.420)	(0.311)	(0.298)
Prison duration in years	3.471	3.351	1.591	1.500	4.847	4.530
	(3.396)	(2.976)	(1.843)	(1.446)	(3.709)	(3.137)
Incarceration ordinal number	1.632	1.720	1.213	1.278	1.828	1.947
	(1.295)	(1.356)	(0.605)	(0.658)	(1.358)	(1.605)
Benefit duration in months	4.754	7.454	4.244	6.518	5.045	8.069
	(1.908)	(4.362)	(2.116)	(4.532)	(1.732)	(4.198)
Benefits exhausted (0/1)	0.603	0.425	0.474	0.312	0.675	0.486
	(0.489)	(0.495)	(0.500)	(0.464)	(0.469)	(0.500)
Unemployment duration in months	9.708	10.864	7.768	9.119	10.355	11.793
	(11.604)	(11.273)	(9.600)	(9.582)	(10.717)	(11.211)
Exit to employment (0/1)	0.411	0.444	0.586	0.622	0.293	0.340
	(0.492)	(0.497)	(0.493)	(0.486)	(0.456)	(0.474)
Reincarcerated (0/1)	0.188	0.215	0.123	0.134	0.232	0.258
	(0.391)	(0.411)	(0.329)	(0.341)	(0.422)	(0.438)
County unemployment rate (% of national avg.)	131.343	165.258	131.917	165.035	131.275	165.353
	(10.107)	(9.665)	(9.932)	(9.342)	(10.040)	(9.773)

Notes: Sample means, standard deviations in parentheses. Sample period 2009–2018, restricted to observations within a bandwidth of 35 around the PBD threshold. Each pathway group excludes workers who also qualified via the other route. Variable definitions as in Table 1.

Table B.2: Optimal bandwidths across outcomes and samples

Outcome	All recipients	Pre-prison work qualifiers	In-prison work qualifiers
Benefit duration	32.05	30.57	33.66
Unemployment duration	30.68	29.39	35.66
Exit to employment	37.42	39.02	35.30
Exit to prison	37.89	35.92	34.05
Average	34.51	33.73	34.67
Grand average		34.30	

Notes: Optimal bandwidths are estimated separately for each outcome within each sample. *Average* is the mean across outcomes within a sample; *Grand average* is the mean of the sample-specific averages. A common bandwidth of 35 is used in the baseline specifications.

Table B.3: Covariate balance at the eligibility threshold

	Age (years) (1)	Female (0/1) (2)	Education above vocational (0/1) (3)	Prison duration (years) (4)	Incarceration ordinal number (5)
Bandwidth			35		
Panel A: All benefit recipients					
RD estimate	0.102	0.006	0.020	-0.186	-0.020
	(0.324)	(0.008)	(0.018)	(0.148)	(0.070)
Mean, PBD = 6	33.842	0.023	0.189	3.471	1.632
Observations			3,026		
Panel B: Pre-prison work qualifiers					
RD estimate	0.329	0.019	0.031	0.184	0.001
	(0.657)	(0.013)	(0.042)	(0.130)	(0.044)
Mean, PBD = 6	33.312	0.024	0.308	1.591	1.213
Observations			917		
Panel C: In-prison work qualifiers					
RD estimate	-0.210	0.006	-0.029	-0.115	-0.033
	(0.479)	(0.013)	(0.019)	(0.193)	(0.105)
Mean, PBD = 6	34.443	0.026	0.108	4.847	1.828
Observations			1,621		

Notes: Each column reports the estimated discontinuity from equation (1) with the listed predetermined characteristic as the dependent variable and no individual controls. Bandwidth 35, county and year fixed effects. *Mean, PBD = 6* reports the mean of the dependent variable below the threshold. Standard errors clustered at the county level in parentheses. Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table B.4: Effects of longer PBD on durations and probabilities

Variation:	6 months longer PBD			
Dependent variable:	Months of		Exit to	
	benefit receipt	unemployment	employment	prison
	(1)	(2)	(3)	(4)
Bandwidth	35			
Panel A: All benefit recipients				
RD estimate	2.325***	1.313**	0.045**	-0.011
	(0.201)	(0.525)	(0.021)	(0.017)
Elasticity	0.489	0.135	-	-
Mean, PBD = 6	4.754	9.708	0.411	0.188
Observations	3,026			
Panel B: Pre-prison work qualifiers				
RD estimate	2.297***	2.016***	0.105**	0.030
	(0.301)	(0.756)	(0.045)	(0.030)
Elasticity	0.541	0.260	-	-
Mean, PBD = 6	4.244	7.768	0.586	0.123
Observations	917			
Panel C: In-prison work qualifiers				
RD estimate	2.512***	1.384**	0.066***	-0.033
	(0.243)	(0.659)	(0.024)	(0.029)
Elasticity	0.498	0.134	-	-
Mean, PBD = 6	5.045	10.355	0.293	0.232
Observations	1,621			

Notes: Estimates from equation (1), individual level, bandwidth 35. Columns (1) and (2) report benefit duration and unemployment duration in months; columns (3) and (4) report indicators for exit to employment and likelihood of reincarceration. *Mean, PBD = 6* reports the mean of the dependent variable below the threshold within the estimation bandwidth. All specifications control for age, gender, education above vocational level, and the sequence number and duration of the most recent incarceration spell, and include county and year fixed effects. Standard errors clustered at the county level in parentheses. Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table B.5: Effects of longer PBD on durations and probabilities — county level analysis

Variation:	6 months longer PBD			
Dependent variable:	Months of		Exit to	
	benefit receipt	unemployment	employment	prison
	(1)	(2)	(3)	(4)
Bandwidth	35			
Panel A: All benefit recipients				
RD estimate	2.305***	1.106*	0.067**	-0.006
	(0.230)	(0.618)	(0.032)	(0.020)
Elasticity	0.491	0.118	-	-
Observations	973			
Panel B: Pre-prison work qualifiers				
RD estimate	2.687***	2.618***	0.096*	0.047
	(0.378)	(0.851)	(0.056)	(0.033)
Elasticity	0.619	0.335	-	-
Observations	574			
Panel C: In-prison work qualifiers				
RD estimate	2.275***	0.508	0.085**	-0.048
	(0.266)	(0.761)	(0.039)	(0.031)
Elasticity	0.456	0.049	-	-
Observations	697			

Notes: As Table B.4, except that the unit of observation is a county-year and all outcomes are annual county-level averages. Controls are county-level averages of the individual characteristics. Standard errors clustered at the county level in parentheses. Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.