The Optimal Timing of Unemployment Benefits: Theory and Evidence from Sweden

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Abstract

This paper provides a simple, yet robust framework to evaluate the time profile of benefits paid during an unemployment spell. We derive sufficient-statistics formulae capturing the marginal insurance value and incentive costs of unemployment benefits paid at different times during a spell. Our approach allows us to revisit separate arguments for inclining or declining profiles put forward in the theoretical literature and to identify welfare-improving changes in the benefit profile that account for all relevant arguments jointly. For the empirical implementation, we use administrative data on unemployment, linked to data on consumption, income and wealth in Sweden. First, we exploit duration-dependent kinks in the replacement rate and find that, if anything, the moral hazard cost of benefits is larger when paid earlier in the spell. Second, we find that the drop in consumption affecting the insurance value of benefits is large from the start of the spell, but further increases throughout the spell. In trading off insurance and incentives, our analysis suggests that the flat benefit profile in Sweden has been too generous overall. However, both from the insurance and the incentives side, we find no evidence to support the introduction of a declining tilt in the profile.

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The key objective of social insurance programs is to provide insurance against adverse events while maintaining incentives. The impact of these adverse events is dynamic and so are the insurance value and incentive cost of social protection against these events. As a consequence, the design of social insurance policies tends to be dynamic as well, specifying a schedule of benefits and taxes that are time-dependent. In the context of unemployment insurance (UI), the UI policy specifies a full benefit profile designed to balance incentives and insurance throughout the unemployment spell. Solving this dynamic problem can prove daunting, especially when adding important features of unemployment dynamics involving selection and non-stationarities. Indeed, there seems to be little consensus in practice on the optimal profile of UI benefits. Unemployment policies vary substantially across countries in the time profile of benefits paid during an unemployment spell, above and beyond differences in the overall generosity. In the US, benefits are paid only during the first six months of unemployment. In other countries, like Belgium and Sweden, the unemployed could receive the same benefit level forever. Recent policy reforms, however, reduced the benefits for the long-term unemployed relative to the short-term unemployed.

This paper proposes and implements an evidence-based framework to characterize the optimal time profile of UI benefits and evaluate the welfare consequences of changes in the profile of existing UI policies. In doing so, this paper aims to bridge three different strands of the literature. There is an influential theoretical literature on optimal dynamic policies, but derived in stylized models that are often difficult to connect to the data (e.g. Shavell and Weiss [1979], Hopenhayn and Nicolini [1997], Werning [2002]). An important empirical literature has analyzed the structural dynamics of unemployment, but without drawing the consequences for dynamic policies (e.g. Van den Berg [1990], Eckstein and Van den Berg [2007]). Finally, a recent, but growing empirical literature started evaluating social insurance design using the so-called *sufficient statistics* approach, but this literature has been mostly silent about the dynamic features of social insurance programs (e.g. Chetty [2008a], Schmieder et al. [2012b]).

In the spirit of the sufficient-statistics approach we derive a characterization of the optimal profile of unemployment benefits based on a limited set of high-level statistics. This simple, yet robust characterization provides new and transparent insights on the forces affecting the optimal trade-off between insurance and incentives costs throughout the unemployment spell. Our approach also identifies the relevant behavioral responses in this dynamic context to evaluate the welfare consequences of (local) changes in the policy. Our analysis therefore provides a clear guide for dynamic policy design and in particular for analyzing how insurance value and incentive cost of unemployment benefits evolve over the unemployment spell. We implement this approach empirically, using Swedish administrative data on unemployment, linked with survey data on consumption and tax register data on income and wealth.

We start by setting up a rich, dynamic model of unemployment that incorporates job search and consumption decisions and which allows for unobservable heterogeneity and duration dependence in job finding rates in addition to unobservable heterogeneity in assets and preferences. Using dynamic envelope conditions, we show that the Baily-Chetty intuition (Baily [1978], Chetty [2006]) generalizes for a dynamic unemployment policy: the UI benefits paid at time t of the unemployment spell should balance the corresponding insurance value with the implied moral hazard (or incentive) cost at the margin. At the optimal policy, the marginal value and cost are equalized for any part of the benefit profile. If they are not, one can identify (local) policy changes that increase welfare.

Like in the original Baily-Chetty formula, the insurance value and moral hazard cost of the dynamic policy can be expressed as a function of identifiable and estimable statistics. The incentive cost of benefits paid at time t of the unemployment spell depends only on the behavioral revenue effect, *i.e.,* the effect of this benefit level on the government expenditures through agents' unemployment responses. This fiscal externality is fully captured by the responses of the survival rate throughout the unemployment spell, weighted by the benefit levels paid. In other words, regardless of the primitives underlying the dynamics of the agents' search behavior (e.g., heterogeneity vs. true duration dependence in exit rates), these survival rate responses are sufficient to evaluate the incentive cost of changes in the benefit profile. From the insurance perspective, the marginal value of benefits paid at time t of the unemployment spell depends only on the average marginal utility of consumption for agents unemployed at time t. To capture this insurance value, we explore the robustness of the so-called consumption implementation approach, which consists in evaluating the marginal utility of consumption using observed consumption patterns over the unemployment spell and calibrated values of risk aversion. We demonstrate how the nature of selection into longer unemployment spells can affect the relative consumption smoothing gain from benefits paid at different time t of the unemployment spell.

The empirical part of this paper provides novel insights on the incentive costs and insurance value of UI benefits over the unemployment spell. We use a unique administrative dataset in Sweden based on unemployment and tax and asset registers for the universe of Swedish individuals from 1999 until 2007, combined with surveys on household consumption for a subset of the population. We first exploit duration-dependent caps on unemployment benefits using a regression kink design. These caps have been affected by several policy reforms, allowing us to estimate non-parametrically how unemployment survival responds to different variations in the benefit profile. The policy variation also offers compelling placebo settings that confirm the robustness of our approach. We then leverage the comprehensive information on income, transfers and wealth from Swedish registers to construct a residual measure of household expenditures, and, linking this measure to unemployment records, we identify how consumption expenditures change with unemployment and the duration of an unemployment spell in particular. We provide complementary and robustness analysis using survey data on consumption expenditures linked to unemployment records.

Our empirical analysis provides the following main results:

First, unemployment durations respond significantly to changes in benefit levels, whether these benefits are paid early or later in the spell. Furthermore, we find that the response to changes in benefits paid earlier in the spell is larger than the response to benefits paid later in the spell. This result may seem surprising. All else equal, the incentive cost from increasing benefits for the long-term unemployed is expected to be larger as it also discourages the short-term unemployed from leaving unemployment when they are forward-looking. Using the same regression-kink design, we do provide clear evidence that exit rates early in the spell respond to benefit changes applying later in the spell, but also that agents become less responsive to comparable changes in the policy later in the spell. Importantly, such non-stationary forces, which may be driven by duration dependence or dynamic selection on returns to search effort over the unemployment spell, are large enough to offset the significant effect of forward-looking incentives.

Second, consumption expenditures drop substantially and early in the spell. We find that expenditures drop on average by 4.4% in the first 20 weeks of unemployment, compared to their pre-unemployment level. This drop deepens to 9.1% on average for those who are unemployed for longer. We also leverage the richness of the data to document the mechanisms underlying the observed patterns of consumption, and how they translate into consumption smoothing gains of UI benefits over the spell. We show that the role of selection effects in explaining the observed consumption patterns is rather limited. We document the role of assets and liquidity constraints in explaining the drop in consumption over the spell, and show the limited role of the the added-worker effect in smoothing the unemployment shock, even for long-term unemployed. The consumption surveys also shed light on the types of consumption goods that individuals adjust over the spell, including substitution towards home production and away from durable goods. Taken together, our evidence consistently indicates that the consumption smoothing value of UI is higher for the long-term unemployed.

Finally, our empirical estimates can be mapped into the sufficient statistics derived in our theoretical analysis, allowing for a transparent local evaluation of the benefit levels in a two-part benefit profile. Our baseline implementation assumes preference homogeneity, separability between consumption and leisure, and the absence of other externalities (besides the described fiscal externality). Since we find that in Sweden the incentive costs are high relative to the drop in consumption throughout the unemployment spell, our implementation suggests that reducing the generosity of either part of the benefit profile increases welfare for reasonable values of risk aversion. The incentive cost, however, decreases over the unemployment spell as do the consumption expenditures of the unemployed. In the absence of offsetting selection on preferences, our estimates suggest a welfare gain of decreasing the marginal krona spent on the short-term unemployed that is more than twice as high as decreasing the marginal krona spent on the long-term unemployed. As the benefit profile was flat during our period of study, this suggests that the introduction of an inclining benefit profile could have increased welfare. We provide a complementary welfare analysis based on a structural model. We use our empirical analysis of mechanisms to inform the choice of primitives of the model, that we then calibrate to match the sufficient statistics underlying our local policy recommendations. The structural analysis allows us to go beyond these local policy recommendations, but relies on the structure of the calibrated model. The calibration exercise indicates that an inclining tilt remains welfare improving when lowering the overall generosity of the policy.

Our paper contributes to several literatures. First, the sufficient-statistics approach has a long tradition in UI starting with Baily [1978], implemented by Gruber [1997], generalized by Chetty [2006] and recently reviewed in Chetty and Finkelstein [2013]. To date, this literature has focused almost entirely on the optimal average generosity of the system.[1] Conversely, the theoretical literature on the optimal time profile of UI has generated results in stationary, representative-agent models, which are hard to take to the data. Our analysis shows how the previously identified forces (e.g., in Hopenhayn and Nicolini [1997] and Shimer and Werning [2008]) come together, but also integrates heterogeneity and duration-dependence (see for example Shimer and Werning [2006], Pavoni [2009]). Second, our empirical analysis of unemployment responses relates to a long literature on labor supply effects of social insurance. This literature has focused on exploiting isolated sources of variation in one part of the benefit profile.[2] We contribute by explicitly using duration-dependent variation in benefits and identifying the welfare-relevant unemployment responses for *multiple parts* of the benefit profile. Our analysis indicates that differences in the timing of the benefit variation could explain different estimates of unemployment responses in the literature. Finally, a large literature has used consumption surveys to analyze consumption drops as a response of income shocks and unemployment in particular (e.g., Gruber [1997]). We provide novel insights on the evolution of consumption as a function of time spent unemployed. We do this using administrative data on income and wealth to construct a residual, registry-based measure of consumption, which allows us to identify moral hazard costs and consumption responses for the very same sample of unemployed.[3]

The remainder of the paper proceeds as follows. Section 1 analyzes the characterization and implementation of sufficient-statistics formulae for the evaluation of local policy changes in a dynamic model of unemployment. Section 2 describes our data and the policy context in Sweden. Section 3 describes our regression kink design and provides estimates of the policy-relevant unemployment elasticities. Section 4 analyzes how consumption evolves during the unemployment spell and how this translates to the consumption smoothing gains of UI. Section 5 analyzes welfare complementing the implementation of the sufficient statistics with a calibration exercise of our structural model. Section 6 concludes.

1 Model

This section sets up a dynamic model of unemployment and identifies the key trade-offs in designing the time profile of the unemployment benefits. We provide a characterization of the optimal profile in a non-stationary environment with heterogeneous agents. In the spirit of the "sufficient-statistics" literature, our approach consists in identifying the minimal level of information necessary

[1]Recent dynamic extensions of the Baily-Chetty formula can be found in Schmieder et al. [2012b], analyzing the potential benefit duration for a given benefit level, and in Spinnewijn [2015], providing a formula for the optimal intercept and slope of a linear benefit profile.

[2]See Krueger and Meyer [2002] for a review on the labor supply effects of social insurance. Recent examples analyzing variation in UI are Rothstein [2011], Valletta and Farber [2011], Landais [2015], Card et al. [2015] and Mas and Jonhston [2015].

[3]See for instance Mogstad and Kostol [2015], Kreiner et al. [2014] and Pistaferri [2015] for a survey of recent developments of consumption analysis using registry data.

for this characterization. Our focus goes beyond the primitives of the environment and the assumptions on agents' behavior in our specific model. Instead, we aim to identify the observable variables that are relevant for policy in a broad class of models and can be estimated empirically.

1.1 Setup

We first describe the set up of our dynamic unemployment model, building on the dynamic model in Chetty [2006], the agents' choices and the unemployment policy. We try to save on notation in the main text, but provide more details in the technical Appendix A. We consider a partial equilibrium framework with a continuum of agents with mass 1. The model is in discrete time t, starts at $t = 1$ and ends at $t = T$.

Each agent i starts unemployed and remains unemployed until she finds work. Once an agent has found work, she remains employed until the end. When employed, the agent earns w, when unemployed she earns 0. Before the start of the model, the government commits to an unemployment policy P providing insurance against the unemployment risk: the policy specifies an unemployment benefit profile depending on the duration of the ongoing unemployment spell (i.e., a benefit level b_t for each time t if the unemployment spell is still ongoing) and a uniform tax τ paid when employed.

Job search Each agent i decides at each time t how much search effort $s_{i,t}$ to exert as long as she is unemployed. This effort level determines the agent's exit probability at time t. We denote the agent's exit rate out of unemployment at time t by $h_{i,t}(s_{i,t})$. We allow this mapping to depend on the type of agent i, capturing heterogeneity in employability across agents, and the time t she has spent unemployed, capturing differences in employment prospects due to the time spent unemployed.[4] The agent's probability to be unemployed after t periods equals the survival probability $S_{i,t} \equiv \prod_{t'=1}^{t-1}(1 - h_{i,t'}(s_{i,t'}))$ with $S_{i,1} = 1$. While we cannot observe an agent's specific survival probability, we can observe the population average of survival probabilities $S_t \equiv \int S_{i,t} di$.

Intertemporal Consumption Each agent i decides at each time t how much to borrow or save (at interest rate r). An agent starts the unemployment spell with asset level $a_{i,1}$, but borrowing constraints prevent her from running down her asset below $\bar{a}_i$ at any time. The agent's savings decisions determine her consumption level throughout the unemployment spell and when reemployed. We denote these levels by $c^u_{i,t}$ and $c^e_{i,t}$ for when unemployed and employed respectively.[5] While we cannot observe an agent's contingent consumption plan, we can observe average levels of consumption, for example at different spell lengths, $\bar{c}^u_t = \int (S_{i,t}/S_t) c^u_{i,t} di$.

[4] Potential reasons for *true* duration-dependence in exit rates are human capital depreciation (see Acemoglu [1995] and Ljungqvist and Sargent [1998a]) and stock-flow sampling (see Coles and Smith [1998]). We assume exogenous exit rate functions that only depend on the agent's search, but do not directly depend on other job seekers' search like in rationing models (e.g., Michaillat [2012b]) or on the unemployment policy like in employer screening models (e.g., Lockwood [1991]). We discuss this further in Section 1.4 and Appendix Section A.2.

[5] The agent's consumption choice at time t will depend on her unemployment history. In particular, even when employed, the agent's unemployment history will have affected her asset accumulation and thus her optimal consumption at time t. We introduce formal notation to denote the relevant state variables in the technical appendix.

Preferences Each agent i has time-separable preferences (with discount factor β) with per-period utility increasing in consumption, but decreasing in search efforts exerted when unemployed. We denote agent i's per-period utility by $v_i^u(c_{i,t}^u, s_{i,t})$ and $v_i^e(c_{i,t}^e)$. We allow for preference heterogeneity and non-separable preferences in the characterization of the optimal policy, but will assume preference homogeneity and separability between consumption and efforts in our baseline implementation. Each agent chooses how much to search and how much to consume in order to maximize her expected utility, taking the unemployment policy P as given. The dynamics of the agent's behavior depend on her assets and the time spent unemployed in addition to the unemployment policy. To reduce notation, we will drop the arguments of the agent's behavior. We denote the agent's value function of her maximization problem by $V_i(P)$, accounting for her optimal consumption and search choices and potentially binding borrowing constraints.

Unemployment Policy We characterize the welfare impact of local deviations from the unemployment policy P. In particular, we consider changes in the benefit b_t paid at time t of the unemployment spell, starting from any benefit profile $\{b_t\}_{t=1}^T$. Our expressions naturally generalize for changes in step-wise policies, paying benefit level $b_{\mathbf{k}}$ for part $\mathbf{k}$) of the unemployment spell (from time $B_{\mathbf{k-1}}$ until $B_{\mathbf{k}}$. Flat benefit profiles with no or few steps are very common in practice. In Sweden, the unemployment policy in Sweden is entirely flat for some workers and exists of two parts for other, with the benefit dropping to a lower (positive) level at twenty weeks of unemployment.[6] Our implementation evaluates the benefit levels of this two-tier benefit profile.

The government's budget depends on the expected benefit payments paid to the unemployed and the expected tax revenues received from the employed. Note that the average *unemployment duration* D simply equals the sum of the survival rates at each duration $\sum_{t=1}^T S_t$. Similarly, $D_{\mathbf{k}} = \Sigma_{B_{\mathbf{k-1}}+1}^{B_{\mathbf{k}}} S_t$ denotes the expected time spent unemployed while receiving benefit $b_{\mathbf{k}}$, which we refer to as the average *benefit duration.* We ignore time discounting in our characterization of the optimal policy (i.e., $1 + r = \beta = 1$), but generalize this in the technical Appendix A.

The government's budget simplifies to

$$G(P) = [T - D]\tau - \Sigma_{t=1}^T S_t b_t. \tag{1}$$

Social welfare associated with an unemployment policy P can be written as the Lagrangian

$$W(P) = \int V_i(P)\, di + \lambda \left[G(P) - \bar{G}\right], \tag{2}$$

where λ equals the Lagrange multiplier on the government's budget constraint and $\bar{G}$ is an exogenous revenue constraint. We assume that the social welfare function is differentiable.

[6] We discuss the details of the Swedish unemployment policy in Section 2.1.

1.2 Dynamic Sufficient Statistics

We consider the welfare impact of local policy deviations, which we decompose into the corresponding consumption smoothing gains and moral hazard cost. Our approach does not provide an explicit characterization of the optimal policy, but is sufficient to test for the (local) optimality of the policy in place. Evaluating local policy changes, away from the optimal policy, is of interest to identify how welfare can be increased and how the policy can be changed towards the optimal policy (if welfare is concave in the policy variables).

Consider now an increase in the benefit level b_t in period t of the unemployment spell. The total impact on welfare depends on how much the unemployed value this increase in benefits b_t relative to its budgetary cost,

$$\frac{\partial W(P)}{\partial b_t} = \int \frac{\partial V_i(P)}{\partial b_t} di + \lambda \frac{\partial G(P)}{\partial b_t}. \tag{3}$$

This welfare effect depends on the agents' behavioral responses to the policy, but only to the extent that the agents' behavior has consequences that they did not internalize themselves. Indeed, an agent's response to a policy change will have only a second order impact on her own welfare $V_i(P)$. Assuming differentiability, this follows from the envelope conditions $\partial V_i / \partial x^z_{i,t'} = 0$, which hold for any behavior $x^z_{i,t'}$ the agent optimizes over, at any time t', when employed ($z = e$) or unemployed ($z = u$) and when the borrowing constraint is binding or not (see Chetty [2006]).[7] So we only need to account for the impact of behavioral responses on the government's budget $G(P)$ and the direct impact of the policy change on agents' welfare, which proves particularly powerful in this dynamic context.

Moral Hazard Consider first the budgetary impact from an increase in b_t. The first effect from increasing the benefit level is mechanical and depends on the share of workers still unemployed after t periods, S_t. The second effect is behavioral and is determined by the budgetary cost of the agents' reduced search in response to the more generous benefit. This depends on the induced change in the average survival rates *throughout* the unemployment spell,

$$\frac{\partial G(P)}{\partial b_t} = -S_t - \Sigma_{t'=1}^{T} \frac{\partial S_{t'}}{\partial b_t} (b_{t'} + \tau) = -S_t \times \left[1 + \Sigma_{t'=1}^{T} \frac{S_{t'} (b_{t'} + \tau)}{S_t b_t} \varepsilon_{t',t} \right] \tag{4}$$

$$\equiv -S_t \times [1 + MH_t]. \tag{5}$$

The moral hazard cost MH_t of an increase in b_t simply equals the weighted sum of the elasticities $\varepsilon_{t',t} = (\partial S_{t'} / \partial b_t) / (S_{t'} / b_t)$ of the average survival rate $S_{t'}$ with respect to the benefit level b_t. The elasticities are weighted by the relative share of the budget spent at different times during the unemployment spell. The budgetary spillover effects of a change in b_t on other parts of the policy

[7] Changes in the choice variables might be discontinuous in response to small policy changes. In principle we can allow for such discontinuous behavioral responses if they average out when integrating across heterogeneous individuals so that the social welfare function is differentiable.

is less relevant the less generous these other parts are. There is, however, a correction for the tax rate because more time spent unemployed also reduces the taxes received from employment.

Evaluated at a flat profile ($b_t = \bar{b}$ for all t), the moral hazard cost of an increase at time t is fully determined by the response in the average duration D, scaled by the survival rate at t,

$$MH_t = \frac{\partial D / \partial b_t}{S_t}\left(\bar{b} + \tau\right) = \frac{D\left(\bar{b} + \tau\right)}{S_t \bar{b}} \varepsilon_{D,b_t}. \tag{6}$$

This average duration response combines the potentially heterogeneous responses by unemployed workers throughout the unemployment spell, including responses earlier in the spell in anticipation of the increase in b_t and selection effects later in the spell due to the increase in b_t.

Consumption Smoothing Let us now turn to the insurance value of an increase in the benefit b_t. Due to the envelope conditions, the welfare increase is completely captured by the marginal utility of consumption at this time of the spell for the agents who are still unemployed,

$$\begin{aligned}\int \frac{\partial V_i\left(P\right)}{\partial b_t} di &= \int S_{i,t} \frac{\partial v_i^u\left(c_{i,t}^u, s_{i,t}\right)}{\partial c_{i,t}^u} di, \\ &\equiv S_t \times E_t^u\left[\frac{\partial v_i^u\left(c_{i,t}^u, s_{i,t}\right)}{\partial c_{i,t}^u}\right].\end{aligned}$$

As defined above, the expectation operator E_t^u takes the weighted average over all individuals' marginal utility of consumption in the t-th period of the unemployment spell (with weights $S_{i,t}/S_t$). By analogy to the budgetary cost, we can write

$$\int \frac{\partial V_i\left(P\right)}{\partial b_t} di / \lambda = S_t \times \left[1 + CS_t\right], \tag{7}$$

where the consumption smoothing gain $CS_t \equiv \{E_t^u\left[\frac{\partial v_i^u\left(c_{i,t}^u, s_{i,t}\right)}{\partial c_{i,t}^u}\right] - \lambda\}/\lambda$. Since the Lagrange multiplier λ equals the shadow cost of the government's budget constraint, the consumption smoothing gains can be interpreted as the return of a government dollar spent to the unemployed in period t of the unemployment spell relative to the value of an unconditional transfer.[8] Importantly, in spite of potential heterogeneity across agents and in their responses to the policy change, the welfare gain is fully captured by the average marginal utility of consumption at time t of the unemployment spell.

Welfare Impact An optimal unemployment policy balances consumption smoothing gains and moral hazard costs. A dynamic benefit profile allows solving this trade-off at each point during

[8] Note that when the government can provide such lump sum transfer, it would be optimally set such that λ equals the average marginal value of resources at the start of this model. More generally, the consumption smoothing gain CS_t corresponds to the net social marginal welfare weight assigned to the unemployed at time t.

the unemployment spell. Combining expressions (3), (5) and (7), we find

$$\frac{\partial W(P)}{\partial b_t} = \lambda S_t \times [CS_t - MH_t].$$

An increase (decrease) in benefit b_t increases welfare as long as the consumption smoothing gains are larger (smaller) than the moral hazard cost. This implies a natural characterization of the optimal policy:

Proposition 1. *Consider an unemployment policy P, charging tax τ to the employed and paying a dynamic benefit profile $\{b_t\}_{t=1}^T$ to the unemployed. Assuming differentiability, an interior, optimal policy needs to satisfy*

$$\frac{E_t^u\left[\frac{\partial v_i^u\left(c_{i,t}^u, s_{i,t}\right)}{\partial c_{i,t}^u}\right] - \lambda}{\lambda} = \Sigma_{t'=1}^T \frac{S_{t'}\left(b_{t'}+\tau\right)}{S_t b_t} \times \varepsilon_{t',t} \text{ for each } t, \tag{8}$$

$$\frac{\lambda - E^e\left[\frac{\partial v_i^e\left(c_{i,t}^e\right)}{\partial c_{i,t}^e}\right]}{\lambda} = \Sigma_{t'=1}^T \frac{S_{t'}\left(b_{t'}+\tau\right)}{\left(T-D\right)\tau} \times \varepsilon_{t',\tau}, \tag{9}$$

and the budget constraint $G(P) = \bar{G}$.

Proof. *See Appendix A.*■

The expectation operator E_t^u is defined as before and takes the weighted average over all individuals unemployed at time t (with weights $S_{i,t}/S_t$). Similarly, E^e takes the weighted average over all individuals and times in employment (with weights $\left(1 - S_{i,t}\right)/\left[T-D\right]$). The conditions for all benefit levels and the tax level in Proposition 1 can be combined to recover the well-known Baily-Chetty formula (Baily [1978] and Chetty [2006]) for a flat benefit profile ($b_t = \bar{b}$),[9]

$$\frac{E^u\left[\frac{\partial v_i^u\left(c_{i,t}^u, s_{i,t}\right)}{\partial c_{i,t}^u}\right] - E^e\left[\frac{\partial v_i^e\left(c_{i,t}^e\right)}{\partial c_{i,t}^e}\right]}{E^e\left[\frac{\partial v_i^e\left(c_{i,t}^e\right)}{\partial c_{i,t}^e}\right]} \cong \frac{\bar{b}+\tau}{\bar{b}} \varepsilon_{D,\bar{b}}. \tag{10}$$

Our analysis extends the "sufficient statistics" approach to the dynamics of the unemployment policy, aiming to provide a simple, yet robust guide for dynamic policy design. As we argue below, the characterization depends on a limited set of empirically implementable moments and is robust to the primitives and specific assumptions of the underlying model. Our dynamic extension also overcomes challenges that have constrained empirical and theoretical work in identifying the key dynamic forces. Empirically, identifying the role of different, non-stationary forces underlying a

[9]The expectation operator E^u in condition (10) takes the weighted average over all unemployment periods (with weights $S_{i,t}/D$). The approximation relies on the unemployment response to taxes to be small. The exact expression for the right-hand side is $[1 + \frac{\bar{b}+\tau}{\bar{b}}\varepsilon_{D,\bar{b}}]/[1 + \frac{\bar{b}+\tau}{\tau}\varepsilon_{T-D,\tau}] - 1$. Note that the standard Baily-Chetty formulation uses the elasticity wrt a budget-balanced increase in the benefit level, joint with an increase in the tax level, and ignores other tax distortions in the economy. Our model allows the tax to cover general expenditures $\bar{G}$ and our expressions are in terms of partial elasticities, which are more transparent for multi-dimensional policies and correspond more directly to the policy variation we exploit in the empirical analysis.

job seeker's environment, including the role of unobserved heterogeneity, proves daunting. Several studies have tried to estimate or calibrate the contribution to the negative duration-dependence of exit rates from dynamic selection effects, *true* duration-dependence in the search environment (e.g., skill-depreciation or stock-flow sampling of vacancies), or an interaction of the two (e.g., duration-based employer screening).[10] Theoretically, it has also proven difficult to derive the optimal benefit profile and, in particular, the impact of non-stationary forces and heterogeneity (see Shimer and Werning [2006], Pavoni [2009]). In contrast, Proposition 1 provides a robust mapping from a non-stationary model with heterogeneous agents into a set of implementable moments to evaluate the benefit profile.[11]

1.3 Implementation

We first consider the implementability of our characterization, which guides our empirical analysis in Sections 3 and 4. Our focus is on the benefit levels of a two-part policy $(b_{\mathbf{1}}, b_{\mathbf{2}}, B)$ paying benefit $b_{\mathbf{1}}$ until time B and $b_{\mathbf{2}}$ thereafter, like in place in Sweden (see Section 2.1) and illustrated in Panels A.I and B.I of Figure 1. This section clarifies additional assumptions and the policy variation required for the implementation we propose.

Moral hazard cost An extensive literature has analyzed unemployment responses to changes in the unemployment policy. Our analysis indicates that it is essential to have variation in unemployment benefits at different times during the unemployment spell.

For a two-part profile, the benefit duration $D_{\mathbf{1}}$ $(D_{\mathbf{2}})$, which denotes the expected time spent receiving benefit $b_{\mathbf{1}}$ $(b_{\mathbf{2}})$, corresponds to the area under the survival function before (after) B, as illustrated in Panels A.II and B.II. The moral hazard cost of changing the benefit level $b_{\mathbf{k}}$ during part $\mathbf{k}$ of the policy fully depends on the response in both benefit durations,

$$MH_{\mathbf{1}} = \frac{b_{\mathbf{1}}+\tau}{b_{\mathbf{1}}}\varepsilon_{D_{\mathbf{1}},b_{\mathbf{1}}} + \frac{D_{\mathbf{2}}\left(b_{\mathbf{2}}+\tau\right)}{D_{\mathbf{1}}b_{\mathbf{1}}}\varepsilon_{D_{\mathbf{2}},b_{\mathbf{1}}} \tag{11}$$

$$MH_{\mathbf{2}} = \frac{D_{\mathbf{1}}\left(b_{\mathbf{1}}+\tau\right)}{D_{\mathbf{2}}b_{\mathbf{2}}}\varepsilon_{D_{\mathbf{1}},b_{\mathbf{2}}} + \frac{b_{\mathbf{2}}+\tau}{b_{\mathbf{2}}}\varepsilon_{D_{\mathbf{2}},b_{\mathbf{2}}} \tag{12}$$

Panel A and Panel B of Figure 1 illustrate how estimating the moral hazard costs requires duration-dependent policy variation. Rather than having benefits change throughout the spell, which would be sufficient to evaluate a flat policy, we need changes in benefits paid only to the short-term $db_{\mathbf{1}}$ or to the long-term unemployed $db_{\mathbf{2}}$.

Our evaluation of the benefit profile is conditional on B, which determines the potential duration of the two parts. Still, our expressions can be used to approximate the moral hazard cost of changing the potential benefit durations, as analyzed in Schmieder et al. [2012b]. Indeed, an increase in

[10] See Ljungqvist and Sargent [1998a] and Machin and Manning [1999] for reviews on the negative duration dependence of exit rates out of unemployment. See Kroft et al. [2013] and Alvarez et al. [2016] for recent examples.

[11] In Section 5.2, we also show how this mapping can be useful to uncover the role of stationary vs. non-stationary forces from the estimated moments and understand their respective role for the optimal timing of benefits.

potential benefit duration from B to $B+1$ corresponds to a (discrete) change of benefit b_{B+1} at $B+1$ (from level $b_{\mathbf{2}}$ to level $b_{\mathbf{1}}$), where the moral hazard cost of a (marginal) change in b_{B+1} equals

$$MH_{b_{B+1}} = \frac{\partial D_{\mathbf{1}}/\partial b_{B+1}}{S_{B+1}} \left(b_{\mathbf{1}} + \tau\right) + \frac{\partial D_{\mathbf{2}}/\partial b_{B+1}}{S_{B+1}} \left(b_{\mathbf{2}} + \tau\right).$$

While there is no such policy variation in the Swedish context, Schmieder and Von Wachter [2016] review recent empirical work that analyzes either duration responses in unemployment or UI benefit receipt to changes in potential benefit duration. The expression above shows that the responses in both the benefit duration $D_{\mathbf{1}}$ and the average duration D (since $D_{\mathbf{2}} = D - D_{\mathbf{1}}$) affect the fiscal externality, unless the tax is small ($\tau \cong 0$) and no benefits are paid after time B ($b_{\mathbf{2}} = 0$). Importantly, any evaluation of the potential benefit duration would be conditional on the benefit levels and does not allow to evaluate the tilt of the profile itself.

Consumption Smoothing Attempts at quantifying the consumption smoothing gains of UI policies have been more scarce as the estimation of differences in marginal utility levels proves difficult in practice. We follow the "consumption implementation" approach (Gruber [1997], Chetty [2006]), relating the difference in marginal utilities to the difference in consumption levels, and extend this approach to our dynamic setting. Using consumption wedges to actually quantify the relevant consumption smoothing gains of UI requires the following assumptions:[12]

First, we rely on approximations of the marginal utility of consumption using Taylor expansions, assuming that third- and higher-order derivatives of the utility function are small. That is,

$$\frac{\partial v_i^u}{\partial c} \left(c_{i,t}^u, s_{i,t}\right) \cong \frac{\partial v_i^u}{\partial c} \left(\tilde{c}, s_{i,t}\right) \times \left[1 - \tilde{\gamma}_{i,t} \times \frac{\tilde{c} - c_{i,t}^u}{\tilde{c}}\right],$$

where $\tilde{\gamma}_{i,t} \equiv \tilde{c}\frac{\partial^2 v_i^u}{\partial c^2} \left(\tilde{c}, s_{i,t}\right) / \frac{\partial v_i^u}{\partial c} \left(\tilde{c}, s_{i,t}\right)$ equals the relative risk aversion.[13]

Second, we assume that preferences over consumption are separable from leisure, i.e., $\partial v_i^u \left(c, s\right) / \partial c = \partial v_i^e \left(c\right) / \partial c = v_i' \left(c\right)$, so that consumption smoothing benefits do not depend on other behavior or the employment status itself, but only on the consumption wedges. This excludes potentially important complementarities between consumption and leisure during unemployment.[14]

Third, we express the consumption smoothing gains from an increase in unemployment benefits relative to an increase in resources just before the onset of the unemployment spell (denoted by

[12] Note that the limitations of the consumption-based implementation have inspired alternative approaches relating the marginal utility gap to observable behavioral responses: Chetty [2008a] decomposes unemployment responses in liquidity and substitution effects, Shimer and Werning [2007] analyze reservation wage responses. The extension of these alternative approaches to a dynamic setting seems promising, but requires that the policy variation used for the static implementation also changes over the unemployment spell.

[13] If the third-order derivative of the utility function is non-negligible, the consumption smoothing gains depend on an additional term that depends on the coefficient of relative prudence, corresponding to precautionary saving motives (see Chetty [2006]). We calculate the magnitude of this approximation error in section 5.3.

[14] In subsection 4.2, we discuss the issues related to this assumption in more detail. One example is the substitution towards home production when no longer employed (e.g., Aguiar and Hurst [2005]).

$v_i'(c_{i,0})$). This normalization emphasizes the insurance value of the policy.[15]

Fourth, we assume that preferences are homogeneous (i.e., $v_i(c) = v(c)$), but we consider the implications of preference-based selection over the unemployment spell in Section 5.2.

Under these four assumptions, we can approximate the consumption smoothing gains by[16]

$$CS_{\mathbf{k}} = \frac{\frac{1}{D_{\mathbf{k}}}\int \Sigma_{B_{\mathbf{k-1}}+1}^{B_{\mathbf{k}}} S_{i,t} v'\left(c_{i,t}^{u}\right) di - \int v'(c_{i,0})\, di}{\int v'(c_{i,0})\, di} \tag{13}$$

$$\cong \frac{v'(\bar{c}_{\mathbf{k}}^{u}) - v'(\bar{c}_0)}{v'(\bar{c}_0)} \cong -\frac{v''(\bar{c}_0)\,\bar{c}_0}{v'(\bar{c}_0)} \times \frac{\bar{c}_0 - \bar{c}_{\mathbf{k}}^{u}}{\bar{c}_0}, \tag{14}$$

where $\bar{c}_0$ and $\bar{c}_{\mathbf{k}}^{u}$ denote the average consumption level before the onset of the spell and during part $\mathbf{k}$ of the spell. The resulting expression directly relates to the original approximation in Baily [1978] and highlights the role of the profile of the average consumption level over the unemployment spell to evaluate the unemployment benefit profile. If the unemployed consume less the longer they are unemployed, ceteris paribus, unemployment benefits are more valuable later in the spell.

For a two-part profile, the implementation thus comes down to calculating the average wedge in consumption for the short-term unemployed and the long-term unemployed, as illustrated in Panel C of Figure 1. Importantly, no policy variation is needed to estimate these consumption wedges.

1.4 Robustness

We briefly consider the robustness of our sufficient-statistics characterization in a dynamic context. As argued before by Chetty [2006], the set of moments in Proposition 1 are sufficient to provide local evaluations of the unemployment policy, independently of the underlying primitives. That is, when different values for our models' parameters map into the same values for the identified moments for a given policy, the local policy recommendations remain the same. In particular, our dynamic model explicitly allows for (exogenous) heterogeneity in exit rate functions across agents ($h_{i,t}(\cdot) \neq h_{j,t}(\cdot)$) and variation in exit rates over the unemployment spell ($h_{i,t}(\cdot) \neq h_{i,t'}(\cdot)$). While separating unobserved heterogeneity and *true* duration-dependence in exit rates is hard, our approach shows that this is unnecessary for estimating the moral hazard cost and its evolution over

[15]In our stylized model, all individuals start unemployed, so the Lagrange multiplier (evaluated at the optimal policy) equals $\lambda = \int \partial V/\partial a_{i,1} di = \int \partial V/\partial b_1 di$. In Appendix Section A.2.4, we consider a more general model where agents can start employed or unemployed and may experience multiple spells. The Lagrange multiplier then corresponds to the average marginal utility of consumption at the start of the model across all individuals. By considering the marginal utility of consumption before the onset of the unemployment spell in our implementation, we are capturing the insurance value of the unemployment policy, while ignoring the value of redistributing between different types of workers who face different layoff risks. Importantly, the evaluation of (budget-balanced) changes in the benefit profile is independent of this normalization.

[16]The first approximation relies on a Taylor expansion of $v'\left(c_{i,t}^{u}\right)$ for each individual i and time t around the average consumption level during the corresponding part $\mathbf{k}$ of the unemployment spell, $\bar{c}_{\mathbf{k}}^{u} \equiv \frac{1}{D_{\mathbf{k}}}\int \Sigma_{B_{\mathbf{k-1}}+1}^{B_{\mathbf{k}}} c_{i,t}^{u} S_{i,t} di$, using

$$\frac{1}{D_{\mathbf{k}}}\int \Sigma_{B_{\mathbf{k-1}}}^{B_{\mathbf{k}}} S_{i,t} v''(\bar{c}_{\mathbf{k}}^{u})\left[c_{i,t}^{u} - \bar{c}_{t}^{u}\right] di = 0.$$

This also applies at time $t = 0$, just before the onset of the unemployment spell. The second approximation simply uses a Taylor expansion of $v'(\bar{c}_{\mathbf{k}}^{u})$ around $\bar{c}_0$.

the unemployment spell. The intuition is that only the survival responses need to be known, as they fully determine the fiscal externality of job seekers' behavior. As is well known, this result criticially relies on the application of the envelope conditions for the job seekers' behavior (see Chetty [2006], Chetty and Finkelstein [2013]). The result also indicates that the foundations of our dynamic model of search and consumption can be further extended without (substantially) changing the characterization of the optimal benefit profile. That is, the same set of moments will continue to determine the local evaluation of the unemployment policy.[17] Our baseline setup illustrates this robustness explicitly for exogenous heterogeneity across agents and across durations, but the intuition generalizes to other models of search and self-insurance.[18]

While our implementation is robust to heterogeneity in employment prospects and assets and the corresponding dynamic selection, our baseline implementation assumes preferences that are homogenous and separable in consumption and search efforts. Under this assumption the relative consumption smoothing gains, for example for short-term and long-term unemployed, simplify to the corresponding relative consumption drops, which thus become sufficient to recommend welfare-improving changes in the tilt of the benefit profile. This avoids the well-known challenge for the consumption-based implementation of translating consumption wedges to welfare (see Chetty and Finkelstein [2013]).[19] We assess this challenging conversion from consumption into welfare in more depth and also gauge the potential for dynamic selection on preferences in Sections 4.2 and 5.2.[20]

Finally, as our characterization critically relies on the application of the envelope theorem, by the same token, the presence of other externalities – not internalized by agents, but relevant for welfare – would affect the optimal policy characterization. Recent work has analyzed the impact of different types of externalities on the characterization of a static unemployment policy.[21,22] These

[17] Chetty [2006] has shown how the simple formula (10) characterizing the flat benefit profile continues to apply with leisure benefits from non-employment, alternative means of self-insurance, spousal labor supply, human capital decisions, etc. See the review chapter by Chetty and Finkelstein [2013] for a more detailed discussion of different advantages and challenges for the sufficient-statistics approach.

[18] In Appendix Section A.2.2 we show how our model can be indeed extended to multiple unemployment spells, allowing for moral hazard on the job and different means of self-insurance, while the same formulae continue to apply. The relevant variables for evaluating the benefit profile are the overall unemployment rate and the survival rates S_t at different unemployment durations, averaged over multiple spells. Layoff responses to UI policy affect the magnitude of the policy-relevant elasticities, but only if moral hazard on-the-job were to be important. In Appendices A and B, we provide and discuss evidence based on the pdf of pre-unemployment wages around a kink in the unemployment policy that indicates that layoff rates do not respond strongly to the unemployment policy in our empirical context.

[19] See Chetty [2008a], or Landais [2015], for the development of alternative methods, exploiting comparative statics of effort choices, in order to to evaluate consumption smoothing gains. These methods could circumvent the issue of having to make assumptions regarding dynamic selection on risk preferences.

[20] Note also that our model assumes a utilitarian social welfare function. See Andrews and Miller [2013] for a discussion on the aggregation of individual welfare gains under preference heterogeneity in the context of the Baily-Chetty formula. With heterogeneous Pareto weights, the dynamic selection based on these weights will matter as well.

[21] For example, Nekoei and Weber [2015], account for the fiscal impact of reservation wage responses, conditional on unemployment duration, which tend to be small relative to the duration responses themselves. Spinnewijn [2015] accounts for internalities due to biased beliefs about employment prospects. Landais et al. [2010] adjust the characterization to account for frictions in the labor market and general equilibrium effects.

[22] In Appendix Section A.2.4, we show how our framework can account for the fiscal externality created by the presence of an income tax used to fund other government expenditures, which we use to analyze the sensitivity of our welfare recommendations.

insights generalize in our dynamic setting, but are of particular relevance for our analysis when the externality depends on the timing of the unemployment benefits. In Appendix Section A.2, we demonstrate this in the context of employer screening (e.g., Lockwood [1991]), which gives rise to negative duration-dependence when employers use unemployment spell length as a negative signal of unobserved productivity. In such a setting, job seekers do not internalize their impact on the hiring probability for other job seekers, which happens through the relative survival rates of different types. As shown by Lehr [2017] for a flat benefit profile, the unemployment policy can affect this hiring externality, but only if the relative survival rate of different productivity types depends on the unemployment policy. We show in appendix that for a dynamic benefit profile, the externality-adjusted moral hazard cost equals

$$MH_t^x \equiv \Sigma_{t'=1}^{T} \frac{S_{t'}}{S_t} \left[\frac{b_{t'} + \tau}{b_t} \varepsilon_{t',t} - E_{t'}^u \left(\omega_{t'}^h \frac{\partial h_{t'}}{\partial b_t} \right) \right], \tag{15}$$

where $\omega_{t'}^h$ corresponds to an agent's welfare gain of finding a job at time t' and $\partial h_{t'}/\partial b_t$ equals the change in the job finding rate due to the employer's hiring response. If the hiring response depends on the timing of the benefits, the externality-adjustment could vary over the spell. This requires the relative survival rate of different productivity types to change with the timing of benefits. In appendix we demonstrate that types with higher returns to search are more responsive to changes in benefits early on, but due to their low survival into longer unemployment spells, can be less responsive to changes in benefits later on. Hence, when types with higher returns to search are also more productive, the hiring externality can be positive for benefits paid early in the spell, but negative for benefits paid late in the spell.

2 Context and Data

To implement our formulae and evaluate the profile of UI benefits, two pieces of empirical evidence are needed. First, one needs to identify and estimate responses of unemployment durations to variations in the benefit profile, i.e., variations in UI benefits at different points of an unemployment spell. Second, one needs to estimate the time profile of consumption to identify how consumption (relative to employment) drops over an unemployment spell.

Our empirical analysis offers contributions on both dimensions by using a unique administrative dataset that we created in Sweden merging unemployment registers, tax registers - with exhaustive information on income and wealth - and household consumption surveys. We present here the institutional background and data used in our empirical implementation.

2.1 Institutional background

In Sweden, displaced workers who have worked for at least 6 months prior to being laid-off are eligible to unemployment benefits, replacing 80% of their earnings up to a cap. In practice, the level of the cap is quite low relative to the earnings distribution and applies to about 50% of unemployed workers. Individuals can receive unemployment benefits indefinitely. To continue receiving benefits after 60 weeks of unemployment, the unemployed must accept to participate in counselling activities and, potentially, active labor market programs set up by the Public Employment Service. Like in other Scandinavian countries, UI in Sweden is administered by different unemployment funds (of which most are affiliated with a labor union) and contributions to the funds are voluntary in principle. Over the period 1999 to 2007, more than 85% of all workers were contributing to an unemployment fund. Our sample focuses on workers with more than 6 months of employment history prior to being laid-off and who contribute to UI funds.

The time profile of benefits has changed during the period we study. Before 2001, the time profile of UI benefits was flat for all unemployed workers.[23] Full-time workers would get daily benefits of 80% of their pre-unemployment daily wage throughout the spell (i.e., for as long as they remain unemployed), with daily benefits capped at 580SEK a day ($\approx$ 63USD a day, or 320USD a week).[24] The cap thus applies for daily wages above 725SEK ($\approx$ 399USD a week).[25] In July 2001, a system of duration-dependent caps was introduced, which created a decreasing time profile of benefits for the unemployed above the threshold wage. The cap for the benefits received during the first 20 weeks of unemployment was increased to 680SEK (daily wage above 850SEK $\approx$ 467.5USD a week) while the cap for benefits received after the first 20 weeks was kept unchanged at 580SEK. In July 2002, the cap for benefits received during the first 20 weeks of unemployment was increased to 730SEK (daily wage above 912.5SEK $\approx$ 500USD a week) and the cap for benefits received after the first 20 weeks was increased to 680SEK.[26]

The 2001 and 2002 reforms introduce variation in the benefit profile which makes it possible to estimate the causal impact of benefits received at different times during the unemployment spell on survival in unemployment. We explain in Section 3 how the 2001 and 2002 variations in the duration-dependent caps can be used in a regression kink design to identify the effects on unemployment durations of UI benefits given in the first 20 weeks of a spell and of benefits given after 20 weeks.

2.2 Data

Unemployment history data come from the HÄNDEL register of the Public Employment Service (PES, Arbetsförmedlingen) and were merged with the ASTAT register from the UI administration (IAF, Inspektionen för Arbetslöshetsförsäkringen) in Sweden. The data contain information from 1999 to 2007 on the date the unemployed registered with the PES (which is a pre-requisite to start

[23] The potential duration of benefits is theoretically infinite in Sweden during our period of interest, and there is no exhaustion point of UI benefits.

[24] We use the following exchange rate: 1SEK $\approx$ 0.11USD.

[25] The daily wage is computed as gross monthly earnings divided by number of days worked in the last month prior to becoming unemployed.

[26] Some unions have launched their own complementary UI-schemes which further increased the cap (by up to 3 times the cap on regular UI) by topping up the regular UI-benefit to 80 percent of the previous wage. Importantly, our regression-kink design analysis focuses on the effect of the 725SEK-kink in the UI schedule, which was removed in 2002 before the introduction of the top-ups, so that all unemployed had to comply to the same kinked schedule of benefits.

receiving UI benefits), eligibility to receive UI benefits, earnings used to determine UI benefits, weekly information on benefits received, unemployment status and participation in labor market programs. We define unemployment as a spell of non-employment, following an involuntary job loss, and during which an individual has zero earnings, receives unemployment benefits and reports searching for a full time job.[27] To define the start date of an unemployment spell, we use the registration date at the PES. The end of a spell is defined as finding any employment (part-time or full-time employment, entering a PES program with subsidized work or training, etc.) or leaving the PES (labor force exit, exit to another social insurance program such as disability insurance, etc.).[28]

These data are linked with the longitudinal dataset LISA which merges several administrative and tax registers for the universe of Swedish individuals aged 16 and above. In addition to socio-demographic information (such as age, family situation, education, county of residence, etc.), LISA contains exhaustive information on earnings, taxes and transfer and capital income on an annual basis. Data on wealth comes from the wealth tax register (Förmögenhetsregistret), which covers the asset portfolios for the universe of Swedish individuals from 1999 to 2007. The register contains detailed information on the stock of all financial assets (including debt) and real assets as of December of each year.[29] For the financial assets, we have information on all savings by asset class (bank accounts, bonds, stocks, mutual funds, private retirement accounts, etc.). The dataset also contains information on total outstanding debt including mortgage debt, consumer credit, student debt, etc. For real estate, we have information on all asset holdings at market value as used for the wealth tax assessment.[30] Data on asset balances as of December is complemented with data on financial asset transactions (KURU) and real estate transactions from the housing registries. The comprehensiveness and detailed nature of both the income and wealth data in Sweden is exceptional, providing a unique opportunity to investigate what means individuals use to smooth consumption (transfers, asset rebalancing, increase in debt, etc.). We take advantage of the richness of the income and wealth data to construct a registry-based measure of annual household consumption expenditures as of December of each year. Our approach builds on previous attempts to measure consumption from registry-data (e.g. Koijen et al. [2014], Browning and Leth-Petersen [2003]) and is closely related to Eika et al. [2017] in exploiting additional information on asset portfolio choices and returns to reduce measurement error and excess dispersion of consumption measures based solely on first-differencing asset stocks. All details regarding the construction of the registry-based

[27]Involuntary job loss is defined as a layoff or a quit following a "valid reason". "Valid reasons" for quitting a job are defined as being sick or injured from working, being bullied at work, or not being paid out one's wage by one's employer. Quits are reviewed by the Public Employment Service at the moment an individual registers a new spell and if the quit is made because of a "valid reason", the individual is eligible for UI and a notification is made in the PES data, allowing us to observe such "quits under valid reasons". Note however that quits are a small fractions of spells in our sample: 95.0% of job separations observed in our data are due to layoffs.

[28]To deal with a few observations without any end date, we censor the duration of spells at two years.

[29]All financial institutions are compelled to report this information directly to the tax administration for the purpose of the wealth tax, which ensures quality and exhaustiveness of the data. The wealth tax was abolished in Sweden in 2007, after which the government collected only limited information on the stock of assets.

[30]All asset holdings are reported to the tax administration at the individual level. We aggregated assets at the household level using household identifiers from the registry data.

measure of consumption are given in Appendix C.1.

The richness of the Swedish administrative data, its universal coverage and panel structure, enable us to identify duration responses and within-household consumption responses to unemployment shocks for the very same sample of individuals, and with a unique degree of precision. An important characteristic of the registry-based measure of consumption though, is that it captures the sum of annual consumption expenditures as of December of each year. While this peculiarity does not represent a serious limitation to identify higher frequency flows of consumption throughout the unemployment spell, as shown in section 4, we complement our data with information on consumption available through the yearly household budget survey (HUT, Hushållens utgifter), which provides direct measures of bi-weekly consumption expenditures at the moment the household is surveyed. From 2003 to 2009, individuals sampled in the HUT can be matched to the registry data, which allows us to reconstruct the full employment history of individuals whose household is surveyed in the HUT. While the HUT has a small sample size, and does not have a panel structure, it provides a flow measure of consumption at the time of the survey, and allows us to explore patterns of consumption responses for different categories of expenditures.

In Appendix Table 1, we provide summary statistics on unemployment, demographics, income and wealth for our sample of unemployed individuals used in the duration analysis of Section 3 and in the consumption analysis of Section 4. The average unemployment spell of unemployed in this sample is 26.8 weeks. The average time spent unemployed during the first twenty weeks of the spell equals $D_1 = 13.5$ weeks. The average time spent unemployed in the second part of the benefit profile (after the first twenty weeks of the spell) equals $D_2 = 13.2$ weeks. The average replacement rate is 72%. Socio-demographic characteristics, reported in Panel II of Table 1, show that the unemployed are relatively young (35 years old on average) and a minority of them are married or cohabiting (39% on average).

Prior to the onset of a spell, the average unemployed in the sample has yearly gross earnings (before any tax or payroll contribution) of 190,300 SEK ($\approx$ 21,000USD).[31] A large fraction of unemployed in the sample starts unemployment with no or negative net wealth. Most wealth is held in the form of real estate. Liquid assets such as bank holdings represent less than 30% of yearly earnings at the start of the spell. Total debt, which mostly comprises mortgage, student loans and credit card debt, is fairly large and represents on average almost 200% of the yearly earnings of an unemployed at the onset of her spell.

3 Duration Responses

This section analyzes unemployment responses to changes in the benefit profile. The presence of duration-dependent caps in the Swedish UI system provides compelling variation in UI benefits at different points in time during the unemployment spell. We exploit this variation using a regression kink (RK) design.

[31] All figures are expressed in constant SEK2003. 1SEK2003 $\approx$ 0.11USD2003.

3.1 Regression-Kink Design: Strategy & Results

The time-dependent caps introduce kinks in the schedule of UI benefits given during the first 20 weeks of unemployment and after 20 weeks of unemployment. Figure 2 shows UI benefits as a function of daily pre-unemployment wages for spells starting between January 1999 and July 2001 (panel A.I), for spells starting between July 2001 and July 2002 (panel B.I) and for spells starting after July 2002 (panel C.I). For spells starting before July 2001, the same cap applies to unemployment benefits given in the first 20 weeks of unemployment ($b_{\mathbf{1}}$) and after 20 weeks of unemployment ($b_{\mathbf{2}}$). The schedule of both $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$ thus exhibits a kink at a daily wage of 725 SEK, generating variation in the policy that allows us to identify the effect of an overall change in the benefit level (i.e., a joint change in $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$) on unemployment durations. For spells starting after July 2001 and before July 2002, the cap for $b_{\mathbf{1}}$ is increased, while the cap in $b_{\mathbf{2}}$ remains unchanged. The relationship between $b_{\mathbf{1}}$ and previous wages therefore becomes linear around the 725 SEK threshold, where the schedule of $b_{\mathbf{2}}$ still exhibits a kink at 725 SEK. This makes it possible to identify the effect on unemployment durations of a change in $b_{\mathbf{2}}$ only. Finally, the cap in $b_{\mathbf{2}}$ is also increased for spells starting after July 2002, so that kinks in the schedule of both $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$ disappear at the 725 SEK threshold. This offers a placebo setting to test for the robustness of our approach at the 725 threshold.

Our identification strategy relies on a RK design. Formally, we consider the general model:

$$Y = y(b_{\mathbf{1}}, b_{\mathbf{2}}, w, \mu),$$

where Y is the duration outcome of interest, μ is (unrestricted) unobserved heterogeneity, and $b_{\mathbf{1}}$, $b_{\mathbf{2}}$ and w (previous daily wage) are endogenous regressors. We are interested in identifying the marginal effect of benefits given during part $\mathbf{k}$ of the spell on the duration outcome Y, $\alpha_{\mathbf{k}} = \frac{\partial Y}{\partial b_{\mathbf{k}}}$. The RK design consists in exploiting the fact that $b_{\mathbf{k}}$ is a deterministic, continuous function of the wage w, kinked at $w = \bar{w}_{\mathbf{k}}$. The RK design relies on two identifying assumptions. First, the direct marginal effect of w on Y should be smooth around the kink point $\bar{w}_{\mathbf{k}}$. Second, the distribution of unobserved heterogeneity μ is assumed to be evolving smoothly around the kink point. This means that the conditional density of the wage ($f_{w|\mu}(\cdot)$) and its partial derivative with respect to the wage ($\partial f_{w|\mu}(\cdot)/\partial w$) are assumed to be continuous in the neighborhood of $\bar{w}_{\mathbf{k}}$. This second assumption implies imperfect sorting around the threshold $\bar{w}_{\mathbf{k}}$, i.e., individuals cannot have perfect control over their assignment in the schedule. We provide in Appendix B various tests to assess the robustness of these identifying assumptions and the validity of our RK design.

Under these two identifying assumptions, $\alpha_{\mathbf{k}}$ can be identified as:

$$\alpha_{\mathbf{k}} = \frac{\lim_{w \to \bar{w}_{\mathbf{k}}^{+}} \frac{\partial E[Y|w]}{\partial w} - \lim_{w \to \bar{w}_{\mathbf{k}}^{-}} \frac{\partial E[Y|w]}{\partial w}}{\lim_{w \to \bar{w}_{\mathbf{k}}^{+}} \frac{\partial b_{\mathbf{k}}}{\partial w} - \lim_{w \to \bar{w}_{\mathbf{k}}^{-}} \frac{\partial b_{\mathbf{k}}}{\partial w}}$$

In practice, we provide estimates $\hat{\alpha}_{\mathbf{k}} = \frac{\hat{\delta}_{\mathbf{k}}}{\nu_{\mathbf{k}}}$ where $\hat{\delta}_{\mathbf{k}}$ is the estimated change in slope between Y

and w at $\bar{w}_{\mathbf{k}}$ and $\nu_{\mathbf{k}}$ is the deterministic change in slope between $b_{\mathbf{k}}$ and w at $\bar{w}_{\mathbf{k}}$. We estimate the former using the following regression model:

$$E[Y|w] = \beta_0 + \beta_1(w - \bar{w}_{\mathbf{k}}) + \delta_{\mathbf{k}}(w - \bar{w}_{\mathbf{k}}) \cdot \mathbb{1}[w \geq \bar{w}_{\mathbf{k}}]. \tag{16}$$

This model is estimated for $|w - \bar{w}_{\mathbf{k}}| \leq h$, where h is the bandwidth size.

Preliminary graphical evidence of a change in slope in the relationship between (total) duration of the unemployment spell and previous daily wage in response to the kink in UI benefits is provided in the right-hand side panels of Figure 2. They plot average unemployment duration in bins of previous daily wage for the three periods of interest. Panel A.II shows a significant change in the relationship between wage and unemployment duration around the 725SEK threshold for spells starting up to July 2001. In this period the schedule of UI benefits exhibits kinks in both $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$ at 725SEK (as shown in Panel A.I). In panel B.II, a significant yet smaller change in slope can be detected at the 725SEK threshold for spells starting between July 2001 and July 2002 when the schedule at 725SEK exhibits a kink in $b_{\mathbf{2}}$ only. Finally, panel C.II shows evidence of perfect linearity in the relationship between wage and unemployment duration around the 725SEK threshold for spells starting after July 2002, when kinks in the schedule at 725SEK are eliminated for both $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$.

RK estimates for the effects of benefits on unemployment duration D are shown in Figure 3, where we report for each policy period the point estimate and 95% robust confidence interval of the change in slope $\hat{\delta}_{\mathbf{k}}$ for the regression model in (16). We use a linear specification and a bandwidth $h = 90$SEK around the 725SEK threshold. Because the change in caps applies to ongoing spells as well, we censor spell duration at their duration as of July 2001 and July 2002, and report here estimates from Tobit models on the right-censored data.[32] In line with the evidence presented in Figure 2, the estimated change in slope is large and significant for spells starting before July 2001. In Figure 3 we also report the implied benefit elasticity of unemployment duration, computed as $\varepsilon_{D,b_{\mathbf{k}}} = \hat{\delta}_{\mathbf{k}} \cdot \frac{\bar{w}_{\mathbf{k}}}{D^{cap}}$, where $\hat{\delta}_{\mathbf{k}}$ is the estimated marginal slope change, and D^{cap} is the observed average duration at the kink.[33] Standard errors on the elasticities are obtained from bootstrapping using 50 replications. The implied elasticity of unemployment duration with respect to an overall change in the benefit level (both $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$) is $\varepsilon_{D,b} = 1.53$ (.13). The change in slope for spells starting between July 2001 and July 2002 is smaller but precisely estimated, and implies an elasticity of unemployment duration with respect to $b_{\mathbf{2}}$ of $\varepsilon_{D,b_{\mathbf{2}}} = .68$ (.13).

The same approach can be used to estimate the effect of benefits on the survival rate in unemployment at any spell duration. Table 2 shows the RK estimates for the effect of the benefit changes on the benefit durations D, $D_{\mathbf{1}}$ and $D_{\mathbf{2}}$, where $D_{\mathbf{1}} = \sum_{t<20wks} S_t$ is the time spent receiving benefit $b_{\mathbf{1}}$ and $D_{\mathbf{2}} = \sum_{t\geq 20wks} S_t$ is the time spent receiving benefit $b_{\mathbf{2}}$. For both $D_{\mathbf{1}}$ and $D_{\mathbf{2}}$,

[32] An alternative solution is to get rid of observations who have an ongoing spell at the moment the schedule changes. In Appendix Figure B-9 we provide evidence showing that the two techniques deliver identical results.

[33] To get this formula for the elasticity $\varepsilon_{D,b_{\mathbf{k}}} = \hat{\delta}_{\mathbf{k}} \cdot \frac{\bar{w}_{\mathbf{k}}}{D^{cap}}$, we simply use the fact that, at the kink, $b_k = .8.\bar{w}_{\mathbf{k}}$, and the fact that the deterministic change in slope in the benefit schedule at the kink is $\nu_{\mathbf{k}} = .8$

we find that the change in slope of the relation between wage and benefit duration is significant, but substantially smaller at the kink in $\boldsymbol{b_2}$ than at the kink in both $\boldsymbol{b_1}$ and $\boldsymbol{b_2}$.

We provide a comparison of these duration responses to estimates of labor supply responses to UI benefits available in the literature in Appendix Section B.6. For these comparisons to be meaningful, we focus on elasticities that use similar variations in benefits. When benchmarked against conceptually similar elasticities, our duration responses prove quite similar to moral hazard estimates in the literature, although probably in the high end of the range of existing estimates.

The most common threat to identification and inference in the RK design is the presence of non-linearity that underlies the relationship between the assignment variable and the outcome, but is unrelated to the effect of the kinked policy schedule. To deal with this threat, we use spells starting after July 2002, for which the UI schedule is linear at the threshold, as a placebo and reject the presence of non-linearity around the 725SEK threshold. In line with the evidence presented in Panel C.II of Figure 2, the estimated change in slope is very close to zero and not statistically significant. This precisely estimated zero effect alleviates the concern that our estimates are spuriously capturing some non-linear functional dependence between wages and unemployment duration around the 725SEK threshold.[34]

We provide several additional tests to assess the validity of the RK design and the robustness of the RK estimates in Appendix B. We start by testing for manipulation of the assignment variable around the kink point $\bar{w}$, which would constitute a clear violation of the second identifying assumption of the RK design. Such manipulation could arise due to selection into job loss, but also due to selection into UI, as UI is voluntary in Sweden. Appendix Figure B-4 displays the probability density function of wages and reports tests in the spirit of McCrary [2008] that confirm continuity of the pdf and of its first derivative around the 725SEK threshold. Appendix Table B-1 further indicates that the removal of the kinks did not significantly affect the distribution of daily wages above and below the kink, suggesting that the presence of kinks in the UI schedule does not significantly affect the distribution of daily wages around the kink. This rules out clear manipulation of the assignment variable in response to the kinked schedule.

We also find smoothness in the relationship between observable characteristics of unemployed workers and wages at the 725SEK threshold plotted in Appendix Figure B-5. This is reassuring, as non-smoothness in the distribution of observable heterogeneity would have cast doubt on the validity of the assumption of smoothness in the distribution of unobservable heterogeneity around the kink.

Appendix B finally provides additional tests for the presence of confounding non-linearities in the relationship between the assignment variable and the outcomes. The sensitivity of the RK estimates to the size of the bandwidth is explored in Appendix Figure B-6. The stability of the RK estimates across bandwidth sizes further alleviates the concern that the RK estimates pick

[34]In Appendix Figure B-3, we also plot the evolution of the estimates of the change in slope year by year from 1999 to 2007. All estimates for the placebo years 2002 to 2007 are close to zero and insignificant. The figure provides clear evidence that our estimated responses are indeed due to the policy changes, and not due to time trends in the distribution of durations around the kink.

up some underlying non-linearity in the relationship between wages and unemployment duration. In Appendix Figure B-10 we perform tests aimed at detecting non-parametrically the presence and location of a kink point in the relationship between unemployment duration and wages, as suggested in Landais [2015]. All these tests strongly support the conclusion that there is a change in slope that occurs right at the actual kink point in the UI schedule. We explore the sensitivity of our inference to alternative strategies in Appendix Table B-2. In particular, we report 95% confidence interval based on permutation tests as in Ganong and Jaeger [2014].[35] Interestingly, due to the linearity in the relationship between unemployment duration and wages across the whole support of the assignment variable, these confidence intervals are much tighter than those based on bootstrapped or robust standard errors. We finally explore in Appendix Table B-3 sensitivity of the RK estimates to variation in the order of the polynomial used to fit the data. The Table displays estimates of the change in slope at the kink for linear, quadratic and cubic specifications, and assesses the model fit for these different specifications. For spells starting between 1999 and July 2001, the estimates are very similar across polynomial orders. For spells starting between July 2001 and July 2002, estimates from the quadratic model are, although imprecisely estimated, somewhat larger in magnitude than estimates using a linear or cubic specification. Yet, the linear specification is having similar root mean squared errors (RMSE) and minimizes the Aikake information criterion (AIC) compared to the quadratic and cubic specification, suggesting that linear estimates should be preferred.

3.2 Implications for Moral Hazard Costs

Our results carry important implications for the moral hazard costs of modifying the time profile of UI benefits. Table 2 displays our elasticity estimates and the implied moral hazard costs with bootstrapped standard errors using 50 replications.

First, the moral hazard cost of the Swedish unemployment policy is large overall. For a flat profile, the moral hazard cost from increasing the benefit level throughout the unemployment spell (i.e., increasing both $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$) equals $MH_{\bar{b}} = \frac{\bar{b}+\tau}{\bar{b}}\varepsilon_{D,b} = 1.64$ (.14). This means that because of behavorial responses when increasing all benefits by 1%, the planner would need to levy from the employed 1.64 times more resources to balance its budget than the implied static cost absent behavioral responses. For these moral hazard computations, we use $\bar{b} = .72$, which corresponds to the observed average replacement rate in the sample, and we assume a tax $\tau = .05$, implying $\frac{\bar{b}+\tau}{\bar{b}} = 1.07$. The tax rate corresponds to the rate required to balance the government's UI budget for the average unemployment rate during the period 1999-2007 (*i.e.*, assuming no other expenditures $\bar{G}$ in equation (1)). Note that when $\bar{G} > 0$, the required tax level to balance the government's total budget will be higher. Moreover, when accounting for the fiscal externality through the rest

[35] Ganong and Jaeger [2014] propose to evaluate whether the true coefficient estimate is larger than those at "placebo" kinks placed away from the true kink. The idea behind their permutation test is that, if the counterfactual relationship between the assignment variable and the outcome (i.e., in the absence of the kink in the budget set) is non-linear, then the curvature in this relationship will result in many of the placebo estimates being large and statistically significant.

of the tax system, the implied moral hazard cost would be even higher. In Appendix Section A.2.4, we show that in the presence of an income tax τ^y, levied on both employed and unemployed, the additional fiscal externality equals $\tau^y \frac{w-\bar{b}}{\bar{b}} \varepsilon_{D,\bar{b}}$, which would increase the total moral hazard cost to $MH_{\bar{b}} = 1.76$ (resp. 1.70) for an effective income tax rate of $\tau^y = .20$ (resp. .10).[36]

Second, our results suggest that the moral hazard cost of an increase in benefits is lower when timed later in the spell. As discussed before, for a flat profile, the moral hazard cost $MH_{\mathbf{k}}$ simplifies to the expression in equation (6). The moral hazard cost of increasing benefits $b_{\mathbf{2}}$ after 20 weeks of unemployment is therefore $MH_{\mathbf{2}} = \frac{\bar{b}+\tau}{\bar{b}} \frac{D}{D_{\mathbf{2}}} \varepsilon_{D,b_{\mathbf{2}}} = 1.44$ (.28). Compared to $MH_{\bar{b}} = 1.64$ (.14), this implies that the incentive cost of increasing benefits after 20 weeks is smaller than the incentive cost from increasing benefits throughout the unemployment spell. A second implication is that the incentive cost is somewhat larger for increasing benefits in the first 20 weeks than after. In fact, we can use our estimates to back out the elasticity of unemployment duration with respect to a change in $b_{\mathbf{1}}$ only.[37] This then implies that $MH_{\mathbf{1}} = 1.82$ (.40).

Our estimates thus indicate that the point estimate of the moral hazard cost of increasing benefits for the first 20 weeks is 26 percent larger than that of increasing benefits after 20 weeks of unemployment. Formal z-tests of equality of $MH_{\mathbf{1}}$ and $MH_{\mathbf{2}}$ provided in Table 2 do not allow to reject that $MH_{\mathbf{1}}$ is equal to $MH_{\mathbf{2}}$. We also conduct inference using permutation tests instead of z-tests. We draw placebo kinks and obtain a placebo distribution of estimates of $MH_{\mathbf{1}}$ and $MH_{\mathbf{2}}$. The procedure is described in Appendix B.5.4, and we report in Table 2 the corresponding p-value (5.98%) from a test of equality of $MH_{\mathbf{1}}$ and $MH_{\mathbf{2}}$, which is much tighter than the p-value from our z-test.

To further assess the robustness of our findings regarding the relative magnitude of $MH_{\mathbf{1}}$ and $MH_{\mathbf{2}}$, we also exploit additional sources of variations in $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$ stemming from variation in the location of the kink in $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$ during the period 1999 to 2007. As explained in detail in Appendix Section B.5, these sources of variations provide two broad strategies to identify the relative moral hazard costs of $b_{\mathbf{1}}$ vs $b_{\mathbf{2}}$, corresponding to four different potential estimates of the ratio $MH_{\mathbf{1}}/MH_{\mathbf{2}}$, summarized in Appendix Table B-4. The first strategy consists in comparing estimates at the same "kink" over time. This approach has the advantage of comparing similar individuals over time at the same level of income. One drawback may be that behavioral elasticities are time varying due to business cycle fluctuations for instance. The second strategy consists in comparing estimates at different "kinks" within the same time period. This second approach has the advantage of comparing individuals within the same time period, therefore controlling for the fact that behavioral elasticities are time varying due to business cycle fluctuations for instance. A potential drawback of this approach is that individuals at different kinks may differ in their responsiveness to the policy. Results from Appendix Table B-4 show that estimates are very robust to the sources of variation used for identification, and suggest that the larger magnitude of $MH_{\mathbf{1}}$ relative to $MH_{\mathbf{2}}$ is a very robust finding, across all identification strategies. [38]

[36] In Sweden, UI benefits are fully included in individuals' income subject to the personal income tax.

[37] We simply use $\varepsilon_{D,b} = \varepsilon_{D,b_{\mathbf{1}}} \frac{b}{b_{\mathbf{1}}} + \varepsilon_{D,b_{\mathbf{2}}} \frac{b}{b_{\mathbf{2}}} = \varepsilon_{D,b_{\mathbf{1}}} + \varepsilon_{D,b_{\mathbf{2}}}$ for $b_{\mathbf{1}} = b_{\mathbf{2}}$.

[38] In Appendix Figure B-13, we also provide inference using the permutation-based approach (taking random draws

4 Consumption Smoothing Over the Unemployment Spell

In this section we provide empirical evidence on the evolution of consumption at job loss and throughout the unemployment spell. We then discuss how this evidence relates to the evolution of consumption smoothing gains of UI.

4.1 Registry-based Measure of Consumption: Estimating Consumption Drops

Annual Consumption Drop We start by providing evidence that annual household consumption drops significantly at unemployment, following an approach similar to Gruber [1997]. We implement an event study approach using the panel dimension of the data. We define event year $n = 1$ as the year an unemployment spell starts. Because our annual measure of consumption is observed as of December, we focus on individuals who are observed unemployed in December of year $n = 1$, (and who were employed in December of year $n = 0$). We estimate the following event study specification:

$$\dot{\mathscr{C}}_{i,n} = \sum_j \gamma_j \cdot \mathbb{1}[n = j] + \sum_l \eta_l \cdot \mathbb{1}[\text{Calendar Year}_{i,n} = l] + \mu_{i,n} \tag{17}$$

where $\dot{\mathscr{C}}_{i,n} = \ln \mathscr{C}_n - \ln \mathscr{C}_{n-1}$, is the change in the log annual consumption of individual i between event years $n-1$ and n. Specification (17) only controls for a set of calendar year fixed effects η_l. In Figure 5, we report estimates of the event year dummies γ_j, where the omitted category is year $n = 0$. Our estimates provide compelling evidence of a sharp consumption drop at unemployment. First, the results show that annual consumption growth is remarkably similar in the five years prior to the start of a spell, which suggest that consumption profiles show limited signs of anticipation effects prior to unemployment. Second, annual consumption drops sharply and significantly in year $n = 1$, when individuals start an unemployment spell. Consumption growth is still negative in year $n = 2$, reflecting the fact that some individuals may still be unemployed through part, or all, of year $n = 2$. Consumption growth then picks up slightly as individuals find new jobs. Yet, as consumption *growth* remains similar to pre-unemployment levels after year $n = 3$, this implies that consumption *levels* remain significantly lower than their pre-unemployment levels, even five years after the start of a spell. This evidence of persistent consumption effects of unemployment is in line with the abundant literature documenting the long lasting earnings consequences of job displacement (e.g., Couch and Placzek [2010]).

Consumption Drop over the Spell To evaluate the consumption smoothing benefits of benefits given in different parts of the UI profile, we need to recover higher frequency consumption measures. Here, we wish to retrieve weekly consumption flows in order to estimate $\Delta C_{\mathbf{1}} = \bar{c}_{\mathbf{1}}^u - \bar{c}_0$,

of placebo kinks) for all four estimates, which confirms the robustness of our conclusion.

the average consumption drop in the first part of the profile (first 20 weeks of an unemployment spell) and $\Delta C_{\mathbf{2}} = \bar{c}^u_{\mathbf{2}} - \bar{c}_0$, the average consumption drop in the second part of the profile (after 20 weeks of unemployment). Our data enables us to estimate non-parametrically $\Delta C_{\mathbf{1}}$ and $\Delta C_{\mathbf{2}}$ with minimal and easily testable assumptions. We leverage two things. First, annual consumption is the sum of higher frequency expenditure flows during the year. Second, at the moment we measure annual consumption (December of each year), individuals differ in the time they have been unemployed, which provides variation to identify how consumption evolves over the unemployment spell. Based on these two simple observations, and defining event time t as the number of weeks since the start of an unemployment spell, the annual household consumption of an individual who is, in December, observed in the t-th week of her unemployment spell is: $\mathscr{C}_t = \sum_{j=t-51}^{t} c_j$ where c_j is the weekly consumption flow in event week j. We also define $\Delta\mathscr{C}_t = \mathscr{C}_t - \mathscr{C}_0$, the drop in annual consumption of an individual observed in her t-th week of a spell, compared to pre-unemployment annual consumption $\mathscr{C}_0$.[39] Under the assumption that consumption profiles are constant prior to unemployment, i.e. $c_j = \bar{c}_0, \forall j < 0$, we have $\Delta\mathscr{C}_t = \sum_{j=t-51, j>0}^{t} (c^u_j - \bar{c}_0)$. The above expression shows that the comparison of drops in annual consumption of individuals observed at different weeks t of their unemployment spell directly reveals information about the drops in flow consumption $c^u_j - \bar{c}_0$ throughout the unemployment spell. Compare for instance the drop in annual consumption of individuals observed in their first week of unemployment $\Delta\mathscr{C}_1 = c^u_1 - \bar{c}_0$ and the drop in annual consumption of individuals observed in their second week of unemployment $\Delta\mathscr{C}_2 = c^u_2 - \bar{c}_0 + c^u_1 - \bar{c}_0$. The difference $\Delta\mathscr{C}_2 - \Delta\mathscr{C}_1 = c^u_2 - \bar{c}_0$ reveals the drop in flow consumption in the second week of unemployment.[40]

It follows that we can easily estimate non-parametrically our two statistics of interest $\Delta C_{\mathbf{1}}$ and $\Delta C_{\mathbf{2}}$ from the observed drops in annual consumption of individuals observed at different points in their unemployment spell. In practice, we have:

$$\Delta C_{\mathbf{1}} = \sum_{j=1}^{20} \frac{S_j}{D_{\mathbf{1}}} (\Delta\mathscr{C}_j - \Delta\mathscr{C}_{j-1}) \tag{18}$$

$$\Delta C_{\mathbf{2}} = \sum_{j=21}^{104} \frac{S_j}{D_{\mathbf{2}}} \big((\Delta\mathscr{C}_j - \Delta\mathscr{C}_{j-1}) + \mathbb{1}[j > 52] \cdot (\Delta\mathscr{C}_{j-52} - \Delta\mathscr{C}_{j-53})\big) \tag{19}$$

where S_j is the survival rate after j weeks of unemployment, and $D_{\mathbf{1}}$ and $D_{\mathbf{2}}$ are the average time

[39] Pre-unemployment annual consumption $\mathscr{C}_0$ is measured as the annual consumption in event year $n = 0$, i.e., the last pre-unemployment year during which the individual is observed being employed. Formally, $\mathscr{C}_0 = \mathscr{C}_{t-51}$ for individuals who are in their first year of unemployment as of December ($t < 52$), and $\mathscr{C}_0 = \mathscr{C}_{t-103}$ for individuals who are in their second year of unemployment as of December ($t \geq 52$).

[40] More generally, for individuals in their first year of unemployment, ($t < 52$), the difference in annual consumption drops between t and $t-1$, $\Delta\mathscr{C}_t - \Delta\mathscr{C}_{t-1} = c^u_t - \bar{c}_0$ directly reveals the flow drop in consumption in week t. For individuals in their second year of unemployment, ($t \geq 52$), the difference in annual consumption drops between t and $t-1$, $\Delta\mathscr{C}_t - \Delta\mathscr{C}_{t-1} = \sum_{j=t-51}^{t} (c^u_j - \bar{c}_0) - \sum_{j=t-52}^{t-1} (c^u_j - \bar{c}_0) = (c^u_t - \bar{c}_0) - (c^u_{t-52} - \bar{c}_0)$ identifies the sum of the flow drop in consumption in week t and the flow drop in consumption in week $t-52$. The drop in flow consumption $c^u_t - \bar{c}_0$ for $t > 52$ is therefore identified by $\big(\Delta\mathscr{C}_t - \Delta\mathscr{C}_{t-1}\big) + \big(\Delta\mathscr{C}_{t-52} - \Delta\mathscr{C}_{t-53}\big)$.

spent in the first and the second part of the benefit schedule.[41]

To implement this strategy, we start by estimating, on the sample of individuals who are unemployed in December of any year between 2000 and 2007, the average annual consumption drops as a function of the time t that individuals have spent unemployed as of December, using the following specification

$$\Delta\mathscr{C}_{i,t} = \sum_j \beta_j \cdot \mathbb{1}[t = j] + \sum_l \eta_l \cdot \mathbb{1}[\text{Calendar Year}_{i,t} = l] + \varepsilon_{i,t} \tag{20}$$

We display in Figure 6 panel A our baseline estimates of the drop in annual consumption where in practice, we group individuals in 10 weeks intervals of unemployment duration for estimation.[42] Panel A reports $\hat{\beta}_t/\mathscr{C}_0 = \widehat{\Delta\mathscr{C}_t}/\mathscr{C}_0$, i.e., we scale our estimates by the average annual consumption in the last year prior to unemployment, to express consumption drops relative to pre-unemployment levels. The Figure shows that the drop in annual consumption increases steadily with time spent unemployed as of December for individuals observed throughout the first year of unemployment, and then stabilizes around 10% for individuals observed in their second year of unemployment. The decrease in the first year of unemployment is partly mechanical as individuals who are observed at a later time t in the spell have spent more time of the year unemployed and thus experienced a larger drop in annual consumption. Still, the shape of the slope of these estimates reveals all the necessary information about the evolution of flow consumption throughout the spell.[43]

Panel A of Figure 6 also reports our non-parametric estimates of the average drops in consumption for both parts of the benefit profile $\widehat{\Delta C_{\mathbf{1}}} = \sum_{j=1}^{20} \frac{\hat{S}_j}{D_{\mathbf{1}}}(\hat{\beta}_j - \hat{\beta}_{j-1})$ and $\widehat{\Delta C_{\mathbf{2}}} = \sum_{j=21}^{104} \frac{\hat{S}_j}{D_{\mathbf{2}}}\big((\hat{\beta}_j - \hat{\beta}_{j-1}) + \mathbb{1}[j > 52] \cdot (\hat{\beta}_{j-51} - \hat{\beta}_{j-52})\big)$. Standard errors are computed using the Delta-method. Our results suggest that the average drop in (flow) household consumption for individuals in the first part of the profile is 4.4%, while it is 9.1% on average for individuals in the second part of the profile. We also report the p-value from a test of equality of $\widehat{\Delta C_1}$ and $\widehat{\Delta C_{\mathbf{2}}}$ which strongly rejects that consumption drops are the same in the first and second part of the benefit profile. Our evidence thus indicates that consumption expenditures do significantly decline over the unemployment spell.[44]

Robustness Our implementation so far has rested on two assumptions, which we will investigate next. First, for simplicity we assumed that $c_j = \bar{c}_0, \forall j < 0$, or in other words, that

[41] Note that we censor durations at two years to deal with a few observations without any end date. Hence, our summation in $\Delta C_{\mathbf{2}}$ only goes up to week 104.

[42] The left hand side variable in specification (20) is a change in levels (and not a log change in consumption). The reason is that we need the change in levels (and not the log changes) to retrieve higher frequency consumption from the annual changes.

[43] Intuitively, a linear slope in the first year would imply that $c^u_j - c_0 = c^u_k - c_0, \forall k, j < 52$, which means that flow consumption drops right at job loss and remains constant after. Concavity in the estimates for the first year would indicate that $c^u_j - c_0 < c^u_k - c_0, \forall k < j < 52$, which means that flow consumption continues to decline throughout the unemployment spell.

[44] Our approach allows to estimate non-parametrically how consumption evolves at, potentially, *any* frequency over the unemployment spell. We report for instance in Appendix Figure C-6 the estimated flow consumption drops over the first year of unemployment at a 10 week frequency.

the consumption profile is flat prior to the start of an unemployment spell. To investigate how consumption profiles evolve prior to the unemployment spell we report in Figure C-1 the annual log household consumption changes $\mathscr{C}_{i,t} - \mathscr{C}_{i,t-52}$ as a function of time t since the onset of a spell (in 5 weeks bins), relative to the last 5 weeks prior to the onset of a spell. We go as far back as 4 years prior to the onset of a spell, to detect longer anticipation effects. The Figure shows two interesting patterns. First, the consumption patterns of households are extremely stable in the 4 years prior to displacement, suggesting no long term anticipation behaviors. Second, there are no sharp consumption changes immediately preceding displacement, suggesting no significant short term anticipation behaviors. Overall, this evidence strongly supports the assumption that flow consumption profiles are flat prior to the onset of an unemployment spell.

The second assumption used in our baseline implementation is that average flow consumption profiles are similar for individuals observed at different time t in their spell. In practice, the sample of individuals who survive until t (i.e. individuals with a completed spell length $L \geq t$) may differ from the sample of individuals who survive until $t-1$ (i.e., individuals with a completed spell length $L \geq t-1$), and this may bias our estimates if individuals who select into different spell lengths L differ in their underlying consumption profiles (dynamic selection on consumption profiles). Note, however, that we can easily relax the assumption that average consumption profiles are independent of actual spell length. To ensure that we compare the same sample of individuals at different time t, we estimate separate flow profiles by completed spell length L using the following specification:

$$\Delta\mathscr{C}_{i,t} = \sum_k \sum_j \beta_{j,k} \cdot \mathbb{1}[t=j] \cdot \mathbb{1}[L=k] + \sum_l \eta_l \cdot \mathbb{1}[\text{Calendar Year}_{i,t} = l] + \varepsilon_{i,t}$$

where we break down spell length into 4 brackets: 1 to 20 weeks, 21 to 40 weeks, 41 to 60 weeks, and longer than 61 weeks. In Panel B of Figure 6 we report our estimates of $\hat{\beta}_{t,L}$, representing the drops in annual consumption as a function of time spent unemployed for the 4 categories of completed spell lengths. Results show that the profile of drops in annual consumption is very similar across spell length, pointing to a limited role for dynamic selection on consumption profiles. We also report the implied estimates of $\Delta C_{\mathbf{1}}$ and $\Delta C_{\mathbf{2}}$, allowing for heterogeneity in profiles across spell length, $\widehat{\Delta C_{\mathbf{1}}} = \sum_k \sum_{j=1}^{20} \frac{\hat{S}_{j,k}}{D_{\mathbf{1}}}(\hat{\beta}_{j,k} - \hat{\beta}_{j-1,k})$ and $\widehat{\Delta C_{\mathbf{2}}} = \sum_k \sum_{j=21}^{104} \frac{\hat{S}_{j,k}}{D_{\mathbf{2}}}\big((\hat{\beta}_{j,k} - \hat{\beta}_{j-1,k}) + \mathbb{1}[j > 52] \cdot (\hat{\beta}_{j-52,k} - \hat{\beta}_{j-53,k})\big)$. Results are very similar to our baseline estimates, and confirm that flow consumption significantly decreases over the unemployment spell. The average drop in (flow) household consumption for individuals in the first part of the profile is 4.8%, while it is 9.6% on average for individuals in the second part of the profile.

Additional Survey Evidence We further investigate the robustness of our results using the household consumption surveys (HUT) in Appendix C.4. While the HUT sample is much smaller, and does not have a panel structure, it directly measures flow consumption at the time of the survey interview. Results from our preferred specification of Table C-3 column (4) suggest that the drop in the first part of the profile $\widehat{\Delta C_{\mathbf{1}}}$ is 4.6% and the drop in the second part of the profile $\widehat{\Delta C_{\mathbf{2}}}$

is 10.8%. Reassuringly, despite much larger standard errors, these point estimates are remarkably similar to our baseline estimates using the registry-based measure of consumption.

4.2 Implications for Consumption Smoothing Gains

The empirical evidence presented above indicates that consumption declines significantly over the unemployment spell. This decline in consumption suggests that the marginal value of unemployment benefits increases over the spell, evaluated for the flat policy in place in Sweden.

An important outstanding issue is that our empirical analysis considers changes in expenditures. The question is to what extent expenditures translate into consumption and thus capture the potential welfare value of unemployment benefits.[45] In particular, unemployed workers may try to re-allocate certain categories of expenditures to smooth the shocks in their consumption. A first example is the substitution towards home production, which has been analyzed extensively in the context of retirement (e.g., Aguiar and Hurst [2005]). A second example is the substitution away from expenditures on durable goods that provide a consumption flow for future periods as well.

The HUT offers insights into the type of consumption goods that households adjust over the spell. In Table C-4, we investigate how various categories of expenditures evolve over the unemployment spell. Consumption of non-durable, uncommitted goods, such as food, recreation, transportation, or restaurants (columns (2),(6),(7) and (8)), drops significantly early on in the spell and further decreases over the spell.[46] More committed expenditures like housing rents paid by renters (column (4)) do not seem to decline significantly, neither early, nor later in the unemployment spell.[47] Interestingly, we find a larger drop in restaurant expenditures than in food expenditures, consistent with substitution towards home production. Substitution towards home production may affect the level of the consumption smoothing gains of UI benefits.[48] However, its effect on the evolution of the consumption smoothing gains of UI over the spell is ambiguous, and will depend on the relative availability of substitution towards home production for households that select into long versus short spells.[49]

Table C-4 also indicates that expenditures on durable goods such as the purchase of new vehicles

[45] See Campos and Reggio [2016].

[46] Gruber [1997] considers only food consumption using U.S. data from the PSID for the period 1968-1987 and finds an average drop of 6.8% in the first year of unemployment, which is very similar to our estimates.

[47] Fixed commitments reduce the ability to smooth consumption and can increase the relevant value of γ to translate consumption drops into welfare (Chetty and Szeidl [2007]).

[48] The sign of the impact of such substitution patterns on the consumption smoothing gains of UI benefits is nevertheless ambiguous. The availability of consumption insurance through home production means that the drop in consumption will be smaller than the actual drop in expenditures, which decreases the marginal utility of unemployment benefits. At the same time, complementarity between expenditures and household production increases the marginal utility of benefits. The relative magnitude of these two effects will determine the sign of the impact on the consumption smoothing gains of UI.

[49] Note that results displayed in Appendix Tables C-5 and C-6 indicate that the expenditure profiles over food, recreation, transportation, or restaurants are not significantly different for households that select into long versus short spells. This suggests that there is no significant dynamic selection over the spell based on the availability of substitution towards home production.

or the purchase of furniture and home appliances (columns (4) and (5)) decline strongly early during the spell, but increase later during the spell, yet remaining largely below their pre-unemployment level. These results suggest that the unemployed can initially smooth the marginal utility of consumption services by shifting spending away from durables, but they lose the capacity to do so after some time.[50] This in turn will tend to further increase the inclining profile of the consumption smoothing value of UI benefits over the unemployment spell.

To sum up, empirical evidence seems to substantiate that, in the Swedish context, and for a flat benefit profile, consumption declines over the spell, and that the consumption smoothing gains of UI benefits are larger after 20 weeks of unemployment than during the first 20 weeks of unemployment. This conclusion seems robust to the presence of dynamic substitution across categories of expenditures.

5 Welfare Analysis

This concluding section brings our theoretical and empirical analysis together to provide an evidence-based assessment of the UI benefit profile in Sweden. We use our empirical estimates to implement our sufficient-statistics formulae and to shed light on forces underlying how the sufficient statistics evolve over the unemployment spell. While our formulae allow for a transparent evaluation of local changes in the policy profile, we also illustrate how a calibrated structural model allows us to go beyond the local recommendations.

5.1 Sufficient-Statistics Implementation

The welfare consequences of a marginal increase in UI benefits are reported in Table 3. The first row examines the consequences of a marginal increase in the benefit level $\bar{b}$ throughout the unemployment spell; the second row examines the consequences of a marginal increase in the benefit $b_{\mathbf{1}}$ during the first 20 weeks of unemployment and the third row examines the consequences of a marginal increase in the benefit $b_{\mathbf{2}}$ after 20 weeks of unemployment. Different columns show the different components of the welfare impact.

The first column repeats our estimates of the moral hazard costs (Panel IV of Table 2). As noted before, the estimated moral hazard cost is high overall (all estimates exceed 1), but the cost is 26 percent higher for benefits paid during the first 20 weeks of unemployment compared to those paid after 20 weeks of unemployment. The second column repeats our estimates of the drops in average consumption during the first 20 weeks of unemployment and after 20 weeks of unemployment, but also averaged over the full unemployment spell. We convert the respective consumption drops into estimates of the consumption smoothing gains $CS_{\mathbf{k}}$, following the implementation in (14), which relies on a Taylor approximation and homogeneous preferences. As the appropriate value of risk

[50]Note again that, although results are relatively imprecise due to the small sample size, Appendix Table C-5 indicates that dynamic selection based on durable expenditure profiles is not significant. The availability of consumption smoothing through shifting expenditures away from durables is not significantly different for households that select into long versus short spells.

aversion in this context is uncertain, the consumption smoothing gains are reported for a range of plausible values (see Chetty [2009], Chetty and Finkelstein [2013]) and range between .05 and .5. Regardless of the level of risk aversion, when risk preferences are homogenous, the estimated consumption smoothing gains are twice as high for benefits paid later in the spell.

Putting the estimates of the CS gains and the MH costs together indicates that the MH costs are substantially larger than the CS gains, even for high risk aversion.[51] This is true both for benefits paid to short-term and long-term unemployed and thus suggests that UI benefits are too generous overall. Indeed, expressing the welfare effects in terms of "benefit-cost" ratio's $CS_{\mathbf{k}}/MH_{\mathbf{k}}$, we find that the return to the marginal krona spent on the unemployed is substantially lower than 1. These estimates are, however, sensitive to our implementation assumptions including the risk preference parameter γ, the tax distortion τ and our normalization relative to the value of krona spent pre-unemployment.[52]

Comparing the benefit-costs ratio's for different parts of the policy allows us to formulate recommendations on the tilt of the benefits profile $b_{\mathbf{1}}/b_{\mathbf{2}}$.[53] We find that the relative moral hazard cost $MH_{\mathbf{1}}/MH_{\mathbf{2}}$ (= 1.26) is more than twice as high as the relative consumption smoothing gain $CS_{\mathbf{1}}/CS_{\mathbf{2}}$ (= .50). The implied return to the marginal krona spent on the short-term unemployed is thus at least twice as low compared to the long-term unemployed, indicating that welfare could be increased by introducing an inclining tilt ($b_{\mathbf{1}} < b_{\mathbf{2}}$) in the flat Swedish benefit profile. Importantly, this evaluation of the tilt is less sensitive to the above assumptions. In particular, $CS_{\mathbf{1}}/CS_{\mathbf{2}}$ does not depend on the preference parameter γ, while $MH_{\mathbf{1}}/MH_{\mathbf{2}}$ does not depend on the tax distortion τ. More generally, our recommendations on the tilt of the benefit profile are robust to implementation errors that are uncorrelated with the unemployment duration.

In sum, our implementation suggests that on average UI benefits should be decreased and especially so for the short-term unemployed. The 2001 reform in Sweden did the exact opposite by increasing the benefit cap for the first 20 weeks of unemployment.

5.2 From Sufficient Statistics to Mechanisms

While the previous local welfare implementation does not need to identify the forces that shape the sufficient statistics, it is in practice interesting to explore and analyze these forces, for at least two reasons. First, from a descriptive point of view, it is interesting to understand what mechanisms can explain the observed patterns of these statistics. Second, from a welfare perspective, in order

[51] We note again that, although our duration elasticities and consumption drops are estimated from different samples, individuals in the two samples are almost identical in terms of observable characteristics and unemployment durations as discussed in Section 2.2.

[52] Note that consumption drops are experienced at the household level. Our analysis only accounts for the consumption smoothing gains for the unemployed individual, hereby underestimating the overall consumption smoothing gain of UI.

[53] Expressing the welfare effects in terms of "cost/benefit"ratio's (see Hendren [2013]) is a common approach to formulate policy recommendations. Whenever $CS_{\mathbf{1}}/MH_{\mathbf{1}} > CS_{\mathbf{2}}/MH_{\mathbf{2}}$, welfare can be increased by increasing the tilt $b_{\mathbf{1}}/b_{\mathbf{2}}$. Moreover, a budget-balanced increase in the tilt $b_{\mathbf{1}}/b_{\mathbf{2}}$ increases welfare if and only if $\frac{1+CS_{\mathbf{1}}}{1+MH_{\mathbf{1}}} > \frac{1+CS_{\mathbf{2}}}{1+MH_{\mathbf{2}}}$, where the ratio's correspond the so-called "marginal value of public funds". We show this in Corollary 1 in Appendix Section A.2.

to extend our recommendations to non-local variations, one needs to impose more structure on the data. Investigating what mechanisms drive the estimated sufficient statistics is a useful way to inform the choice of structural assumptions to impose on the underlying model used to simulate the welfare consequences of non-local variations, as we do in subsection 5.3.

In order to shed light on mechanisms, we can leverage the fact that different assumptions regarding the underlying structure of the model map into different dynamic behaviors of our sufficient statistics.

5.2.1 Dynamic Profile of Sufficient Statistics

Our sufficient-statistics characterization indicates that, everything else equal, unemployment benefits should decline over the spell when the consumption smoothing gains are lower for benefits paid later in the spell or when their moral hazard cost is higher. The optimal benefit profile has been the topic of a series of seminal papers, mostly studying single-agent, stationary environments (see Shavell and Weiss [1979], Hopenhayn and Nicolini [1997] and Shimer and Werning [2008]). The relevant forces in a stationary setting and how they affect the optimal benefit profile are well understood from this literature. We can re-express the impact of these forces on the unemployment policy through their impact on the gradient of the moral hazard costs and consumption smoothing gains:[54]

Proposition 2. *Consider a flat benefit profile ($b_t = \bar{b} < w - \tau$ for $\forall t$) in a single-type, stationary environment ($h_{i,t}(\cdot) = \bar{h}(\cdot)$ for $\forall i, t$) with $\beta(1+r) = 1$, $T = \infty$ and assuming differentiability:*

(i) when agents are borrowing-constrained throughout the unemployment spell ($a_{i,t} = 0$ for $\forall i, t$),

$$MH_t \leq MH_{t'} \text{ and } CS_t = CS_{t'} \text{ for } t < t'.$$

(ii) when agents are not borrowing-constrained, but preferences are separable ($\partial v_i^u(c,s)/\partial c = \partial v_i^e(c)/\partial c = v'(c)$),

$$CS_t < CS_{t'} \text{ for } t < t'.$$

The force underlying the increasing moral hazard cost is the forward-looking behavior of job seekers. Increasing unemployment benefits later in the spell discourages forward-looking job seekers already early in the spell and is therefore always more costly than increasing benefits earlier in the spell. This causes the optimal benefit profile to be declining, as shown before in Shavell and Weiss [1979] and Hopenhayn and Nicolini [1997]. The force underlying the increasing consumption smoothing gain is that, when available, unemployed job seekers use liquid assets to smooth their marginal utility of consumption. As assets are depleted while unemployed, the marginal utility of consumption and thus the value of unemployment benefits is always higher later in the spell. This

[54] Appendix Section A.3 provides the proof (closely following **?** and Shimer and Werning [2008]) and gauges the robustness of Proposition 2 within a stationary environment.

causes the optimal benefit profile to be inclining, as shown before in Shavell and Weiss [1979] and Shimer and Werning [2008].[55]

In practice, however, unemployment dynamics are non-stationary. Non-stationary forces, and dynamic selection in particular, can mitigate the forces underlying Proposition 2 and, when strong enough, actually reverse the gradient of the sufficient statistics. We discuss this below in relation to our empirical estimates, but also develop this argument formally in Appendix Section A.4.[56] While characterizing the impact of non-stationarities on the optimal policy is challenging and previous work has mostly relied on calibrated models (e.g., Shimer and Werning [2006] and Pavoni [2009]), our framework provides an alternative approach, which is to analyze how non-stationary forces affect the gradient of the moral hazard costs and consumption smoothing gains respectively.

In the end, how both the consumption smoothing gains and moral hazard costs evolve over the unemployment spell remains an empirical question. Yet, because different assumptions regarding the underlying structure of dynamic models of unemployment map into different dynamic behaviors of our sufficient statistics, we can use our estimated sufficient statistics to shed interesting light on the relative importance of different mechanisms in shaping consumption smoothing and moral hazard over the spell.

5.2.2 Moral Hazard Cost

Our empirical analysis indicates that the moral hazard costs increase over the unemployment spell in Sweden, suggesting the importance of non-stationary forces dominating the forward-looking incentives. To illustrate how these forces shape the dynamics of moral hazard costs over the spell, we can use the earlier expression (11) to re-write the moral hazard cost of an increase in benefit level $b_{\mathbf{2}}$ after 20 weeks of unemployment, as the sum of two parts,

$$MH_{\mathbf{2}} = \frac{\bar{b}+\tau}{\bar{b}} \times \varepsilon_{\tilde{D}_B, b_{\mathbf{2}}} + \frac{\bar{b}+\tau}{\bar{b}} \times [\frac{D_{\mathbf{1}}}{D_{\mathbf{2}}}\varepsilon_{D_{\mathbf{1}}, b_{\mathbf{2}}} + \varepsilon_{S_B, b_{\mathbf{2}}}].$$

The first part captures the response in the *remaining duration* of unemployment, conditional on still being unemployed at time $B = 20$ weeks ($\tilde{D}_B \equiv \Sigma_{s=0}^{T-B} S_{B+s}/S_B$). In a stationary, single-agent environment, this first part is the same for any time B and equal to the moral hazard cost of an overall increase in the benefit level $MH_{\bar{b}}$. The second part captures the forward-looking incentives by which an increase in benefits in the second part of the policy reduces the exit out of unemployment in the first part of the spell ($D_{\mathbf{1}}$) and increases the survival rate into the second

[55] Due to the opposing forces coming from moral hazard and consumption smoothing, even in a stationary environment, deriving the optimal timing of benefits is difficult. Werning [2002] and Shimer and Werning [2008] analyze these two opposing forces in models with both search and savings being endogenous and show that they exactly cancel out in case of CARA preferences; a flat benefit profile is optimal conditional on the unemployed having access to liquidity.

[56] Appendix Section A.4 starts from a specific stationary, search environment with borrowing-constrained agents (subsection A.4.2). We then consider depreciation in the returns to search over the unemployment spell (subsection A.4.3) and unobservable heterogeneity in the returns to search (subsection A.4.4). We finally allow for unobservable savings and consider heterogeneity in assets or preferences. We show how the extensions can mitigate and/or reverse the difference in moral hazard costs or consumption smoothing gains at different times during unemployment spell.

part of the spell (S_B). This second part converges to zero as B goes to 1. Following the argument in Shavell and Weiss [1979], the forward-looking incentives drive the result that in a stationary environment the moral hazard cost of increasing unemployment benefits is higher when timed later in the spell (see Proposition 2). We find the exact opposite in the Swedish context, which necessitates the presence of significant non-stationary forces that reduce the responsiveness of job seekers to changes in the policy later in the spell.

Evidence of Forward-looking Behavior We first note that our results unambiguously show that unemployed individuals are indeed forward-looking. We already reported the estimated elasticity of $D_{\mathbf{1}}$, the time spent on the first part of the profile, with respect to benefits $b_{\mathbf{2}}$ received in the second part of the profile is positive and significant ($\varepsilon_{D_{\mathbf{1}},b_{\mathbf{2}}} = .60$ (.11), see Table 2). We can also study the effect of UI benefits on the hazard rates out of unemployment. Figure B-1 in Appendix Section B.1 reports the corresponding RKD estimates and clearly shows that benefits $b_{\mathbf{2}}$ received after 20 weeks do have a negative effect on the hazard rate in the first 20 weeks. Unemployed individuals are thus not fully myopic: they react early in the spell to variation in benefits paid later in the spell.

Evidence on Non-stationary Forces To highlight the non-stationary forces underlying the reversed pattern of moral hazard costs in the Swedish context, we compare at different spell lengths t, the elasticity of the remaining duration $\tilde{D}_t \equiv \Sigma_{s=0}^{T-t} S_{t+s}/S_t$ with respect to the overall benefit level $\bar{b}$. In a stationary, single-agent environment, these elasticities would remain constant and be equal to $\varepsilon_{D,\bar{b}}$.[57] Instead, our estimates, reported in Figure 4, show that the elasticity of the remaining duration strongly declines as a function of t. Five months into the spell the elasticity is only one third of the elasticity at the onset of the spell. This evidence of strong non-stationarities is corroborated in Appendix Figure B-1 which shows that the effect of UI benefits $\bar{b}$ on hazard rates is mostly concentrated in the first two to three months. After three months, the effect of UI benefits on the hazard rate is small and almost always insignificant.

Both depreciation and heterogeneity in the returns to search can explain why the responsiveness of exit rates decreases substantially over the unemployment spell. The reduced responsiveness to changes in the benefit profile, either through depreciation or through selection, makes it less costly to increase benefits later in the spell. While depreciation by itself cannot, selection - when strong enough - can in fact reverse the decline of moral hazard costs over the spell, as shown in Appendix Section A.4. Separating true depreciation from selection on heterogeneity in returns to search is notoriously difficult and beyond the scope of this paper, but our analysis nevertheless suggests the importance of embedding such non-stationary features in dynamic models of unemployment.

[57] Because exit rates only depend on the continuation policy in a stationary environment and there are no selection effects in a single-agent environment in response to policy variation earlier in the spell, the elasticity of the remaining duration $\tilde{D}_t$ with respect to an overall change in the benefit level is independent of t.

5.2.3 Consumption Responses

We now turn to the dynamic profile of the consumption smoothing gains. We find that the average consumption drop is more pronounced later in the spell. We therefore revisit mechanisms that can explain this pattern and whether they necessarily translate in consumption smoothing gains that increase over the spell.

Assets and Liquidity Constraints As stated above, a single-agent model of intertemporal consumption predicts that the marginal utility of consumption is weakly increasing during the unemployment spell. Intuitively, long-term unemployment implies a larger shock in resources, so we expect the unemployed to run down their assets and consume less the longer they are unemployed. Or they become liquidity-constrained and start consuming "hand-to-mouth".

We find clear evidence of the role of assets and liquidity constraints in determining the consumption smoothing over the spell. First, following the same methodology as in Figure 6, we report the average annual consumption drops by time spent unemployed as of December, but now breaking down the sample by the level of net wealth of the household at the start of the spell. As shown in Appendix Figure C-2, individuals with higher net wealth experience a lower drop in consumption. Second, we provide direct evidence that individuals do use their liquid assets to smooth consumption over the spell. Appendix Figure C-3 displays the relative change in liquid bank asset holdings by time spent unemployed. The Figure shows that individuals use their liquid assets to smooth consumption, especially earlier on in the spell, by depleting their bank accounts or reducing their savings. Finally, evidence from registry data also indicates that debt does not offer much help in smoothing consumption over the unemployment spell, hinting at the presence of liquidity constraints. In Appendix Figure C-4, we provide evidence of a reduction in the use of non mortgage-related credit over the unemployment spell. This suggests that as the duration of the spell increases, access to credit becomes harder and consumption out of debt falls significantly.

Overall, this analysis confirms that assets and debt may offer means to smooth consumption over short spells. But as the unemployment spell continues, these means get quickly exhausted. Households in the second part of the profile seem therefore closer to being hand-to-mouth.

Spousal Labor Supply Within a household, the labor supply of other members of the household may help reduce the drop in household consumption over the spell. In Figure C-5 we investigate how the labor supply of other members of the household affects the drop in household consumption over the spell. Following the same methodology as in Figure 6, we report the average change in total earnings and in total disposable income of all other members of the household as a function of time spent unemployed, scaled by the annual household consumption level prior to unemployment. Results show that within-household changes in the earnings and disposable income of all other members of the household are small and almost never significantly different from zero throughout the unemployment spell. This suggests that in our context, labor supply responses of other household members is not playing a significant role in increasing household consumption in response to an unemployment shock, even for long-term unemployed.

Dynamic Selection In practice, agents are heterogeneous and selection into longer unemployment spells may affect consumption responses over the spell and the gradient of consumption smoothing gains. As shown in Appendix Section A.4.5), if those selecting into longer unemployment spells suffer less from a drop in consumption, the positive gradient of consumption smoothing benefits due to the falling consumption is reduced and can potentially be reversed.

In theory, as long as preferences are homogeneous, we can use the approximation in condition (14) to translate the average consumption wedge into the consumption smoothing gains of increasing benefits. When preferences are heterogeneous, the consumption smoothing gains are approximated (using Taylor expansions of $v'\left(c_{i,t}^u\right)$ around $c_{i,0}$ for each individual) by

$$CS_t \cong \frac{E_t^u\left[v_i'\left(c_{i,0}\right)\right]-E_0\left[v_i'\left(c_{i,0}\right)\right]}{E_0\left[v_i'\left(c_{i,0}\right)\right]}-\frac{E_t\left[v_i''\left(c_{i,0}\right)\left(c_{i,0}-c_{i,t}^u\right)\right]}{E_0\left[v_i'\left(c_{i,0}\right)\right]}, \tag{21}$$

where E_t^u takes the weighted average among the unemployed at time t (with weights $S_{i,t}/S_t$) and E_0 takes the average before the onset of the unemployment spell. The first part of this expression indicates that the consumption smoothing gains of long-term benefits will be reduced if individuals with lower marginal utility of consumption remain unemployed for longer. The second part of this expression indicates that the consumption smoothing gains are reduced if households with lower risk aversion and or lower consumer drops remain unemployed for longer.[58]

Our analysis of the consumption profiles by completed spell length already suggested that the dynamic selection on consumption profiles was limited (see Figure 6, Section 4.1). To further assess the potential magnitude of dynamic selection, we investigate in Appendix Table C-1 how various observable characteristics that have been shown to correlate with consumption and risk preferences are distributed across short term and long term unemployed. The patterns of selection are generally small in magnitude and often ambiguous in sign. Income levels and net wealth levels have quantitatively small and non-monotonic effects on the probability to select into longer spells. Compared to households with no or negative net wealth, households with some small wealth (<500kSEK≈ 55kUSD) have a slightly lower probability to be long term unemployed. But individuals with high net wealth have a slightly higher probability to select into long spells. This result suggests that in our context, individuals with better means to smooth consumption do not unambiguously select into longer spells, which corroborates the evidence of no clear patterns of selection on consumption profiles. Also portfolio characteristics (i.e., the fraction of portfolio wealth invested in stocks, and leverage defined as total debt divided by gross assets), which have been shown to be correlated with risk preferences, have small and non-monotonic impacts on the probability to experience a long unemployment spell. In contrast, the probability of experiencing

[58] Note that if preferences are heterogeneous within the group of unemployed at a given time during the unemployment spell, any negative correlation between the drop in consumption and risk aversion, would further reduce the consumption smoothing gains (see Andrews and Miller [2013]).

long unemployment spells ($\mathbb{1}[L > 20 \text{ wks}]$) is significantly and monotonically correlated with age. However, existing evidence from the literature suggests a U-shape relationship between age and risk aversion (Cohen and Einav [2007]), so the dynamic selection on age has an ambiguous effect on the evolution of risk preferences over the spell.

Overall, the observed dynamic selection patterns on consumption and risk preferences do not suggest that relative consumption smoothing gains would be significantly different than our estimates based on the average drops in consumption.

5.3 Beyond Local Recommendations

Our local recommendation regarding the policy and the tilt in particular may not be informative about how the optimal policy should look like.[59] To go beyond our local recommendations, we need to know how the sufficient statistics vary with the policy parameters. We briefly illustrate how this can be explored using the structure of a model. That is, we calibrate a rich, non-stationary structural model to match the sufficient statistics for local policy evaluation and simulate how these sufficient statistics change when moving away from the local policy.

We describe the details of the calibration and simulation in Appendix D. We consider a model with CRRA preferences and additive search costs like in Hopenhayn and Nicolini [1997], Lentz and Traneas [2005] and Chetty [2008a], but motivated by the evidence discussed in the previous subsection we allow for an exit rate function with heterogeneous returns to search that depreciate exponentially and for a heterogeneous asset distribution at the start of the unemployment spell. As our calibration targets our sufficient statistics estimates, the structural model delivers the same local recommendations around the existing policy. That is, the flat benefit $\bar{b} = .72$ is too generous overall and especially so for the short-term unemployed.

Figure 7 illustrates how the moral hazard costs and consumption smoothing gains change when reducing the overall generosity of the policy in our calibrated model. Two findings emerge. First, the consumption smoothing gains increase while the moral hazard costs decrease as we reduce the benefit level.[60] In particular, $CS_{\bar{b}}$ and $MH_{\bar{b}}$ are equalized for a flat replacement of $\bar{b} = .58$. Second, as we decrease the overall generosity, the consumption smoothing gains remain higher for benefits paid to the long-term unemployed ($CS_{\mathbf{2}} > CS_{\mathbf{1}}$), while the moral hazard costs remain lower ($MH_{\mathbf{2}} < MH_{\mathbf{1}}$). The introduction of an inclining tilt ($b_{\mathbf{2}} > b_{\mathbf{1}}$) thus remains welfare-improving for lower replacement rates. In particular, our model predicts that welfare would be maximized by setting $b_{\mathbf{1}} = .48$ for the short-term unemployed and $b_{\mathbf{2}} = .68$ for the long-term unemployed (while keeping the tax rate unchanged).

In Appendix D we further illustrate the complementary value of the structural approach and sufficient-statistics approach. In particular, we explore how our findings depend on the assump-

[59] Of course, this issue may also arise for a one-dimensional change in benefit level $b_{\mathbf{k}}$. Only if the policy problem is strictly concave in all benefit levels, can we be certain that a welfare-increasing change in the benefit level $b_{\mathbf{k}}$ actually moves the policy closer to the optimal policy.

[60] Note that we keep the asset distribution at the start of unemployment fixed, while this is likely to respond to the generosity of the unemployment policy.

tions of the non-stationary features of the search environment and the assumptions made for our consumption-based implementation of the consumption smoothing gains.[61]

6 Conclusion

This paper has offered a simple, general and empirically implementable framework to evaluate the optimal time profile of unemployment benefits. Our theoretical approach proves that, independent of the underlying primitives of the model, the dynamic problem of balancing insurance value and incentive costs can be characterized in a transparent way as a series of simple trade-offs involving just a few estimable statistics. Putting this simple characterization to the data, our empirical implementation has shown that it is not at all obvious that declining benefit profiles are always optimal. Despite the forward-looking behavior of job seekers, important non-stationary forces can make the moral hazard costs of benefits offered to the long-term unemployed higher than the costs of benefits offered early in the spell. The limited access to consumption smoothing opportunities that we document among the unemployed in Sweden also makes cutting benefits particularly costly for the long-term unemployed.

We have presented a framework that is easily replicable and our hope is that it will trigger new empirical work that analyzes the relevant statistics for policy evaluation in other contexts where labor market conditions, access to credit or the unemployment policy may be very different. Our analysis has shown that the empirical analysis of labor supply responses to UI should pay particular attention to the timing of benefits in order to produce estimates that can be meaningful from a welfare perspective. In terms of assessing the value of UI benefits, our analysis shows that fruitful avenues of research are being opened by administrative and/or proprietary data on wealth and expenditures matched with UI records. Most importantly, the tools developed in this paper can be applied to other dynamic contexts. An important area for future work will be to develop such simple, yet robust characterization of various other dynamic policies, including the design of retirement pensions or parental leave policies.

[61] In line with Chetty [2006], we find that for CRRA preferences the Taylor approximation of the marginal utilities substantially underestimates the consumption smoothing gains (by more than 20% for $\gamma = 2$). The approximation error on the relative consumption smoothing gains, however, is much smaller (equal to 5% for $\gamma = 2$), leaving the recommendation on the tilt basically unaffected.

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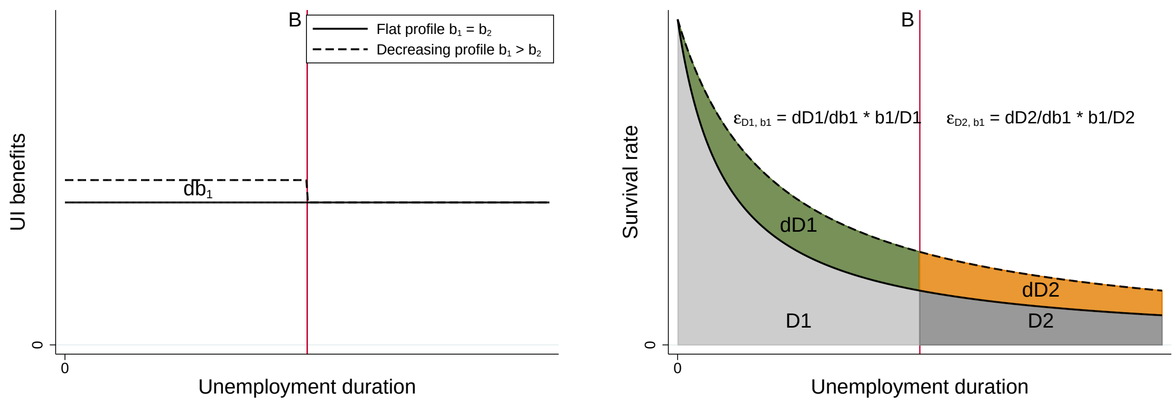
Appendix

Figure 1: Sufficient statistics for welfare analysis of two-part policy

A. Unemployment response wrt b_1

A.I Policy variation | A.II Survival Function



B. Unemployment response wrt b_2

B.I Policy variation | B.II Survival Function

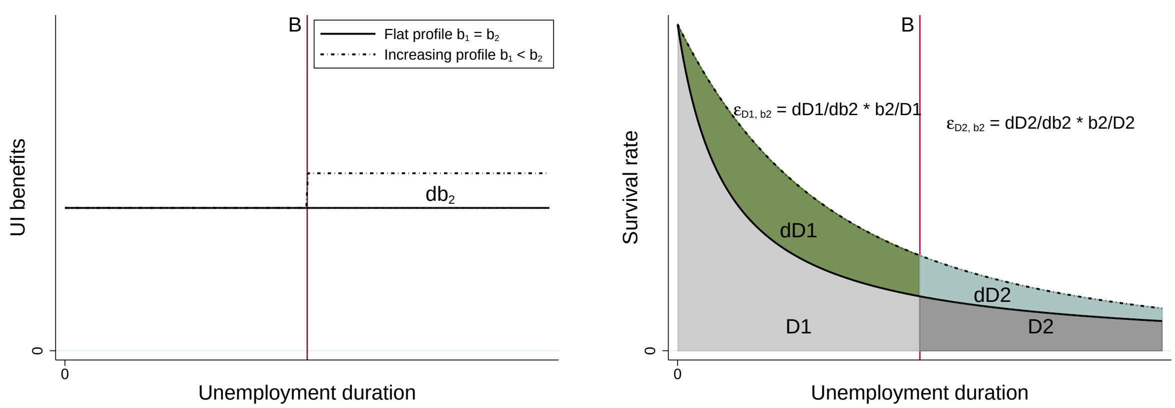
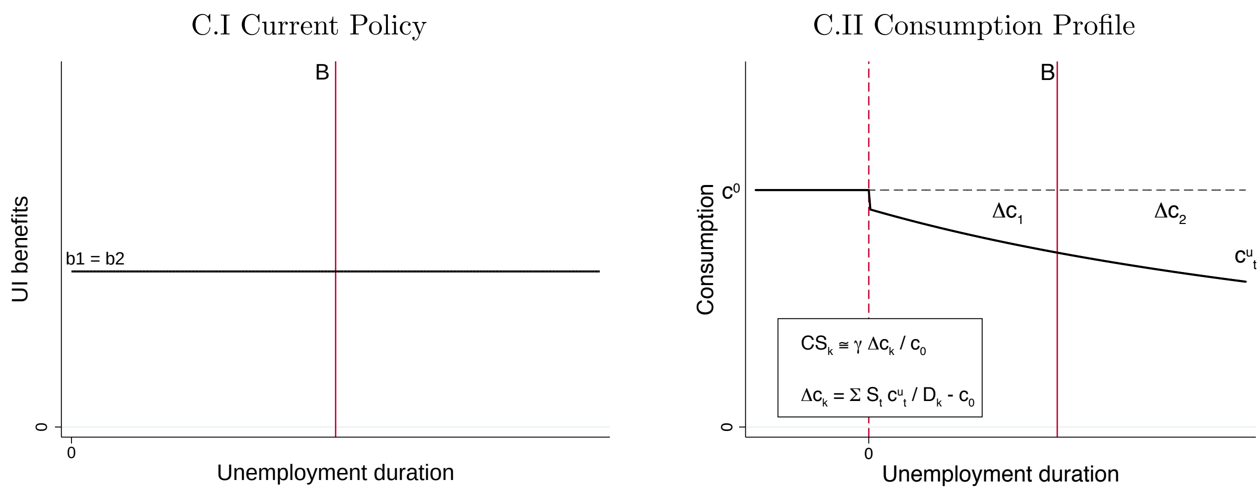


Figure 1: Sufficient statistics for welfare analysis of two-part policy (*continued)*

C. Consumption profile for current policy

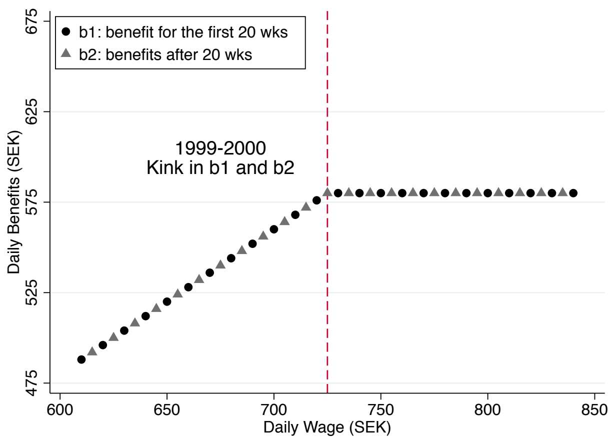


Notes: The figure summarizes the policy variation and statistics needed to characterize an optimal two-part profile giving $b_{\mathbf{1}}$ for the first B weeks and $b_{\mathbf{2}}$ afterwards. In Panel A, we display policy variation $db_{\mathbf{1}}$ in benefits given for the first B weeks that allows evaluating the moral hazard cost of a change in $b_{\mathbf{1}}$, $MH_{\mathbf{1}}$. The moral hazard cost depends on the responses of the duration spent in the first part of the profile $D_{\mathbf{1}}$ and in the second part of the profile $D_{\mathbf{2}}$, as exemplified in the right panel. These responses enable the identification of the cross-duration elasticities $\varepsilon_{D_{\mathbf{1}},b_{\mathbf{1}}}$ and $\varepsilon_{D_{\mathbf{2}},b_{\mathbf{1}}}$ entering the RHS of the corresponding dynamic Baily-Chetty formula (11). Since we start from a flat profile, as is the case in our empirical application in Sweden, the response in total unemployment duration D is in principle sufficient, following equation (6). In Panel B, we display policy variation that allows to evaluate the moral hazard cost of a change in $b_{\mathbf{2}}$. Like in Panel A, we start from a flat profile and display variation $db_{\mathbf{2}}$ in benefits given after B weeks, which enables the identification of the cross-duration elasticities $\varepsilon_{D_{\mathbf{1}},b_{\mathbf{2}}}$, $\varepsilon_{D_{\mathbf{2}},b_{\mathbf{2}}}$ entering the RHS of the corresponding dynamic Baily-Chetty formula. To evaluate the consumption smoothing gains of the two-part policy, following the implementation in (14), the planner requires the average drop in consumption $\Delta c_{\mathbf{1}}$ for individuals in the first part of the profile receiving $b_{\mathbf{1}}$, and the average drop in consumption $\Delta c_{\mathbf{2}}$ for individuals in the second part of the profile receiving $b_{\mathbf{2}}$. This can be calculated based on the profile of consumption as a function of time spent unemployed as depicted in Panel C. Note that these consumption statistics need to be evaluated at the current profile, and do not require any policy variation.

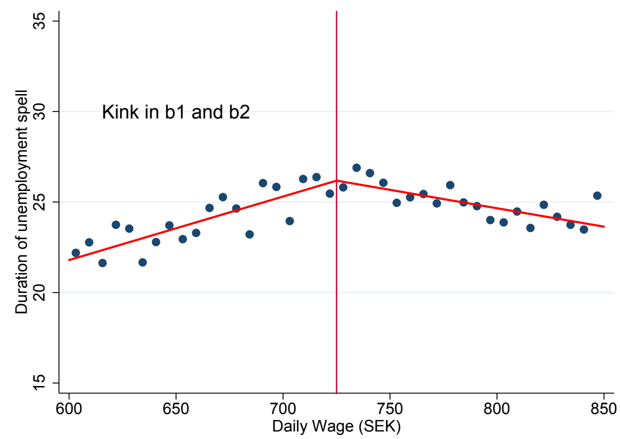
Figure 2: UI benefits and unemployment duration as a function of daily wage around the 725SEK threshold

A. 1999 - 2000

A.I UI schedule

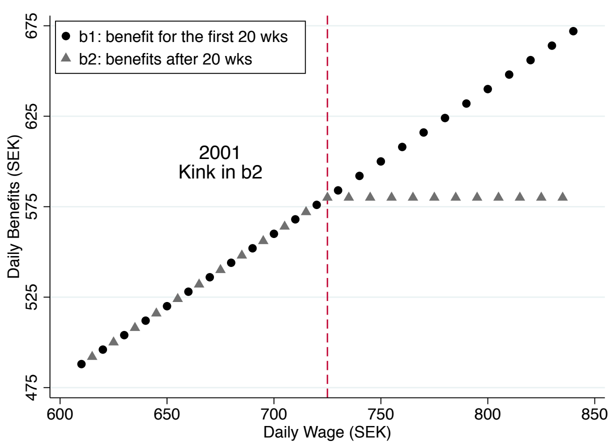


A.II Unemployment duration



B. 2001

B.I UI schedule



B.II Unemployment duration

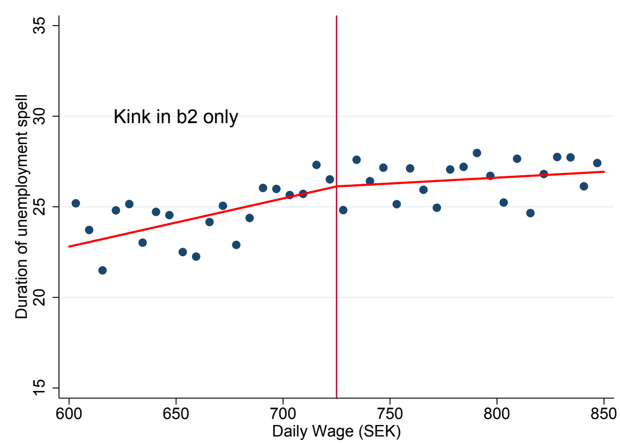
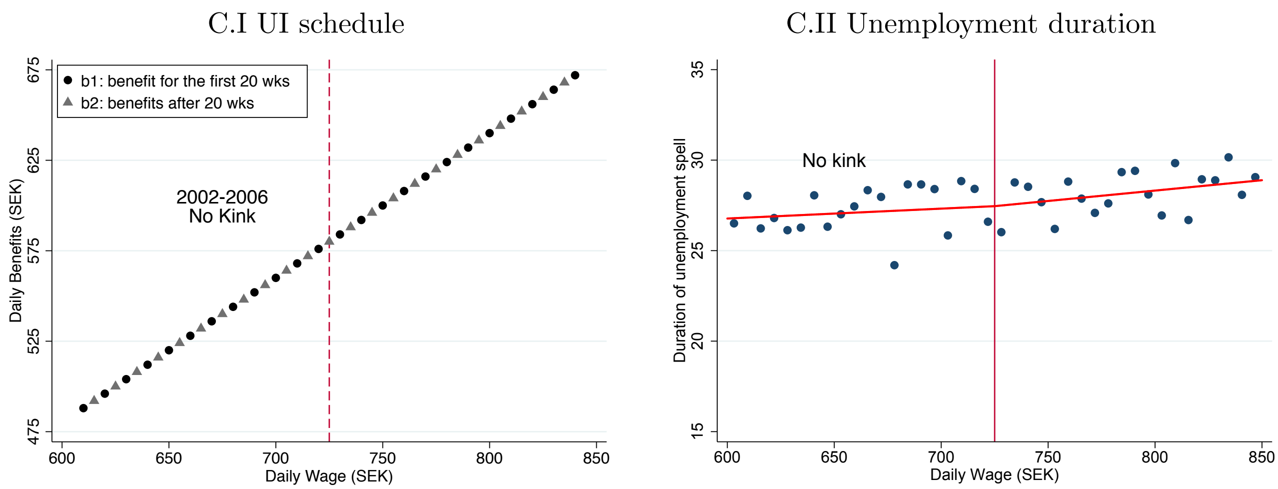


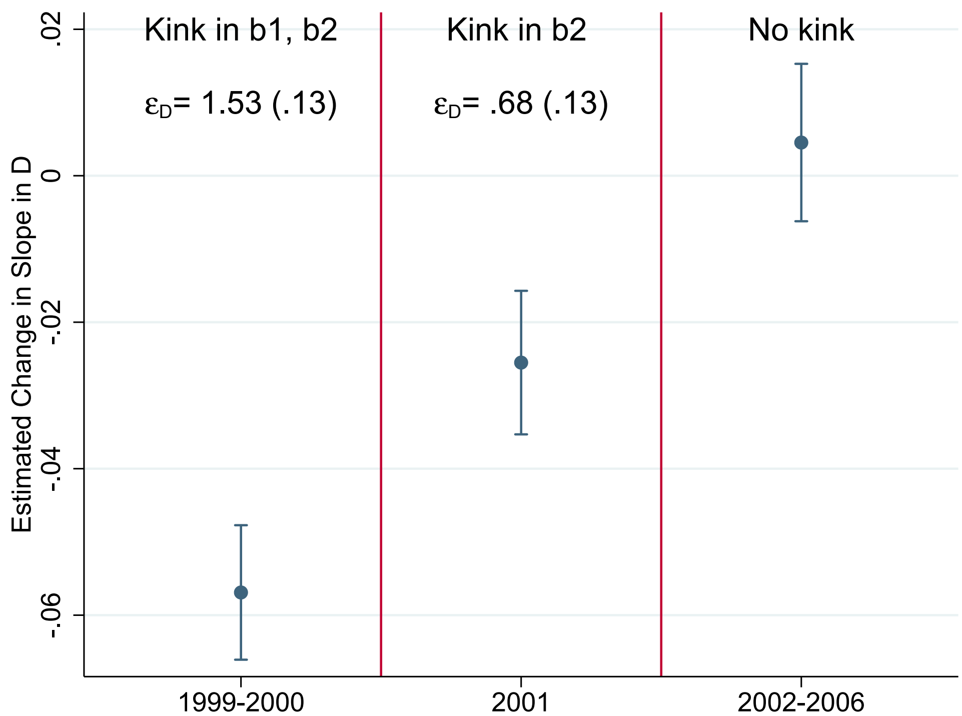
Figure 2: UI benefits and unemployment duration as a function of daily wage around the 725SEK threshold (*continued*)

C. 2002-2007



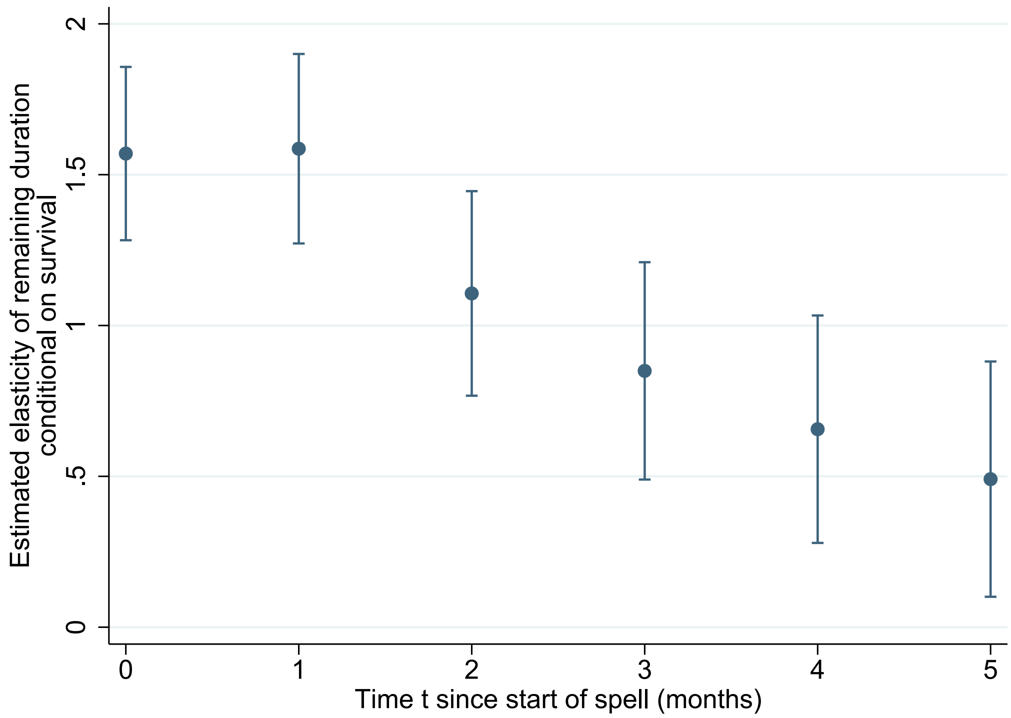
Notes: The left panels display the UI benefit level received during the first 20 weeks of unemployment (b_1) and after 20 weeks of unemployment (b_2) as a function of daily wage prior to becoming unemployed. For spells starting before July 2001 (A.I), the schedule exhibits a kink in both b_1 and b_2 at the 725SEK threshold, which can be used to identify the effect of both b_1 and b_2 on unemployment duration. For spells starting between July 2001 and July 2002 (B.I), the schedule exhibits a kink in b_2 only at the 725SEK threshold, which can be used to identify the effect of b_2 on unemployment duration. Finally, for spells starting after July 2002 (C.I), the schedule is linear for both b_1 and b_2 at the 725SEK threshold, which offers a placebo setting to assess the validity of the RK design at the 725SEK threshold. The right panels plot average unemployment duration in bins of previous daily wage for the three periods of interest. Unemployment duration is defined as the number of weeks between registration at the PES and exiting the PES or finding any employment (part-time or full-time employment, entering a PES program with subsidized work or training, etc.). Unemployment duration is capped at two years. Sample is restricted to unemployed individuals with no earnings who report being searching for full-time employment. The graphs provide graphical evidence of a change in slope in the relationship between unemployment duration and previous daily wage in response to the kink in UI benefits. The change in slope is larger for spells starting before July 2001, when both b_1 and b_2 are capped at the 725SEK threshold (A.II). The magnitude of the change in slope decreases for spells starting between July 2001 and July 2002 when only b_2 is capped at the 725SEK threshold. Finally, there is no significant change in slope for spells after July 2002, when the schedule is linear for both b_1 and b_2 at the threshold, which is supportive of the identifying assumptions of the RK design. Formal estimates of the change in slope using linear specifications of the form of equation (16) are displayed in Table 2. The red lines display predicted values of the regressions in the linear case.

Figure 3: RKD ESTIMATES AT THE 725SEK THRESHOLD



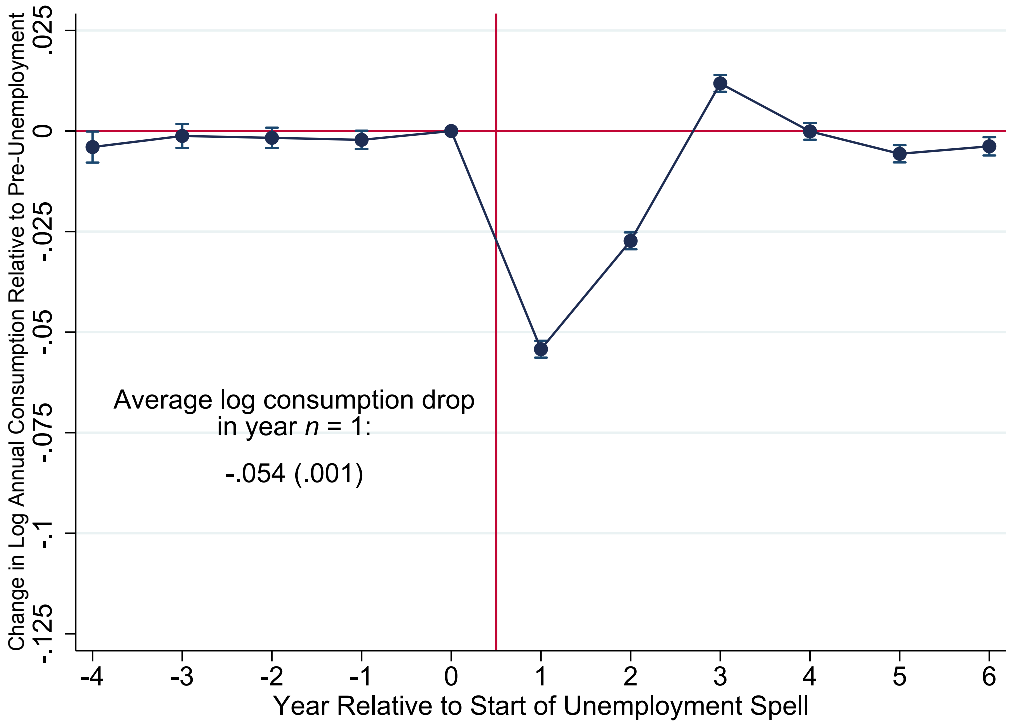
Notes: The figure reports estimates of the change in slope with 95% robust confidence interval in the relationship between unemployment duration and daily wage at the 725SEK threshold using linear regressions of the form of equation (16) with a bandwidth size h = 90SEK. These estimates are reported for three periods of interest: 1999-2000 (i.e., spells starting before July 2001), 2001 (i.e., spells starting after July 2001 and before July 2002) and 2002- (i.e., spells starting after July 2002). Unemployment duration is defined as the number of weeks between registration at the PES and exiting the PES or finding any employment (part-time or full-time employment, entering a PES program with subsidized work or training, etc.). Unemployment duration is capped at two years. Sample is restricted to unemployed individuals with no earnings who report being searching for full-time employment. The figure also reports the corresponding elasticities of unemployment duration with respect to $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$ (for period 1999-2000) and with respect to $b_{\mathbf{2}}$ only (period 2001). Bootstrapped standard errors computed using 50 replications are in parentheses. Formal estimates of the change in slope using linear specifications are displayed in Table 2.

Figure 4: Testing for stationarity: elasticity of the remaining duration of unemployment, conditional on surviving until t, with respect to changes in the flat benefit level



Notes: The figure reports RKD estimates (with 95% robust confidence interval) of the elasticity of the remaining duration of unemployment conditional on surviving until t with respect to changes in the flat benefit level $\bar{b}$. Estimates use the presence, for spells starting before July 2001, of a kink in the benefit schedule of the flat benefit $\bar{b}$ at the 725SEK wage threshold. We use polynomial regressions of the form of equation (16) with a bandwidth size h = 100SEK. The remaining duration $\tilde{D}_t$ is the unemployment duration D minus t, conditional on being still unemployed after t months. Unemployment duration is defined as the number of weeks between registration at the PES and exiting the PES or finding any employment (part-time or full-time employment, entering a PES program with subsidized work or training, etc.). Sample is restricted to unemployed individuals with no earnings who report being searching for full-time employment. In a stationary environment, the elasticity of $\tilde{D}_t$ with respect to the flat benefit $\bar{b}$ should be constant with t. As the estimated elasticities strongly decline with t, our results suggest the presence of strong non-stationary forces (i.e., dynamic selection, duration-dependence, etc.).

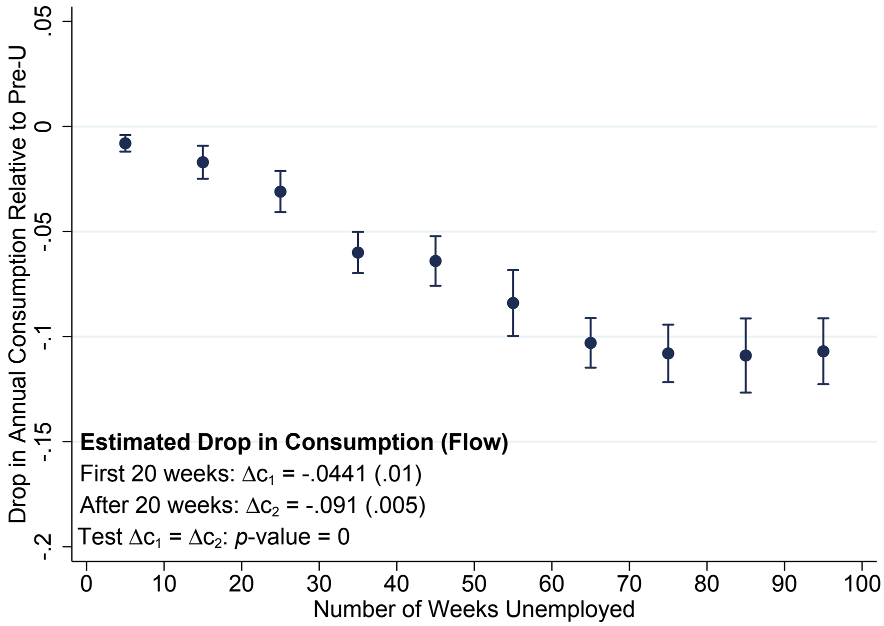
Figure 5: Estimated Change in Log Annual Consumption as a Function of Time Since Start of Unemployment Spell



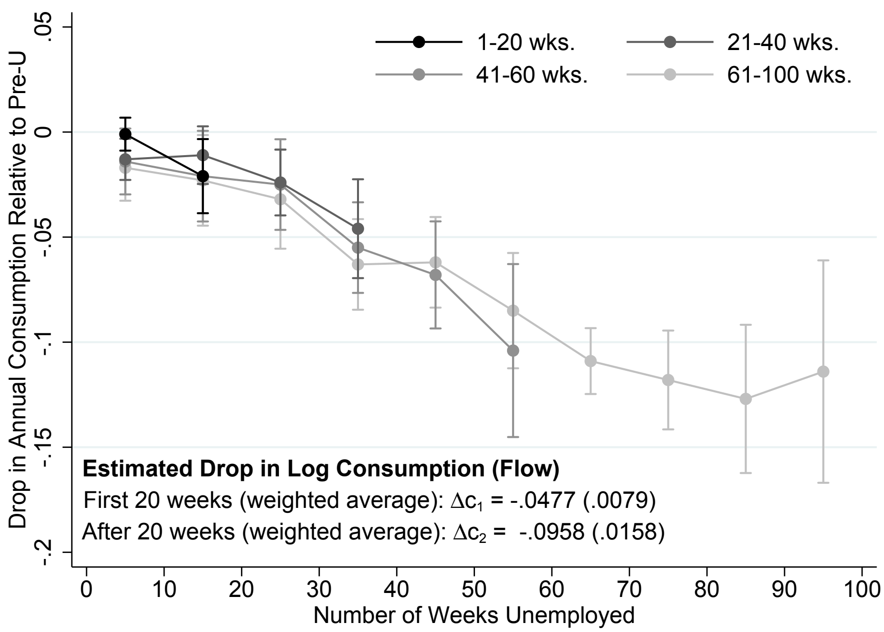
Notes: The Figure analyzes changes in log annual household consumption from our registry-based measure of consumption expenditures at unemployment, following the standard approach in the literature (e.g. Gruber [1997]). We define event year $n = 1$ as the year an unemployment spell starts, and focus on individuals who are observed unemployed in December of year $n = 1$, and who were employed in December of year $n = 0$. We report coefficient estimates from the event year dummies γ_j from event study specification (17). Coefficients are relative to event year $n = 0$. See text for details.

Figure 6: Estimated Drop in Annual Consumption Relative to Pre-Unemployment as a Function of Time Spent Unemployed

A. Baseline Estimates

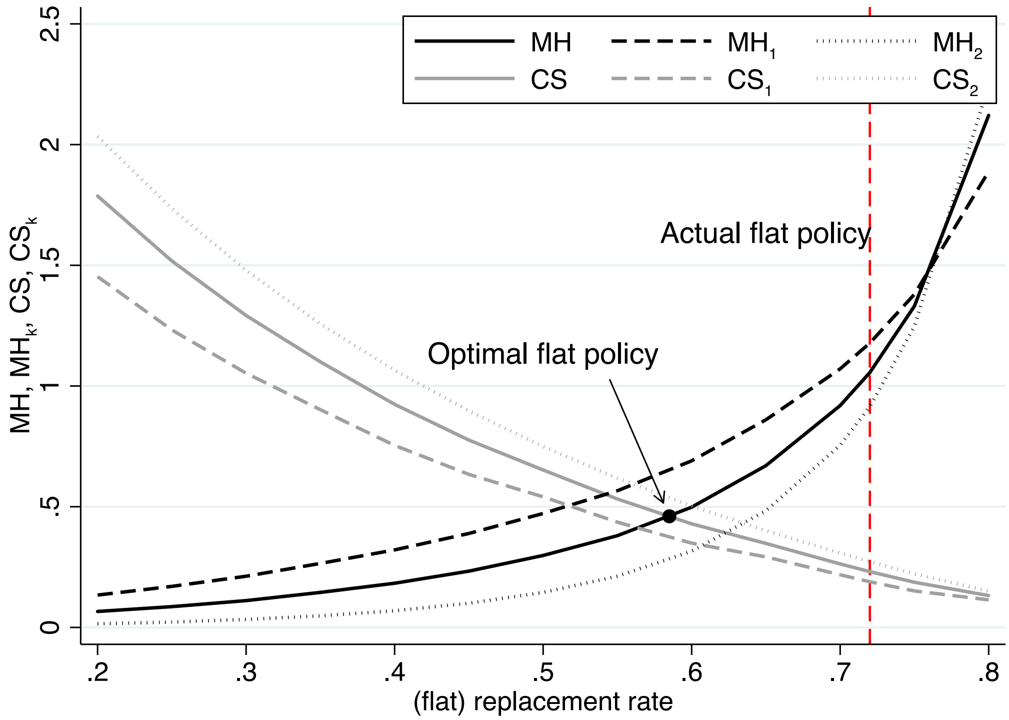


B. Profile Heterogeneity by Completed Spell Length



Notes: This Figure shows average annual consumption drops compared to pre-unemployment $\Delta\mathscr{C}_t$ by time t spent unemployed as of December (when annual consumption is observed in the registry data), and uses this variation to non-parametrically estimate the evolution of higher frequency household consumption throughout the unemployment spell. Panel A reports $\hat{\beta}_t/\mathscr{C}_0 = \widehat{\Delta\mathscr{C}_t}/\mathscr{C}_0$, i.e. estimates from equation (20) scaled by the average annual consumption in the last year prior to unemployment, so that all consumption drops are expressed relative to pre-unemployment levels. The Figure also reports non-parametric estimates of the average drops in consumption in each parts of the benefit profile $\widehat{\Delta C_\mathbf{1}}$ and $\widehat{\Delta C_\mathbf{2}}$ following the methodology explained in subsection 4.1. Standard errors are computed using the Delta-method. We also report the p-value from a test of equality of $\widehat{\Delta C_\mathbf{1}}$ and $\widehat{\Delta C_\mathbf{2}}$. In Panel B, we estimate separate profiles $\Delta\mathscr{C}_{t,k=1,.,4}$ for 4 categories of total completed spell length L, and report estimates of $\widehat{\Delta C_1}$ and $\widehat{\Delta C_2}$ allowing for profile heterogeneity, following the methodology explained in section 4.1. Both panels provide compelling evidence of a significant drop in average flow consumption over the unemployment spell.

Figure 7: Structural Model: Welfare Effects for Different Benefit Levels



Notes: The figure illustrates how the moral hazard costs and consumption smoothing gains change for different levels of the flat benefit profile in our structural model. The model is calibrated to match our sufficient-statistics estimates, evaluated at the policy in place, which corresponds to a flat profile with average replacement rate of .72 as indicated by the vertical dashed line. We report the simulated moral hazard costs and consumption smoothing gains for an overall change in the flat benefit profile $\bar{b}$, for an increase in the benefit level in the first 20 weeks of unemployment, and for an increase in the benefit level after 20 weeks of unemployment. The simulated values of $CS_{\bar{b}}$ and $MH_{\bar{b}}$ (labeled CS and MH in the figure) are equalized for $\bar{b} = .58$. The consumption smoothing gains remain higher for benefits paid after 20 weeks ($CS_{\mathbf{2}} > CS_{\mathbf{1}}$), while the moral hazard costs remain lower ($MH_{\mathbf{2}} < MH_{\mathbf{1}}$), indicating that the introduction of an inclining tilt ($b_{\mathbf{2}} > b_{\mathbf{1}}$) remains welfare improving for lower replacement rates.

Table 1: Summary statistics: Duration Response and Consumption Response Sample

	Mean	P10	P50	P90
	I. Unemployment			
Duration of spell (wks)	26.78	3	16	73.43
Duration on $b_{\mathbf{1}}$ (wks)	13.52	3	16	20
Duration on $b_{\mathbf{2}}$ (wks)	13.26	0	0	53.43
Replacement rate	.72	.64	.78	.8
	II. Demographics			
Age	34.88	23	32	52
Fraction men	.48	0	0	1
Fraction married	.39	0	0	1
Fraction with higher educ	.25	0	0	1
	III. Income and Wealth, SEK 2003(K)			
Gross earnings (individual)	190.3	171.3	191.3	207.52
Household net wealth	481.2	-222.4	20.6	1318.1
Household bank holdings	52.9	0	0	139.1
Household real estate	631.2	0	163.9	1605.6
Household debt	385.1	0	176.2	935.9

Notes: The table provides summary statistics for our main sample of unemployed individuals used in the RKD analysis of Section 3 and in the consumption response analysis of Section 4. The sample is drawn from the universe of unemployed individuals from the unemployment registers (PES - IAF) in Sweden from 1999 to 2007. We kept in the sample all individuals with daily wage in a bandwidth of 200SEK around the kink point in the benefit schedule. All earnings, income and asset level measures are from wealth and income registers, and are yearly measures aggregated at the household level in constant k2003SEK for the last calendar year of full employment prior to the start of the unemployment spell. All financial assets are estimated at their market value. Real estate is gross of debt and assessed at market value. Debt includes student loans, mortgage, credit card debt, etc. Note that 1SEK2003 $\approx$ 0.11USD2003.

Table 2: RKD ESTIMATES AT THE 725SEK THRESHOLD

	(1) Unemployment Duration D	(2) Duration $D_{\mathbf{1}}$ (< 20 weeks)	(3) Duration $D_{\mathbf{2}}$ (≥ 20 weeks)
	I. 1999-2000: Kink in b_1 and b_2		
$\delta_{\mathbf{k}}$	-.0569	-.0246	-.0299
	(.0050)	(.0012)	(.0039)
$\varepsilon_{D_{\mathbf{k}},\bar{b}}$	1.530	1.319	1.615
	(.1300)	(.0645)	(.1986)
N	187518	187518	187518
	II. 2001: Kink in b_2 only		
$\delta_{\mathbf{k}}$	-.0255	-.0115	-.0105
	(.0049)	(.0020)	(.0030)
$\varepsilon_{D_{\mathbf{k}},b_2}$	.6765	.6015	.5921
	(.1312)	(.1061)	(.1642)
N	65545	65545	65545
	III. 2002-. : Placebo		
$\delta_{\mathbf{k}}$	.0045	-.0016	.006
	(.0055)	(.0011)	(.0049)
N	172645	172645	172645
	IV. Moral Hazard Estimates		
$MH = \frac{\bar{b}+\tau}{\bar{b}} \cdot \varepsilon_{D,\bar{b}}$	1.637	(.1391)	
$MH_{\mathbf{2}} = \frac{\bar{b}+\tau}{\bar{b}} \cdot \frac{D}{D_{\mathbf{2}}} \cdot \varepsilon_{D,b_{\mathbf{2}}}$	1.445	(.2829)	
$MH_{\mathbf{1}} = \frac{\bar{b}+\tau}{\bar{b}} \cdot \frac{D}{D_{\mathbf{1}}} \cdot (\varepsilon_{D,\bar{b}} - \varepsilon_{D,b_{\mathbf{2}}})$	1.819	(.4032)	
Hypotheses testing : $\mathbf{MH_1 = MH_2}$		z-stat	p-value
Z-test		-0.57	0.569
Permutation Test			0.059

Notes: The table reports estimates of the change in slope $\delta_{\mathbf{k}}$, at the 725SEK threshold, in the relationship between daily wage and the total duration of unemployment D (col. (1)), the time $D_{\mathbf{1}}$ spent on the first part of the Swedish UI profile (col. (2)) and the time $D_{\mathbf{2}}$ spent on the second part of the Swedish UI profile (col. (3)). Estimates are obtained from linear regressions of the form of equation (16) with a bandwidth size h = 90SEK. Panel I reports estimates for spells starting before July 2001. Panel II reports estimates for spells starting after July 2001 and before July 2002. Panel III reports estimates for spells starting after July 2002. Unemployment duration is capped at two years. We report implied elasticities, computed as $\varepsilon_{D_{\mathbf{k}},b_{\mathbf{k}}} = \hat{\delta}_k \cdot \frac{\bar{w}_{\mathbf{k}}}{D_b^{cap}}$, where $\hat{\delta}_{\mathbf{k}}$ is the estimated marginal slope change, and D_b^{cap} is the observed average duration at the kink. In Panel IV we also report implied moral hazard costs estimates defined in equation (6). MH is the moral hazard cost of increasing benefits throughout the unemployment spell. $MH_{\mathbf{1}}$ is the moral hazard cost of increasing benefits given for the first 20 weeks of the spell. $MH_{\mathbf{2}}$ is the moral hazard cost of increasing benefits given after the first 20 weeks of the spell. Computations assume $\tau = .05$ which balances the UI budget on average during the period 1999-2007. It follows that $\frac{\bar{b}+\tau}{\bar{b}} = 1 + .05/.72 = 1.07$. See text for details. Standard errors are obtained from bootstrapping using 50 replications.

Table 3: Welfare evaluation of actual profile using estimated sufficient statistics

	(1) Moral hazard hazard costs	(2) Average consumption drop	(3) Consumption smoothing gains $CS_{\mathbf{k}}$	(4)	(5)	(6) Benefit-Cost ratio
	$MH_{\mathbf{k}}$	$\Delta c_{\mathbf{k}}$	$\gamma = 1$	$\gamma = 2$	$\gamma = 5$	$CS_{\mathbf{k}}/MH_{\mathbf{k}}$
Benefits given throughout the spell: $\bar{b}$	1.64	.08	.08	.16	.40	$\gamma \times .049$
	(.14)	(.01)	(.01)	(.02)	(.04)	
Benefits given for the first 20 weeks: $b_{\mathbf{1}}$	1.82	.05	.05	.10	.24	$\gamma \times .026$
	(.40)	(.01)	(.01)	(.02)	(.04)	
Benefits given after first 20 weeks: $b_{\mathbf{2}}$	1.45	.10	.10	.19	.48	$\gamma \times .066$
	(.28)	(.02)	(.02)	(.03)	(.08)	

Notes: The Table reports estimates of the sufficient statistics needed to evaluate the benefit profile in place in Sweden. The first row analyzes the welfare consequences of an increase in benefits $\bar{b}$ throughout the unemployment spell. The second row analyzes the welfare consequences of an increase in benefits $b_{\mathbf{1}}$ during the first 20 weeks of the unemployment spell. The third row analyzes the welfare consequences of an increase in benefits $b_{\mathbf{2}}$ after 20 weeks of unemployment. In each case, column (1) repeats our estimates of the moral hazard cost of an increase in benefits (Panel IV of Table 2). The second column repeats our estimates of the average consumption drop (column 1 of Table C-3). To estimate the average consumption drop over the entire spell, we run the same regression as in (29) but with only one dummy $\mathbb{1}[t > 0]$ that indicates being observed while unemployed. We convert the respective consumption drops into estimates of the consumption smoothing gains $CS_{\mathbf{k}}$, following the implementation in (14), which relies on a Taylor approximation and homogeneous preferences. The consumption smoothing gains are reported for a range of plausible values of the relative risk aversion γ in columns (3) to (5). Column (6) shows the ratio of consumption smoothing gains to moral hazard costs, $CS_{\mathbf{k}}/MH_{\mathbf{k}}$, depending on the uniform relative risk aversion γ. This ratio corresponds to the marginal value of a (tax-funded) kroner spent on unemployment benefits, accounting for the unemployment responses. Bootstrapped standard errors in parentheses.

FOR ONLINE PUBLICATION - Appendix A: Technical Appendix

This Appendix provides the technical details of our model setup and the proofs of the Propositions. We also consider extensions of the baseline model to gauge the robustness of our characterization of the optimal benefit profile. We finally compare the evolution of consumption smoothing gains and moral hazard in stationary and non-stationary environments.

A.1 Setup

We closely follow the setup in Chetty [2006], but allow for heterogeneous agents and non-stationarities. Let $\omega_{i,t}$ denote a vector of state variables that contain all relevant information up to time t in determining an agent i's employment status and behavior at time t. Let $F_{i,t}(\omega_{i,t})$ denote the unconditional distribution of $\omega_{i,t}$ given information available at time 0. We assume that $F_{i,t}$ is a smooth function and let Ω denote the maximal support of $F_{i,t}$ for $\forall i, \forall t$. In our stylized model, the vector of state variables $\omega_{i,t}$ includes only the asset level, time and the employment status.

In each period t, an agent decides how much to consume from her income and assets. In our stylized model, an agent earns $w-\tau$ when employed and receives b when unemployed. The law of motion of assets in the employment and unemployment state are respectively,

$$a_{i,t+1} = ra_{i,t} + w - \tau - c^e_{i,t} \tag{22}$$

$$a_{i,t+1} = ra_{i,t} + b_t - c^u_{i,t}, \tag{23}$$

but are constrained to be above $a_{i,t+1} \geq \bar{a}_i$ for each agent i and any time t. We denote the Lagrange multipliers on these constraints by $\mu^e_{i,t}(\omega_{i,t}), \mu^u_{i,t}(\omega_{i,t})$ and $\mu^a_{i,t}(\omega_{i,t})$ respectively.

Let $\theta_{i,t}(\omega_{i,t})$ denote an agent's employment status at time t in state $\omega_{i,t}$. If $\theta = 1$, the agent is employed, and if $\theta = 0$, the agent is unemployed. In each period t, an unemployed agent chooses a level of search effort $s_{i,t}$ as well. This search effort level determines the probability to leave unemployment for employment in that period. This mapping may be agent-specific and change depending on the length of the unemployment spell.

Each agent i chooses a program (s_i, c^u_i, c^e_i) with

$$\begin{aligned} s_i &= \{s_{i,t}(\omega_{i,t})\}_{t\in\{1,2,..T\},\omega_{i,t}\in\Omega,\theta(\omega_{i,t})=0}\,, \\ c^u_i &= \{c^u_{i,t}(\omega_{i,t})\}_{t\in\{1,2,..T\},\omega_{i,t}\in\Omega,\theta(\omega_{i,t})=0}\,, \\ c^e_i &= \{c^e_{i,t}(\omega_{i,t})\}_{t\in\{1,2,..T\},\omega_{i,t}\in\Omega,\theta(\omega_{i,t})=1}\,, \end{aligned}$$

to solve

$$\begin{aligned} V_i(P) = \max \Sigma_{t=1}^T \beta^{t-1} &\int \{v^u_i\left(c^u_{i,t}(\omega_{i,t}), s_{i,t}(\omega_{i,t})\right)[1-\theta_{i,t}(\omega_{i,t})] + v^e_i\left(c^e_{i,t}(\omega_{i,t})\right)\theta_{i,t}(\omega_{i,t})\} dF_{i,t}(\omega_{i,t}) \\ + \Sigma_{t=1}^T \beta^{t-1} &\int \mu^u_{i,t}(\omega_{i,t})\left[ra_{i,t}(\tilde{\omega}_{i,t-1}) + b_t - c^u_{i,t}(\omega_{i,t}) - a_{i,t+1}(\omega_{i,t})\right][1-\theta_{i,t}(\omega_{i,t})]\, dF_{i,t}(\omega_{i,t}) \\ + \Sigma_{t=1}^T \beta^{t-1} &\int \mu^e_{i,t}(\omega_{i,t})\left[ra_{i,t}(\tilde{\omega}_{i,t-1}) + w - \tau - c^e_{i,t}(\omega_{i,t}) - a_{i,t+1}(\omega_{i,t})\right]\theta_{i,t}(\omega_{i,t})\, dF_{i,t}(\omega_{i,t}) \\ + \Sigma_{t=1}^T \beta^{t-1} &\int \mu^a_{i,t}(\omega_{i,t})\left[\bar{a}_i - a_{i,t+1}(\omega_{i,t})\right] dF_{i,t}(\omega_{i,t})\,, \end{aligned}$$

where we use the short-hand notation $\tilde{\omega}_{i,t-1}$ to denote the vector of state variables at time $t-1$ that preceded the vector of state variables $\omega_{i,t}$ at time t. Following Chetty [2006], we assume that lifetime utility is smooth, increasing and strictly quasi-concave in (c^u_i, c^e_i, s_i) and that the value function $V_i(P)$ is differentiable such that the Envelope

Theorem applies. This implies that

$$\begin{aligned}\frac{\partial V_i(P)}{\partial b_t} &= \beta^{t-1}\int \mu_{i,t}^u(\omega_{i,t})\left[1-\theta_{i,t}(\omega_{i,t})\right]dF_{i,t}(\omega_{i,t})\\ &= \beta^{t-1}\int \frac{\partial v_i^u\left(c_{i,t}^u(\omega_{i,t}),s_{i,t}(\omega_{i,t})\right)}{\partial c_{i,t}^u}\left[1-\theta_{i,t}(\omega_{i,t})\right]dF_{i,t}(\omega_{i,t}).\end{aligned}$$

The second equality uses the optimality of the consumption choice $c_{i,t}^u(\omega_{i,t})$, which does not depend on the borrowing constraint being binding or not.

In our stylized model, the agent starts unemployed and remains employed until T once she finds a job. The agent's exit rate out of unemployment at time t only depends on her search effort at time t. The (unconditional) probability to be unemployed at time $t+1$ therefore simplifies to

$$\Pr(\theta_{i,t+1}=0)=\int(1-h_{i,t}(s_{i,t}(\omega_{i,t})))\left[1-\theta_{i,t}(\omega_{i,t})\right]dF_{i,t}(\omega_{i,t}).$$

This simplifying assumption makes that on the optimal path an agent's unemployment consumption $c_{i,t}^u(\omega_{i,t})$ only varies with time t, which coincides with the number of periods she is currently unemployed. The agent's employment consumption $c_{i,t}^e(\omega_{i,t})$, however, depends on both time t and the number of periods she has been unemployed.

We now turn to the policy. We can express the present value of the government's budget as

$$G(P)=\Sigma_{t=1}^T[1+r]^{-(t-1)}\int\int\{-b_t\left[1-\theta_{i,t}(\omega_{i,t})\right]+\tau\theta_{i,t}(\omega_{i,t})\}dF_{i,t}(\omega_{i,t})\,di,$$

which simplifies to (1) when $r=0$.

The government solves

$$\max\int V_i(P)\,di+\lambda\left[G(P)-\bar{G}\right],$$

where λ is the Lagrange multiplier on the government's budget constraint and $\bar{G}$ is an exogenous revenue constraint. Our characterization is based on local policy changes and thus only allows for local tests and recommendations. For the local recommendations to translate globally, we would need the program to be strictly concave in P.[62] To provide tractable expressions of the local welfare implications we assume that the social welfare function is differentiable.

A.2 Dynamic Unemployment Policy

A.2.1 Proof of Proposition 1

The welfare impact of a change in benefit level b_t of policy P equals

$$\frac{\partial W(P)}{\partial b_t}=\int\frac{\partial V_i(P)}{\partial b_t}di+\lambda\frac{\partial G(P)}{\partial b_t},$$

where, using $S_t^r\equiv S_t/[1+r]^{(t-1)}$ and $\varepsilon_{t',t}^r=\frac{\partial S_{t'}^r}{\partial b_t}\frac{b_t}{S_{t'}^r}$,

$$\frac{\partial G(P)}{\partial b_t}=-S_t^r-\Sigma_{t'=1}^T(b_{t'}+\tau)\frac{\partial S_{t'}^r}{\partial b_t}=-S_t^r\times\left[1+\Sigma_{t'=1}^T\frac{S_{t'}^r(b_{t'}+\tau)}{S_t^r b_t}\varepsilon_{t',t}^r\right],$$

[62] Chetty (2006) provides regularity conditions such that the government's problem is strictly concave in case of flat unemployment policies (i.e., $b_k=\bar{b}$).

which simplifies to expression (5) for $r = 0$, and

$$\int \frac{\partial V_i(P)}{\partial b_t} di = \int \int \beta^{t-1} \frac{\partial v_i^u\left(c_{i,t}^u(\omega_{i,t}), s_{i,t}(\omega_{i,t})\right)}{\partial c_{i,t}^u} \left[1 - \theta_{i,t}(\omega_{i,t})\right] dF_{i,t}(\omega_{i,t})\} di$$
$$= \beta^{t-1} S_t E\left(\frac{\partial v_i^u\left(c_{i,t}^u(\omega_{i,t}), s_{i,t}(\omega_{i,t})\right)}{\partial c_{i,t}^u} | t, \theta_{i,t}(\omega_{i,t}) = 0\right).$$

This expression simplifies to (7) for $\beta = 1 + r = 1$. The expectation operator $E_t^u(\cdot)$ thus *averages over all potential states in which the agent is unemployed at time t. In our stylized setup (which assumes that the agent starts unemployed and remains employed once she finds a job), the agent's unemployment consumption $c_{i,t}^u(\omega_{i,t})$ only depends on the length of the ongoing unemployment spell. The weight of agent i's marginal utility in calculating the average marginal utility among the unemployed at time t is scaled by $S_{i,t}/S_t$.*

Combining the two expressions, we find

$$\frac{\partial W(P)}{\partial b_t} = 0 \Leftrightarrow \frac{\int \frac{\partial V_i(P)}{\partial b_t} di - \lambda}{\lambda} = \Sigma_{t'=1}^T \frac{S_{t'}^r(b_{t'} + \tau)}{S_t^r b_t} \varepsilon_{t',t}^r.$$

In the same way, we find

$$\frac{\partial G(P)}{\partial \tau} = \left[\Sigma_{t=1}^T (1 - S_t)/[1+r]^{(t-1)}\right] \times \left[1 + \Sigma_{t'=1}^T \frac{S_{t'}^r(b_{t'} + \tau)}{\Sigma_{t=1}^T \{(1 - S_t)/[1+r]^{(t-1)}\}\tau} \varepsilon_{t',\tau}^r\right],$$
$$\int \frac{\partial V_i(P)}{\partial \tau} di = \Sigma_{t=1}^T \beta^{t-1} (1 - S_t) E\left(\frac{\partial v_i^e\left(c_{i,t}^e(\omega_{i,t})\right)}{\partial c_{i,t}^e} | t, \theta_{i,t}(\omega_{i,t}) = 1\right),$$

and, hence,

$$\frac{\partial W(P)}{\partial \tau} = 0 \Leftrightarrow \frac{\lambda - \int \frac{\partial V_i(P)}{\partial \tau} di}{\lambda} = \Sigma_{t'=1}^T \frac{S_{t'}^r(b_{t'} + \tau)}{\Sigma_{t=1}^T \{(1 - S_t)/[1+r]^{(t-1)}\}\tau} \varepsilon_{t',\tau}^r,$$

which simplifies to expression (9) for $\beta = 1 + r = 1$, with $(T - D)$ equal to the expected time spent employed $\Sigma_{t=1}^T (1 - S_t)$. *The expectation operator $E^e(\cdot)$* in *(9) is over all employment states* and periods t. *Compared to consumption during unemployment, employment consumption $c_{i,t}^e(\omega_{i,t})$ depends on the unemployment history and not just on time t. Hence, we need to calculate the average marginal utility when employed at time t for any agent i and scale the weight in calculating the average marginal utility among the employed at time t by $(1 - S_{i,t})/(1 - S_t)$. We then average over all periods t using weights $(1 - S_t)/(T - D)$.*

The $n + 1$ first-order conditions stated in the Proposition, jointly with the budget constraint, are necessary considtions for an interior, optimal policy.■

A.2.2 Robustness of Characterization

We briefly show how the optimal tax formulae continue to apply in a model with multiple unemployment spells where an agent i's layoff probability $l_i(e_{i,t})$ at time t depends on her effort on the job $e_{i,t}$. We still assume that $\omega_{i,t}$ contains all relevant information up to time t in determining an agent i's employment status and behavior at time t. Let $\theta_{i,t}(\omega_{i,t})$ still denote an agent's employment status at time t in state $\omega_{i,t}$. If $\theta = 1$, the agent is employed, and if $\theta = 0$, the agent is unemployed.

From the consumption smoothing perspective, the agent's marginal utility when employed can now depend on the effort on the job, $\partial v_i^e\left(c_{i,t}^e(\omega_{i,t}), e_{i,t}(\omega_{i,t})\right)/\partial c_{i,t}^e$. From the moral hazard perspective, the (unconditional) probability to be unemployed now equals

$$Pr\left(\theta_{i,t+1}=0\right)=\int\left\{\left(1-h_{i}\left(s_{i,t}\left(\omega_{i,t}\right),\omega_{i,t}\right)\right)\left[1-\theta_{i,t}\left(\omega_{i,t}\right)\right]+l_{i}\left(e_{i,t}(\omega_{i,t})\right)\theta_{i,t}\left(\omega_{i,t}\right)\right\}dF_{i,t}\left(\omega_{i,t}\right).$$

We introduce the indicator functions $I_{i,t}^{\tilde{t}}\left(\omega_{i,t}\right)$ which take value 1 if the length of the ongoing unemployment spell equals $\tilde{t}$ and 0 otherwise. Hence,

$$Pr\left(I_{i,t+1}^{1}=1\right)=\int l_{i}\left(e_{i,t}(\omega_{i,t})\right)\theta_{i,t}\left(\omega_{i,t}\right)dF_{i,t}\left(\omega_{i,t}\right),$$
$$Pr\left(I_{i,t+1}^{\tilde{t}}=1\right)=\int(1-h_{i}\left(s_{i,t}\left(\omega_{i,t}\right),\omega_{i,t}\right))I_{i,t}^{\tilde{t}-1}\left(\omega_{i,t}\right)dF_{i,t}\left(\omega_{i,t}\right).$$

The budget constraint still depends on the survival rate at each unemployment duration $S_{\tilde{t}}^{r}$, but now potentially spread over multiple spells. That is,

$$S_{\tilde{t}}^{r}=\Sigma_{t=1}^{T}\left[1+r\right]^{-(t-1)}\int\int I_{i,t}^{\tilde{t}}\left(\omega_{i,t}\right)dF_{i,t}\left(\omega_{i,t}\right)di.$$

Hence, the optimal formulae in Proposition 1 remain exactly the same (with the marginal utility of consumption when employed depending on effort on the job). The policy-relevant elasticity should account for potential responses in the layoff rate to a change in the unemployment policy. In our context, however, we find no significant responses in the layoff rates to changes in UI benefits.[63]

We refer to Chetty [2006] for a detailed treatment of other extensions of the model (including private insurance arrangements, spousal labour supply, etc.) which do not affect the optimal tax formulae due to envelope conditions. This remains true when extending his analysis to a dynamic benefit profile. For example, we can introduce alternative sources of income $z_{i,t}\left(x_{i,t},\omega_{i,t}\right)$ into the agent's budget constraints (22) and/or (23), with the income level depending on the agent's choice variable $x_{i,t}$, which may enter the agent's utility function when employed and/or unemployed. As long as there are no externalities related to this alternative source of income, envelope conditions imply that the welfare impact of a policy change is still captured by the same statistics.

A.2.3 Characterization with Employer Screening

We consider a reduced-form model of employer screening based on Lockwood [1991] and Lehr [2017]. The job finding rate rate $h_{i,t}\left(s_{i,t},\mathbf{S}_{t}\right)=\lambda_{i}\left(s_{i,t}\right)\times\mu_{t}(\mathbf{S}_{t})$ is determined not only by the probability that agent i finds a vacant position $\lambda_{i}\left(s_{i,t}\right)$, depending on her own search effort, but also by the probability that the matching firm hires the agent $\mu_{t}(\mathbf{S}_{t})$ where $\mathbf{S}_{t}=\left\{S_{i',t}\right\}_{i'}$. The hiring probability at time t depends on the relative survival rates by agents with different productivity. In a two-type version of the model (H and L), where an agent i's type affects both her productivity θ_i and the probability of finding a vacancy $\lambda_{i}\left(s_{i,t}\right)$, the firm's optimal hiring rate when matched with a job seeker who has been unemployed for t periods increases in the relative survival rate of the high type at time t. In particular, if the firm's profit of hiring an agent equals $\theta_i - w$, where $\theta^H > \theta^L = 0$, the firm's optimal hiring decision equals $\mu_t = 1$

[63]First, if layoff rates respond to the unemployment policy, this has implications regarding the pdf of daily wages around the kink in our empirical setting. The presence of a kink in benefits should create bunching at the kink if there is moral hazard on the job with convex costs of shirking. We show in subsection B.4 of Appendix B that we cannot detect any bunching at the kink. Furthermore, if layoffs are responsive to UI benefits this should also affect the pdf of daily wages when the kink in the schedule is removed. We show in subsection B.4 of Appendix B that we cannot detect such changes in the pdf of daily wages after the removal of the kinks in the schedule. While this evidence is far from definitive, it suggests that layoff rates do not seem to strongly respond to UI benefits in our context.

if $\frac{S_t^H}{S_t^L}\theta^H \geq w$ and 0 otherwise. As a consequence, with the more productive type leaving unemployment at a faster rate, the firm would not hire job seekers who have been unemployed for longer than $\bar{t}$ where $\frac{S_{\bar{t}}^H}{S_{\bar{t}}^L}\theta^H = w$.

In the employer screening model, an agent's search effort will affect the job finding probability of any other agent, positively or negatively depending on her type, but no agent internalizes this effect. For simplicity we focus on job seekers' welfare and ignore the impact on firms' profits. Note that the setup can in principle also encompass richer models with rationing (e.g., Michaillat [2012b]) and employer ranking (e.g., Blanchard and Diamond [1994]), in which job seekers' search effort crowd out the job finding rate of other job seekers. This is analyzed in Landais et al. [2010] who account for firms' profits and labor-demand behavior more generally and show how the distinction between micro and macro elasticities becomes relevant for the characterization of the optimal (static) unemployment policy in general equilibrium.

The impact of a policy change on the agents' welfare equals

$$
\begin{aligned}
\int \frac{\partial V_i(P)}{\partial b_t} di &= \int S_{i,t} \frac{\partial v_i^u(c_{i,t}^u, s_{i,t})}{\partial c_{i,t}^u} di + \Sigma_{t'=1}^{\infty} [\int S_{i,t'} \lambda_i(s_{i,t'}) \frac{\partial \mu_{t'}}{\partial \mathbf{S}_{t'}} \frac{\partial \mathbf{S}_{t'}}{\partial b_t} \beta^{t'} \left[V_{i,t'}^e - V_{i,t'}^u\right] di] \\
&= \lambda S_t \{[1 + CS_t] + \Sigma_{t'=1}^{\infty} \frac{S_{t'}}{S_t} [\int \frac{S_{i,t'}}{S_{t'}} \frac{\partial \lambda_i(s_{i,t'})\mu_{t'}}{\partial \mathbf{S}_{t'}} \frac{\partial \mathbf{S}_{t'}}{\partial b_t} \beta^{t'} \left[V_{i,t'}^e - V_{i,t'}^u\right] / \lambda di]\} \\
&\equiv \lambda S_t \left\{ [1 + CS_t] + \Sigma_{t'=1}^{\infty} \frac{S_{t'}}{S_t} E_{t'}^u \left(\frac{\partial h_{t'}}{\partial b_t} \omega_{t'}^h \right) \right\}.
\end{aligned}
$$

We can correct the moral hazard cost for this new externality so that

$$\frac{\partial W(P)}{\partial b_t} = \lambda S_t \times [CS_t - MH_t^x],$$

with

$$MH_t^x \equiv \Sigma_{t'=1}^{T} \frac{S_{t'}}{S_t} \left[\frac{b_{t'} + \tau}{b_t} \varepsilon_{t',t} - E_{t'}^u \left(\omega_{t'}^h \frac{\partial h_{t'}}{\partial b_t} \right) \right],$$

The welfare impact of an increase in the exit rate rate is positive, $\omega_{i,t'}^h = \beta^{t'} \left[V_{i,t'}^e - V_{i,t'}^u\right] / \lambda$. The change in the exit rate $\partial h_{t'} / \partial b_t$ (through the change in hiring) will depend on the change in the relative survival rates at time t' in response to the change in benefits at time t. This corresponds to the correction proposed by Lehr [2017] for a flat benefit profile. A change in the unemployment policy won't change hiring, if the survival rate response of types with different productivity responds is the same. That is, $\varepsilon_{t',t}^i = \varepsilon_{t',t}^{i'}$.

Embedding this in our framework allows to assess the impact on the benefit profile as well. For the hiring externatity to change the gradient of the moral hazard costs, we need a change in the benefit level b_t to cause a different response in the relative survival rate (scaled by $1/S_t$) depending on the timing of the change. In our stylized two-type example the response in relative survival rate $S_{\bar{t}}^H / S_{\bar{t}}^L$ at the threshold duration $\bar{t}$ determines whether hiring increases or decreases. Indeed, the firm will hire job seekers with longer unemployment duration than $\bar{t}$ if the productive type is more responsive than the less productive type to a change in b_t, i.e., $\varepsilon_{\bar{t},t}^H > \varepsilon_{\bar{t},t}^L$. The externality response would be positive and thus causes $MH_t^x < MH_t$. In the model with heterogeneity in the returns to search, discussed in subsection 5.2.1 and considered below, we show that $\varepsilon_{\bar{t},t}^i$ can be increasing in the return to search for benefit levels paid early in the spell, but at the same decreasing for benefit levels paid late in the spell. Intuitively, the increase in the returns to search reduces the survival rate into longer unemployment spells and thus reduces the responsiveness to changes in benefits timed later on. Hence, with heterogeneity in the returns to search, the gradient of the moral hazard cost could become steeper when adjusted for the hiring externality.

A.2.4 Characterization with Income Taxation

We briefly illustrate the role of other fiscal externalities beyond the one introduced by the unemployment policy. In previous work on the Baily-Chetty formula, the only tax distortion in the economy comes from the unemployment policy. That is, no other revenue requirement exists ($\bar{G}=0$) and the government imposes a lump sum contribution τ on the employed to balance the UI expenditures. Our model allows for taxes to fund an additional revenue requirement $\bar{G}>0$. In practice, however, general government expenditures are funded through an income tax that is levied on both the employed and the unemployed.[64] Consider the case with a proportional income tax τ^y in addition to a lump sum UI contribution τ^u paid by employed workers. The (integrated) government budget can be rewritten as

$$G(P)-\bar{G}=[T-D](\tau^u+\tau^y w)-\Sigma S_t(b_t-\tau^y b_t)-\bar{G},$$

where

$$\begin{aligned}\frac{\partial G(P)}{\partial b_t} &= -S_t^r(1-\tau^y)-\Sigma_{t'=1}^T(b_{t'}-\tau^y b_{t'}+\tau^u+\tau^y w)\frac{\partial S_{t'}^r}{\partial b_t}\\ &= -S_t^r(1-\tau^y)-\Sigma_{t'=1}^T(b_{t'}+\tau^u)\frac{\partial S_{t'}^r}{\partial b_t}-\Sigma_{t'=1}^T\tau^y(w-b_{t'})\frac{\partial S_{t'}^r}{\partial b_t}\\ &= -D_k^r\left[(1-\tau^y)+\Sigma_{t'=1}^T\frac{S_{t'}^r}{S_t^r}\frac{b_{t'}+\tau^u}{b_t}\varepsilon_{t',t}^r+\Sigma_{t'=1}^T\frac{S_{t'}^r}{S_t^r}\frac{\tau^y(w-b_{t'})}{b_t}\varepsilon_{t',t}^r\right].\end{aligned}$$

The first two terms capture the standard mechanical and behavioral effect of an increase in the benefit level on the expenditures and revenues related to the unemployment policy. The third term captures the fiscal externality through the income tax, accounting for the reduction in income tax revenues when increasing unemployment. For a flat profile, this effect is proportional to $\tau^y\frac{w-\bar{b}}{\bar{b}}$ and thus small when the average effective income tax rate is small or the replacement rate is high. It is a standard simplification in related work to ignore these fiscal spillover effects across different government policies. Note also that from the consumption smoothing perspective, the difference in marginal utilities remains sufficient.

A.2.5 Welfare Impact of Change in Tilt

Corollary 1. *Whenever $\frac{CS_1}{MH_1}>\frac{CS_2}{MH_2}$, welfare can be increased by increasing the tilt b_1/b_2. A budget-balanced increase in the tilt b_1/b_2 increases welfare if and only if $\frac{1+CS_1}{1+MH_1}>\frac{1+CS_2}{1+MH_2}$.*

Proof: *By implicit differentiation, we find that when increasing b_1 and decreasing b_2 at rate*

$$\frac{db_2}{db_1}|_1=-\frac{D_1(1+MH_1)}{D_2(1+MH_2)}, \tag{24}$$

the policy budget remains balanced. The welfare impact of this budget-balanced increase in the tilt b_1/b_2 equals

$$\begin{aligned}\frac{\partial W(P)}{\partial b_1}-\frac{\partial W(P)}{\partial b_2}\frac{db_2}{db_1}|_1 &= \lambda D_1[1+CS_1-1-MH_1]-\lambda D_2[1+CS_2-1-MH_2]\frac{D_1(1+MH_1)}{D_2(1+MH_2)}\\ &= \lambda D_1(1+MH_1)\times\{\frac{1+CS_1}{1+MH_1}-\frac{1+CS_2}{1+MH_2}\}.\end{aligned}$$

This proves the second part of the corollary. Consider now an increase in b_1 jointly with a decrease in b_2 at rate

$$\frac{db_2}{db_1}|_2=-\frac{D_1MH_1}{D_2MH_2}$$

[64] In Sweden, UI benefits are fully included in individuals' taxable income to the personal income tax.

The welfare impact of such increase in the tilt b_1/b_2 *equals*

$$\frac{\partial W(P)}{\partial b_1} - \frac{\partial W(P)}{\partial b_2}\frac{db_2}{db_1}|_2 = \lambda D_1 [CS_1 - MH_1] - \lambda D_2 [CS_2 - MH_2] \frac{D_1 MH_1}{D_2 MH_2}$$
$$= \lambda D_1 MH_1 \times \left\{ \frac{CS_1}{MH_1} - \frac{CS_2}{MH_2} \right\}.$$

Hence, whenever CS_1/MH_1 *exceeds* CS_2/MH_2*, such increase in the tilt* b_1/b_2 *increases welfare and vice versa.*■

A.3 Dynamic Sufficient Statistics in Stationary Environment

A.3.1 Proof of Proposition 2

We consider a flat benefit profile $b_t = \bar{b} < w - \tau$ *for* $\forall t$ *in a single-type, stationary environment* $h_{i,t}(\cdot) = \bar{h}(\cdot)$ *for* $\forall i, t$*. We also assume* $\beta(1+r) = 1$ *and* $T = \infty$*. We compare the impact of an increase in the benefit level at time* t *and at time* $t+1$*.*

We analyze first the moral hazard costs. We assume that the agent is borrowing constrained and thus consumes hand-to-mouth when unemployed and employed ($c_t^u = b_t$ *and* $c_t^e = w - \tau$*). This set up follows Hopenhayn and Nicolini [1997]. Using notation* $S_t^r = (1+r)^{-(t-1)} S_t$*, we find*

$$\frac{\partial G(P)}{\partial b_t} = -S_t^r - \Sigma_{j=1}^T (b_j + \tau) \frac{\partial S_j^r}{\partial b_k}$$
$$= -S_t^r \times [1 + \frac{b+\tau}{b} \frac{D^r}{S_t^r} \varepsilon_{D^r, b_t}].$$

For an increase in b_{t+1}*, we find*

$$\frac{\partial G(P)}{\partial b_{t+1}} = -S_{t+1}^r \times [1 + \frac{b+\tau}{b} \frac{D^r}{S_{t+1}^r} \varepsilon_{D^r, b_{t+1}}].$$

Using

$$D^r = \Sigma_{j=1}^T S_j^r = 1 + D_2^r = 1 + S_2^r \tilde{D}_2^r,$$

where $D_2^r = \Sigma_{j=2}^T S_j^r$ *and* $\tilde{D}_2^r = \left[\Sigma_{j=2}^T S_j^r / S_2^r\right]$*, we can write*

$$\varepsilon_{D^r, b_{t+1}} = \frac{\partial [1 + D_2^r]}{\partial b_{t+1}} \frac{b}{D^r} = \frac{\partial D_2^r}{\partial b_{t+1}} \frac{b}{D_2^r} \frac{D_2^r}{D^r}$$
$$= \left[\varepsilon_{S_2^r, b_{t+1}} + \varepsilon_{\tilde{D}_2^r, b_{t+1}}\right] \frac{D_2^r}{D^r}.$$

Since the environment is stationary and the agent is borrowing-constrained, the agent's search behavior remains the same over the unemployment spell (conditional on the continuation policy being the same). Starting from a flat profile, an increase in b_t *has the same impact on the continuation policy evaluated at time 1 as an increase in* b_{t+1} *has on the continuation policy evaluated at time 2, conditional on being still unemployed then. The impact of the policy changes at time* t *and* $t+1$ *on the remaining duration at time 1 and time 2 respectively is the same. Hence, we have* $\varepsilon_{\tilde{D}_2^r, b_{t+1}} = \varepsilon_{D^r, b_t}$ *for* $T = \infty$*. Denoting the constant exit rate for the flat profile by* h*, we have* $D^r = \frac{1+r}{r+h}$ *and* $D_2^r = \frac{1-h}{1+r}\frac{1+r}{r+h}$*, while* $S_{t+1}^r = \frac{1-h}{1+r} S_t^r$*. This implies*

$$\frac{D_2^r}{S_{t+1}^r} = \frac{D^r}{S_t^r}.$$

Using this equality and the expression for $\varepsilon_{D^r,b_{t+1}}$, we can re-write

$$\begin{aligned}\frac{\partial G(P)}{\partial b_{t+1}} &= -S^r_{t+1}\times[1+\frac{b+\tau}{b}\frac{D^r}{S^r_{t+1}}\varepsilon_{D^r,b_{t+1}}] \\ &= -S^r_{t+1}\times[1+\frac{b+\tau}{b}\frac{D^r_{\mathbf{2}}}{S^r_{t+1}}\left[\varepsilon_{S^r_2,b_{t+1}}+\varepsilon_{D^r,b_t}\right]] \\ &= -S^r_{t+1}\times[1+\frac{b+\tau}{b}\frac{D^r}{S^r_t}\left[\varepsilon_{S^r_2,b_{t+1}}+\varepsilon_{D^r,b_t}\right]].\end{aligned}$$

This implies that

$$MH_{t+1}=\frac{b+\tau}{b}\frac{D^r}{S^r_t}\varepsilon_{S^r_2,b_{t+1}}+\frac{b+\tau}{b}\frac{D^r}{S^r_t}\varepsilon_{D^r,b_t}\geq MH_t,$$

since $\varepsilon_{S^r_2,b_{t+1}}\geq 0$. Starting from a flat profile, the moral hazard cost is thus higher for any benefit increase that is timed later during the spell.

We now analyze the consumption smoothing gains. In our stylized setup (which assumes that the agent starts unemployed and remains employed once she finds a job), an optimizing agent's unemployment consumption $c^u_t(\omega_t)$ (and search effort $s_t(\omega_t)$) only depends on the length of the ongoing unemployment spell. Hence, we have

$$\int\frac{\partial V_i(P)}{\partial b_t}di=\beta^{t-1}S_t\frac{\partial v^u(c^u_t,s_t)}{\partial c^u_t}.$$

When the agent is borrowing constrained, the agent is hand-to-mouth $c^u_t=b_t$ and the marginal utility of consumption (and thus CS_t) remains constant for a flat benefit profile. When not borrowing constrained, an agent who is unemployed at time t increases her consumption by depleting her assets to equalize the marginal utility of consumption at time t with the expected marginal utility of consumption at $t+1$. The unemployment consumption level c^u_t at time t, the consumption level upon finding employment c^e_{t+1} at time $t+1$ and the consumption level when still being unemployed c^u_{t+1} at time $t+1$ satisfy a standard Euler condition,

$$\frac{\partial v^u(c^u_t,s_t)}{\partial c^u_t}=h_t(s_t)\frac{\partial v^e(c^e_{t+1})}{\partial c^e_{t+1}}+(1-h_t(s_t))\frac{\partial v^u(c^u_{t+1},s_{t+1})}{\partial c^u_{t+1}}$$

for $\beta(1+r)=1$. With separable concave preferences, $\partial v^u(c,s)/\partial c=\partial v^e(c)/\partial c=v'(c)$,and, benefits lower than the after-tax wage $b<w-\tau$, for any given asset level, an agent has higher expected lifetime income when employed than when unemployed. The marginal value of an increase in assets is lower when employed than when unemployed, i.e., $\partial V^e_{t+1}/\partial a_{t+1}<\partial V^u_{t+1}/\partial a_{t+1}$. This implies the marginal utility of consumption is lower when employed than when unemployed at $t+1$. Hence, by the Euler condition, $v'(c^u_{t+1})>v'(c^e_{t+1})$ implies $v'(c^u_{t+1})>v'(c^u_t)$. On the optimal path, the marginal utility of consumption is increasing over the spell and the consumption gains are thus always higher for benefits timed later during the unemployment spell.■

The stationary forces and how they affect the optimal benefit profile are well known in the literature and arguably robust. Our set up with the borrowing-constrained agent follows Hopenhayn and Nicolini [1997]. The assumption that the agent is borrowing constrained is restrictive, but guarantees that search behaviour remains the same over the unemployment spell and thus simplifies the derivations. Note that search behaviour remains the same in a model with savings when the agent has CARA preferences with monetary cost of search efforts (i.e., $v^u(c,s)=-\exp(-\sigma[c-\psi(s)])$ as in Spinnewijn [2015], again simplifying the derivation of the optimal benefit profile. It is also clear from the proof that relaxing the borrowing constraint would not change the conclusion regarding the gradient of the moral hazard costs when $\varepsilon_{\tilde{D}^r_{\mathbf{2}},b_{t+1}}\geq\varepsilon_{D^r,b_t}$ and $h_{t+1}\geq h_t$ (so that $\frac{D^r_{\mathbf{2}}}{S^r_{t+1}}>\frac{D^r}{S^r_t}$) for any t. We analyze this further for a specific search environment with non-stationary features. The result that the marginal utility of consumption is increasing over the spell continues to hold for unconstrained job seekers when the benefit profile is not flat but $b_t<w-\tau$ for all t. The assumption that the agent's preferences are separable is also more restrictive than necessary. The proof highlights that it suffices for the marginal value of an increase in assets to be lower when employed than when unemployed.

A.4 Dynamic Sufficient Statistics in a Non-stationary Environment

We now specify particular functions for the search environment and introduce non-stationary features in our model. We allow for depreciation in search efficacy, heterogeneity in search efficacy, and heterogeneity in assets. We study how these forces affect the predicted increase in MH_t and CS_t throughout the unemployment spell from Proposition 2. We allow for all these non-stationary forces simultaneously in our structural model in Appendix D.

We establish three results: (i) in a model with depreciation in the return-to-search parameter (i.e., $h_t(s_{i,t}) = h_0 + \theta^t h_1 s_{i,t}^{\rho}$), the moral hazard cost cost of benefit changes that start later in the spell can be arbitrarily close to the moral hazard cost of benefit changes that start earlier, (ii) in a model with heterogeneity in the return-to-search parameter (i.e., $h_t(s_{i,t}) = h_0 + h_1^i s_{i,t}^{\rho}$), the moral hazard costs can actually be lower for benefit changes timed later in the spell,, (iii) in a model with asset heterogeneity, negatively correlated with exit rates, the consumption smoothing gains can actually be higher earlier in the spell.

A.4.1 Preliminaries

For our analysis of moral hazard costs, we assume that agents are borrowing constrained throughout the unemployment spell (i.e., unemployment consumption equals UI benefits), that preferences are separable in consumption and search, $u(c,s) = u(c) - s$, and that the exit rate function has the following form,

$$h(s_{i,t}) = h_0 + h_1 s_{i,t}^{\rho} \text{ for } \forall i, t.$$

For tractability, we assume that the optimal search effort is interior and thus the resulting exit rate is between 0 and 1.

Each individual has a value function for the employed and unemployed state shown below:

$$\begin{aligned} V_{i,t}^e &= u(w-\tau) + \beta V_{i,t+1}^e \\ V_{i,t}^u &= u(b_t) - s_{i,t} + \beta h_{i,t}(s_{i,t}) [V_{i,t+1}^e - V_{i,t+1}^u] + \beta V_{i,t+1}^u, \end{aligned}$$

Since the employment state is absorbing, we have $V_t^e = \frac{u(w-\tau)}{1-\beta}$.The optimal level of effort equals

$$s_{i,t} = (\rho\beta h_{1,i,t}[V_{i,t+1}^e - V_{i,t+1}^u])^{\frac{1}{1-\rho}}.$$

We start from a flat benefit profile and compare a permanent benefit rise in $t = 2$ (denoted by $b_{2\to\infty}$) and in $t = 3$ (denoted by $b_{3\to\infty}$) respectively. Note that $t = 1$ is the first period that an agent exerts effort, but this is unaffected by the benefit level b_1. The moral hazard cost of raising benefits permanently in period t, starting from a flat profile, is given by:

$$MH_{t\to\infty} = \frac{\frac{\partial D^r}{\partial b_{t\to\infty}}}{D_{t\to\infty}^r}(b+\tau).$$

To save on notation we consider instead $D = \sum_{t'=1}^{\infty} S_{t'}$ and $D_{t\to\infty} = \sum_{t'=t}^{\infty} S_{t'}$, corresponding to D^r and $D_{t\to\infty}^r$ for $r = 0$, but we make sure our conclusions are robust to $\beta \to 1$.

A.4.2 Stationary Environment

We first confirm that for this specific search environment, in the absence of non-stationary features, the moral hazard cost of increasing the UI benefits is always higher when this increase is timed later in the spell, in line with Proposition 2. This will help highlighting why non-stationary features can affect this result.

Consider first an increase in $b_{2\to\infty}$. In this scenario, we have that $V_t^u = V_{t+1}^u = V_2^u$ if $t \geq 2$. As a result we have only two value functions when unemployed:

$$V_1^U = u(b) - s + \beta(h_0 + h_1 s^\rho)[V^e - V_2^u] + \beta V_2^u$$
$$V_2^u = u(b_{2\to\infty}) - s + \beta(h_0 + h_1 s^\rho)[V^e - V_2^u] + \beta V_2^u$$

and one level of effort:

$$s = (\rho\beta h_1[V^e - V_2^u])^{\frac{1}{1-\rho}}.$$

Hence, we can write

$$S_t = (1 - h_0 - h_1 s^\rho)^{t-1},$$
$$D = \frac{1}{h_0 + h_1 s^\rho}$$
$$D_{2\to\infty} = \frac{1 - h_0 - h_1 s^\rho}{h_0 + h_1 s^\rho}$$

Since we evaluate the benefit change for a flat profile, we will use the fact that before differentation $V_1^u = V_2^u = V^u$. We calculate the effect of the benefit rise on the average unemployment duration, which in turn depends on the change in effort, which in turn depends on the change in the value of being unemployed:

$$\frac{\partial D}{\partial b_{2\to\infty}} = -\frac{\rho h_1 s^{\rho-1} \frac{\partial s}{\partial b_{2\to\infty}}}{(h_0 + h_1 s^\rho)^2},$$
$$\frac{\partial s}{\partial b_{2\to\infty}} = -\frac{s}{1-\rho}[V^e - V^u]^{-1} \frac{\partial V_2^u}{\partial b_{2\to\infty}},$$
$$\frac{\partial V_2^u}{\partial b_{2\to\infty}} = \frac{u'(\cdot)}{1 - \beta(1 - h_0 - h_1 s^\rho)}.$$

Consider now an increase in $b_{3\to\infty}$. Note that $V_t^u = V_{t+1}^u = V_3^U$ if $t \geq 3$. Therefore, there are only three value functions when unemployed:

$$V_1^u = u(b) - s_1 + \beta(h_0 + h_1 s_1^\rho)[V^e - V_2^u] + \beta V_2^u$$
$$V_2^u = u(b) - s_2 + \beta(h_0 + h_1 s_2^\rho)[V^e - V_3^u] + \beta V_3^u$$
$$(1-\beta)V_3^U = u(b_{3\to\infty}) - s_2 + \beta(h_0 + h_1 s_2^\rho)[V^e - V_3^u],$$

and two levels of effort:

$$s_1 = (\rho\beta h_1[V^e - V_2^u])^{\frac{1}{1-\rho}}$$
$$s_2 = (\rho\beta h_1[V^e - V_3^u])^{\frac{1}{1-\rho}}.$$

Similar to before, we find

$$\frac{\partial D}{\partial b_{3\to\infty}} = \frac{-\rho h_1 s^{\rho-1} \frac{\partial s_1}{\partial b_{3\to\infty}}(h_0 + h_1 s^\rho) - \rho h_1 s^{\rho-1} \frac{\partial s_2}{\partial b_{3\to\infty}}(1 - h_0 - h_1 s^\rho)}{(h_0 + h_1 s^\rho)^2},$$

which is composed of the following derivatives:

$$\frac{\partial s_1}{\partial b_{3\to\infty}} = -\frac{s}{1-\rho}[V^e - V^u]^{-1} \frac{\partial V_2^u}{\partial b_{3\to\infty}}$$
$$\frac{\partial s_2}{\partial b_{3\to\infty}} = -\frac{s}{1-\rho}[V^e - V^u]^{-1} \frac{\partial V_3^u}{\partial b_{3\to\infty}},$$

which are, in turn, composed of the following derivatives:

$$
\begin{aligned}
\frac{\partial V_3^U}{\partial b_{3\to\infty}} &= \frac{u'(\cdot)}{1-\beta(1-h_0-h_1 s^\rho)}, \\
\frac{\partial V_2^U}{\partial b_{3\to\infty}} &= -\beta(h_0+h_1 s^\rho)\frac{\partial V_3^u}{\partial b_{3\to\infty}} + \beta\frac{\partial V_3^u}{\partial b_{3\to\infty}}.
\end{aligned}
$$

Putting everything together, we find for $b_{2\to\infty}$ and $b_{3\to\infty}$:

$$
\frac{\partial D}{\partial b_{3\to\infty}} = \frac{\partial D}{\partial b_{2\to\infty}}(1-h_0-h_1 s^\rho)[1+\beta(h_0+h_1 s^\rho)],
$$

while

$$
D_{3\to\infty} = D_{2\to\infty}(1-h_0-h_1 s^\rho).
$$

The impact of an increase in $b_{3\to\infty}$ on the time spent unemployed is smaller than the impact of an increase in $b_{2\to\infty}$, since a smaller part of the unemployment policy is affected, as captured by the scalar $[1-h_0-h_1 s^\rho]$ in the both expressions above, for the duration responses and the durations respectively. However, while the increase in $b_{3\to\infty}$ starts later, it will reduce the exit rates earlier in the spell as well, as captured by the scalar $[1+\beta(h_0+h_1 s^\rho)]$ in the expression for the duration responses. These are the forward-looking incentives identified before in Shavell and Weiss [1979]. Indeed, in line with Proposition 2, we find

$$
\frac{MH_{2\to\infty}}{MH_{3\to\infty}} = \frac{1}{1+\beta(h_0+h_1 s^\rho)} < 1.
$$

This intuition generalizes for any changes $b_{t\to\infty}$ and $b_{t+1\to\infty}$ respectively and is robust to $\beta \to 1$.

A.4.3 Depreciation in Search Efficacy

We now assume that the returns to search depreciate at a geometric rate,

$$
h_t\left(s_{i,t}\right) = h_0 + \theta^{t-1} h_1 s_{i,t}^\rho = \theta^{t-1} h s_{i,t}^\rho \text{ for } \theta \in [0,1].
$$

To simplify the expressions below, we assume that the exit rate is zero when no search is exerted (i.e., $h_0 = 0$), but the argument below continues to apply when relaxing this assumption. From the value of being unemployed at time t and the effort level at time t, we can derive the following derivatives:

$$
\begin{aligned}
\frac{\partial V_t^u}{\partial b_{\bar{t}\to\infty}} &= \beta(1-h\theta^{t-1}s_t^\rho)\frac{\partial V_{t+1}^u}{\partial b_{\bar{t}\to\infty}} \;\forall\; 0<t<\bar{t} \\
\frac{\partial V_t^u}{\partial b_{\bar{t}\to\infty}} &= u'\left(b\right) + \beta(1-h\theta^{t-1}s_t^\rho)\frac{\partial V_{t+1}^u}{\partial b_{\bar{t}\to\infty}} \;\forall\; t\geq\bar{t} \\
\frac{\partial s_t}{\partial b_{\bar{t}\to\infty}} &= \frac{-s_t}{1-\rho}[V^e - V_{t+1}^u]^{-1}\frac{\partial V_{t+1}^u}{\partial b_{\bar{t}\to\infty}},
\end{aligned}
$$

which can be used in order to derive an expression for the derivative of the average unemployment duration:

$$
\begin{aligned}
\frac{\partial D}{\partial b_{\bar{t}\to\infty}} &= \frac{\partial}{\partial b_{\bar{t}\to\infty}}[1+(1-hs_1^\rho)+(1-hs_1^\rho)(1-h\theta s_2^\rho)+(1-hs_1^\rho)(1-h\theta s_2^\rho)(1-h\theta^2 s_3^\rho)+\ldots] \\
&= -\rho h\left[s_1^{\rho-1}\frac{\partial s_1}{\partial b_{\bar{t}\to\infty}}\frac{D_{2\to\infty}}{1-hs_1^\rho} + \theta s_2^{\rho-1}\frac{\partial s_2}{\partial b_{\bar{t}\to\infty}}\frac{D_{3\to\infty}}{1-h\theta s_2^\rho}+\ldots\right] \\
&= -\rho h\sum_{t'=1}^{\infty}\theta^{t'-1}s_{t'}^{\rho-1}\frac{D_{t'+1\to\infty}}{1-h\theta^{t'-1}s_{t'}^\rho}\frac{\partial s_{t'}}{\partial b_{\bar{t}\to\infty}} \\
&= \frac{\rho}{1-\rho}h\sum_{t'=1}^{\infty}\theta^{t'-1}s_{t'}^\rho[V^e - V_{i+1}^u]^{-1}\frac{D_{t'+1\to\infty}}{1-h\theta^{t'-1}s_{t'}^\rho}\frac{\partial V_{t'+1}^u}{\partial b_{\bar{t}'\to\infty}}
\end{aligned}
$$

We then use the following feature:

$$\frac{\partial V_t^u}{\partial b_{\bar{t}\to\infty}} = \frac{\partial V_t^u}{\partial b_{\hat{t}\to\infty}} \ \forall\ t \geq \max[\bar{t}, \hat{t}]$$

and

$$\begin{aligned}\frac{\partial V_2^u}{\partial b_{2\to\infty}} &= u'(\cdot) + \beta(1 - h\theta s_2^\rho)\frac{\partial V_3^U}{\partial b_{2\to\infty}} \\ &= u'(\cdot) + \frac{\partial V_2^U}{\partial b_{3\to\infty}}\end{aligned}$$

to re-express the ratio of duration responses as

$$\begin{aligned}\frac{\frac{\partial D}{\partial b_{2\to\infty}}}{\frac{\partial D}{\partial b_{3\to\infty}}} &= \frac{s_1^\rho[V^e - V_2^u]^{-1}\frac{D_{2\to\infty}}{1-hs_1^\rho}\frac{\partial V_2^u}{\partial b_{2\to\infty}} + \sum_{t'=2}^{\infty}\theta^{t'-1}s_{t'}^\rho[V^e - V_{t'+1}^u]^{-1}\frac{D_{t'+1\to\infty}}{1-h\theta^{t'-1}s_{t'}^\rho}\frac{\partial V_{t'+1}^u}{\partial b_{2\to\infty}}}{s_1^\rho[V^e - V_2^u]^{-1}\frac{D_{2\to\infty}}{1-hs_1^\rho}\frac{\partial V_2^u}{\partial b_{3\to\infty}} + \sum_{t'=2}^{\infty}\theta^{t'-1}s_{t'}^\rho[V^e - V_{t'+1}^u]^{-1}\frac{D_{t'+1\to\infty}}{1-h\theta^{t'-1}s_{t'}^\rho}\frac{\partial V_{t'+1}^u}{\partial b_{3\to\infty}}} \\ &= \frac{u'(\cdot) + \frac{\partial V_2^u}{\partial b_{3\to\infty}} + A}{\frac{\partial V_2^u}{\partial b_{3\to\infty}} + A},\end{aligned}$$

where

$$A = \sum_{t'=2}^{\infty}\theta^{t'-1}\frac{s_{t'}^\rho}{s_1^\rho}\frac{V^e - V_2^u}{V^e - V_{t'+1}^U}\frac{D_{t'+1\to\infty}}{D_{2\to\infty}}\frac{1-hs_1^\rho}{1-h\theta^{t'-1}s_{t'}^\rho}\frac{\partial V_{t'+1}^u}{\partial b_{j\to\infty}} \text{ for } j = 2, 3,$$

capturing the response in exit rates from time 2 onwards. Hence, we can find an explicit expression for the ratio of moral hazard costs,

$$\frac{MH_{2\to\infty}}{MH_{3\to\infty}} = \frac{u'(\cdot) + \frac{\partial V_2^u}{\partial b_{3\to\infty}} + A}{\frac{\partial V_2^u}{\partial b_{3\to\infty}} + A}\frac{D_{3\to\infty}}{D_{2\to\infty}}.$$

Note that when we set $\theta = 1$, we return to the stationary model and find that $MH_{2\to\infty}/MH_{3\to\infty} < 1$. However, in this non-stationary setting, the ratio depends crucially on θ and can be made arbitrarily close to 1 for sufficiently low θ. That is, using $D_{3\to\infty} = D_{2\to\infty} - (1 - hs_1^\rho)$,

$$\begin{aligned}MH_{2\to\infty}/MH_{3\to\infty} &\approx 1 \Leftrightarrow \\ u'(\cdot)[D_{2\to\infty} - (1 - hs_1^\rho)] &\approx (1 - hs_1^\rho)\left[\frac{\partial V_2^u}{\partial b_{3\to\infty}} + A\right] \Leftrightarrow \\ u'(\cdot)D_{3\to\infty} &\approx (1 - hs_1^\rho)\left[\beta(1 - h\theta s_2^\rho)\frac{\partial V_3^u}{\partial b_{3\to\infty}} + A\right] \Leftrightarrow \\ u'(\cdot)D_{3\to\infty} &\approx (1 - hs_1^\rho)\left[\beta(1 - h\theta s_2^\rho)\left(u'(\cdot) + \beta(1 - h\theta^2 s_3^\rho)\frac{\partial V_4^U}{\partial b_{3\to\infty}}\right) + A\right]\end{aligned}$$

Iterating the substitution of $\frac{\partial V_t^u}{\partial b_{3\to\infty}} = u'(\cdot) + \beta(1 - h\theta s_t^\rho)\frac{\partial V_{t+1}^U}{\partial b_{3\to\infty}}$, we find

$$u'(\cdot)D_{3\to\infty} \approx \left[\beta(1 - hs_1^\rho)(1 - h\theta s_2^\rho)u'(\cdot) + \beta^2(1 - hs_1^\rho)(1 - h\theta s_2^\rho)(1 - h\theta^2 s_3^\rho)u'(\cdot) + \ldots\right] + (1 - hs_1^\rho)A.$$

The bracketed term in the RHS converges to $u'(\cdot)D_{3\to\infty}$ for $\beta \to 1$. (In particular, when properly discounting the survival rates to calculate the moral hazard costs, the two terms would coincide for $\beta\,(1 + r) = 1$.) This shows the importance of the term A, determined by the exit rate responses later in the spell, in driving the wedge between the moral hazard costs. This wedge is still positive, like in the stationary model, since $A > 0$. However, the wedge disappears when A converges to 0. Now by setting θ arbitrarily small we can make A arbitrarily small, since the terms in the summation are scaled by $\theta^{t'-1}$ and thus converge to 0, while all other factors can be bounded from above. Hence, for small enough θ, we have that $MH_{2\to\infty} \approx MH_{3\to\infty}$.

The intuition underlying this result is that depreciation in the returns to search reduces the responsiveness in the exit rates later in the unemployment spell to changes in the UI benefits. While this force cannot reverse the relative magnitude of the moral hazard cost, it mitigates the weight on the forward-looking incentives in driving this wedge. We now turn to a case where the relative magnitudes can actually be reversed.

A.4.4 Heterogeneity in Search Efficacy

We now consider heterogeneity in search efficacy, allowing for two types of agents, type y and type z. Type-y agents have higher return to their search effort,

$$h_1^y > h_1^z.$$

The proportion of y-types at the start of the unemployment spell equals α.

Our approach is different from the stationary case and the case with search depreciation in which we derived an explicit expression for $(\partial D/\partial b_{t\to\infty})/D_{t\to\infty}$. Instead we follow the approach in the proof of Proposition 2 and Section 5.2. We decompose the moral hazard cost of raising benefits in period 3 permanently into the response to forward-looking incentives and the response in the *remaining duration* of unemployment, conditional on still being unemployed in period 3,

$$MH_{3\to\infty} \times \frac{b}{b+\tau} = \frac{D_{1\to 2}}{D_{3\to\infty}} \varepsilon_{D_{1\to 2},b_{3\to\infty}} + \varepsilon_{S_3,b_{3\to\infty}} + \varepsilon_{\tilde{D}_{3\to\infty},b_{3\to\infty}}, \tag{25}$$

where $\tilde{D}_{3\to\infty} \equiv \frac{D_{3\to\infty}}{S_3}$ and $D_{1\to 2} \equiv S_1 + S_2$. In a single-agent model without heterogeneity, the latter response corresponds to the moral hazard cost of an overall increase in benefits, $\varepsilon_{\tilde{D}_{3\to\infty},b_{3\to\infty}} = \varepsilon_{D,b_{1\to\infty}}$. With heterogeneity, the magnitude and the weights attached to the different elasticities depend on the different y- and z-types and their respective survival.

We first show that, for a given type, all three terms in (25) are increasing in search efficacy h_1. Define the following common component amongst all three terms

$$B = \frac{\rho}{1-\rho} \frac{u'(b)}{(1-\beta(1-h_0-h_1 s^\rho))} \frac{b}{[V^e - V^u]},$$

We then have

$$\begin{aligned}
\varepsilon_{\tilde{D}_{3\to\infty},b_{3\to\infty}} &= \varepsilon_{D,b_{1\to\infty}} = B \frac{h_1 s^\rho}{h_0 + h_1 s^\rho} \\
\varepsilon_{S_3,b_{3\to\infty}} &= B \frac{h_1 s^\rho [1 + \beta(1-h_0-h_1 s^\rho)]}{1-h_0-h_1 s^\rho} \\
\frac{D_{1\to 2}}{D_{3\to\infty}} \varepsilon_{D_{1\to 2},b_{3\to\infty}} &= B \frac{h_1 s^\rho \beta (h_0 + h_1 s^\rho)}{1-h_0-h_1 s^\rho}.
\end{aligned}$$

For tractability, we continue under the assumption that

$$\frac{\partial B}{\partial h_1} = -B \cdot \frac{\rho}{1-\rho} \frac{\beta s^\rho (1-\beta)}{(1-\beta(1-h_0-h_1 s^\rho))^2} \approx 0,$$

which follows from $1-\beta \approx 0$. We also have

$$\begin{aligned}
\frac{\partial [h_0 + h_1 s^\rho]}{\partial h_1} &= s^\rho + h_1 \rho s^{\rho-1} \frac{\partial s}{\partial h_1} \\
&= s^\rho + \frac{\rho(1-\beta)s^\rho}{(1-\rho)(1-\beta(1-h_0-h_1 s^\rho))} \approx s^\rho,
\end{aligned}$$

using again $1-\beta \approx 0$. As a consequence, all the above terms are increasing in search efficacy h_1.

The above elasticities are derived for a given type. This aggregrates up as follows:

$$\varepsilon_{\bar{D},b_{1\to\infty}} = \alpha\frac{D^y}{\bar{D}}\varepsilon_{D^y,b_{1\to\infty}} + (1-\alpha)\frac{D^z}{\bar{D}}\varepsilon_{D^z,b_{1\to\infty}},$$

$$\varepsilon_{\bar{S}_3,b_{3\to\infty}} = \alpha\frac{D^y_{3\to\infty}}{\bar{D}_{3\to\infty}}\varepsilon_{S^y_3,b_{3\to\infty}} + (1-\alpha)\frac{D^z_{3\to\infty}}{\bar{D}_{3\to\infty}}\varepsilon_{S^z_3,b_{3\to\infty}},$$

$$\frac{\bar{D}_{1\to 2}}{\bar{D}_{3\to\infty}}\varepsilon_{\bar{D}_{1\to 2},b_{3\to\infty}} = \alpha\frac{D^y_{1\to 2}}{\bar{D}_{3\to\infty}}\varepsilon_{D^y_{1\to\infty},b_{3\to\infty}} + (1-\alpha)\frac{D^z_{1\to 2}}{\bar{D}_{3\to\infty}}\varepsilon_{D^z_{1\to\infty},b_{3\to\infty}},$$

To emphasize the difference, we have introduced the upper-bar notation to refer to aggregates. We now wish to show that $MH_{1\to\infty} > MH_{3\to\infty}$ is true in the presence of sufficient heterogeneity. This is equivalent to:

$$\varepsilon_{\bar{D},b_{1\to\infty}} > \frac{\bar{D}_{1\to 2}}{\bar{D}_{3\to\infty}}\varepsilon_{\bar{D}_{1\to 2},b_{3\to\infty}} + \varepsilon_{\bar{S}_3,b_{3\to\infty}} + \varepsilon_{\bar{D}_{3\to\infty},b_{3\to\infty}}.$$

Substituting for the aggregate elasticities and re-arranging, we find

$$\alpha\left[\frac{D^y}{\bar{D}} - \frac{D^y_{3\to\infty}}{\bar{D}_{3\to\infty}}\right]\varepsilon_{D^y,b} + (1-\alpha)\left[\frac{D^z}{\bar{D}} - \frac{D^z_{3\to\infty}}{\bar{D}_{3\to\infty}}\right]\varepsilon_{D^z,b} >$$
$$\alpha\frac{D^y_{1\to 2}}{\bar{D}_{3\to\infty}}\varepsilon_{D^y_{1\to 2},b_{3\to\infty}} + (1-\alpha)\frac{D^z_{1\to 2}}{\bar{D}_{3\to\infty}}\varepsilon_{D^z_{1\to 2},b_{3\to\infty}} + \alpha\frac{D^y_{3\to\infty}}{\bar{D}_{3\to\infty}}\varepsilon_{S^y_3,b_{3\to\infty}} + (1-\alpha)\frac{D^z_{3\to\infty}}{\bar{D}_{3\to\infty}}\varepsilon_{S^z_3,b_{3\to\infty}}.$$

Using

$$\alpha\left[\frac{D^y}{\bar{D}} - \frac{D^y_{3\to\infty}}{\bar{D}_{3\to\infty}}\right] + (1-\alpha)\left[\frac{D^z}{\bar{D}} - \frac{D^z_{3\to\infty}}{\bar{D}_{3\to\infty}}\right] = 0,$$

we can re-write the inequality as

$$\alpha\left[\frac{D^y}{\bar{D}} - \frac{D^y_{3\to\infty}}{\bar{D}_{3\to\infty}}\right]\left[\varepsilon_{D^y,b} - \varepsilon_{D^z,b}\right] >$$
$$\alpha\frac{D^y_{3\to\infty}}{\bar{D}_{3\to\infty}}\left[\frac{D^y_{1\to 2}}{D^y_{3\to\infty}}\varepsilon_{D^y_{1\to 2},b_{3\to\infty}} + \varepsilon_{S^y_3,b_{3\to\infty}}\right] + (1-\alpha)\frac{D^z_{3\to\infty}}{\bar{D}_{3\to\infty}}\left[\frac{D^z_{1\to 2}}{D^z_{3\to\infty}}\varepsilon_{D^z_{1\to 2},b_{3\to\infty}} + \varepsilon_{S^z_3,b_{3\to\infty}}\right].$$

At this point we can see the mechanism at work. The LHS of the inequality can be made larger by increasing h^y_1 relative to h^z_1. The y-type agents are more responsive to changes in benefits (i.e., $\partial\varepsilon_{D^y,b}/\partial h^y_1 > 0$) and spend relatively less time unemployed later in the spell (i.e., $\partial(D^y/D^y_{3\to\infty})/\partial h^y_1 > 0$). At the same time, we can make the right-hand side arbitrarily small by increasing the heterogeneity. As we increase h^y_1 and decrease h^z_1, the forward looking elasticities of the y-types increase while the same elasticities decrease for the z-types. However, increasingly little weight (converging to zero for $h_0 + h^y_1 s^\rho_y \to 1$) gets placed on the y-types' elasticity. More weight gets placed on the forward-looking elasticities of the z-types, but these are low and converge to zero for $h^z_1 \to 0$.

Hence, with sufficient heterogeneity, we have that $MH_{1\to\infty} > MH_{3\to\infty}$.

A.4.5 Heterogeneity in Assets

Having introduced heterogeneity in exit rates, it is straightforward to reverse the prediction on the gradient of the consumption smoothing gains as well. This requires individuals with lower marginal utility of consumption to select into longer unemployment spells in a way that the dynamic selection offsets the increase in marginal utility for a given individual due to the depletion of assets. This can be obtained for example by heterogeneity in assets where an agent's asset holdings are negatively correlated with her exit rate. The same argument applies with heterogeneity in preferences.

To illustrate this, consider our two-type setup where type y is borrowing constrained and has high exit rate h^y - potentially induced by the constrained consumption when unemployed - and type z who has access to assets and low exit rate $h^z < h^y$. To obtain $CS_t > CS_{t'}$, we need

$$\alpha \frac{S_t^y}{S_t} u'\left(c_t^y\right) + (1-\alpha) \frac{S_t^z}{S_t} u'\left(c_t^z\right) > \alpha \frac{S_{t+1}^y}{S_{t+1}} u'\left(c_{t+1}^y\right) + (1-\alpha) \frac{S_{t+1}^z}{S_{t+1ds}} u'\left(c_{t+1}^z\right)$$

$\Leftrightarrow$

$$\alpha \left[\frac{S_t^y}{S_t} - \frac{S_{t+1}^y}{S_{t+1}}\right] u'(b) + (1-\alpha) \left[\frac{S_t^z}{S_t} - \frac{S_{t+1}^z}{S_{t+1}}\right] u'\left(c_t^z\right) + (1-\alpha) \frac{S_{t+1}^z}{S_{t+1}} \left[1 - \frac{u'\left(c_{t+1}^z\right)}{u'\left(c_t^z\right)}\right] u'\left(c_t^z\right) > 0$$

$\Leftrightarrow$

$$\alpha \left[\frac{S_t^y}{S_t} - \frac{S_{t+1}^y}{S_{t+1}}\right] \left[u'(b) - u'\left(c_t^z\right)\right] + (1-\alpha) \frac{S_{t+1}^z}{S_{t+1}} \left[u'\left(c_t^z\right) - u'\left(c_{t+1}^z\right)\right] > 0$$

Now notice that with $h_1^y > h_1^z$, the relative survival rate of agents of type y is decreasing over the spell,

$$\begin{aligned}
\frac{S_t^y}{S_t} - \frac{S_{t+1}^y}{S_{t+1}} &= \frac{(1-h^y)^{t-1}}{\alpha (1-h^y)^{t-1} + (1-\alpha)(1-h^z)^{t-1}} - \frac{(1-h^y)^t}{\alpha(1-h^y)^t + (1-\alpha)(1-h^z)^t} \\
&= \frac{\left[\frac{S_{t+1}}{S_t} - (1-h^y)\right](1-h^y)^{t-1}}{\alpha(1-h^y)^t + (1-\alpha)(1-h^z)^t} > 0
\end{aligned}$$

since $S_{t+1}/S_t > (1-h^y)$. Moreover, the difference in marginal utility of consumption is positive, $u'(b) - u'\left(c_t^z\right) > 0$, and more so the higher the asset level of agents of type z. Hence, with sufficient heterogeneity in the exit rates and the asset levels, the selection effect can offset the increase in marginal utility for agents of type z, $u'\left(c_t^z\right) - u'\left(c_{t+1}^z\right) < 0$, and as such make the consumption smoothing gains higher for benefits timed earlier in the spell.

A.4.6 Relative Survival Rate Response

We use the model with heterogeneity in search efficacy to show that the relative survival rate in a two-type model can increase in response to benefits paid early in the spell and decrease in response to benefits paid later in the spell. Embedding this in a model with employer screening considered in would imply that the gradient of the moral hazard cost could become steeper when adjusted for the hiring externality.

In particular, we are interested in

$$\frac{\partial h_{t'}}{\partial b_t}/S_t \propto \frac{\partial \left[S_{t'}^H / S_{t'}^L\right]}{\partial b_t}/S_t = S_{t'}^H / S_{t'}^L \frac{\frac{\partial S_{t'}^H}{\partial b_t}/S_{t'}^H - \frac{\partial S_{t'}^L}{\partial b_t}/S_{t'}^H}{S_t}.$$

We now analyze how $\frac{\partial S_{t'}^i}{\partial b_t}/S_{t'}^i$ changes when increasing search efficacy $h_{1,i}$. For simplicity, we assume $h_0 = 0$ and denote $h_{1,i} = h_i$. Using similar steps as before and starting again from a flat profile, we find

$$\frac{\partial S_{t'}^i}{\partial b_t}/S_t^i = \frac{h_i s_i^\rho}{\beta} \sum_{j=0}^{t'} [\beta(1-h_i s_i^\rho)]^{t-j-2} \frac{B}{b}$$

for $t > t'$. Like before, assuming that $\partial B/\partial h_i \approx 0$ and $\partial h_i s_i^\rho / \partial h_i \approx s_i^\rho$, which follows from $1-\beta \approx 0$, we find

$$\frac{\partial}{\partial h_i}\left[\frac{\partial S_{t'}^i}{\partial b_t}/S_t^i\right] \approx \left[s^\rho \sum_{j=0}^{t'} [\beta(1-hs^\rho)]^{t-j-2} - s^\rho \frac{hs^\rho}{1-hs^\rho} \sum_{j=0}^{t'} (t-j-2)[\beta(1-hs^\rho)]^{t-j-2}\right] B/b$$

We now want to see if this term can be negative for high enough t, but positive for low enough t. First,

$$\begin{aligned}
\frac{\partial}{\partial h_i}\left[\frac{\partial S_{t'}^i}{\partial b_t}/S_t^i\right] &< \left[x^\rho \sum_{j=0}^{t'} [\beta(1-hs^\rho)]^{t-j-2} - x^\rho \frac{hs^\rho}{1-hs^\rho}(t-t'-2) \sum_{j=0}^{t'} [\beta(1-hs^\rho)]^{t-j-2}\right] B/b \\
&= 1 - \frac{hs^\rho}{1-hs^\rho}(t-t'-2) < 0.
\end{aligned}$$

The last inequality holds for high enough t (provided $hs^\rho > 0$). Second,

$$\begin{aligned}\frac{\partial}{\partial h_i}\left[\frac{\partial S^i_{t'}}{\partial b_t}/S^i_t\right] &> \left[s^\rho \sum_{j=0}^{t'} [\beta(1-hs^\rho)]^{t-j-2} - s^\rho \frac{hs^\rho}{1-hs^\rho}(t-2)\sum_{j=0}^{t'}[\beta(1-hs^\rho)]^{t-j-2}\right] B/b \\ &= 1 - \frac{hs^\rho}{1-hs^\rho}(t-2) > 0.\end{aligned}$$

The last inequality now holds for low enough t (provided hs^ρ is small).

FOR ONLINE PUBLICATION - Appendix B: Additional results and robustness of the RK design

This Appendix presents additional results on the duration responses to benefits and various robustness checks of the RK design.

B.1 Additional Results: Hazard Rate Responses

To further investigate the non-stationary patterns in unemployment responses, Figure B-1 reports the RKD estimates of the effect of UI benefits on the hazard rates out of unemployment.

Since hazard rates are quite noisy at very high frequency, we have defined hazard rates by periods of 5 weeks. Blue dots represent the marginal effect of a change in both b_1 and b_2, estimated in the regression kink design for spells starting between 1999 and July 2001. Red dots represent the marginal effect of a change in b_2 only, estimated in the regression kink design for spells starting between July 2001 and July 2002. In both cases, 95% confidence interval around the point estimates, from robust standard errors, are displayed. The figure conveys quite clearly a series of interesting findings.

First, the graph shows that the effect of UI benefits is mostly concentrated in the first 10 to 15 weeks. After 15 weeks, the effect of UI benefits on the hazard rate is small and almost always insignificant.

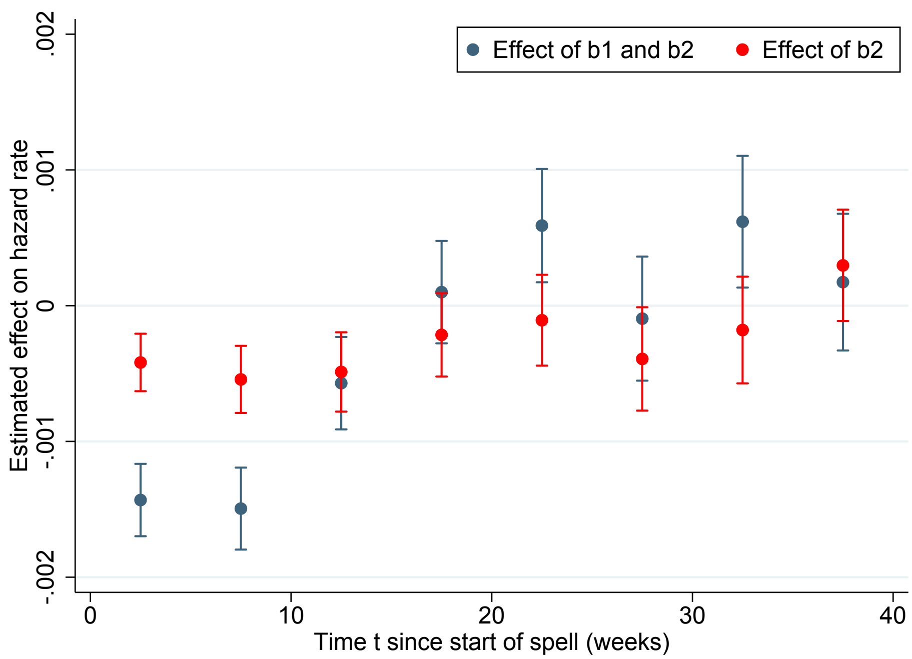
Second, the graph shows that b_2 (benefits received after 20 weeks) do have an effect on the hazard rate in the first 10 weeks. This confirms that unemployed individuals are forward-looking. b_2 does have a somewhat negative effect on contemporaneous hazard rates (after 20 weeks), but this effect is small and almost always insignificant.

The effect of b_1 can easily be inferred as it is the difference, for each hazard rate, between the effect of b_1 and b_2, and the effect of b_2 only. From the figure, we can easily see that the effect of b_1 is almost twice as large as the effect of b_2 early on in the spell. Because hazard rates are very responsive to b_1 in the spell, b_1 is having a large effect on the probability to survive into unemployment after 20 weeks. This creates a large mechanical effect of b_1 on D_2, the average duration spent in the second part of the benefit profile.

The total effect of b_1 on D_2 is the sum of the mechanical effect on survival plus the effect of b_1 on hazard rates after 20 weeks. Interestingly, the figure shows that the latter effect is positive (though small) for some hazard rates after 20 weeks. This is an indication of some (positive) dynamic selection going on: individuals who remain unemployed due to higher b_1 have a slightly higher hazard rate later in the spell. Yet, this dynamic selection effect is not large enough to undo the large mechanical effect that a much larger fraction of individuals survive into the second part of the benefit profile.

The figure therefore provides some intuition for why b_1 has a MH cost that is somewhat larger than b_2. b_1 increases D_1 more than b_2 because it strongly affects hazard rates early in the spell. This in turn has a large mechanical effect on D_2 since more individuals survive into the second part of the benefit profile. The effects of b_1 (positive) and b_2 (negative) on hazard rates after 20 weeks are too small and insignificant to undo, in the MH costs estimates, the effects on hazard rates early in the spell.

Figure B-1: RKD estimates on hazard rates at the SEK725 kink



Notes: The figure reports the RKD estimates of the effect of UI benefits on the hazard rates out of unemployment. Empirical hazard rates are the observed fraction of individuals exiting unemployment in period t conditional on surviving until the start of period t, and are defined by periods of 5 weeks. Blue dots represent the marginal effect of a change in both $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$, estimated in the regression kink design for spells starting between 1999 and July 2001. Red dots represent the marginal effect of a change in $b_{\mathbf{2}}$ only, estimated in the regression kink design for spells starting between July 2001 and July 2002. All estimates are from linear specifications using the changes in the UI schedule at the 725SEK kink with a 90SEK bandwidth. 95% confidence intervals around the point estimates, from robust standard errors, are displayed.

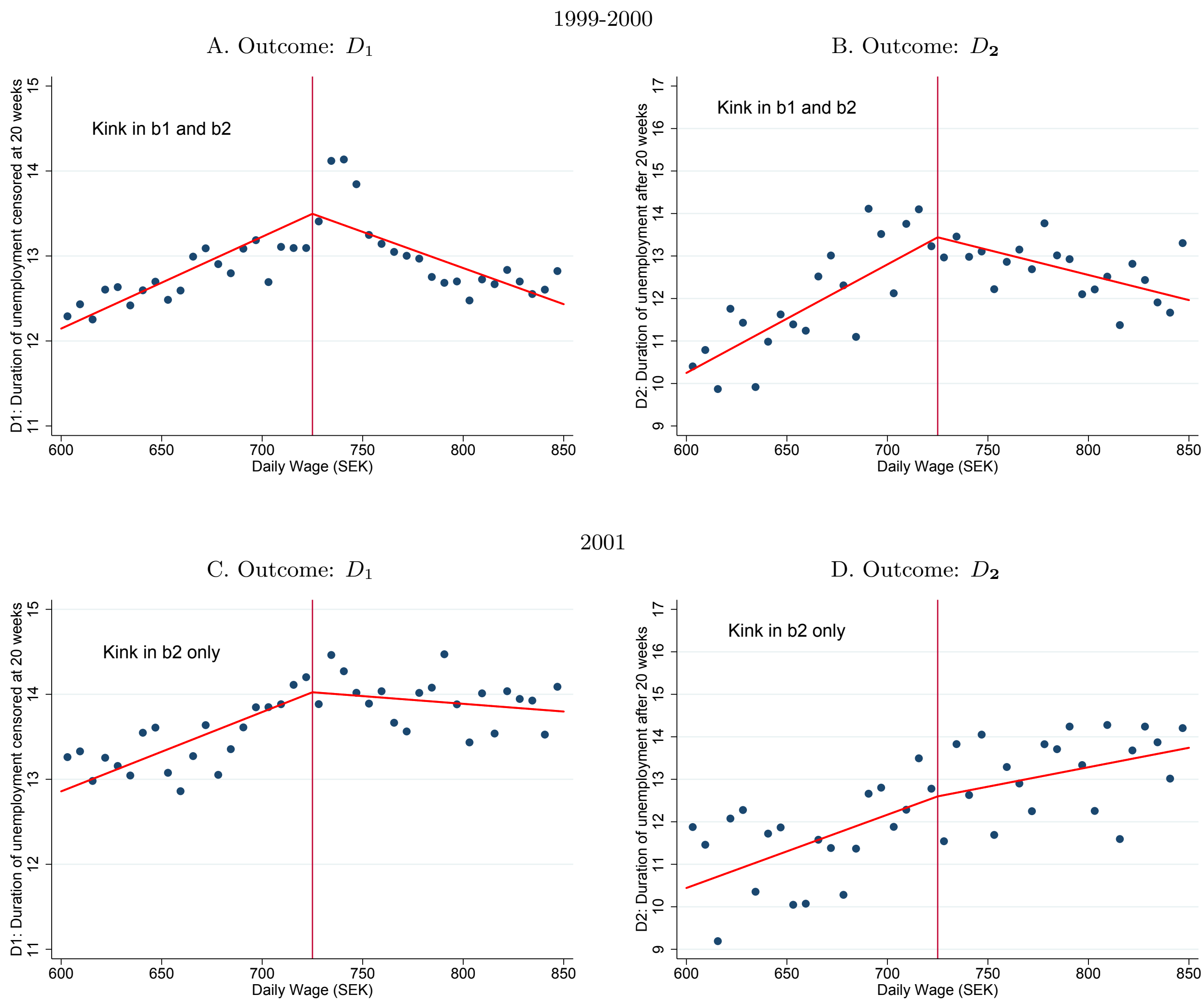
B.2 RK design for D_1 and D_2

To assess the validity of the RK design for unemployment duration D_1 spent on the first part of benefit profile and unemployment duration D_2 spent in the second part of the benefit profile, Figure B-2 below displays the raw data, replicating for D_1 and D_2 what Figure 2 was doing for total unemployment duration D. The graphs provide graphical evidence of a change in slope in the relationship between both D_1 and D_2 and previous daily wage in response to the kink in UI benefits. The change in slope is larger for spells starting before July 2001, when both b_1 and b_2 are capped at the 725SEK threshold. The magnitude of the change in slope decreases for spells starting between July 2001 and July 2002 when only b_2 is capped at the 725SEK threshold. Formal estimates of the change in slope using linear specifications of the form of equation (16) are displayed in Table 2. The red lines display predicted values of the regressions in the linear case.

B.3 Year by year RKD estimates

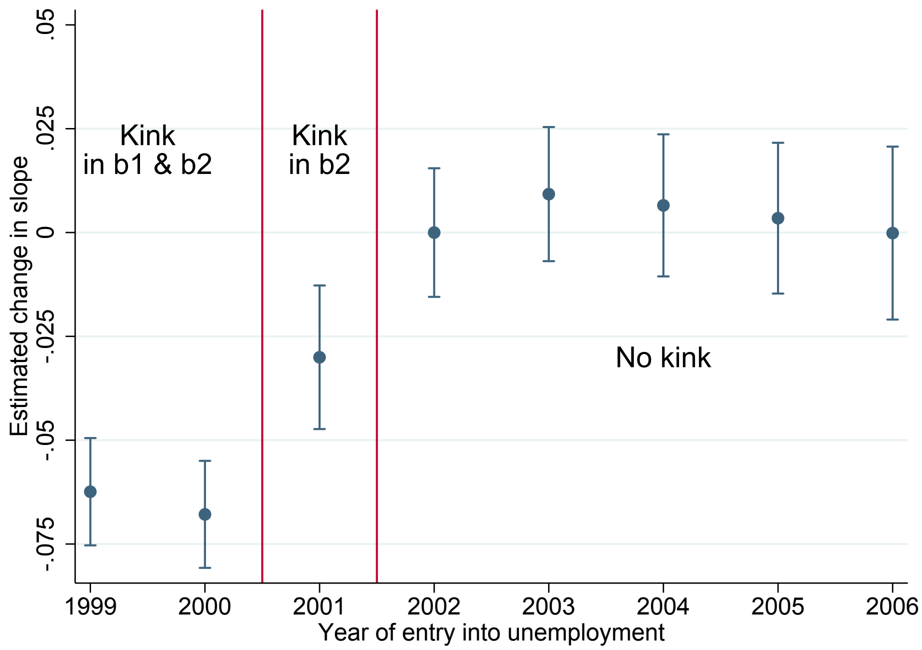
Figure B-3 plots the year-by-year evolution of the estimates of the change in slope in the relationship between total unemployment duration D and pre-unemployment daily wages from 1999 to 2007. The figure provides clear evidence that our estimated responses are indeed due to the policy changes, and not due to time trends in the distribution of durations around the kink.

Figure B-2: RK design at the SEK725 threshold for D_1 and D_2



Notes: The Figure plots average unemployment duration D_1 spent on the first part of benefit profile and average unemployment duration D_2 spent on the second part of the benefit profile, in bins of previous daily wage for the two periods of interest. Sample is restricted to unemployed individuals with no earnings who report being searching for full-time employment. The graphs provide graphical evidence of a change in slope in the relationship between both D_1 and D_2 and previous daily wage in response to the kink in UI benefits. The change in slope is larger for spells starting before July 2001, when both b_1 and b_2 are capped at the 725SEK threshold. The magnitude of the change in slope decreases for spells starting between July 2001 and July 2002 when only b_2 is capped at the 725SEK threshold. Formal estimates of the change in slope using linear specifications of the form of equation (16) are displayed in Table 2. The red lines display predicted values of the regressions in the linear case.

Figure B-3: RKD ESTIMATES ON UNEMPLOYMENT DURATION D AT THE SEK725 KINK BY YEAR OF ENTRY



Notes: The figure reports the RKD estimates of the effect of UI benefits on total duration of unemployment by year of entry into unemployment, at the 725SEK kink. Entry into unemployment in Year N is defined as starting a spell between of July 1st of Year N and July 1st of Year $N+1$. Spells starting before 2001 are therefore subject to a kink in both $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$. Spells starting in 2001 are subject to a kink in $b_{\mathbf{2}}$ only. Spells starting in 2002 and after do not face any kink in the schedule and represent a placebo. All estimates are from linear specifications using the changes in the UI schedule at the 725SEK kink with a 90SEK bandwidth. 95% confidence intervals around the point estimates, from robust standard errors, are displayed. The figure provides clear evidence that estimated responses in the RK design are indeed due to the policy changes, and not due to time trends in the distribution of durations around the kink.

B.4 Additional robustness analysis of the RK design

This subsection presents various additional robustness checks of the RK design. We start by restating the two fundamental identifying assumptions of the RK design, and then propose various tests to assess their potential validity, by looking for clear violations of these assumptions.

We consider the general model:

$$Y = y(b_{\mathbf{1}}, b_{\mathbf{2}}, w, \mu),$$

We are interested in identifying the marginal effect of benefits $b_k, k = 1, 2$ on the duration outcome Y, $\alpha_{\mathbf{k}} = \frac{\partial Y}{\partial b_{\mathbf{k}}}$. $b_{\mathbf{k}}$ is a deterministic, continuous function of the wage w, kinked at $w = \bar{w}_{\mathbf{k}}$. Identification of $\alpha_{\mathbf{k}}$ in the RK design relies on two assumptions:

Assumption 1: the direct marginal effect of the assignment variable w on Y is assumed to be smooth around the kink point $\bar{w}_{\mathbf{k}}$. This means that $\frac{\partial y(b_{\mathbf{1}}, b_{\mathbf{2}}, w, \mu)}{\partial w}$ is assumed to be continuous in the neighborhood of the kink point.

Assumption 2: the distribution of unobserved heterogeneity μ is assumed to be evolving smoothly around the kink point. This means that the conditional density ($f_{w|\mu}(\cdot)$) and its partial derivative with respect to w, ($\partial f_{w|\mu}(\cdot)/\partial w$) are assumed to be continuous in the neighbourhood of the kink point.

These identifying assumptions are, by definition, untestable. Yet, we can use the various "experiment arms" of our quasi-experimental setting as well as sensitivity analysis to try to detect clear violations of these assumptions and to provide some sense of the potential robustness of these identifying assumptions and the validity of our RK design.

Testing for clear violations of Assumption 2: manipulation The most obvious violation of the assumption of smooth distribution of heterogeneity at the kink arises if individuals are able to locate their daily wage strategically around the kink point. A few tests can help assess the robustness of this assumption.

First, Figure B-4 plots the density of the daily wage and shows graphically the smoothness of the distribution of the assignment variable at the kink point in the UI schedules. The graph shows the probability density function of the daily wage around the 725SEK threshold and displays two formal tests. The first is a standard McCrary test of the discontinuity of the pdf of the assignment variable. We report the difference in height of the pdf at the threshold. The second is a test for the continuity of the first derivative of the p.d.f. We report the coefficient estimate of the change in slope of the pdf in a regression of the number of individuals in each bin on polynomials of the assignment variable interacted with a dummy for being above the threshold. Both tests suggest smoothness of the assignment variable around the threshold

The continuity in the pdf of the assignment variable indicates that there is no bunching at the kink point. Such bunching would have constituted proof of the ability of individuals to manipulate their location on the UI schedule, which would have been a clear violation of Assumption 2. Absence of bunching at the kink is not a sufficient condition to rule out that individuals respond to the kinked schedule in their earnings decision, which would question the validity of Assumption 2. The absence of bunching could be driven by optimization frictions which attenuate the ability to bunch at the kink, or by the fact that the compensated elasticity of daily wage with respect to marginal tax rates is small. Even if the compensated elasticity of the daily wage is small, income effects could still be large, and would affect earnings decisions as we move further away from the kink. This would then be picked up by variations in the slope of the pdf at the kink. The fact that we do not detect any change in the first derivative of the pdf of daily wage at the kink point, as reported in Figure B-4, is reassuring.

Interestingly, because the kinks in the schedule of $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$ are removed in July 2001 and July 2002, we can actually directly estimate whether the distribution of daily wages reacts to the removal of the kink and therefore get a direct test of whether the pdf of the assignment variable is affected by the presence of the kink. In Table B-1 below, we report the results of a difference-in-difference model where we look at the evolution of log wages above and below the kink, before July 2001 (when both kinks were in place) and after July 2001 (when one kink is removed).

The wages of individuals who had optimally chosen their daily wages at or above the kink, will be affected by the removal of the kink. To the contrary, individuals who had optimally chosen daily wages below the kink should not be affected by the removal of the kink. If individuals' daily wages respond to the kinked UI schedule, we therefore expect a differential change in the average log wages above the kink after July 2001 relative to log wages below the kink. Estimates, reported in Table B-1 indicate that the removal of the kinks did not significantly affect the distribution of daily wages above and below the kink. There is no differential change in the daily wage below and above the kink after July 2001. This in turn suggests that the presence of kinks in the UI schedule does not significantly affect the distribution of daily wages around the kink.

Testing for clear violations of Assumption 2: observable heterogeneity To further investigate the evolution of the distribution of heterogeneity at the kink, the panels in Figure B-5 show how the mean values of different covariates (age, fraction of men, highly educated and foreigners) evolve with the daily wage around the kink. We do not find any non-linearity around the kink. This is also reassuring, as non-smoothness in the distribution of observable heterogeneity would have cast doubt on the validity of the assumption of smoothness in the distribution of unobservable heterogeneity around the kink.

Testing for underlying non-linearities: Bandwidth size The panels in Figure B-6 report our RKD estimates for different bandwidth sizes. For all periods we consider, the estimates remain stable for bandwidths above $h =$ 60SEK.

Testing for underlying non-linearities: Permutation tests Ganong and Jaeger [2014] suggest that it can be helpful to assess whether the true coefficient estimate is larger than those at "placebo" kinks placed away from the true kink. The idea behind their permutation test is that, if the counter-factual relationship between the assignment variable and the outcome (i.e., in the absence of the kink in the budget set) is non-linear, then the curvature in this relationship will result in many of the placebo estimates being large and statistically significant. In Table B-2, we report 95% confidence interval based on this permutation procedure and compare them to bootstrapped standard errors and robust standard errors.

Testing for underlying non-linearities: Non-parametric detection of kink point Figure B-10 shows the R-squared when we run the RKD regression in (16) for "placebo" kinks placed in 10SEK increments from the true location of the threshold. This procedure, proposed in Landais [2015], and inspired from the time series literature on detection of trend breaks, enables to non-parametrically detect where a true kink is most likely to be located in the data, by looking at the placebo kink where the R-squared is maximized. The figure shows that the R-squared is maximized at the location of the actual kink point, again supporting the evidence that there is in fact a change in slope that occurs at the actual kink point. In both panels A and B, the preferred location of the kink is extremely close to the true kink and the relationship between the placebo kink location and R-squared of the model exhibits a clear concave shape. In panel C, reassuringly, when there is no true kink at 725SEK, this relationship is perfectly flat.

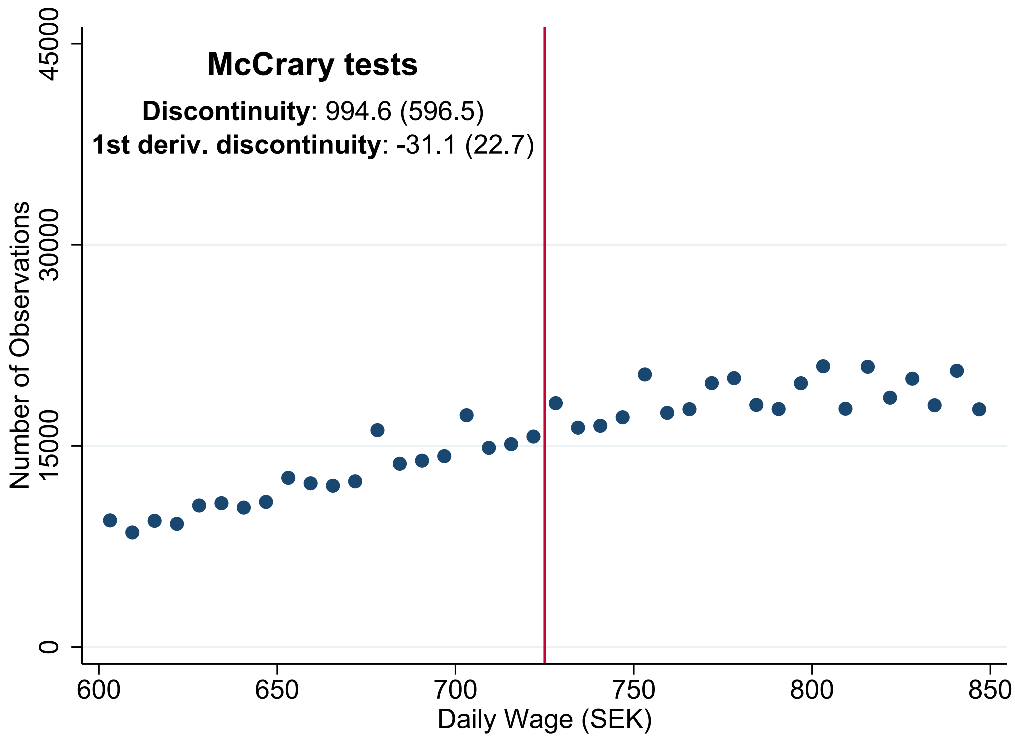
Polynomial order Table B-3 shows estimates of the change in slope at the kink for linear, quadratic and cubic specifications, assessing the model fit for these different specifications.

For the 1999-2000 period, the estimates are very similar across polynomial orders. For the 2001 period, estimates vary across polynomial orders, and estimates from the quadratic model are larger in magnitude than estimates using a linear specification. Yet, model fit analysis suggests that linear estimates should be preferred. The linear specification is having similar root mean squared errors (RMSE) and minimizes the Aikake information criterion (AIC). Note also that, although larger, the point estimates on the quadratic specification are very imprecisely estimated, so that we cannot actually reject that they are equal to the estimates from the linear model.

We also plot below in Figure B-8 the prediction from the linear and quadratic specifications on top of the raw data to see how these models fit the data. For the period 1999-2000, panel A shows that both the quadratic and the linear model fit the data equally well and deliver extremely similar results for the change in slope at the kink. For the period 2001, the quadratic model delivers a larger change in slope at the kink compared to the linear fit. But this is driven by a higher curvature so that the linear model overall does deliver a better fit of the data, as indicated by the root squared mean error and the AIC reported on the graph.

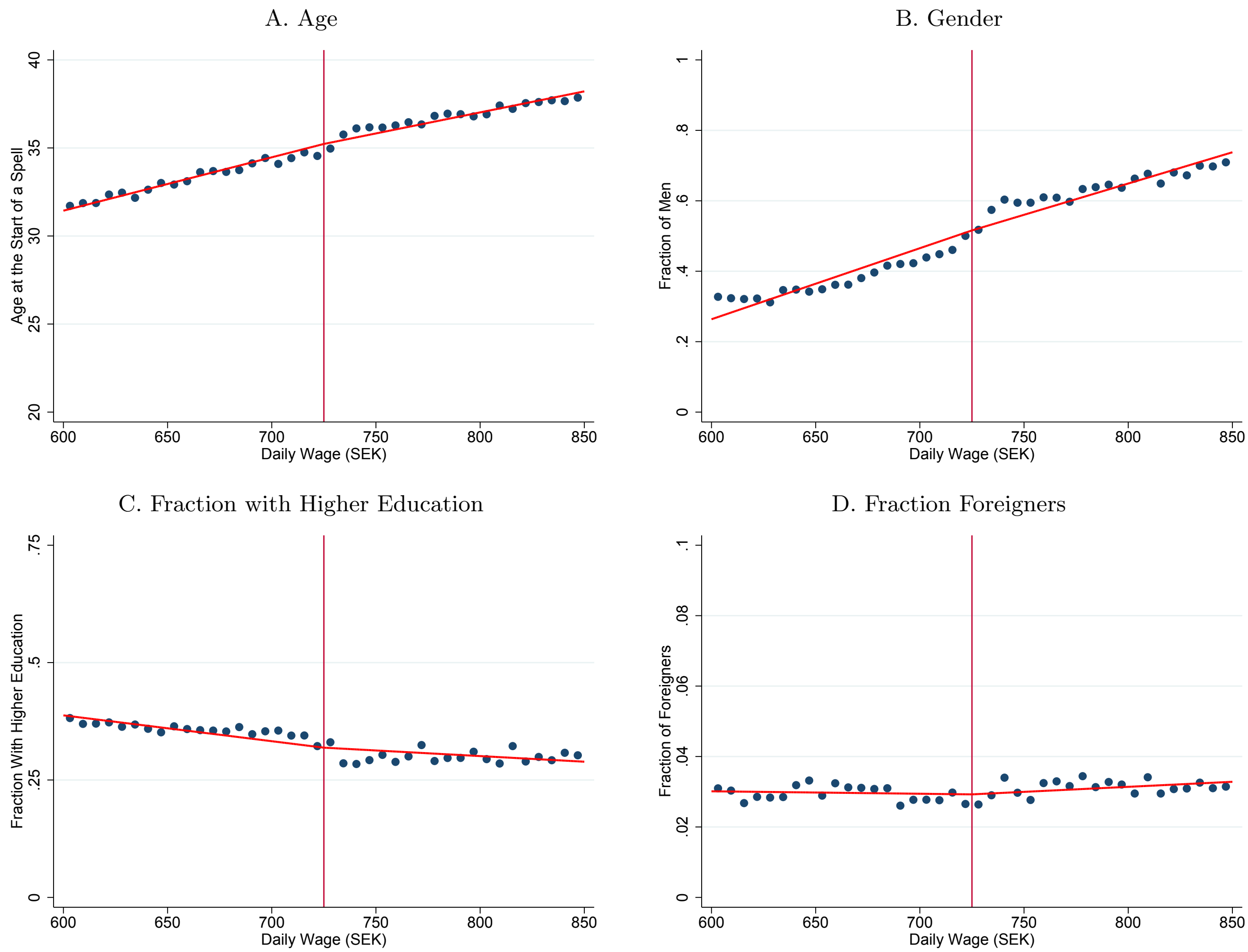
Right-censoring When the schedule of benefits changes, individuals with ongoing spells are transferred to the new schedule. To control for this, two solutions can be envisaged. First, one may get rid of observations who have an ongoing spell at the moment the schedule changes. An alternative solution is to treat the duration of these observations as censored at the moment when these individuals transfer the new schedule. One can then estimate a Tobit model on the right-censored data. In Figure B-9 below, we report the estimates for the estimated change in slope in $D_{\mathbf{1}}$ and $D_{\mathbf{2}}$ for censored and uncensored models, as a function of the RKD bandwidth. The Figure shows that censored and uncensored models deliver identical results, and that the point estimates of the two models are never statistically significantly different. The uncensored model proves a little less precise though, as we end up throwing away some observations. As a consequence, we have decided to focus on the estimates from these censored models for our baseline results.

Figure B-4: Robustness of the RK design: p.d.f of daily wage



Notes: The figure tests graphically the smoothness of the distribution of the assignment variable at the kink point in the UI schedules to assess the validity of the local random assignment assumption underlying the RK design. The Panel shows the probability density function of the daily wage around the 725SEK threshold. We also display two formal tests of the identifying assumptions of the RKD. The first is a standard McCrary test of the discontinuity of the p.d.f of the assignment variable. We report the difference in height of the p.d.f at the threshold. The second is a test for the continuity of the first derivative of the p.d.f. We report the coefficient estimate of the change in slope of the p.d.f in a regression of the number of individuals in each bin on polynomials of the assignment variable interacted with a dummy for being above the threshold. Both tests suggest smoothness of the assignment variable around the threshold, in support of the identifying assumptions of the RK design.

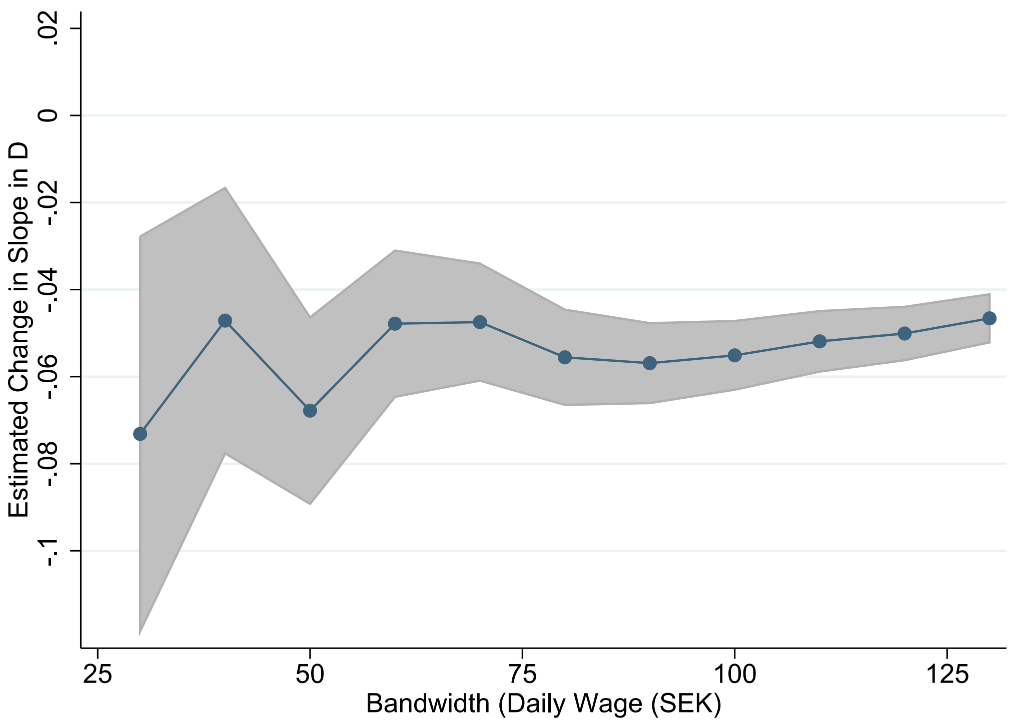
Figure B-5: Robustness of the RK design: Covariates



Notes: The figure tests the validity of the smoothness assumptions of the RK design. Each panel shows the mean values of a different covariate in bins of the assignment variable around the 725SEK threshold. The red lines display predicted values of polynomial regressions of the form of equation (16) in order to detect potential non-linearity around the threshold. The sample is restricted to all spells starting before July 2002, when kinks in the UI schedule are active at the 725SEK threshold. The graphs show evidence of smoothness in the evolution of all covariates at the kink, in support of the RKD identification assumptions.

Figure B-6: RKD estimates as a function of bandwidth size

A. **1999 - 2000**



B. **2001**

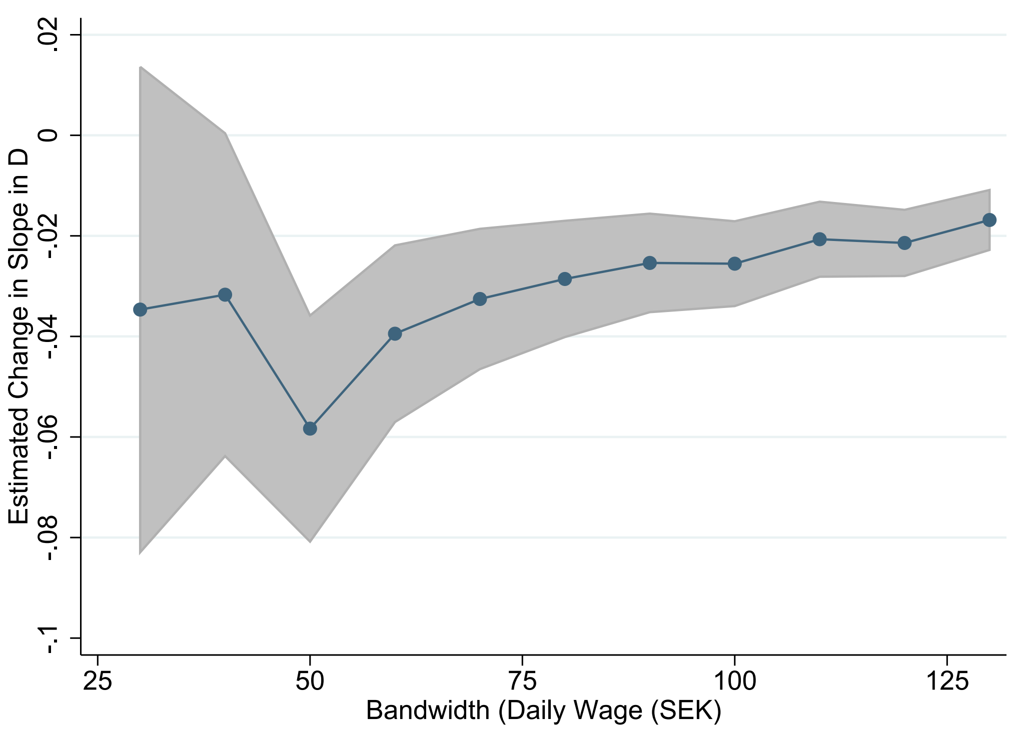
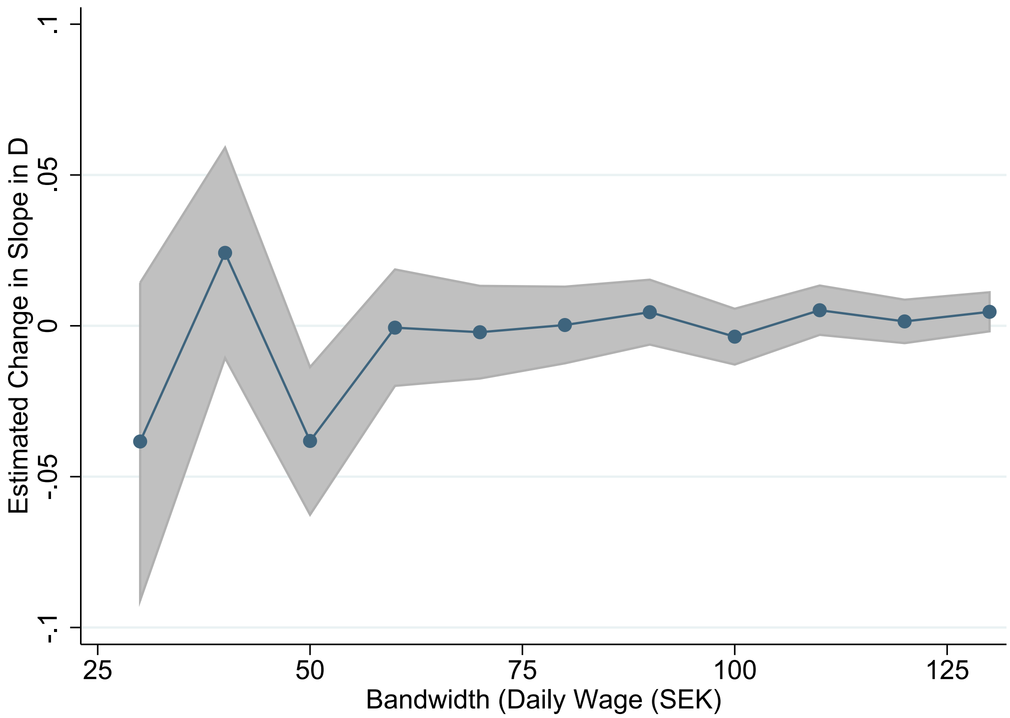


Figure B-7: RKD estimates as a function of bandwidth size (*continued)*

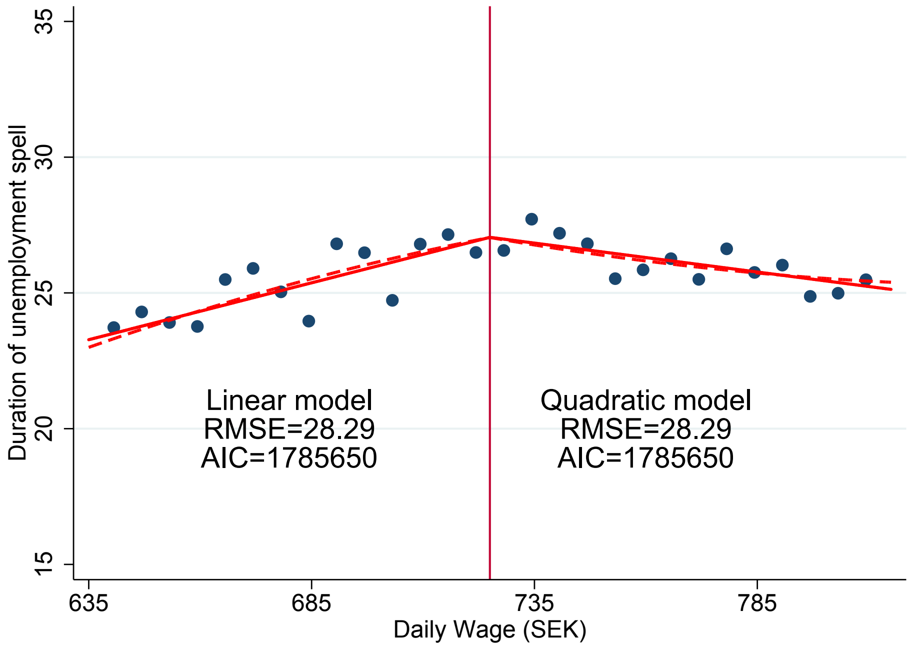
A. **2002-2005**



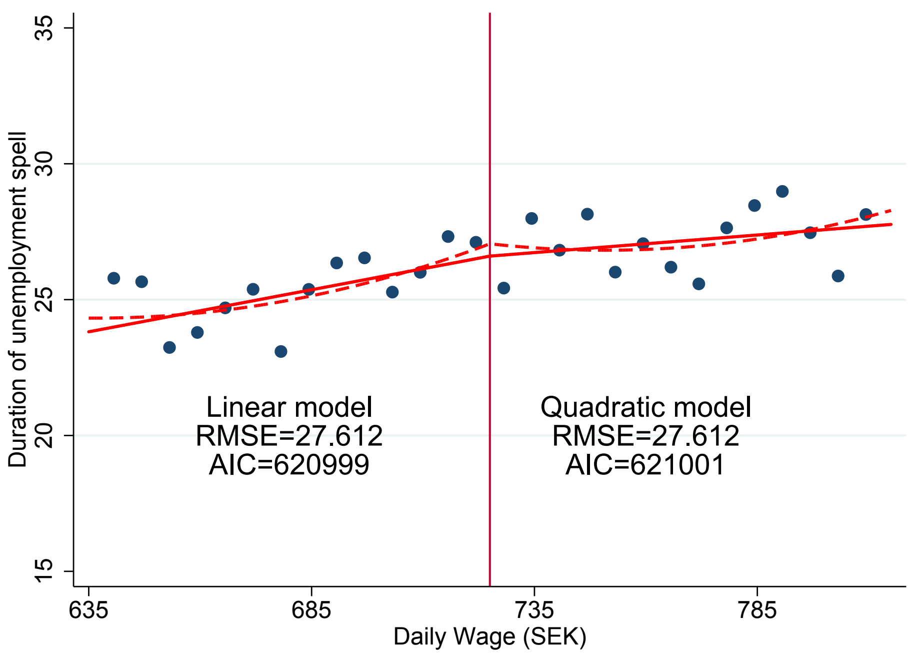
Notes: The figure reports estimates of the change in slope with 95% robust confidence interval in the relationship between unemployment duration and daily wage at the 725SEK threshold using linear regressions of the form of equation (16) as a function of bandwidth size h. These estimates are reported for three periods of interest: 1999-2000 (i.e., spells starting before July 2001), 2001 (i.e., spells starting after July 2001 and before July 2002) and 2002- (i.e., spells starting after July 2002). Unemployment duration is defined as the number of weeks between registration at the PES and exiting the PES or finding any employment (part-time or full-time employment, entering a PES program with subsidized work or training, etc.). Unemployment duration is capped at two years. Sample is restricted to unemployed individuals with no earnings who report being searching for full-time employment.

Figure B-8: Unemployment duration as a function of daily wage around the 725SEK kink, and linear and quadratic model fits

A. 1999-2000

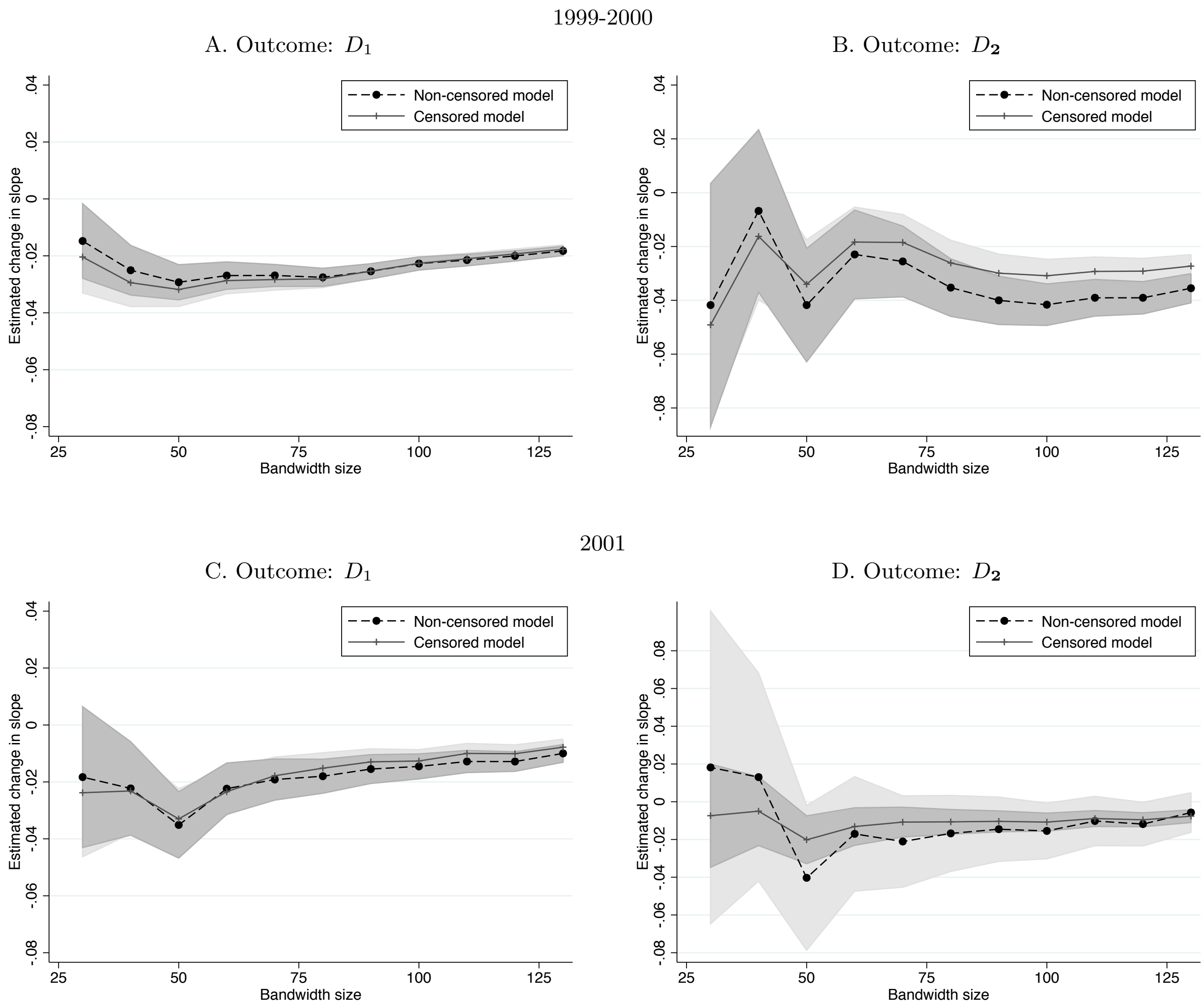


B. 2001



Notes: The figure plots average unemployment duration in bins of previous daily wage for spells starting before July 2001 (panel A) and for spells starting between July 2001 and July 2002 (panel B). On top of the raw data, the figure also displays predictions from linear and quadratic regressions of the form of equation (16) with a bandwidth size $h = 90$SEK. To further assess model fit, we report for each specification the root mean squared error (RMSE) as well as the Aikake information criterion (AIC).

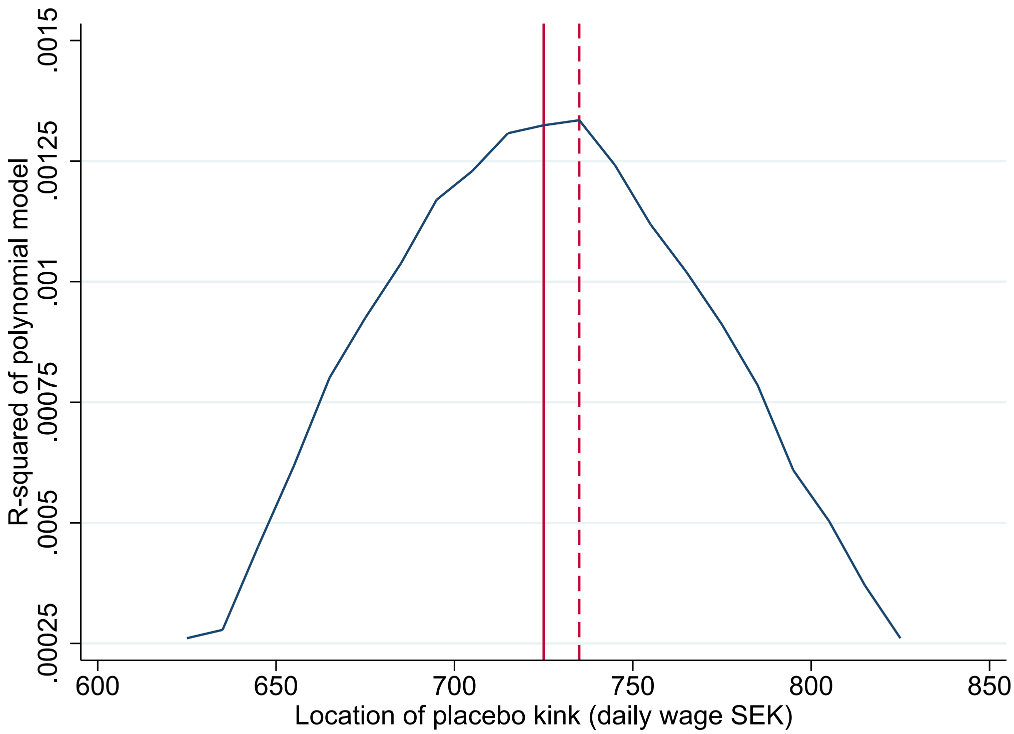
Figure B-9: RKD estimates of the change in slope at the SEK725 kink for OLS model and for the censored regression model



Notes: The figure reports estimates of the change in slope with 95% robust confidence interval in the relationship between unemployment duration and daily wage at the 725SEK threshold using linear regressions of the form of equation (16) as a function of bandwidth size h. When the schedule of benefits changes, individuals who have ongoing spells are transferred to the new schedule. The Figure compares results for two different solutions to account for this. First, one may estimate OLS regressions on a sample where observations who have an ongoing spell at the moment the schedule changes are thrown out (non-censored model). An alternative solution is to treat the duration of these observations as censored at the point when these individuals get in the new schedule. One can then estimate a Tobit model on the right-censored data (censored model). The Figure compares estimates from these two solutions. These estimates are reported for two periods of interest: 1999-2000 (i.e., spells starting before July 2001) and 2001 (i.e., spells starting after July 2001 and before July 2002).

Figure B-10: Non-parametric detection of kink location

A. 1999 - 2000



B. 2001

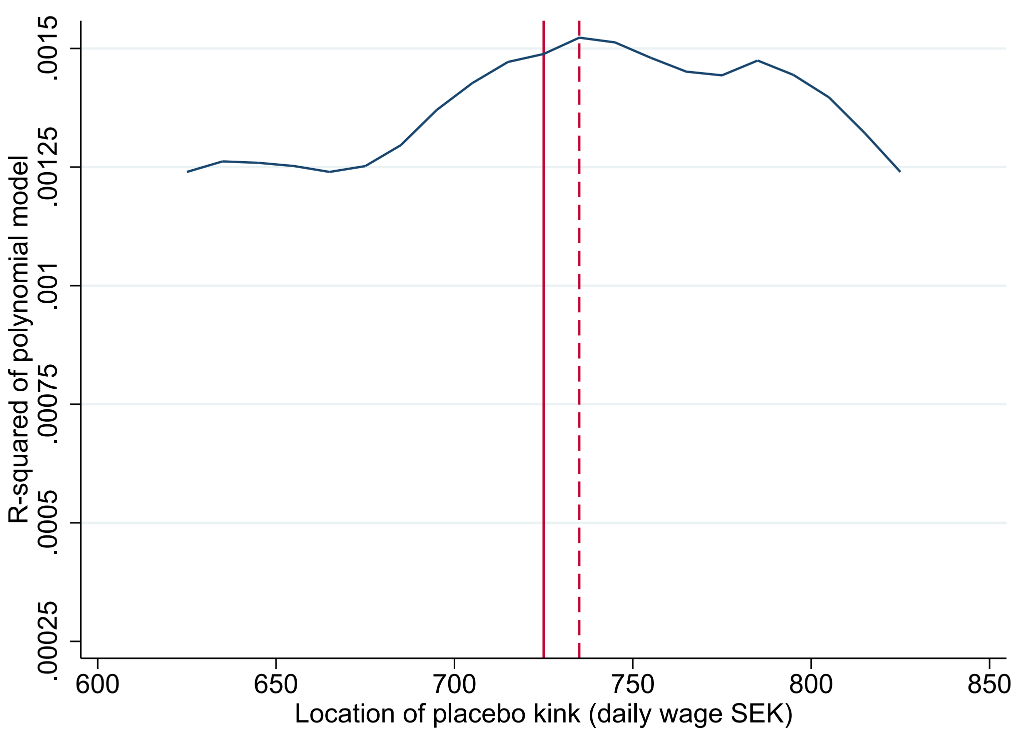
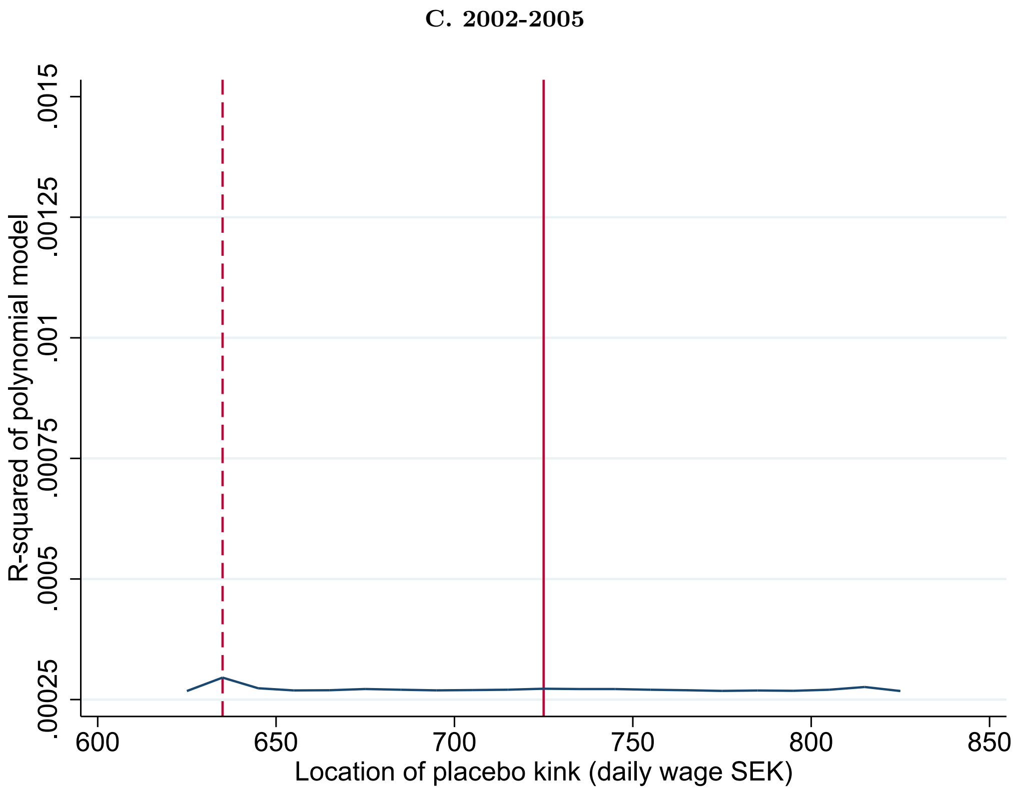


Figure B-11: Non-parametric detection of kink location (*continued)*



Notes: The figure reports the R-squared of polynomial regressions of the form of equation (16) for alternative (placebo) locations of the kink point $\bar{w}_k$ for all observations with wages between 625SEK and 825SEK. The red line indicates the location of the true kink in the schedule. The dashed red line indicates the preferred location of the kink non-parametrically detected in the data, maximizing the R-squared of the model. These estimates are reported for three periods of interest: 1999-2000 (i.e., spells starting before July 2001), 2001 (i.e., spells starting after July 2001 and before July 2002) and 2002- (i.e., spells starting after July 2002). In both panels A and B, the preferred location of the kink is extremely close to the true kink and the relationship between the placebo kink location and R-squared of the model exhibits a clear concave shape. In panel C, when there is no true kink at 725SEK, this relationship is perfectly flat.

Table B-1: EVOLUTION OF DAILY WAGES BELOW AND ABOVE THE 725SEK POINT, BEFORE AND AFTER KINKS IN THE UI SCHEDULE ARE REMOVED

	(1)	(2)	(3)
Dependent variable:		log of daily wage	
$\mathbb{1}[w > \bar{w}]$	0.120***	0.119***	0.119***
	(0.000146)	(0.000146)	(0.000147)
$\mathbb{1}[\text{Spell} > \text{July2001}]$	0.00402***	0.00438***	0.00465***
	(0.000164)	(0.000164)	(0.000165)
$\mathbb{1}[\text{Spell} > \text{July2001}] \times \mathbb{1}[w > \bar{w}]$	0.000305	0.000393	0.000519*
	(0.000222)	(0.000222)	(0.000222)
Age F-E		×	×
Education F-E			×
Gender F-E			×
Industry F-E			×
N	424309	424309	424309

Notes: Standard errors in parentheses.
The Table tests for changes in the distribution of daily wages above and below the kink, as kinks in the schedule of b_1 and b_2 are removed in July 2001 and July 2002. The table reports the results of a difference-in-difference model of the form:

$$\log w = \beta_0 + \beta_1 \mathbb{1}[w > \bar{w}] + \beta_2 \mathbb{1}[\text{Spell} > \text{July2001}] + \beta_3 \mathbb{1}[\text{Spell} > \text{July2001}] \times \mathbb{1}[w > \bar{w}] + \eta$$

The wages of individuals who had optimally chosen their daily wages at or above the kink, will be affected by the removal of the kink. To the contrary, individuals who had optimally chosen daily wages below the kink should not be affected by the removal of the kink. If individuals' daily wages respond to the kinked UI schedule, we therefore expect a differential change in the average log wages above the kink after July 2001 relative to log wages below the kink, captured by β_3. Estimates indicate that the removal of the kinks did not significantly affect the shape of the distribution of daily wages above and below the kink.

Table B-2: RKD estimates at the 725SEK threshold: Inference

	(1) Unemployment Duration D	(2) Duration $D_{\mathbf{1}}$ ($<$ 20 weeks)	(3) Duration $D_{\mathbf{2}}$ ($\geq$ 20 weeks)
	I. 1999-2000: Kink in b_1 and b_2		
Linear - $\delta_{\mathbf{k}}$	-.0569	-.0246	-.0299
Robust s.e.	(.0047)	(.0013)	(.0036)
Bootstrapped s.e.	(.0050)	(.0012)	(.0039)
95% CI - permutation test	[-.0595 ; -.0566]	[-.0319 ; -.0189]	[-.0402 ; -.019]
	II. 2001: Kink in b_2 only		
Linear - $\delta_{\mathbf{k}}$	-.0255	-.0115	-.0105
Robust s.e.	(.005)	(.0021)	(.0028)
Bootstrapped s.e.	(.0049)	(.0020)	(.0030)
95% CI - permutation test	[-.0325 ; -.0190]	[-.0127 ; -.0103]	[-.0115 ; -.0091]
	III. 2002-. : Placebo		
Linear -$\delta_{\mathbf{k}}$	.0045	-.0016	.006
Bootstrapped s.e.	(.0048)	(.0011)	(.0041)
Robust s.e.	(.0055)	(.0011)	(.0049)
95% CI - permutation test	[.0017 ; .0075]	[-.0021 ; -.0011]	[.0053 ; .0071]

Notes: The table reports estimates of the change in slope $\delta_{\mathbf{k}}$, at the 725SEK threshold, in the relationship between daily wage and the total duration of unemployment D (col. (1)), the time $D_{\mathbf{1}}$ spent on the first part of the Swedish UI profile (col. (2)) and the time $D_{\mathbf{2}}$ spent on the second part of the Swedish UI profile (col. (3)). $D_{\mathbf{1}} = \sum_{t<20wks} S_t$ corresponds to duration censored at 20 weeks of unemployment. $D_2 = \sum_{t\geq 20wks} S_t$ corresponds to unconditional duration spent unemployed after 20 weeks of unemployment (i.e., not conditional on having survived up to 20 weeks). Estimates are obtained from linear regressions of the form of equation (16) with a bandwidth size $h = 90$SEK. These estimates are reported for three periods of interest. Panel I reports estimates for spells starting before July 2001. Panel II reports estimates for spells starting after July 2001 and before July 2002. Panel III reports estimates for spells starting after July 2002. Unemployment duration is capped at two years. We report for each estimate $\delta_{\mathbf{k}}$ the White robust standard errors, the bootstrapped standard errors using 50 replications, as well as 95% confidence intervals using the permutation test method of Ganong and Jaeger [2014].

Table B-3: RKD estimates at the 725SEK threshold: Sensitivity to polynomial order

	(1) Unemployment Duration D	(2) Duration D_1 (< 20 weeks)	(3) Duration D_2 ($\geq$ 20 weeks)
	I. 1999-2000: Kink in b_1 and b_2		
Linear - δ_k	-.0569	-.0246	-.0299
	(.0047)	(.0013)	(.0036)
RMSE	28.285	7.049	23.972
AIC	1785650.8	1264546	1723601.1
Quadratic - δ_k	-.0474	-.0344	-.0183
	(.0185)	(.0049)	(.0143)
RMSE	28.285	7.048	23.971
AIC	1785650.5	1264518.9	1723588.4
Cubic - δ_k	-.0527	-.0291	-.0221
	(.0455)	(.0122)	(.0351)
MSE	28.284	7.046	23.971
AIC	1785644.8	1264394.7	1723590
	II. 2001: Kink in b_2 only		
Linear - δ_k	-.0255	-.0115	-.0105
	(.0050)	(.0021)	(.0028)
RMSE	27.612	6.863	23.512
AIC	620999.2	438509.8	599929.6
Quadratic -δ_k	-.0579	-.0299	-.0151
	(.0196)	(.0098)	(.011)
MSE	27.612	6.863	23.512
AIC	621001.3	438509.9	599932.5
Cubic - δ_k	-.0192	-.0268	.0068
	(.0485)	(.0201)	(.0274)
MSE	27.612	6.863	23.512
AIC	621003.5	438508.6	599934.5

Notes: The table reports estimates of the change in slope δ_k, at the 725SEK threshold, in the relationship between daily wage and the total duration of unemployment D (col. (1)), the time D_1 spent on the first part of the Swedish UI profile (col. (2)) and the time D_2 spent on the second part of the Swedish UI profile (col. (3)). Estimates are obtained from polynomial regressions of the form of equation (16) with a bandwidth size h = 90SEK. Estimates are reported for three different polynomial orders: the linear specification, the quadratic specification and the cubic specification. For each polynomial order, we report model fit diagnostics: the root mean squared error (RMSE) as well as the Aikake information criterion (AIC). Panel I reports estimates for spells starting before July 2001. Panel II reports estimates for spells starting after July 2001 and before July 2002. White robust standard errors are in parentheses.

B.5 Exploiting Other Kinks to Assess the Robustness of $MH_1 \geq MH_2$

The Swedish system offers during the period 1999 to 2007 various sources of variations that can be used to identify $MH_{\mathbf{1}}$ and$MH_{\mathbf{2}}$. There are three simple reasons why the baseline estimates focus only on the kink at 725 SEK. First, for expositional convenience: since all statistics of interest can be identified using the same RK design, this made presenting the source of variation and the results particularly convenient. Second, for internal validity: comparing estimates at the same kink ensures that we are comparing behavioral responses for comparable individuals over time. And finally, for precision: there is more density around the 725SEK kink, which enables a higher degree of statistical precision for the RK estimates

In this section, we investigate the robustness of our results to the use of other sources of variations. Indeed, one may be worried that comparing individuals at the same kink over time may be problematic if behavioral elasticities are prone to varying over time (due to business cycle variations for instance). In this sense, there is a trade-off between comparing individuals at the same kink over time and comparing different individuals but in the same time period. In the first case, one may worry that time affects behavioral elasticities, while in the second case, one may worry that individuals at different kinks are different as they have different pre-unemployment incomes to start with.

In what follows, we review the different kinks and the sources of identification they provide, present our strategy and estimates for each kink, and summarize the conclusions that we can draw from this evidence on the relative magnitude of $MH_{\mathbf{1}}$ vs $MH_{\mathbf{2}}$. The bottom-line is that our estimates are very robust to using these alternative sources of identification and in particular the larger magnitude of $MH_{\mathbf{1}}$ relative to $MH_{\mathbf{2}}$ is a very robust finding, irrespective of the source of identification used.

B.5.1 850 SEK kink

In July 2001, the cap in $b_{\mathbf{1}}$, the UI benefits received during the first 20 weeks of unemployment, was increased to 680SEK, which created a kink in the relationship between $b_{\mathbf{1}}$ and the daily wage at a wage level of 850SEK. This gives us the possibility to identify in the RK design the effect of $b_{\mathbf{1}}$ on duration outcomes. In July 2002, the cap in $b_{\mathbf{2}}$, the UI benefits received after the first 20 weeks of unemployment, was increased to 680SEK, which created a kink in the relationship between $b_{\mathbf{2}}$ and the daily wage at a wage level of 850SEK.

Figure B-11 reports the evolution of the RKD estimates of the change in slope in the relationship between unemployment duration and daily wage at the 850 SEK kink, by year of entry into unemployment. Spells starting before 2001 are subject to a linear schedule with no kink in either $b_{\mathbf{1}}$ nor $b_{\mathbf{2}}$ and represent the placebo. Spells starting in 2001 are subject to a kink in $b_{\mathbf{1}}$ only. Spells starting in 2002 and after are subject to a kink in $b_{\mathbf{2}}$ only. The graph provides clear evidence of a break in the relative slopes of the relationship between duration and wage on both sides of the kink after the introduction of the kink in $b_{\mathbf{1}}$ in 2001. It also provides evidence of a slight decrease in the change in slope as the kink in $b_{\mathbf{1}}$ is replaced by a kink in $b_{\mathbf{2}}$ after 2002.

Based on this evidence, we implement a DD-RKD in order to get estimates of the elasticities of duration with respect to $b_{\mathbf{1}}$ and with respect to $b_{\mathbf{2}}$ at the 850SEK kink. Our DD-RKD specification is the following:

$$\begin{aligned} E[Y|w,P] = \quad & \beta_0 + \beta_1(w-\bar{w}) + \delta_0(w-\bar{w})\cdot \mathbb{1}[w \geq \bar{w}] \\ & + \mathbb{1}[P=1]\cdot \Big(\beta_2 + \beta_3(w-\bar{w}) + \delta_1(w-\bar{w})\cdot \mathbb{1}[w\geq \bar{w}]\Big) \\ & + \mathbb{1}[P=2]\cdot \Big(\beta_4 + \beta_5(w-\bar{w}) + \delta_2(w-\bar{w})\cdot \mathbb{1}[w\geq \bar{w}]\Big). \end{aligned} \tag{26}$$

where P denotes the time period in which the unemployment spell started. $P = 0$ for spells starting before July 2001, $P = 1$ for spells starting between July 2001 and July 2002 and $P = 2$ for spells starting after July 2002.

From this specification, we compute the implied benefit elasticities of unemployment duration. The elasticity of duration with respect to $b_{\mathbf{1}}$, estimated at the 850SEK kink, is $\varepsilon_{D,b_{\mathbf{1}}} = \hat{\delta}_{\mathbf{1}} \cdot \frac{850}{D^{cap}}$, where $\hat{\delta}_{\mathbf{1}}$ is the estimated marginal slope change for spells starting in period $P = 1$, and D^{cap} is the observed average duration at the kink for spells starting in period $P = 1$. We find $\varepsilon_{D,b_{\mathbf{1}}} = 1.92$ (.31). The elasticity of duration with respect to b_2, estimated at the

850SEK kink, is $\varepsilon_{D,b_2} = \hat{\delta}_2 \cdot \frac{850}{D^{cap}}$, where $\hat{\delta}_2$ is the estimated marginal slope change for spells starting in period $P = 2$, and D^{cap} is the observed average duration at the kink for spells starting in period $P = 2$. We find $\varepsilon_{D,b_2} = 1.50 \ \ (.24)$.

B.5.2 912.5 SEK kink

In July 2002, the cap for b_1, benefits received during the first 20 weeks of unemployment, was increased to 730SEK, which created a kink in the schedule of benefits at a daily wage level of 912.5SEK. This gives us the possibility to identify the effect of b_1 on duration outcomes, using a RK design at the 912.5SEK kink.

As done previously at the 725SEK and 850SEK kink, Figure B-12 reports the evolution of the RKD estimates of the change in slope in the relationship between unemployment duration and daily wage at the 912.5SEK kink, by year of entry into unemployment. Spells starting before 2002 are subject to a linear schedule with no kink in either b_1 nor b_2 and represent the placebo. Spells starting in 2002 and after are subject to a kink in b_1 only. The graph provides clear evidence of a break in the relative slopes of the relationship between duration and wage on both sides of the kink after the introduction of the kink in b_1 in 2002. However, estimates are much less precise than for the other two kinks, as there is much less density around the 912.5SEK kink.

Based on this evidence, we also implement a DD-RKD in order to get estimates of the elasticities of duration with respect to b_1 at the 912.5SEK kink. Our DD-RKD specification is the following:

$$\begin{aligned} E[Y|w,P] = \quad & \beta_0 + \beta_1(w-\bar{w}) + \delta_0(w-\bar{w})\cdot \mathbb{1}[w \geq \bar{w}] \\ & + \mathbb{1}[P=2]\cdot \Big(\beta_4 + \beta_5(w-\bar{w}) + \delta_2(w-\bar{w})\cdot \mathbb{1}[w \geq \bar{w}]\Big). \end{aligned} \tag{27}$$

where P denotes the time period in which the unemployment spell started, with $P = 2$ for spells starting after July 2002.

From this specification, we compute the implied benefit elasticities of unemployment duration. The elasticity of duration with respect to b_1, estimated at the 912SEK kink is $\varepsilon_{D,b_1} = \hat{\delta}_2 \cdot \frac{912.5}{D^{cap}}$, where $\hat{\delta}_1$ is the estimated marginal slope change for spells starting in period $P = 2$, and D^{cap} is the observed average duration at the kink for spells starting in period $P = 2$. We find $\varepsilon_{D,b_1} = 2.15 \ \ (.70)$.

B.5.3 Combining estimates to identify the relative magnitude of MH_1 vs MH_2

Based on these sources of variations, we now have four different potential estimates of the relative moral hazard costs of b_1 vs b_2, summarized in Table B-4 below

The first strategy consists in comparing estimates at the same "kink" over time. This approach, as mentioned earlier, has the advantage of comparing similar individuals over time at the same level of income. One drawback may be that behavioral elasticities are time varying due to business cycle fluctuations for instance. Two kinks, the 725SEK kink and the 850SEK kink, allow us to implement this first strategy, giving us two different estimates of the moral hazard costs ratio, displayed in panel A and panel B of Table B-4. Results using estimates at the 725SEK kink represent our baseline strategy, displayed in panel A, and compare estimates from 1999-2000 versus 2001, taking advantage of the fact that at the 725SEK kink, we can identify $\varepsilon_{D,b}$ in 1999-2000 and ε_{D,b_2} in 2001. From this baseline strategy, we find a ratio $MH_1/MH_2 = 1.264$, as shown in panel A of Table B-4

Results using estimates at the 850SEK kink over time are displayed in panel B, and compare estimates from 2001 versus 2002-2007, taking advantage of the fact that at the 850SEK kink, we can identify ε_{D,b_1} in 2001 and ε_{D,b_2} in 2002-2007. From this strategy, we find a ratio $MH_1/MH_2 = 1.155$, as shown in panel B of Table B-4. Estimates of the moral hazard cost ratio from panel A and B are very similar. This confirm that the relative magnitude of the moral hazard costs is very robust across kinks, alleviating the concern that the moral hazard cost ratio is extremely sensitive to the location of the income distribution at which it is estimated.

The second strategy consists in comparing estimates at different "kinks" within the same time period. This second approach has the advantage of comparing individuals within the same time period, therefore controlling for the fact that behavioral elasticities may be time varying due to business cycle fluctuations for instance. A potential drawback of this approach is that individuals at different kinks may differ in their responsiveness to the policy.

Two time periods, 2001 and 2002-2007 allow us to implement this second strategy, giving us two additional estimates of the moral hazard costs ratio, displayed in panel C and panel D of Table B-4. For the 2001 period, we have estimates of $MH_{\mathbf{1}}$ from the 850SEK kink and estimates of $MH_{\mathbf{2}}$ from the 725SEK kink. Combining these estimates, we get a ratio $MH_{\mathbf{1}}/MH_{\mathbf{2}} = 2.695$, as shown in panel C of Table B-4. For the 2002-2007 period, we have estimates of $MH_{\mathbf{1}}$ from the 912.5SEK kink and estimates of $MH_{\mathbf{2}}$ from the 850SEK kink. Combining these estimates, we get a ratio $MH_{\mathbf{1}}/MH_{\mathbf{2}} = 1.343$, as shown in panel C of Table B-4. Again, both estimates of the MH cost ratio confirm that the moral hazard cost of $b_{\mathbf{1}}$ is larger than the moral hazard cost of $b_{\mathbf{2}}$ in our context.

Results from Table B-4 therefore all strongly suggest that our finding that $MH_{\mathbf{1}} > MH_{\mathbf{2}}$ is robust to the sources of variations used to estimate the effects of $b_{\mathbf{1}}$ and $b_{\mathbf{2}}$ on unemployment durations. We believe these results to be an important piece of additional evidence on the validity and robustness of our findings on the relative magnitude of moral hazard costs over the unemployment spell.

B.5.4 Inference on estimates of the relative magnitude of $MH_{\mathbf{1}}$ vs $MH_{\mathbf{2}}$

To provide inference on our estimates of the relative magnitude of $MH_{\mathbf{1}}$ vs $MH_{\mathbf{2}}$, we adopt a permutation-test approach. Since our estimation approach relies on a comparison of two kinks, either over time, or within period, we replicate this comparison for placebo kinks drawn at random in the distribution of daily wage in regions where the UI benefit schedule does not exhibit any kink.

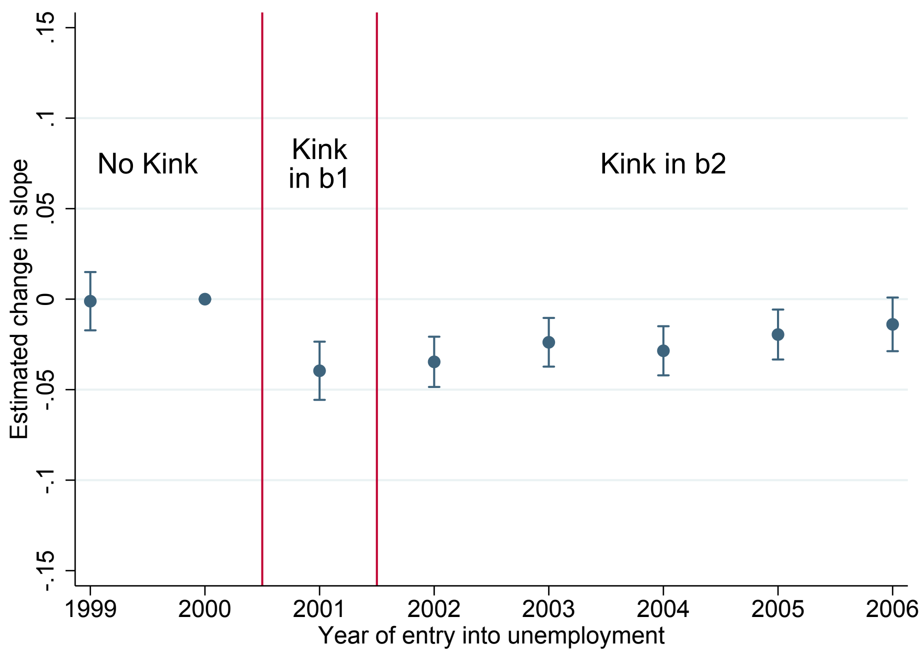
For panel A, we draw 250 kinks at random. For each kink, we estimate the change in slope in 1999-2000 and 2001 and from this obtain a distribution of 250 placebo estimates of the moral hazard $MH_{\mathbf{1}}$ and $MH_{\mathbf{2}}$. In Figure B-13 panel A, we draw the distribution of the placebo estimates of the difference $MH_{\mathbf{1}} - MH_{\mathbf{2}}$ obtained from this procedure. As can be seen, this distribution is centered around 0, and our true estimate obtained from the 725SEK kink lies far in the upper tail of this distribution of placebo estimates. This procedure gives us a p-value of 5.98% for our estimate of the relative magnitude of $MH_{\mathbf{1}} - MH_{\mathbf{2}}$. In other words, this procedure is providing compelling evidence that our estimates of the relative moral hazard costs are not picking up some random variation in the slope of the relationship between durations and daily wage over time.

Similarly, for panel B, we draw again 250 kinks at random. For each kink, we estimate the change in slope in 2001 vs 2002-2007 and from this obtain a distribution of 250 placebo estimates of the moral hazard $MH_{\mathbf{1}}$ and $MH_{\mathbf{2}}$. In Figure B-13 panel B., we draw the distribution of the placebo estimates of the difference $MH_{\mathbf{1}} - MH_{\mathbf{2}}$ obtained from this procedure. Again, this procedure confirms that our estimates of the relative moral hazard costs is not picking up some random variation in the slope of the relationship between durations and daily wage over time. From this distribution of placebo kinks, the p-value for our estimate in Table B-4 panel B is 0.00%.

Similarly, for panel C, we draw 250 kinks at random from which we can generate 31,000 ($= \frac{250^2}{2} - 250$) pairs of RKD estimates for the period July 2001 to July 2002. For each pair of kink, we estimate the changes in slope in 2001 and from this obtain a distribution of 31,000 placebo estimates of the moral hazard $MH_{\mathbf{1}}$ and $MH_{\mathbf{2}}$. In Figure B-13 panel C, we draw the distribution of the placebo estimates of the difference $MH_{\mathbf{1}} - MH_{\mathbf{2}}$ obtained from this procedure. Results again provide compelling evidence that our estimates of the relative moral hazard costs in panel C of Table B-4 are not picking up some random variation in the slope of the relationship between durations and daily wage across different kinks in the 2001 period. The p-value for our estimate in Table B-4 panel C that we obtain from this permutation procedure is 0.00%.

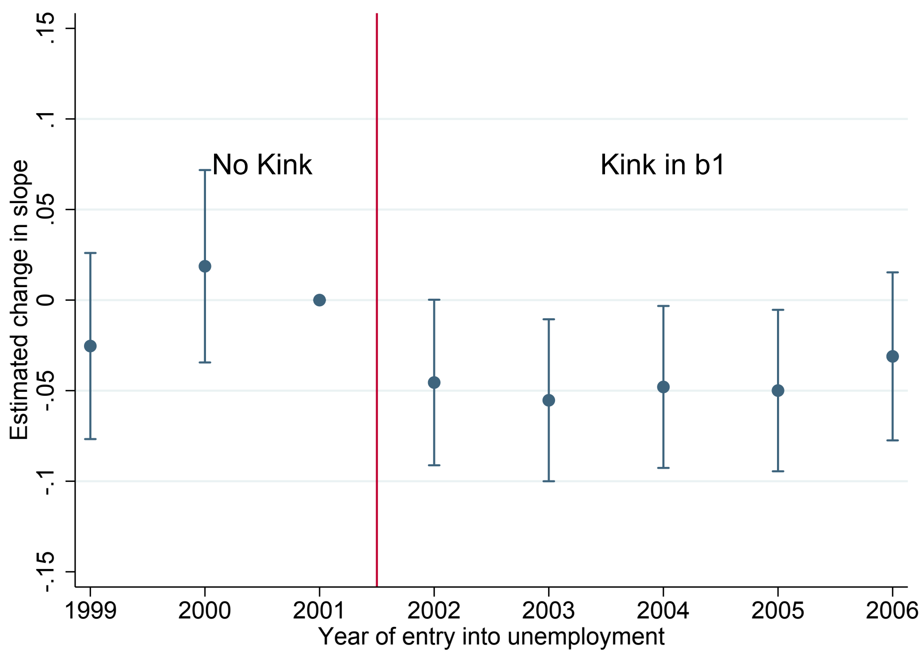
Finally, for panel D, we draw 250 kinks at random from which we can generate 31,000 ($= \frac{250^2}{2} - 250$) pairs of RKD estimates for the period 2002 to 2007. For each pair of kink, we estimate the changes in slope in 2002-2007 and from this obtain a distribution of 31,000 placebo estimates of the moral hazard $MH_{\mathbf{1}}$ and $MH_{\mathbf{2}}$. In Figure B-13 panel D., we draw the distribution of the placebo estimates of the difference $MH_{\mathbf{1}} - MH_{\mathbf{2}}$ obtained from this procedure. This gives us a p-value for our estimate in Table B-4 panel D of 2.63%.

Figure B-11: RKD ESTIMATES ON UNEMPLOYMENT DURATION D AT THE 850SEK KINK BY YEAR OF ENTRY



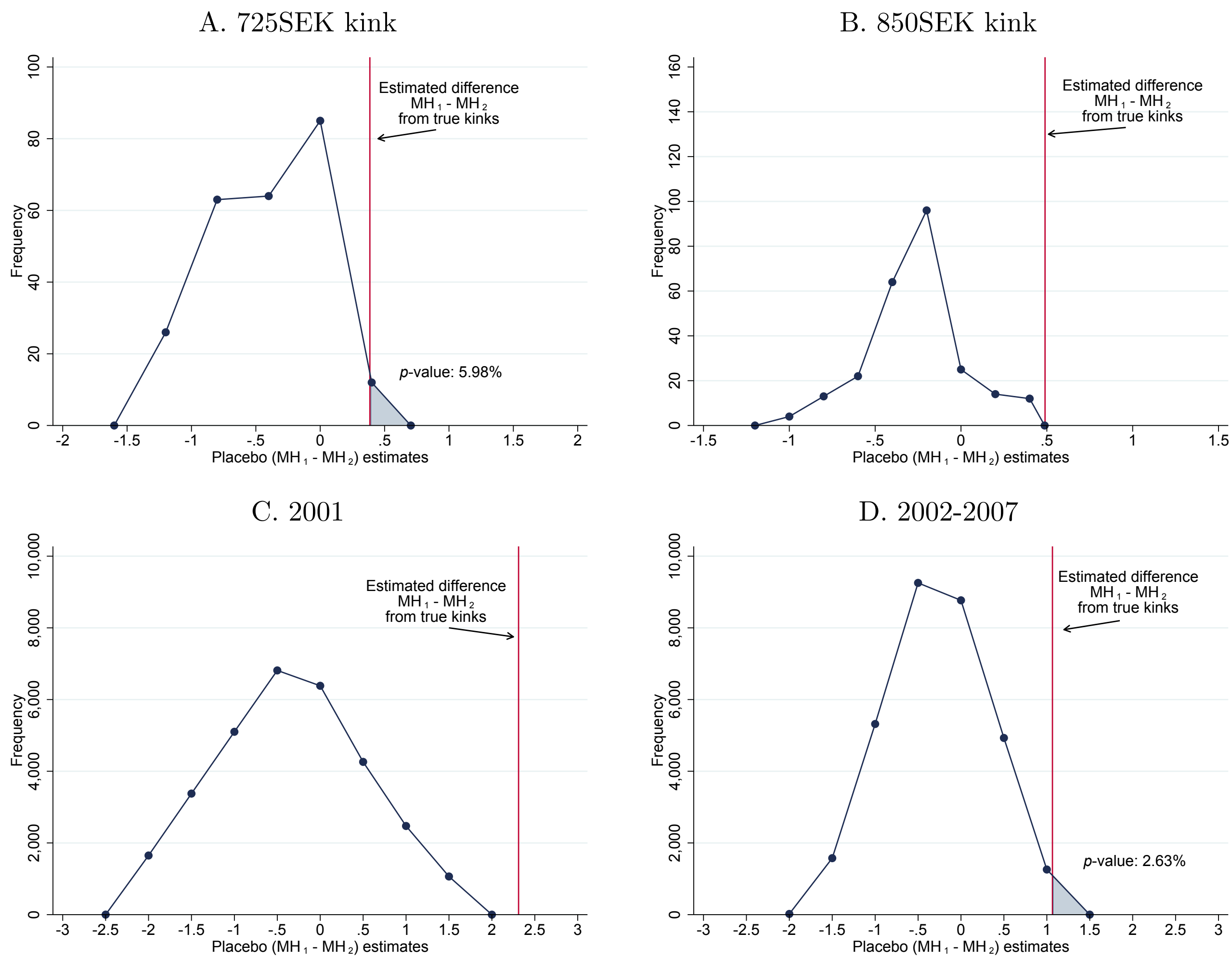
Notes: The figure reports the RKD estimates of the effect of UI benefits on total duration of unemployment by year of entry into unemployment, at the 850SEK kink. Entry into unemployment in Year N is defined as starting a spell between July 1st of Year N and July 1st of Year $N + 1$. Spells starting before 2001 are therefore subject to a smooth linear schedule with no kink in either b_1 nor b_2 and represent the placebo. Spells starting in 2001 are subject to a kink in b_1 only. Spells starting in 2002 and after are subject to a kink in b_2 only. All estimates are from linear specifications using the changes in the UI schedule at the 850SEK kink with a 90SEK bandwidth, and are relative to year 2000, which is the baseline. 95% confidence interval around the point estimates, from robust standard errors, are displayed. The figure provides clear evidence that estimated responses in the RK design are indeed due to the policy changes, and not due to time trends in the distribution of durations around the kink.

Figure B-12: RKD estimates on unemployment duration D at the 912.5SEK kink by year of entry



Notes: The figure reports the RKD estimates of the effect of UI benefits on total duration of unemployment by year of entry into unemployment, at the 912.5SEK kink. Entry into unemployment in Year N is defined as starting a spell between of July 1st of Year N and July 1st of Year $N+1$. Spells starting before 2002 are subject to a smooth linear schedule with no kink in either b_1 nor b_2 and represent the placebo. Spells starting in 2002 and after are subject to a kink in b_1 only. All estimates are from linear specifications using the changes in the UI schedule at the 912.5SEK kink with a 60SEK bandwidth, and are relative to year 2001, which is the baseline. 95% confidence interval around the point estimates, from robust standard errors, are displayed.

Figure B-13: Permutation tests approach to inference on estimates of the relative magnitude of MH_1 vs MH_2



Notes: To provide inference on our estimates of the relative magnitude of $\boldsymbol{MH_1}$ vs $\boldsymbol{MH_2}$, we adopt a permutation-test approach. Since our estimation approach relies on a comparison of two kinks, either over time, or within period, we replicate this comparison for placebo kinks drawn at random in the distribution of daily wage in regions where the UI benefit schedule does not exhibit any kink. For each panel, we show the distribution of placebo estimates of the difference between moral hazard $\boldsymbol{MH_1}$ and $\boldsymbol{MH_2}$ from placebo kinks drawn at random using variation over time at the same kink (panel A, comparing 1999-2000 vs 2001 and panel B, comparing 2001 vs 2002-2007) or across kinks within the same period (panel C in 2001, and panel D in 2002-2007). The red vertical line in each panel displays the estimated value of the difference between moral hazard $\boldsymbol{MH_1}$ and $\boldsymbol{MH_2}$ from the true kinks (displayed in Table B-4).

Table B-4: Summary of Moral Hazard Cost Ratio Estimates

	I. Using Same Kink, Over Time	
	A. 725 SEK - Kink	B. 850 SEK - Kink
δ_{b_1}	-0.036	-0.060
	(0.007)	(0.009)
Variation used	*'99-'00 vs 2001 kink*	*2001 kink*
δ_{b_2}	-.0255	-0.047
	(.0049)	(0.007)
Variation used	*2001 kink*	*2002-2007 kink*
$\frac{MH_1}{MH_2}$	**1.264**	**1.155**
p-value	5.98%	0.00%
	II. Using Different Kinks Within Same Time Period	
	C. 2001	D. 2002-2007
δ_{b_1}	-0.060	-0.062
	(0.009)	(0.02)
Variation used	*850SEK kink*	*912.5SEK kink*
δ_{b_2}	-0.023	-0.047
	(0.008)	(0.007)
Variation used	*725SEK kink*	*850SEK kink*
$\frac{MH_1}{MH_2}$	**2.695**	**1.343**
p-value	0.00%	2.63%

Notes: The Table reports all the different estimates of the moral hazard cost ratios that can be drawn from the systematic exploitation of the kinks in the UI benefit schedule in Sweden over the period 1999 to 2007. Panel I reports estimates from the strategy consisting in comparing estimates at the same "kink" over time. This approach has the advantage of comparing similar individuals over time at the same level of income. Two kinks, the 725SEK kink and the 850SEK kink, allow us to implement this strategy, giving us two different estimates of the moral hazard costs ratio, displayed in panel A and panel B. Panel II reports estimates from the strategy consisting in comparing estimates at different "kinks" in the daily wage distribution within the same time period. This second approach has the advantage of comparing individuals within the same time period, therefore controlling for the fact that behavioral elasticities might be time varying due to business cycle fluctuations for instance. Two time periods, 2001 and 2002-2007 allow us to implement this second strategy, giving us two additional estimates of the moral hazard costs ratio, displayed in panel C and panel D. For each estimate of the moral hazard cost ratio, we report the RKD estimate of the effect on duration D of the change in slope for variation in b_1 (δ_{b_1}) and the estimate of the change in slope for variation in b_2 (δ_{b_2}), along with their standard errors in parenthesis. Below each RKD estimate δ_{b_1} and δ_{b_2}, we display the source of variation used for identification. Below each estimate of the MH cost ratio, we display p-values from a permutation based test, using placebo kinks drawn at random in the distribution of daily wage in regions where the UI benefit schedule does not exhibit any kink. See text for details.

B.6 Comparison to Duration Response Estimates in the Literature

How do our duration response estimates compare to existing estimates of labor supply responses to UI benefits available in the literature? To answer this question, we first need to be very precise about the source of variation of benefits used for the estimates against which we want to benchmark our estimates. As our conceptual analysis makes very clear, elasticity of duration w.r.t benefits paid early during the spell or w.r.t benefits paid later during the spell are conceptually different, and will likely be very different empirically, especially in the presence of non-stationary forces.

The elasticities $\varepsilon_{D_{\mathbf{k}},\bar{b}}$ that we report in Table 2 column (1) in the main text are unique, compared to estimates in the literature, because they measure duration responses to a unique form of benefit variation. To understand this, remember that the potential duration of benefits being infinite in the Swedish system during our period of analysis, individuals can collect these benefits indefinitely. This means that the elasticity $\varepsilon_{D_{\mathbf{k}},\bar{b}}$ that we report measures the response of duration to a change in benefits **forever**. We do not know of any other paper using similar source of variation. The reported elasticities with respect to this unique source of variation appear quite large at first glance, but it might not be that surprising, and it is hard to gauge these magnitudes using available benchmarks.

To appreciate the magnitude of our duration responses, it is therefore better to focus on elasticities that use variations in benefits that are similar to the ones used in the previous literature. Our elasticity of time spent in the first part of the profile $D_{\mathbf{1}}$ with respect to $b_{\mathbf{1}}$, the benefits that you receive in the first 20 weeks is, in this respect, probably the best candidate. It can for instance be compared to estimates of the elasticity of paid unemployment duration (unemployment duration up to exhaustion) with respect to a variation in UI "benefit level" in the US, given individuals receive these benefits only in the first 26 weeks in the US. Put it differently, the elasticity of paid unemployment duration with respect to a variation in UI "benefit level" in the US is conceptually equivalent to our estimate of $\varepsilon_{D_{\mathbf{1}},b_{\mathbf{1}}}$, with the slight difference that $b_{\mathbf{1}}$ is given for 20 weeks in Sweden instead of 26 weeks in the US, and D_1 is the duration up to 20 weeks while the duration of paid unemployment is the duration up to 26 weeks in the US.[65] Our estimate $\varepsilon_{D_{\mathbf{1}},b_{\mathbf{1}}}$ from the 725 SEK kink is .71 (.10). This estimate is in line with available estimates of the elasticity of paid unemployment duration with respect to a variation in UI "benefit level" in the US. The classic study is for instance Meyer [1990], who found an elasticity of .56.[66] Landais [2015] indeed finds a slightly smaller elasticity (around .4). Kroft and Notowidigdo [2016] find an elasticity of .63 (.33) at the average unemployment rate in the US (p.20, based on estimates in Table 2 column 1) and conclude that such estimate "is broadly similar to the previous literature (Moffitt 1985, Meyer 1990, Chetty 2008)".

This evidence suggests that when benchmarked against conceptually similar elasticities, our duration responses prove quite similar to moral hazard estimates available in the literature.

[65] Of course, even if they are conceptually similar, the two elasticities do not need to be the same. One obvious reason they may differ is that these elasticities are potentially endogenous to $b_{\mathbf{2}}$, and $b_{\mathbf{2}} = 0$ in the US, while $b_{\mathbf{2}} = b_{\mathbf{1}}$ in Sweden.

[66] See Meyer [1990], Table VI, column (7). Coefficient estimates for log(b) in the proportional hazard models can be interpreted as the elasticity of the hazard rate s with respect to the benefit level. However, under the assumption that the hazard rate is somewhat constant, these elasticities can be interpreted as elasticities of unemployment duration, since $D \approx 1/s$ so that $\varepsilon_D \approx -\varepsilon_s$

FOR ONLINE PUBLICATION - Appendix C: Additional results and robustness of consumption analysis

This Appendix presents the construction of our residual measure of consumption expenditures based on registry data. It also presents additional results on consumption patterns over the unemployment spell and dynamic selection. It finally compares our results to results obtained using household consumption surveys (HUT), for which we can also explore patterns of consumption over the spell across different categories of expenditures.

C.1 Registry-based measures of consumption

We start by describing the construction of a registry-based consumption expenditure measure and we explain how it can be used it to present complementary analysis of how consumption evolves during the unemployment spell.

The registry-based measure of consumption is based on exhaustive administrative information on income, transfers and wealth in Sweden accounting for all income sources and changes in assets. The measure offers the advantage of being computable for the universe of unemployed households from 1999 to 2007. The measure captures annual expenditures between December of each year.

We start from the accounting identity that expenditures in year n are the sum of all income and transfers received in period n, minus the change in assets between year $n-1$ and year n,

$$expenditures_n = income_n - \Delta assets_n.$$

As a result of the comprehensiveness of the longitudinal administrative dataset that we assembled including all earnings, income, taxes, transfers and wealth, we have precise third-party reported information on all the components needed to construct such residual measure of yearly expenditures for the universe of Swedish individuals and households for years 1999 to 2007.

Our approach builds on previous attempts to measure consumption from registry-data (e.g. Browning and Leth-Petersen [2003]) and in particular on Koijen et al. [2014] who constructed a similar measure in Sweden for years 2003 to 2007 using a smaller subset of individuals, and confirmed its consistency with HUT data. Our approach extends these previous attempts and is closely related to Eika et al. [2017] in exploiting additional information on asset portfolio choices and returns to reduce measurement error and excess dispersion of consumption measures based solely on first-differencing asset stocks.

For interested readers, Kolsrud et al. [2017], (in preparation for a special issue of the *Journal of Public Economics* dedicated to the use of new data sources for consumption analysis), offers all the details of the construction of our residual measure of consumption. In particular, it provides information on all the registers, variables and programs needed to construct a reliable registry-based measure of consumption from the Swedish administrative data, and explains how to use flow transaction information on assets (real estate assets and financial assets) rather than first-differences in asset stocks over time to reduce measurement error.

In practice, we compute consumption in year n as:

$$C_n = y_n + T_n + \tilde{C}_n^b + \tilde{C}_n^d + \tilde{C}_n^v + \tilde{C}_n^h,$$

where:

- y_n represents all earnings and is computed from the tax registers, which contain third-party reported earnings for all employment contracts, including all fringe benefits and severance payments.[67]
- T_n accounts for all income taxes and transfers, including unemployment insurance, disability insurance, sick pay, housing and parental benefits, etc.

[67]Note that self-employed, for whom earnings are in large part self-reported, are excluded from the analysis as they are part of a different UI system.

- $\tilde{C}_n^b = y_n^b - \Delta b_n$ equals consumption out of bank holdings. It is equal to interests earned on these bank holdings during year n, y_n^b, minus the change in the value of bank holdings between year $n-1$ and year n, $\Delta b_n = b_n - b_{n-1}$.
- $\tilde{C}_n^d = -y_n^d + \Delta d_n$ is consumption out of debt, which includes student loans, credit card debt, mortgages, etc., and is third-party reported by financial institutions to the tax authority. It is equal to the change in the stock of debt Δd_n, minus all interests paid on the existing stock of debt y_n^d.
- $\tilde{C}_n^v = y_n^v - \Delta v_n$ is consumption out of financial assets (other than liquid holdings in bank accounts). It is equal to all income from financial assets y_n^v minus the change in the value of the portfolio of financial assets Δv_n. The return on financial assets y_n^v includes interests, dividends and any price change $\Delta p_n^v \times q_n^v$. Such price change would be exactly offset by a change in the value of assets, included in Δv_n, unless the return is realized by selling the asset. In practice, to separate the contribution of price changes from contribution of asset rebalancing, we use detailed data collected by Statistics Sweden from banks and financial institutions on all financial securities held by individuals, which contain information on quantities, ISIN numbers, and transaction values and dates.
- $\tilde{C}_n^h = y_n^h - \Delta h_n$ constitutes consumption out of real estate wealth. It is equal to all income derived from holding real estate assets y_n^h minus the change in the value of real estates Δh_n. Detailed information on the stock of real estate wealth, estimated at market value as of December 31 of each year, is available from the tax authority. The return to holding real estate y_n^h includes rents, but also imputed rents for homeowners, as well as price changes in the value of real estates. Such price change would again be exactly offset by a change in the value of real estate assets, included in Δh_n, unless the return is realized by selling the asset. To separate the contribution of price changes from contribution of real estate buying/selling, we use the housing transaction register which collects information on all real estate transactions operated in Sweden.

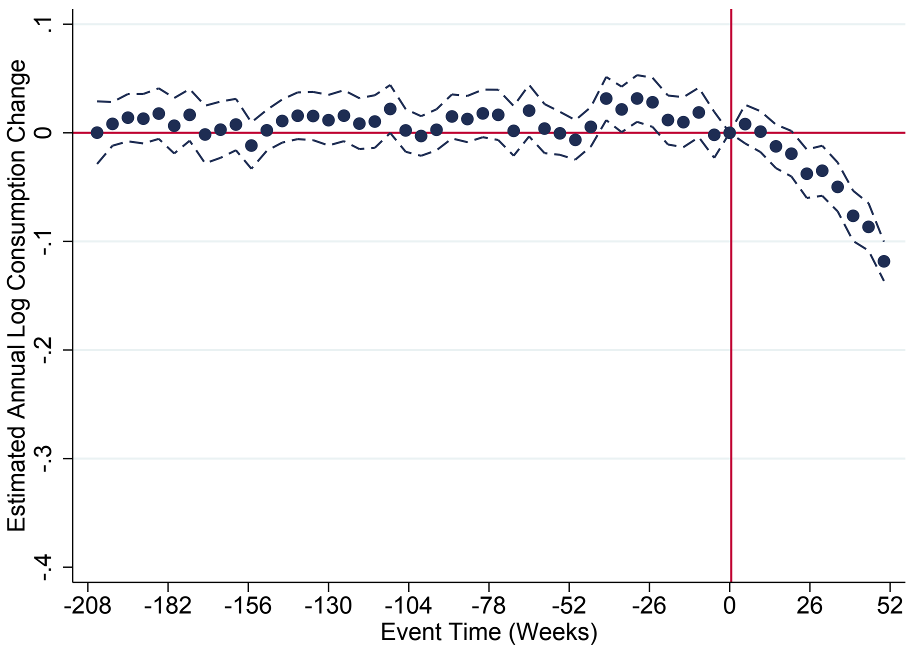
All income, transfers and asset positions are reported to the tax administration (and observed in the data) at the individual level.[68] We have aggregated consumption measures at the household level using household identifiers constructed by the Swedish National Statistical Office (Statistics Sweden).

C.2 Consumption patterns prior to the onset of a spell

To recover the drops in the first and second part of the profile from drops in annual consumption expenditures, our baseline implementation in section 4.1 makes the assumption that $c_j = \bar{c}_0, \forall j < 0$, or in other words, that the consumption profile is flat prior to the start of an unemployment spell. To investigate how consumption profiles evolve prior to the unemployment spell we report in Figure C-1 the annual log household consumption changes as a function of time t since the onset of a spell, $\mathscr{C}_{i,t} - \mathscr{C}_{i,t-52}$ in 5 weeks bins, relative to the last 5 weeks prior to the onset of a spell. We go as far back as 4 years prior to the onset of a spell, to detect longer anticipation effects. The Figure shows two interesting patterns. First, the consumption patterns of households are extremely stable in the 4 years prior to displacement, suggesting no long term anticipation behaviors. Second, there does not seem to be any sharp consumption changes immediately preceding displacement, suggesting no significant short term anticipation behaviors. Overall, this evidence strongly supports the assumption that flow consumption profiles are flat prior to the onset of an unemployment spell.

[68]It is possible that our measure misses some financial wealth held in foreign banks to the extent that these banks do not comply with the requirement to transmit information on the financial wealth of their Swedish customers to the Swedish tax authority. Yet, the fraction of foreign-held assets is low (about 3% of household assets according to the Savings Barometer of Statistics Sweden), and likely to be held by much wealthier households than our sample of unemployed.

Figure C-1: Change in log annual consumption as a function of time since onset of an unemployment spell: pre-unemployment trends



Notes: The Figure investigates the validity of the assumption of a flat profile of consumption pre-unemployment. We report the annual log household consumption changes as a function of time t since the onset of a spell, $\ln \mathscr{C}_{i,t} - \ln \mathscr{C}_{i,t-52}$ in 5 weeks bins, relative to the last 5 weeks prior to the onset of a spell. Estimates provide compelling evidence that consumption profiles are flat prior to the onset of an unemployment spell.

C.3 Additional Results on Consumption Smoothing Means & Dynamic Selection

Assets and liquidity constraints In this subsection, we discuss the evidence of the role of assets and liquidity constraints in consumption smoothing over the spell. First, we report in Appendix Figure C-2, and following the same methodology as in Figure 6, the average annual consumption drops by time spent unemployed as of December, breaking down the sample by the level of net wealth of the household at the start of the spell. The Figure shows that households with higher net wealth experience a lower drop in consumption conditional on unemployment duration. These households tend to have a smoother consumption profiles during unemployment, especially earlier on in the spell.

Second we provide direct evidence that individuals do use their liquid assets to smooth consumption over the spell. In Appendix Figure C-3 panel A, we estimate average change in liquid bank holdings compared to pre-unemployment by time t spent unemployed as of December (when bank holding stock is observed in the registry data). Estimates of the change in bank holdings $\mathscr{B}$ are scaled by the average annual consumption in the last year prior to unemployment, so that all changes in bank holdings are expressed relative to pre-unemployment household consumption levels. The Figure provides evidence that households use their liquid assets to smooth consumption over the unemployment spell, by depleting their bank accounts or reducing their savings. This source of consumption smoothing remains small in magnitude, as even after more than a year of unemployment, the change in bank holdings represent around 2% of pre-unemployment consumption levels of the household.

In panel B of Appendix Figure C-3 we report similar estimates for all other assets. The Figure shows that the change in total net assets other then bank holdings represent $\approx$2.5% of pre unemployment consumption after a year in unemployment.

Finally, evidence from registry data also indicates that debt does not offer much help in smoothing consumption over the unemployment spell, hinting at the presence of liquidity constraints. In Appendix Figure C-4, we provide evidence of a reduction in the use of non mortgage-related credit over the unemployment spell among households with no real estate and no mortgage debt. This Figure shows average change in debt compared to pre-unemployment by time t spent unemployed as of December (when debt level is observed in the registry data). The sample is restricted to individuals with no real estate wealth throughout the sample period. Because we cannot precisely separate mortgage debt from other types of credit in the data, this sample restriction is a direct way to identify how non-mortgage related debt evolves over the unemployment spell. The Figure provides evidence that households reduce their debt level rather than increase it as they become unemployed. Instead of contributing positively to consumption smoothing, debt contributes negatively to consumption over the unemployment spell. This suggests that as the duration of the spell increases, access to credit becomes harder and consumption out of debt falls significantly.

Overall, this analysis confirms that households may have means to smooth consumption over short spells. But as the duration of the spell increases, these means get quickly exhausted. Households in the second part of the profile seem therefore closer to being hand-to-mouth.

Spousal Labor Supply Within a household, the labor supply of other members of the household may help reduce the drop in household consumption over the spell. In Figure C-5 we investigate how labor supply of other members of the household affects the drop in household consumption over the spell. Following the same methodology as in Figure 6, we report the average change in total gross earnings and in total disposable income of all other members of the household as a function of time spent unemployed, scaled by the annual household consumption level prior to unemployment. Individual disposable income include all individual taxes and transfers, as well as individual capital gains and losses.

Results show that within-household changes in the earnings and disposable income of all other members of the household are small and almost always not significantly different from zero throughout the unemployment spell. This suggests that in our context, the labor supply of other members of the household does not play a significant role in smoothing household consumption even for long-term unemployed. This may be because the added-worker effect is not playing a significant role in increasing household consumption in response to an unemployment shock, or because its effect is mitigated by correlated earnings shocks across household members.

Dynamic Selection In practice, agents are heterogeneous and selection into longer unemployment spells may affect consumption responses over the spell and the gradient of consumption smoothing gains. To assess the potential magnitude of dynamic selection on consumption and risk preferences, we investigate in Table C-1 how various observable characteristics that have been shown to correlate with consumption and risk preferences are distributed across short term and long term unemployed.

To do so, we restrict the sample to all individuals about to become unemployed in the next quarter and estimate a linear probability model where the outcome is an indicator variable for experiencing a future spell longer than 20 weeks. The default age category is 18 to 30 years old. Income refers to individual disposable income and results are relative to the first quintile. Net wealth results are relative to individuals with zero or negative net wealth at the start of the spell. We also investigate the effect of two portfolio characteristics, that, conditional on net wealth, are traditionally correlated with risk preferences. First, we look at the fraction of total wealth invested in stocks, and results are relative to the first two quartile of this distribution (50% of the sample have zero stocks prior to becoming unemployed). Second, we look at leverage defined as total debt divided by gross assets, and results are relative to the first quartile of leverage.

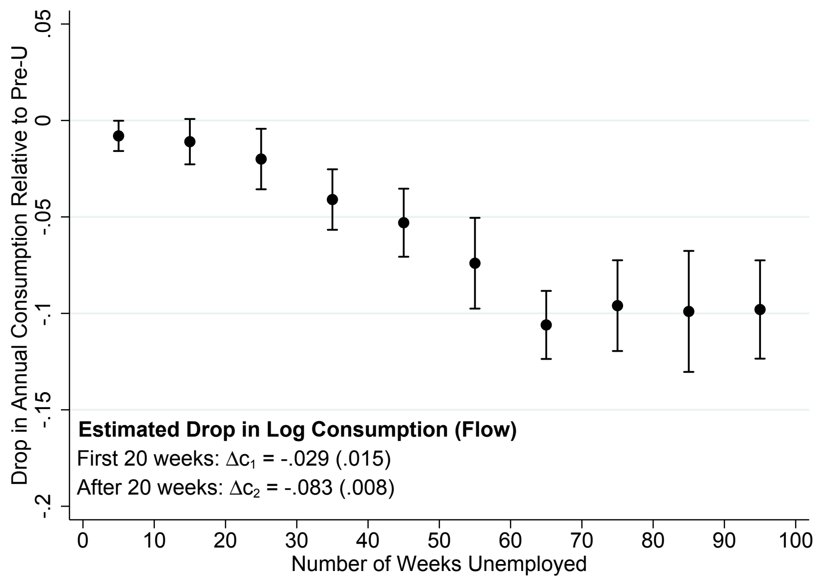
The patterns of selection are generally small in magnitude and often ambiguous in sign. Income levels and net wealth levels have quantitatively small and non-monotonic effects on the probability to select into longer spells. Compared to households with no or negative net wealth, households with some small wealth (<500kSEK) have a slightly lower probability to be long term unemployed. But individuals with high net wealth have a slightly higher probability to select into long spells. This result suggests that in our context, individuals with better means to smooth consumption do not unambiguously select into longer spells, which corroborates the evidence of no clear patterns of selection on consumption profiles displayed in subsection 4.1. Also portfolio characteristics (i.e., the fraction of portfolio wealth invested in stocks, and leverage defined as total debt divided by gross assets), which have been shown to be correlated with risk preferences, have small and non-monotonic impacts on the probability to experience a long unemployment spell. In contrast, the probability of experiencing long unemployment spells ($\mathbb{1}[L > 20 \text{ wks}]$) is significantly and monotonically correlated with age. However, existing evidence from the literature suggests a U-shape relationship between age and risk aversion (Cohen and Einav [2007]), so the dynamic selection on age has an ambiguous effect on the evolution of risk preferences over the spell.

Overall, the observed dynamic selection patterns on consumption and risk preferences do not suggest that relative consumption smoothing gains would be significantly different than our estimates based on the average drops in consumption.[69]

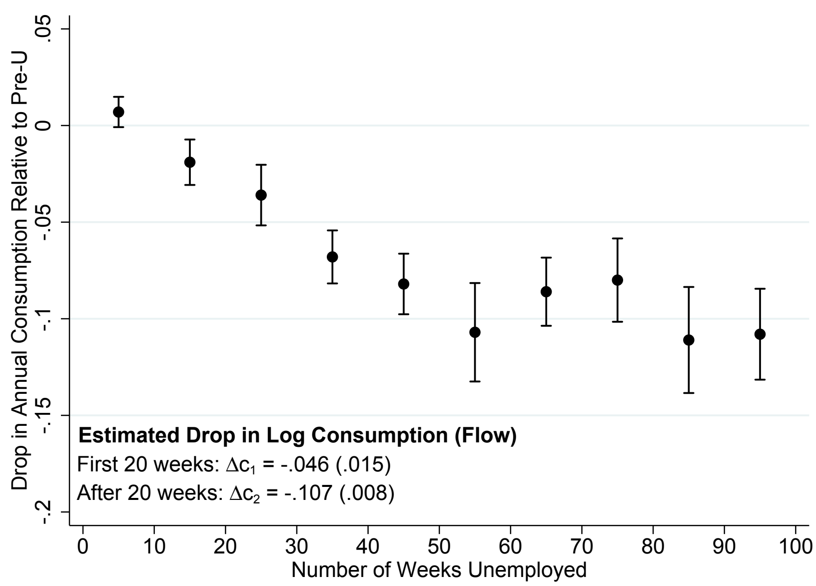
[69] Yet, other methods exploiting comparative statics of effort choices as in Chetty [2008a], or Landais [2015], could also be developed to evaluate the evolution of consumption smoothing gains over the spell. These methods could circumvent the issue of having to make assumptions regarding dynamic selection on risk preferences.

Figure C-2: Estimated Drop in Annual Consumption Relative to Pre-Unemployment as a Function of Time Spent Unemployed: Heterogeneity By Level of Wealth

A. Above Median Wealth Level Pre-Unemployment

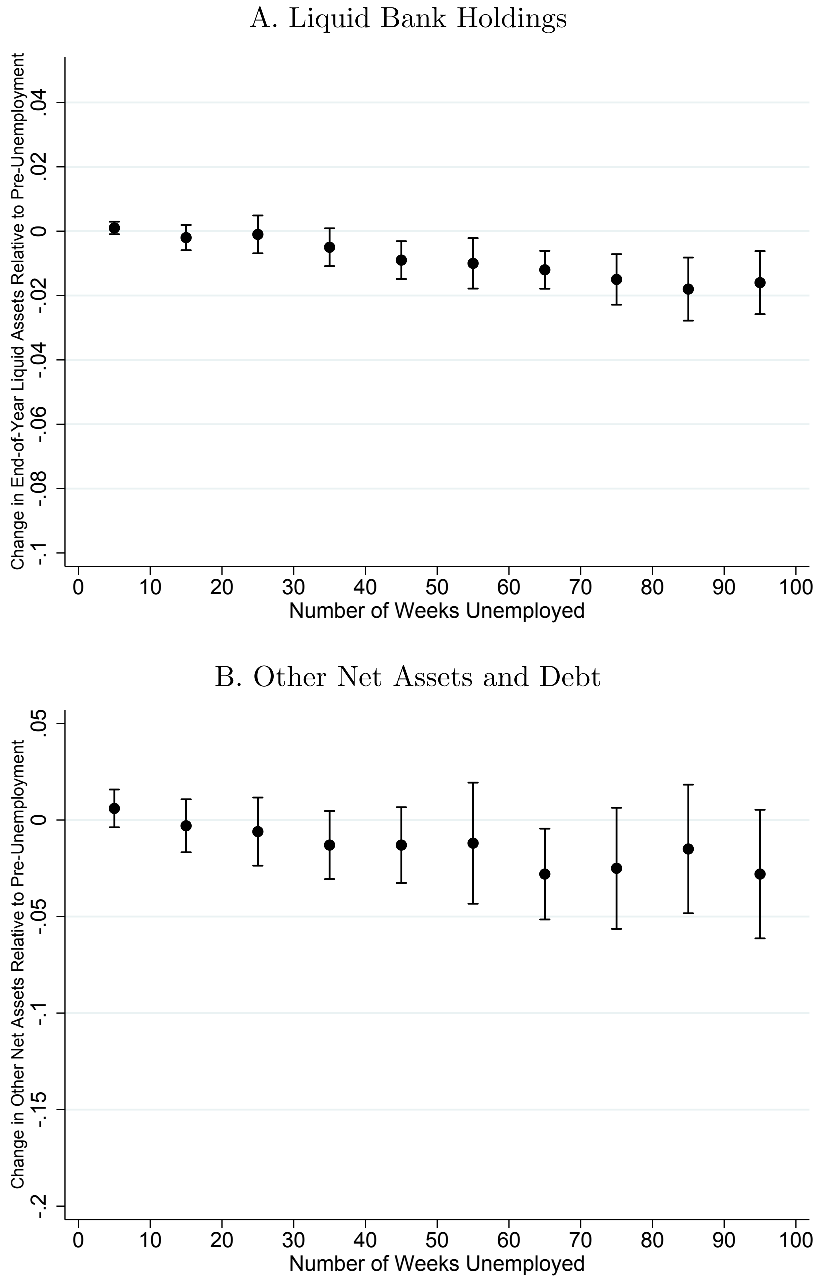


B. Below Median Wealth Level Pre-Unemployment



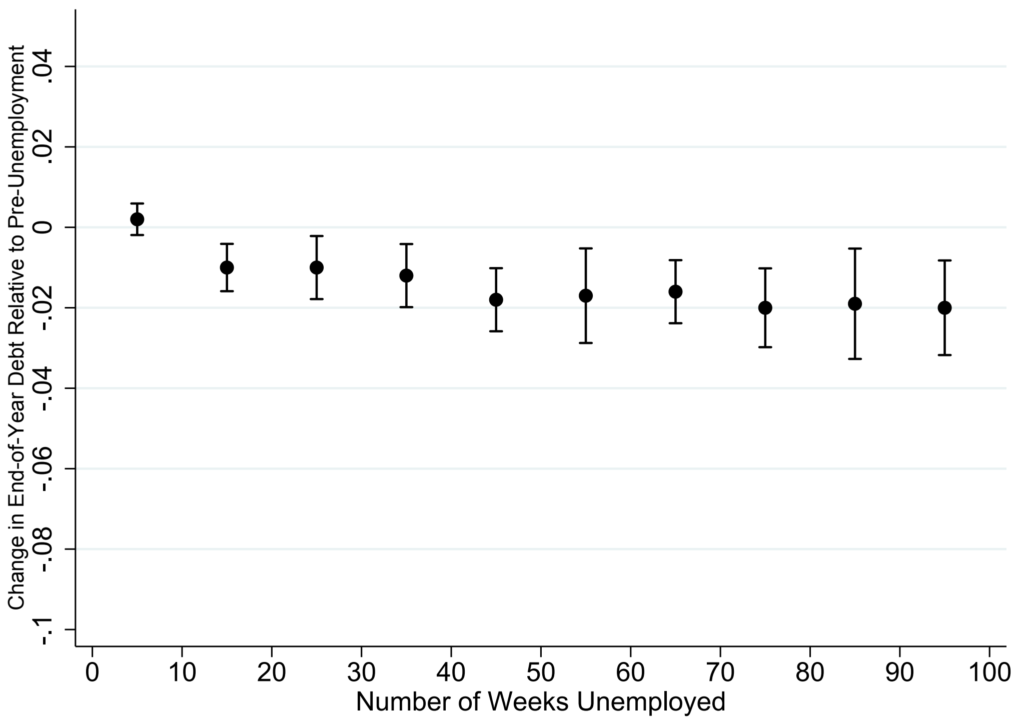
Notes: This Figure shows average annual consumption drops compared to pre-unemployment $\Delta\mathscr{C}_t$ by time t spent unemployed as of December (when annual consumption is observed in the registry data), and compares consumption profiles for households according to their level of net wealth prior to becoming unemployed. In both panels, we report $\hat{\beta}_t/\mathscr{C}_0 = \widehat{\Delta\mathscr{C}_t}/\mathscr{C}_0$, i.e. estimates from equation (20) scaled by the average annual consumption in the last year prior to unemployment, so that all consumption drops are expressed relative to pre-unemployment levels. The Figure also reports non-parametric estimates of the average drops in consumption in each parts of the benefit profile $\widehat{\Delta C_1}$ and $\widehat{\Delta C_2}$ following the methodology explained in subsection 4.1. Standard errors are computed using the Delta-method. Panel A reports estimates for households whose net wealth level in the last year prior to becoming unemployed is above the median wealth level in the sample. Panel B reports estimates for households whose net wealth level in the last year prior to becoming unemployed is below the median wealth level in the sample. The Figure provides evidence that households with higher net wealth experience smaller consumption drops conditional on unemployment duration.

Figure C-3: Estimated Change in Liquid Bank Holdings and Other Net Assets Relative to Pre-Unemployment as a Function of Time Spent Unemployed



Notes: This Figure shows average change in household liquid bank holdings $\mathscr{B}$ (panel A.) and other net assets $\mathscr{A}$ (panel B.) compared to pre-unemployment by time t spent unemployed as of December (when bank holding stock and other assets are observed in the registry data). We follow the same specification as for consumption drops (from equation (20)). In panel A, we report $\hat{\beta}_t/\mathscr{C}_0 = \widehat{\Delta\mathscr{B}_t}/\mathscr{C}_0$, i.e. estimates of the change in bank holdings $\mathscr{B}$ scaled by the average annual consumption in the last year prior to unemployment, so that all changes in bank holdings are expressed relative to pre-unemployment household consumption levels. In panel B we report $\hat{\beta}_t/\mathscr{C}_0 = \widehat{\Delta\mathscr{A}_t}/\mathscr{C}_0$. The Figure provides evidence that households use their assets to smooth consumption over the unemployment spell. This source of consumption smoothing remains small in magnitude. After more than a year of unemployment, the change in bank holdings represent 2% of pre-unemployment consumption levels of the household. Other net assets provide an additional 2.5% consumption smoothing after a year of unemployment.

Figure C-4: Estimated Change in Non-Mortgage Debt Relative to Pre-Unemployment as a Function of Time Spent Unemployed

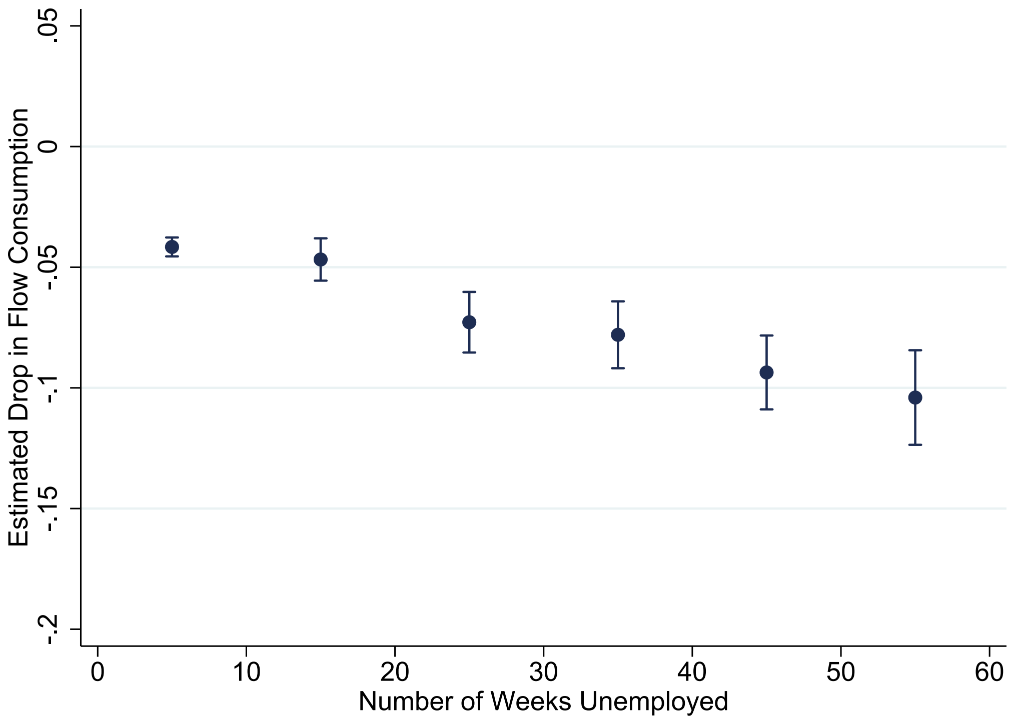


Notes: This Figure shows average change in debt compared to pre-unemployment by time t spent unemployed as of December (when debt level is observed in the registry data). The sample is restricted to individuals with no real estate wealth throughout the sample period. Because we cannot precisely separate mortgage debt from other types of credit in the data, this sample restriction is a direct way to identify how non-mortgage related debt evolves over the unemployment spell. We follow the same specification as for consumption drops (from equation (20)) and report $\hat{\beta}_t/\mathscr{C}_0 = \widehat{\Delta\mathscr{D}_t}/\mathscr{C}_0$, i.e. estimates of the change in debt $\mathscr{D}$ scaled by the average annual consumption in the last year prior to unemployment, so that all changes in debt are expressed relative to pre-unemployment household consumption levels. The Figure provides evidence that households reduce their debt level rather than increase it as they become unemployed. Instead of contributing positively to consumption smoothing, debt contributes negatively to consumption over the unemployment spell, which is suggestive of the presence of credit constraints among unemployed individuals.

Figure C-5: Estimated Change in Earnings and Disposable Income of all other members of the household as a function of unemployment duration

Notes: This figure explores the role of other household members (i.e. all members of the household excluding the unemployed individual) in smoothing household consumption over the spell. We follow the same specification as for consumption drops (from equation (20)) and report $\hat{\beta}_t/\mathscr{C}_0 = \widehat{\Delta\mathscr{Y}_t}/\mathscr{C}_0$, i.e. estimates of the change in gross earnings or disposable income $\mathscr{Y}$ of all the other household members scaled by the average annual consumption in the last year prior to unemployment, so that all changes are expressed relative to pre-unemployment household consumption levels. Disposable income includes individual taxes and transfers, as well as capital gains/losses. Results show that within-household changes in earnings and disposable income of all other members of the household are extremely small throughout the unemployment spell.

Figure C-6: Estimated drop in flow consumption over the first year of the unemployment spell



Notes: The Figure reports non-parametric estimates of the average drops in flow consumption at 10 week frequency following the methodology explained in subsection 4.1. Standard errors are computed using the Delta-method.

Table C-1: Pre-unemployment characteristics of individuals with spells longer than 20 weeks. Linear probability model estimates

	(1)	(2)	(3)	(4)	(5)
	Duration of future spell $\geq$ 20 weeks				
Age: 30 to 39	0.129***	0.118***	0.116***	0.119***	0.120***
	(0.00237)	(0.00250)	(0.00251)	(0.00305)	(0.00311)
Age: 40 to 49	0.164***	0.153***	0.153***	0.162***	0.163***
	(0.00277)	(0.00293)	(0.00295)	(0.00357)	(0.00363)
Age: 50+	0.272***	0.261***	0.265***	0.281***	0.282***
	(0.00288)	(0.00307)	(0.00319)	(0.00367)	(0.00371)
Married	0.0289***	0.0283***	0.0287***	0.0185***	0.0190***
	(0.00243)	(0.00243)	(0.00244)	(0.00280)	(0.00281)
Gender: Female	-0.00226	-0.00209	-0.00279	-0.0146***	-0.0135***
	(0.00192)	(0.00193)	(0.00193)	(0.00230)	(0.00230)
# of children	-0.0329***	-0.0315***	-0.0288***	-0.0318***	-0.0311***
	(0.00193)	(0.00193)	(0.00193)	(0.00231)	(0.00233)
2nd quintile of income		0.0412***	0.0436***	0.0321***	0.0319***
		(0.00319)	(0.00319)	(0.00409)	(0.00412)
3rd quintile of income		0.0842***	0.0885***	0.0850***	0.0849***
		(0.00319)	(0.00319)	(0.00403)	(0.00415)
4th quintile of income		0.0471***	0.0532***	0.0532***	0.0545***
		(0.00328)	(0.00329)	(0.00404)	(0.00421)
5th quintile of income		0.0453***	0.0518***	0.0589***	0.0635***
		(0.00341)	(0.00345)	(0.00411)	(0.00431)
$0<$ Net wealth $\leq$ 200k			-0.0503***	-0.0116***	-0.0122***
			(0.00234)	(0.00271)	(0.00315)
200k $<$ Net wealth $\leq$ 500k			-0.0466***	-0.0146***	-0.0114***
			(0.00324)	(0.00350)	(0.00425)
500k $<$ Net wealth $\leq$ 5M			-0.0186***	0.00576*	0.00774*
			(0.00300)	(0.00336)	(0.00418)
Net wealth $>$ 5M			0.0731***	0.0852***	0.0866***
			(0.0173)	(0.0172)	(0.0174)
Fraction of portfolio in stocks					
3rd quartile				-0.000542	
				(0.00787)	
4th quartile				0.0303***	
				(0.00259)	
Leverage: debt / assets					
2nd quartile					0.0153***
					(0.00390)
3rd quartile					-0.0120***
					(0.00322)
4th quartile					-0.00629*
					(0.00361)
R^2	0.0465	0.0490	0.0511	0.0624	0.0620
N	269931	269931	269931	190176	190176

Notes: The Table assesses the robustness of our welfare conclusions to dynamic selection on risk preferences over the unemployment spell. We investigate how various observable characteristics correlate with the probability to experience a long unemployment spell. To do so, we restrict the sample to all individuals about to become unemployed in the next quarter and estimate a linear probability model where the outcome is an indicator variable for experiencing a future spell longer than 20 weeks. The default age category is 18 to 30 years old. Income refers to individual disposable income and results are relative to the first quintile. Net wealth results are relative to individuals with zero or negative net wealth at the start of the spell. We also investigate the effect of two portfolio characteristics, that, conditional on net wealth, are traditionally correlated with risk preferences. First, we look at the fraction of total wealth invested in stocks, and results are relative to the first two quartile of this distribution (50% of the sample have zero stocks prior to becoming unemployed). Second, we look at leverage defined as total debt divided by gross assets, and results are relative to the first quartile of leverage.

C.4 Consumption profiles over the spell using household consumption surveys (HUT)

In order to analyze the evolution of consumption as a function of time spent unemployed, we have also merged the household consumption surveys (HUT) with the universe of administrative UI records. This enables us to reconstruct the full employment history of individuals whose household is surveyed in the HUT. We observe employment status of all individuals prior, during and after their HUT interview.

We restrict the sample to households where an individual is either unemployed at the time of the interview, or who will become unemployed some time in the next two years following the interview. This leaves us with a pseudo-panel of households for which we can correlate flow measures of consumption with time since (or until) the onset of the unemployment spell. Note that this sample is a pseudo-panel and not a panel *stricto sensu*, as households are surveyed only once in the HUT. However, because we observe the full unemployment history of individuals irrespective of the time they are surveyed, we can fully control for selection issues arising from differences between households who select into spells of different lengths.

In Table C-2, we provide summary statistics for unemployed individuals from the HUT sample. Unemployment duration patterns are almost identical in the HUT sample and in the RKD sample, with an average duration of unemployment of 26.6 weeks, and an average time spent in the first (resp. second) part of the profile of 12.9 weeks (resp. 13.8 weeks). Average replacement rates are also identical at 72%. Interestingly, unemployed individuals have similar demographics in the two samples. The distribution of age, gender, marital status and education levels are almost identical. Average earnings are also similar in the two samples, but the distribution of earnings is by construction much more spread out in the HUT sample than in the RKD sample, because the RKD sample is focused on unemployed with daily wages in the neighborhood of 725SEK. Finally, the two samples have very comparable distribution of assets and debts. In both samples, unemployed individuals have a household net wealth equal to 2.5 times their yearly earnings at the start of the spell. Overall, samples of unemployed individuals in the HUT and RKD analysis are well-balanced in terms of all observable characteristics and unemployment durations.

We start by providing graphical evidence for the evolution of average consumption as a function of time spent unemployed. We report in Figure C-7 the estimated coefficients β_k from the regression model:

$$c_{i,t} = \sum_{k=-3}^{+4} \beta_k \cdot \mathbb{1}[t \in [26 * k, 26 * \{k+1\})] + X_i'\gamma + \varepsilon_{i,t} \tag{28}$$

where $c_{i,t}$ is log household consumption in the HUT survey observed at the time of the HUT interview. t is event time in weeks since the start of the unemployment spell. We aggregate event time in bins of 6 months periods (26 weeks), so that $\mathbb{1}[t \in [26 * k, 26 * \{k+1\})]$ is an indicator for being observed in the k-th 6 months period since the start of an unemployment spell.[70] We include in this regression a set of controls X, which consists of year dummies, calendar month dummies and a set of dummies for family status. The estimated coefficients β_k plotted in Figure C-7 represent the average log household consumption for households with an individual observed in her $26 * k$-th to $26 * (k+1)$-th week of unemployment, relative to the average consumption level of households with an individual observed just prior to becoming unemployed.

The graph provides strong evidence corroborating our findings from the registry-based measure of consumption, that average household consumption drops significantly when unemployed. The average consumption of households where a member has been unemployed for more than a year is almost 15% lower than prior to the unemployment spell. Furthermore, Figure C-7 indicates that consumption declines significantly throughout the unemployment spell. The consumption drop is still limited early on in the spell, but average consumption decreases sharply as duration increases, and seems to be reaching a plateau after about a year in the unemployment spell. The last interesting finding is the lack of anticipation prior to the unemployment spell. Household consumption is flat in the year

[70] $k = 0$ is the baseline and represents consumption in the last 6 months prior to becoming unemployed. HUT surveys collect information on bi-weekly household consumption expenditures at the time of the interview. $c_{i,t}$ is then annualized by multiplying bi-weekly consumption by 26.

preceding the onset of the unemployment spell. This also confirms evidence from the registry data and suggests that unemployment shocks are relatively unanticipated, or that households have little propensity to change consumption prior to becoming unemployed to smooth out the upcoming earnings shock.

To recover the average consumption drops in the first and second part of the unemployment spell and to investigate the role of selection, we move in Table C-3 to a regression setting. In column (1), we start by running the simple regression:

$$c_{i,t} = \eta_1 \cdot \mathbb{1}[0 < t <= 20 \text{ wks}] + \eta_2 \cdot \mathbb{1}[t > 20 \text{ wks}] + X_i'\gamma + \varepsilon_{i,t}, \tag{29}$$

where $c_{i,t}$ is log consumption of the household and η_1 (resp. η_2) captures the effect of having a member of the household unemployed for less than 20 weeks (resp. more than 20 weeks) at the time of the interview, relative to households where one individual will become unemployed in the next 6 months. The baseline specification reported in column (1) controls for year dummies and calendar month dummies. Results suggest that relative to pre-unemployment levels, total average household consumption drops by 6% in the first 20 weeks of unemployment, and by 13% after 20 weeks of unemployment.

In column (2) of Table C-3, we add non-parametric controls for marital status, household size and age, and find similar results with an initial drop of 4% in the first 20 weeks and 13% after 20 weeks of unemployment. In columns (3) and (4), we analyze the role of selection in explaining these consumption patterns. If households with long unemployment spells are inherently different from households with short unemployment spells, the estimated drops in consumption after 20 weeks in column (1) and (2) may partly pick up these differences. To investigate such selection effects, we estimate in column (3) a model of the form

$$c_{i,t} = \eta_1 \cdot \mathbb{1}[0 < t <= 20 \text{ wks}] + \eta_2 \cdot \mathbb{1}[t > 20 \text{ wks}] + \alpha \cdot \mathbb{1}[L > 20 \text{ wks}] + X_i'\gamma + \varepsilon_{i,t}, \tag{30}$$

where $\mathbb{1}[L > 20 \text{ wks}]$ is a dummy variable for the completed length of the unemployment spell being longer than 20 weeks. In other words, because we observe in the UI records the total duration of each spell irrespective of the time when the household is surveyed in the HUT, we can control for consumption differences at any point in time between households with long versus short spells. Results indicate that between households who select into spells longer than 20 weeks versus households who select into spells shorter than 20 weeks, differences in log consumption levels are small and not significant.

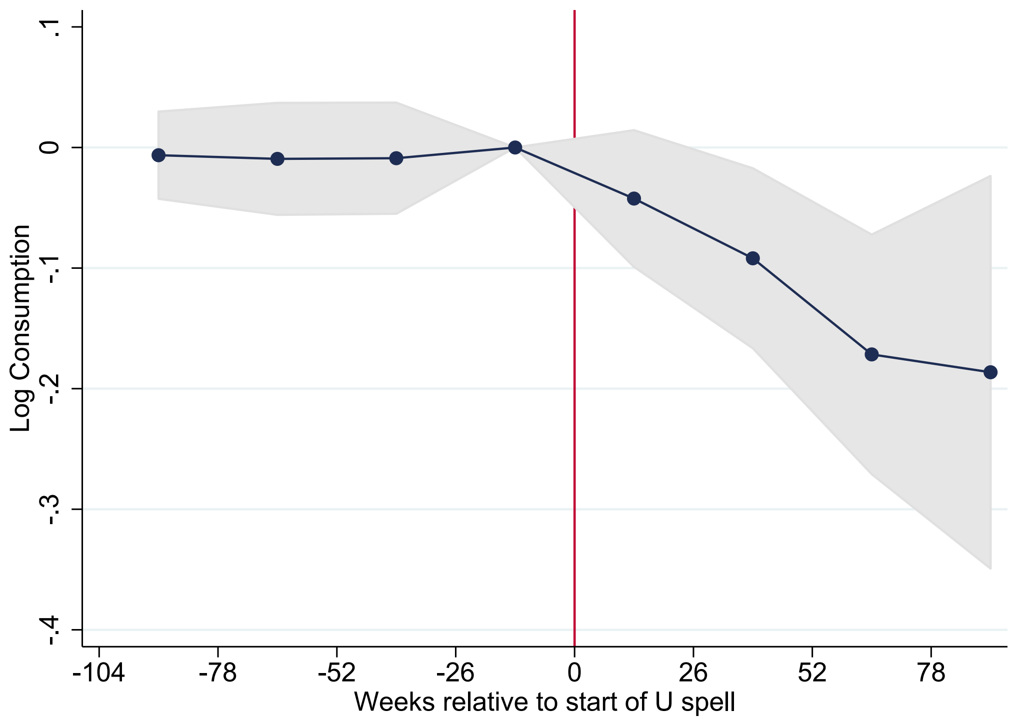
In column (4), we estimate not only differences in levels but also differences in consumption profiles between households who select into spells longer than 20 weeks versus households who select into spells shorter than 20 weeks. To do so, we fully interact $\mathbb{1}[L > 20 \text{ wks}]$ with the event time dummies to estimate two separate consumption profiles over the unemployment spell: one for households with long spells and one for households with short spells. Results reported in column (4) indicate that the consumption drop in the first 20 weeks for households who select into long spells is 1.3% smaller but not statistically different from the consumption drop in the first 20 weeks of households who select into short spells[71]. This evidence suggest that differences in consumption profiles are relatively small and insignificant, so that selection effects are not significantly driving the observed patterns of average household consumption over the unemployment spell.

Tables C-5 and C-6 replicate the analysis of, respectively, column 3 and 4 of Table C-3 for various categories of consumption available in the HUT data. These tables are aimed at gauging how the difference between expenditures and actual consumption may affect our conclusions regarding the evolution of consumption smoothing gains over the unemployment spell, as explained in section 4.2 . The goal is to detect whether individuals who select into long versus short spells have access to different means of smoothing consumption using substitution across different expenditure categories. Results indicate that the expenditure levels and expenditure profiles over food, recreation, transportation, or restaurants are not significantly different for households that select into long versus short spells. This suggests that there is no significant dynamic selection over the spell based on the availability of substitution towards home production. Equivalently, expenditure levels and expenditure profiles over durable goods are not significantly different

[71] To control for length-biased sampling, we also implemented all specifications using reweighting approaches, and with more flexible controls for longer durations L. Results are identical, and available upon request.

for households that select into long versus short spells. The availability of consumption smoothing through shifting expenditures away from durables is not significantly different for households that select into long versus short spells.[72]

Figure C-7: Log household consumption relative to last semester prior to the unemployment spell



Notes: The figure correlates household consumption measured in the HUT survey with the time t since (or until) the onset of the unemployment spell observed in the administrative UI records. Bi-weekly household consumption levels measured at the moment of the HUT interview are annualized and expressed in constant SEK2003. The figure follows from regression model (28) and plots the estimated coefficients β_k for the set of indicators $\mathbb{1}[t \in [26*k, 26*\{k+1\})]$ for being observed in the k-th 6 months period since the onset of one's spell. We also plot the 95% confidence interval from robust standard errors. We include in this regression a set of controls X, which consists of year dummies, calendar months dummies and a set of dummies for family status. The graph provides evidence that average household consumption drops over the unemployment spell. The average consumption of households where a member has been unemployed for a full year is more than 15% lower than the average consumption of household with a member at the start of her spell.

[72] It should be noted that the small sample size in the HUT offers limited statistical power and that it would be interesting to get further evidence on these various expenditure patterns.

Table C-2: Summary statistics at start of unemployment spell: HUT sample

	Mean	P10	P50	P90
	I. Unemployment			
Duration of spell (wks)	26.64	2.86	13.43	65.29
Duration on b_1 (wks)	12.87	2.86	13.43	20
Duration on b_2 (wks)	13.78	0	0	45.29
Replacement rate	.72	.59	.79	.8
	II. Demographics			
Age	34.12	21	33	51
Fraction men	.49	0	0	1
Fraction married	.39	0	0	1
Fraction with higher educ	.26	0	0	1
	III. Income and Wealth, SEK 2003(K)			
Gross earnings (individual)	202.9	9.8	172.6	386.2
Household disposable income	354.4	116.9	330.1	585.3
Household consumption	343	150.3	305.1	572.6
Household net wealth	510.1	-258.3	0	1691.6
Household bank holdings	65.6	0	0	149.8
Household real estate	770.7	0	44	1948.3
Household debt	427.2	0	193.3	1154.3

Notes: The table provides summary statistics for our sample of households unemployed from the HUT consumption surveys 2003 to 2009. To create this sample, we merged the household consumption surveys with the universe of administrative UI records to reconstruct the full employment history of individuals whose household is surveyed in the HUT. We then restrict the sample to households where an individual is either unemployed at the time of the interview, or who will become unemployed some time in the next two years following the interview. This leaves us with a pseudo-panel of about 6,500 households. All earnings, income and asset level measures are from wealth and income registers, and are yearly measures aggregated at the household level in constant k2003SEK for the last calendar year of full employment prior to the start of the unemployment spell. Disposable income is gross earnings, plus capital income minus all taxes plus transfers received. Transfers include unemployment insurance, disability insurance, sick pay, and all housing and parental benefits. All financial assets are estimated at their market value. Real estate is gross of debt and assessed at market value. Debt includes student loans, mortgage, credit card debt, etc. Note that 1SEK2003 $\approx$ 0.11USD2003.

Table C-3: Household log-consumption as a function of time spent unemployed relative to pre-unemployment consumption: consumption survey estimates

	(1)	(2)	(3)	(4)
$\mathbb{1}[0 < t \leq 20 \text{ weeks}]$ (η_1)	-0.0606*	-0.0444	-0.0408	-0.0465
	(0.0316)	(0.0311)	(0.0314)	(0.0413)
$\mathbb{1}[t > 20 \text{ weeks}]$ (η_2)	-0.130***	-0.129***	-0.111***	-0.108***
	(0.0328)	(0.0332)	(0.0386)	(0.0414)
$\mathbb{1}[L > 20 \text{ weeks}]$			-0.0294	-0.0342
			(0.0308)	(0.0378)
$\mathbb{1}[0 < t \leq 20 \text{ weeks}] \times \mathbb{1}[L > 20 \text{ weeks}]$				0.0134
				(0.0629)
Year F-E	×	×	×	×
Calendar months F-E	×	×	×	×
Marital status		×	×	×
Family size		×	×	×
Age group F-E		×	×	×
Test $\eta_1 = \eta_2$				
P-value	**0.06**	**0.02**	**0.08**	**0.10**
R^2	0.0493	0.0866	0.0872	0.0872
N	1551	1548	1548	1548

Notes: Robust standard errors in parentheses. * p<.10, ** p<.05, *** p<.01.
The Table reports estimates of the drop in household consumption over the unemployment spell in the HUT surveys, following model of equation (29). Because the HUT surveys collect information on household consumption expenditures at the time of the interview, HUT estimates can directly recover flow (bi-weekly) measures of consumption c_t. We restrict the sample to households where, at the date of the interview, one (and only one) individual is unemployed, or where, at the date of the interview, one (and only one) individual will become unemployed in the following 6 months. $\mathbb{1}[0 < t <= 20 \text{ weeks}]$ is an indicator for having a member of the household unemployed since less than 20 weeks at the time of the interview. $\mathbb{1}[t > 20 \text{ weeks}]$ is an indicator for having a member of the household unemployed for more than 20 weeks at the time of the interview. $\mathbb{1}[L > 20 \text{ weeks}]$ is an indicator for the total duration of the unemployment spell being longer than 20 weeks. We also report the P-value of a test of equality of η_1 (the drop in consumption in the first 20 wks) and η_2 (the drop in consumption after 20 wks).

Table C-4: Consumption as a function of time spent unemployed relative to pre-unemployment consumption: consumption survey estimates for various expenditure categories

	(1) Total expenditures	(2) Food	(3) Rents	(4) Purchase of new vehicles	(5) Furniture & house appliances	(6) Trans-portation	(7) Recre-ation	(8) Restau-rant
$\mathbb{1}[0 < t \leq 20 \text{ weeks}]$	-0.0606*	-0.0441	-0.0404	-0.418**	-0.160	-0.0788	-0.106	-0.0807
	(0.0316)	(0.0388)	(0.0380)	(0.187)	(0.102)	(0.0661)	(0.0649)	(0.0876)
$\mathbb{1}[t > 20 \text{ weeks}]$	-0.130***	-0.0823*	0.0430	-0.252	-0.0883	-0.348***	-0.189***	-0.165*
	(0.0328)	(0.0441)	(0.0310)	(0.176)	(0.0884)	(0.0803)	(0.0719)	(0.0888)
Year fixed effects	×	×	×	×	×	×	×	×
Calendar months F- E	×	×	×	×	×	×	×	×
R^2	0.0493	0.0650	0.0365	0.0205	0.00975	0.0208	0.0252	0.0154
N	1551	1548	798	982	1548	1488	1543	1119

Notes: Robust standard errors in parentheses. * p<.10, ** p<.05, *** p<.01.
The Table reports estimates of the drop in household consumption over the unemployment spell in the HUT surveys, following model of equation (29). Because the HUT surveys collect information on household consumption expenditures at the time of the interview, HUT estimates can directly recover flow (bi-weekly) measures of consumption c_t. We restrict the sample to households where, at the date of the interview, one (and only one) individual is unemployed, or where, at the date of the interview, one (and only one) individual will become unemployed in the following 6 months. $\mathbb{1}[0 < t <= 20 \text{ weeks}]$ is an indicator for having a member of the household unemployed since less than 20 weeks at the time of the interview. $\mathbb{1}[t > 20 \text{ weeks}]$ is an indicator for having a member of the household unemployed for more than 20 weeks at the time of the interview. $\mathbb{1}[L > 20 \text{ weeks}]$ is an indicator for the total duration of the unemployment spell being longer than 20 weeks.

Table C-5: Log-consumption as a function of time spent unemployed relative to pre-unemployment consumption: consumption survey estimates for various expenditure categories, selection on levels

	(1) Total expenditures	(2) Food	(3) Rents	(4) Purchase of new vehicles	(5) Furniture & house appliances	(6) Trans-portation	(7) Recre-ation	(8) Restau-rant
$\mathbb{1}[0 < t \leq 20$ weeks$]$	-0.0379	-0.0139	-0.0212	-0.192	-0.0657	-0.0596	-0.0948	-0.0765
	(0.0305)	(0.0366)	(0.0368)	(0.172)	(0.0893)	(0.0658)	(0.0649)	(0.0876)
$\mathbb{1}[t > 20$ weeks$]$	-0.113***	-0.0782*	0.0143	-0.387*	-0.0463	-0.330***	-0.146*	-0.141
	(0.0379)	(0.0463)	(0.0356)	(0.217)	(0.102)	(0.0925)	(0.0850)	(0.104)
$\mathbb{1}[L > 20$ weeks$]$	-0.0294	-0.00917	0.0239	0.0106	-0.0394	-0.0223	-0.0775	-0.0208
	(0.0300)	(0.0364)	(0.0312)	(0.159)	(0.0848)	(0.0659)	(0.0654)	(0.0811)
Year fixed effects	×	×	×	×	×	×	×	×
Marital status	×	×	×	×	×	×	×	×
Family size	×	×	×	×	×	×	×	×
R^2	0.139	0.194	0.157	0.0686	0.0442	0.0525	0.0567	0.0369
N	1548	1545	796	606	1523	1485	1540	1116

Notes: Robust standard errors in parentheses. * p<.10, ** p<.05, *** p<.01.
The Table reports estimates of the drop in household consumption over the unemployment spell in the HUT surveys, following model of equation (29). Because the HUT surveys collect information on household consumption expenditures at the time of the interview, HUT estimates can directly recover flow (bi-weekly) measures of consumption c_t. We restrict the sample to households where, at the date of the interview, one (and only one) individual is unemployed, or where, at the date of the interview, one (and only one) individual will become unemployed in the following 6 months. $\mathbb{1}[0 < t <= 20$ weeks$]$ is an indicator for having a member of the household unemployed since less than 20 weeks at the time of the interview. $\mathbb{1}[t > 20$ weeks$]$ is an indicator for having a member of the household unemployed for more than 20 weeks at the time of the interview. $\mathbb{1}[L > 20$ weeks$]$ is an indicator for the total duration of the unemployment spell being longer than 20 weeks.

Table C-6: Log-consumption as a function of time spent unemployed relative to pre-unemployment consumption: consumption survey estimates for various expenditure categories, selection on profile

	(1) Total expenditures	(2) Food	(3) Rents	(4) Purchase of new vehicles	(5) Furniture & house appliances	(6) Trans-portation	(7) Recre-ation	(8) Restau-rant
$\mathbb{1}[0 < t \leq 20 \text{ weeks}]$	-0.0465	-0.0320	-0.00595	-0.288	0.0324	-0.0457	-0.126	-0.107
	(0.0413)	(0.0527)	(0.0456)	(0.223)	(0.114)	(0.0920)	(0.0879)	(0.112)
$\mathbb{1}[t > 20 \text{ weeks}]$	-0.108***	-0.0666	0.00371	-0.347	-0.0976	-0.328***	-0.127	-0.117
	(0.0414)	(0.0528)	(0.0442)	(0.218)	(0.116)	(0.0935)	(0.0881)	(0.115)
$\mathbb{1}[L > 20 \text{ weeks}]$	-0.0342	-0.0258	0.0440	-0.0582	0.0441	-0.0147	-0.0967	-0.0410
	(0.0378)	(0.0482)	(0.0409)	(0.193)	(0.105)	(0.0848)	(0.0804)	(0.102)
$\mathbb{1}[L > 20 \text{ weeks}] \times \mathbb{1}[0 < t \leq 20 \text{ weeks}]$	0.0134	0.0265	-0.0509	0.022	-0.024	-0.0239	0.0750	0.0990
	(0.0629)	(0.0800)	(0.0691)	(0.342)	(0.175)	(0.141)	(0.134)	(0.172)
Year fixed effects	×	×	×	×	×	×	×	×
Marital status	×	×	×	×	×	×	×	×
Family size	×	×	×	×	×	×	×	×
R^2	0.0872	0.107	0.0765	0.0689	0.0273	0.0373	0.0348	0.0246
N	1548	1545	796	606	1523	1485	1540	1116

Notes: Robust standard errors in parentheses. * p<.10, ** p<.05, *** p<.01.
The Table reports estimates of the drop in household consumption over the unemployment spell in the HUT surveys, following model of equation (29). Because the HUT surveys collect information on household consumption expenditures at the time of the interview, HUT estimates can directly recover flow (bi-weekly) measures of consumption c_t. We restrict the sample to households where, at the date of the interview, one (and only one) individual is unemployed, or where, at the date of the interview, one (and only one) individual will become unemployed in the following 6 months. $\mathbb{1}[0 < t <= 20 \text{ weeks}]$ is an indicator for having a member of the household unemployed since less than 20 weeks at the time of the interview. $\mathbb{1}[t > 20 \text{ weeks}]$ is an indicator for having a member of the household unemployed for more than 20 weeks at the time of the interview. $\mathbb{1}[L > 20 \text{ weeks}]$ is an indicator for the total duration of the unemployment spell being longer than 20 weeks.

FOR ONLINE PUBLICATION - Appendix D: Structural Welfare Analysis

This Appendix provides more detail on the specification, approach and results of our calibration exercise.

D.1 Calibration

Our structural model builds on the search environment considered in Hopenhayn and Nicolini [1997], Lentz and Traneas [2005] and Chetty [2008a]. We consider agents with CRRA preferences with additive, iso-elastic search costs,

$$\begin{aligned} v^u_{i,t}\left(c_{i,t}, s_{i,t}\right) &= c^{1-\gamma}_{i,t} / \left(1-\gamma\right) - \psi_0 \left(s_{i,t}\right)^{\psi_1}, \\ v^e_{i,t}\left(c_{i,t}\right) &= c^{1-\gamma}_{i,t} / \left(1-\gamma\right). \end{aligned}$$

In contrast with the previous models, we allow for individual-specific, duration-dependent exit rate functions

$$h_{i,t}\left(s_{i,t}\right) = h_0 + \left[1 + \exp\left(-\theta t\right)\right] \kappa_i h_1 \times s^{\rho}_{i,t},$$

where h_1 determines the return to search, θ is the exponential rate at which these returns depreciate, κ_i is an individual-specific scalar, drawn from a gamma distribution $\Gamma\left(a, b\right)$, and h_0 is a baseline exit rate (at zero effort).

Each household starts the unemployment spell with an asset level $a_{i,1}$, drawn from a Singh-Maddala distribution $F(a|\alpha_{SM}, c_{SM}, k_{SM})$ (Singh and Maddala [1976]). Each household can draw down their asset to increase consumption as long as $a_{i,t} \geq \bar{a}$, where $\bar{a}$ is a uniform asset limit. We choose the distribution parameters to match the average, 25-th and 90-th percentile of the household distribution of liquid assets.

We consider a flat benefit policy that replaces 72% of the pre-unemployment earnings, normalized to 1. When employed, the agent pays a (uniform) tax τ equal to 5%, which balances the expected revenues and expenditures for the average Swedish unemployment rate between 1999 and 2006 of .069. Our calibration targets household consumption levels, we therefore augment the individual's income with a household income of .7, which corresponds to the average share of other household members' earnings relative to the earnings of the unemployed in our sample (see Table 2).

We set the parameter of CRRA $\gamma = 2$, discount factor $\beta = 1$ and interest rate $r = 0$. All other parameters are used to calibrate the model. We solve the model using an asset grid of 500 grid points and 10 different deciles of the distribution of individual-specific returns to search. For each decile we solve the model backwards starting in week $T = 520$ and then simulate the economy forward for each asset level. We then aggregate up the time paths for consumption and the survival rates using the joint distribution over assets and individual-specific returns to search. We compute the elasticities and average durations implied by the model.

D.2 Model Fit

Table D-1 lists the targets of the calibration, which are the variables underlying our sufficient statistics. The targets include the benefit duration elasticities $\varepsilon_{D,b_{\mathbf{k}}}$, average benefit durations $D_{\mathbf{k}}$ and the average consumption drops $\left[\bar{c}_{\mathbf{k}} - \bar{c}_0\right] / \bar{c}_0$. The benchmark consumption level $\bar{c}_0$ is the average level of consumption when agents would have been employed from the start of the model. For each of the target variables, we target the level for the first part of the policy, as well as the ratio for the two parts and we put higher weight on the latter. Our minimization routine searches for the vector of parameter values that minimizes the relative difference between the targeted moments and the simulated values of these moments generated by our model and is robust to different starting values for the parameters. Table D-1 compares the targeted and estimates values of these moments. The calibration slightly overestimates the consumption drop in the first 20 weeks and underestimates the consumption drop thereafter. The simulated drops are well within the 95% confidence intervals of our empirical estimates, but imply that we underestimate the $CS_{\mathbf{2}}/CS_{\mathbf{1}}$ compared to our sufficient-statistics implementation. Our calibrated model also underestimates the unemployment durations and elasticities, but closely matches the relative moral hazard costs $MH_{\mathbf{2}}/MH_{\mathbf{1}}$. The model underestimates the unemployment durations and elasticities, but matches the ratios very well. The calibrated parameters for the returns to search imply both strong depreciation of the exit rates and heterogeneity in exit rates. The average exit rate equals .21 in the first week of the spell and .03 in the 20-th week of unemployment, while the exit rate of a given job seeker is about half as large in the 20-th week compared to the first week. The model also predicts that 53% of the unemployed are liquidity constrained and consume hand-to-mouth from the start of the unemployment spell.

Table D-1: Model Calibration: Targeted and Estimated values

Targets	Description	Targeted Value	Estimated Value
$a_{.25}$	25th percentile of asset distribution	0.00	0.000
$a_{.90}$	90th percentile of asset distribution	0.42	0.553
a_{mean}	mean of asset distribution	0.19	0.203
$\Delta c_{\mathbf{1}}$	consumption drop ST unemployed	0.060	0.072
$\Delta c_{\mathbf{1}}/\Delta c_{\mathbf{2}}$	ratio of consumption drops	0.462	0.658
$D_{\mathbf{1}}$	ST benefit duration	12.600	6.675
$D_{\mathbf{1}}/D_{\mathbf{2}}$	ratio of benefit duration	0.940	0.991
$\varepsilon_{D_{b_{\mathbf{1}}}}$	elasticity wrt ST benefits	0.854	0.549
$\varepsilon_{D_{b_{\mathbf{1}}}}/\varepsilon_{D_{b_{\mathbf{2}}}}$	ratio of elasticities	1.262	1.271

Notes: The table lists the variables we target in our calibration exercise. These are the moments underlying our sufficient statistics such that the structural model provides the same policy recommendations as implied by the sufficient-statistics approach. For each of the target variables, we target the level for the first part of the policy, as well as the ratio for the two parts. The parameters of the asset distribution are set to match the mean, 25-th and 90-th percentile of the household distribution of liquid assets (expressed relative to annual household income). We also set the parameter of CRRA $\gamma = 2$, discount factor $\beta = 1$ and interest rate $r = 0$. The remaining model parameters are calibrated to minimize the relative distance between the targeted moments and the simulated values of these moments generated by our model.

D.3 Counterfactual Analysis

The calibrated model allows analyzing the moral hazard costs and consumption smoothing gains for different unemployment policies, keeping all parameters constant, including the asset distribution at the start of the unemployment spell. The government's budget is affected by the changes in the benefit levels (as we keep the tax rate fixed), but these budgetary changes are traded off against the agents' welfare, using the normalization of the Lagrange multiplier on the government's constraint, $\lambda \equiv E_0\left(v'\left(c_{i,0}\right)\right)$, where $c_{i,0}$ is the agent's level of consumption when she would have been employed from the start of the model. Figure 7 plots the respective consumption smoothing gains and moral hazard costs. $CS_{\bar{b}}$ and $MH_{\bar{b}}$ have a unique intersection for $\bar{b} = .58$, which corresponds to the optimal flat benefit level for the given tax rate and Lagrange multiplier. Similarly, we find that $CS_{\mathbf{1}} = MH_{\mathbf{1}}$ and $CS_{\mathbf{2}} = MH_{\mathbf{2}}$ jointly hold for the unique pair $(b_{\mathbf{1}}, b_{\mathbf{2}}) = (.48, .68)$.

We repeat the previous analysis for two variations of our baseline model. In the model with only heterogeneity in the returns to search, we set the search depreciation parameter $\theta = 0$, but re-calibrate the return to search parameter h_1 to maintain the same average duration. All other parameters remain the same. In the model with only depreciation of the returns to search, we assume a uniform scalar $\kappa_i = \bar{\kappa}$ and re-calibrate h_1 to maintain the same average duration. Figure D-1 illustrates the respective consumption smoothing gains and moral hazard costs. The simulated implications for the flat profile are similar, except that the unemployment elasticities and thus the moral hazard costs are higher in the model with only depreciation (but no heterogeneity). In both alternative models, $MH_{\bar{b}}$ decreases while $CS_{\bar{b}}$ increases when the replacement rate is reduced. They are equalized for $\bar{b} = .60$ in the model with only heterogeneity and $\bar{b} = .40$ in the model with only depreciation. Regarding the tilt, the model with only depreciation of exit rates increases $MH_{\mathbf{2}}$ relative to $MH_{\mathbf{1}}$.[73] This calibration pushes towards a declining tilt and more so as the overall generosity is reduced. However, the model with only heterogeneity continues to prescribe an inclining tilt and this remains true for lower replacement rates. Importantly, conditional on the values of the sufficient statistics, the underlying depreciation and heterogeneity do not affect the local policy recommendations. The differential impact on how the sufficient statistics change with the unemployment policy, however indicates the complementary value of a structural approach and disentangling the two non-stationary forces to go beyond our local recommendations.

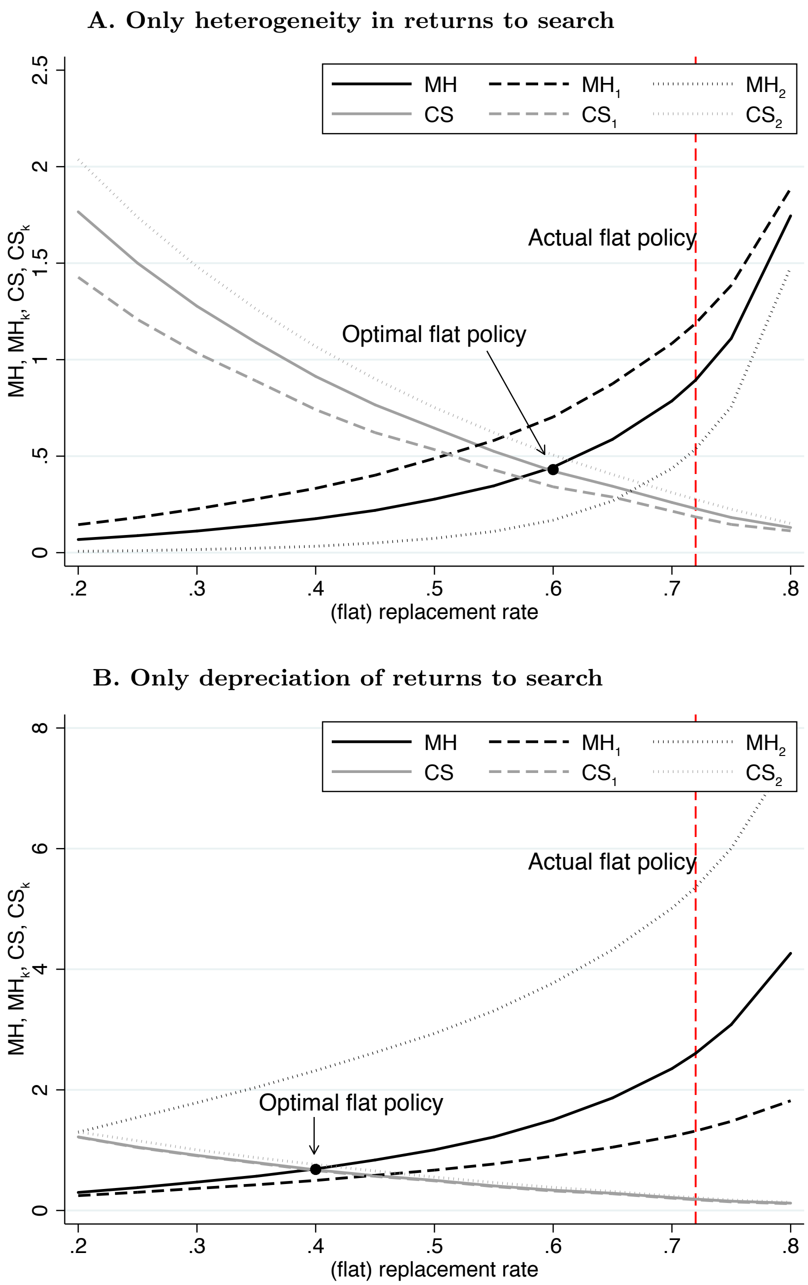
We also note that the structural model can be exploited to assess the potential importance of some implementation assumptions. We simulate the consumption drops and calculate the implied consumption smoothing gains for different levels of risk aversion, reported in Table D-2, while keeping all other parameters constant. The Taylor approximation of the marginal CRRA utilities tends to underestimate the consumption smoothing gains, especially when the level of risk aversion is high. This was already noted in Chetty [2006]. Interestingly, the approximation error on the relative consumption smoothing gains is very small and thus the recommendation on the tilt unaffected, regardless of the level of risk aversion.

We finally use the structural model to gauge the impact of heterogeneity in preferences on the estimated consumption smoothing gains. We simulate a model with share α of job seekers with risk aversion $\gamma = 1$ and share $1 - \alpha$ with risk aversion $\gamma = 3$. We set $\alpha = .58$ such that the average risk aversion among the unemployed, accounting for type-specific unemployment durations, equals $\gamma = 2$. Since the consumption drops hardly change for agents with different preferences, the average consumption drops in the model with heterogeneity remain very similar compared to the baseline model. The dynamic selection on preferences could in principle affect the relative consumption smoothing gains, as shown in subsection 5.2.3. Note that in our model with CRRA preferences and additive search cost (a standard specification also used in Hopenhayn and Nicolini [1997] and Chetty [2008a]), individuals with higher risk aversion have lower marginal utility of consumption (relative to the disutility of search), therefore search less and leave unemployment more slowly. As illustrated in equation (21) in subsection 1.3, dynamic selection on both the marginal utility of consumption $E_k\left[v_i'\left(c_{i,0}\right)\right]/E_0\left[v_i'\left(c_{i,0}\right)\right]$ and the consumption smoothing gains $E_k\left[v_i''\left(c_{i,0}\right)\left(c_{i,0} - c_{i,t}^u\right)\right]/E_0\left[v_i'\left(c_{i,0}\right)\right]$ affect the value of CS_k. Hence, the net impact is ambiguous. In our simulation with heterogeneity, the aggregate consumption smoothing gains $CS_{\mathbf{1}}$ and $CS_{\mathbf{2}}$ are lower relative to the baseline

[73] Note that in Appendix C we demonstrated that depreciation in the search returns can increase $MH_{\mathbf{1}}$ relative to $MH_{\mathbf{2}}$. This comparative static was derived without keeping the average duration constant.

scenario. The ratio $CS_{\mathbf{1}}/CS_{\mathbf{2}}$ decreases as well, which further increases the value of introducing an inclining tilt ($b_{\mathbf{2}} > b_{\mathbf{1}}$). We note though that with heterogeneous preferences, the calculation of $CS_{\mathbf{k}}$ critically depends on the normalization $\lambda \equiv E_0 \left(v' \left(c_{i,0}\right)\right)$. We can avoid this normalisation by directly considering the relative consumption smoothing gains $\left[1 + CS_{\mathbf{1}}\right] / \left[1 + CS_{\mathbf{2}}\right] = E_1 \left(v_i' \left(c_{i,t}\right)\right) / E_2 \left(v_i' \left(c_{i,t}\right)\right)$, relevant for evaluating a budget-balanced change in the tilt. This ratio changes from .934 in the baseline calibration to .935 in the simulation with heterogeneity (cf. Table D-2) Overall, the introduction of preference heterogeneity seems to limited impact on the relative consumption smoothing gains in our simulations and, if anything, increases the welfare gain from introducing an inclining tilt ($b_{\mathbf{2}} > b_{\mathbf{1}}$).

Figure D-1: Alternative Model Specifications: Welfare Effects for Different Benefit Levels



Notes: The figure analyzes the robustness of the findings in Figure 7 for different model specifications. The two panels plot the moral hazard costs and consumption smoothing gains for different levels of the flat benefit profile. Panel A uses a model without depreciation of the returns to search and only heterogeneity in the returns to search. Panel B uses a model with only depreciation of the returns to search and no heterogeneity in the returns to search. Compared to our baseline model, we re-calibrate the uniform return to search parameter to maintain the same average exit rate. The actual policy is a flat profile with average replacement rate of .72 and is indicated by the vertical dashed line. We report the simulated moral hazard costs and consumption smoothing gains for an overall change in the flat benefit profile $\bar{b}$, for an increase in the benefit level in the first 20 weeks of unemployment, and for an increase in the benefit level after 20 weeks of unemployment.

Table D-2: Implementation of consumption smoothing gains

		Consumption Smoothing Gain		
	$\Delta c_k/c$	$E_k\left[v_i'\left(c_{i,t}\right)\right]$	CS_k	$\gamma\Delta c_k/c$
	Baseline Calibration: γ=2			
Throughout spell	0.091	0.443	0.231	0.181
First 20 weeks	0.072	0.428	0.189	0.143
After 20 weeks	0.109	0.458	0.273	0.218
Before 20wks / after 20 wks	0.658	0.934	0.693	0.658
	Simulation with heterogenous preferences			
Throughout spell	0.091	0.485	0.108	0.182
First 20 weeks	0.072	0.469	0.070	0.144
After 20 weeks	0.109	0.501	0.144	0.217
Before 20wks / after 20 wks	0.664	0.935	0.483	0.664
	Simulation with low risk aversion γ=1			
Throughout spell	0.093	0.666	0.109	0.093
First 20 weeks	0.074	0.653	0.088	0.074
After 20 weeks	0.112	0.677	0.129	0.112
Before 20wks / after 20 wks	0.658	0.964	0.682	0.658
	Simulation with high risk aversion γ=3			
Throughout spell	0.088	0.296	0.368	0.264
First 20 weeks	0.071	0.282	0.307	0.212
After 20 weeks	0.105	0.309	0.429	0.316
Before 20wks / after 20 wks	0.671	0.914	0.715	0.671

Notes: This Table illustrates alternative implementations of the consumption smoothing gains for different underlying risk preferences. The top panel reports the simulated consumption drops and consumption smoothing gains for the baseline model with homogeneous CRRA γ=2. The third and fourth panel report the simulated values for lower CRRA γ=1 and higher CRRA γ=3, but still assuming homogenous preferences. The second panel reports the simulated values for a model with a mixture of these two preference types, keeping the average CRRA among the unemployed equal to 2. This makes the results comparable to the baseline model. The first column in each panel shows the drops in consumption during the respective parts of the unemployment spell relative to the consumption levels when employed at the start of the model. The second column shows the average marginal utility of consumption during the respective parts of the unemployment spell. The third column reports $CS_{\mathbf{k}}$, which expresses the average marginal utility of consumption in the second column relative to the average marginal utility of employment consumption at the start of the model. The fourth column provides an approximation of $CS_{\mathbf{k}}$ following the implementation in (14), which relies on a Taylor approximation and homogeneous preferences. For this fourth column, the second panel with heterogeneous preferences simply multiplies the consumption drops by γ=2. The rows of each panel report the consumption drops and consumption smoothing gains over the full unemployment spell, during the first 20 weeks of unemployment and after 20 weeks of unemployment respectively. The last row shows the ratio of the first part to the second part of the unemployment spell.

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