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 recent introduction of a declining tilt in the profile.

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1 Introduction

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.

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 [2008], Schmieder et al. [2012]). We aim to bridge these different strands of the literature.

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, 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 behavioral revenue effect 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 depend 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 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 take advantage of the survey data on consumption expenditures, linked to the detailed unemployment records, to identify how consumption expenditures change with unemployment and the length of an unemployment spell in particular. We provide complementary and robustness analysis using comprehensive information on income, transfers and wealth from Swedish tax registers.

Our empirical analysis provides the following main results:

First, unemployment durations are very responsive overall to changes in the benefit level. Furthermore, we find that the response to changes in benefits does not increase for benefits paid later in the spell compared to benefits paid early 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. To the contrary, we estimate the moral hazard cost of increasing the benefit level to be 26 percent higher for benefits paid in the first 20 weeks of unemployment than for benefits paid after 20 weeks, although our estimates cannot reject equality. 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 later in the spell to comparable changes in the policy. Importantly, such non-stationary forces, which may be driven by duration dependence or dynamic selection 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 6% in the first 20 weeks of unemployment, compared to their pre-unemployment level. This drop deepens to on average 13% for those who are unemployed for longer. These consumption patterns are not driven by selection effects. We estimate small and insignificant differences in consumption levels and profiles for households selecting into longer unemployment spells. This is consistent with the evidence of limited dynamic selection on observables, including characteristics related to risk preferences that would affect the mapping of the consumption drops into the marginal valuation of unemployment benefits. 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.

Our registry data on household income and wealth allows us to provide a complementary account of the underlying smoothing mechanisms. Overall, we indeed find a limited ability to smooth consumption in addition to the government transfers, which are generous throughout the unemployment spell. Most unemployed individuals have few assets prior to becoming unemployed, not sufficient to smooth longer spells of unemployment, and cannot increase their consumption by taking out (more) debt. The added-worker effect is not playing a significant role either in smoothing the unemployment shock, not even for 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 profile in Sweden. We also provide a complementary welfare analysis, by calibrating our structural model to match the sufficient statistics entering the formulae for the optimal benefit profile. This allows us to assess the role of our implementation assumptions and go beyond the local policy recommendations, now relying on the structure of the calibrated model. We find that the incentive costs are high relative to the drop in consumption throughout the unemployment spell. Our model therefore suggests that any reduction in the generosity of the unemployment policy 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 long-term unemployed that is more than twice as high as decreasing the marginal krona spent on the short-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. Our calibration exercise indicates that such an inclining tilt remains welfare improving when lowering the overall generosity of the policy. Strikingly, the Swedish government substantially reduced the generosity of the unemployment policy in 2007, but mainly for the long-term unemployed. Our analysis provides no support for the introduction of such negative tilt in the benefit profile.

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 contributes to a long literature on labor supply effects of social insurance by explicitly using duration-dependent variation in benefits and identifying the welfare-relevant unemployment responses.[2] Our analysis also 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 smoothing of income shocks and unemployment in particular (e.g., Gruber [1997]). Merging these consumption surveys with unemployed registers, gives us the opportunity to analyze the evolution of consumption as a function of time spent unemployed. We provide complementary insights using administrative data on income and wealth and confirm our results using a residual, registry-based measure of consumption expenditures.[3]

The remainder of the paper proceeds as follows. Section 2 analyzes the characterization and implementation of sufficient statistics formulae for the evaluation of local policy changes in a dynamic model of unemployment. Section 3 describes our data and the policy context in Sweden. Section 4 describes our regression kink design and provides estimates of the policy-relevant unemployment elasticities. Section 5 analyzes how consumption evolves during the unemployment spell and how this translates to the consumption smoothing gains of UI. Section 6 analyzes welfare complementing the implementation of the sufficient statistics with a calibration exercise of our structural model. Section 7 concludes.

[1] Recent extensions of the Baily-Chetty formula can be found in Schmieder et al. [2012], analyzing the potential benefit duration, but for a given benefit level. Spinnewijn [2015] provides 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 these recent developments.

2 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 model that allows for non-stationary forces (e.g., duration-dependence in exit rates) and heterogeneity (e.g., heterogenous exit rates, assets, preferences). In the spirit of the "sufficient statistics" literature, our approach consists in identifying the minimal level of information necessary 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. Our characterization also provides a robust evidence-based approach to evaluate local departures from the existing unemployment policy.

2.1 Setup

We first describe the set up of our dynamic unemployment model, the agents' preferences and choices they can make, and the unemployment policy. We try to save on notation in the main text, but provide more details in the technical appendix.

We consider a partial equilibrium framework with a continuum of agents with mass 1. The model is in discrete time t and ends at T. Each agent i starts unemployed at time $t = 1$ and remains unemployed until she finds work. Once she has found work, she remains employed until T. When employed, the agent earns w, when unemployed she earns 0. The government designs an unemployment policy P providing insurance against the unemployment risk: the policy specifies an unemployment benefit profile depending on the length 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 agent's type, capturing heterogeneity in employability across agents, and the time she has spent unemployed, capturing differences in employment prospects due to the time spent unemployed. The agent's probability to be unemployed after t periods equals the survival probability $S_{i,t} \equiv \prod_{x=1}^{t-1}(1 - h_{i,x}(s_{i,x}))$ 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 $\bar{a}_i$. The agent's savings decisions determine her consumption level throughout the unemployment spell and when re-employed. We denote these levels by $c^u_{i,t}$ and $c^e_{i,t}$ for when unemployed and employed respectively.[4] While we cannot observe an agent's contingent consumption plan, we can observe average levels of consumption, for example at different spell lengths, $\bar{c}^u_t = \int S_{i,t}c^u_{i,t}di$.

Preferences We consider time-separable preferences (with discount factor β). Per-period utility is increasing in consumption, but decreasing in search efforts exerted when unemployed. We allow for heterogeneous preferences and denote agent i's instantaneous utility by $v^u_i(c^u_{i,t}, s_{i,t})$ and $v^e_i(c^e_{i,t})$. Taking the unemployment policy P as given, each agent chooses how much to search and how much to consume in order to maximize her expected utility. The dynamics of the agent's behavior depend on her asset level and the duration of the ongoing unemployment spell 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 consider n-part policies that pay benefit b_1 for the first B_1 periods, b_2 for the next $B_2 - B_1$ periods, and so on.[5] We take the potential benefit durations B_k as given, but our characterization generalizes for a fully flexible policy with $n = T$.

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 $\sum_{t=1}^{T} S_t$. We use $D_k = \Sigma^{B_k}_{B_{k-1}} S_t$ to denote the expected time spent unemployed while receiving benefit b_k, so that $D = \Sigma^n_{k=1} D_k$.

We ignore time discounting in our optimal policy characterization (i.e., $1 + r = \beta = 1$), but generalize this in the technical appendix. The government's budget then simplifies to

$$G(P) = [T - D]\tau - D_1 b_1 - D_2 b_2 - .. - D_n b_n. \tag{1}$$

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.

[4] 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.

[5] As it will be clear which case applies, we use (with some abuse of notation) the subindex to either refer to time t or to part k of the unemployment policy.

2.2 Local Policy Changes

Our approach is to consider the welfare impact of local deviations from the unemployment policy P, which we decompose into the corresponding consumption smoothing gains and moral hazard cost.

Consider an increase in the benefit level b_k during part k of the unemployment policy (between period B_{k-1} and B_k). The total impact on welfare depends on how much the unemployed value this increase in benefits b_k relative to its budgetary cost,

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

This effect depends also 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 $V_i(P)$. This follows from the envelope conditions $\partial V_i / \partial x_{i,t} = 0$, which hold for any behavior $x_{i,t}$ the agent optimizes over (consumption and search in our model), also when borrowing constraints are binding (see Chetty [2006]). The application of the envelope theorem thus critically simplifies the characterization as we only need to account for the impact of the behavioral responses on the government's budget $G(P)$.

The policy P can only be optimal if any local change of the policy cannot increase welfare. Considering local policy changes is sufficient to characterize the optimal policy if social welfare is concave in the policy variables. Importantly, evaluating local policy changes is also of interest, away from the optimal policy, to identify how welfare can be increased and the policy changed towards the optimal policy.

While our characterization focuses on one-dimensional changes in the policy, these expressions can be combined to evaluate multi-dimensional changes in the policy as well.

Moral Hazard. We first consider the budgetary impact from an increase in b_k and analyze how this evolves during the unemployment spell. Even absent behavioral responses, a benefit increase reduces the government's budget. This mechanical effect depends on the expected time the unemployed spend on part k of the policy, D_k. A second effect is behavioral and captures the budgetary cost of the agents' reduced search in response to more generous benefits. This moral hazard cost depends on how the reduced incentives change the time D_l spent on each part l of the unemployment policy,

$$\frac{\partial G(P)}{\partial b_k} = -D_k \times \left[1 + \Sigma_{l=1}^{n} \frac{D_l (b_l + \tau)}{D_k b_k} \varepsilon_{D_l, b_k}\right] \tag{4}$$

$$\equiv -D_k \times [1 + MH_k]. \tag{5}$$

The moral hazard cost MH_k of an increase in b_k thus simply equals a weighted sum of the elasticities $\varepsilon_{D_l,b_k} = (\partial D_l / \partial b_k) / (D_l / b_k)$ for each benefit duration D_l with respect to the benefit level b_k.

The first important observation is that the moral hazard cost of a local change in the policy depends only on the average benefit durations and their elasticities. Conditional on the value of these sufficient statistics, the welfare impact is not affected by differences in the underlying primitives and heterogeneity across agents in particular.

We also note that MH_k does not only depend on the increased time the unemployed spend receiving b_k, captured by ε_{D_k,b_k}. First, by increasing the share of people still unemployed at time B_k, a higher b_k also increases the time spent receiving benefits later in the spell ($\varepsilon_{D_l,b_k} \geq 0$ for $l > k$). Second, a higher b_k reduces the incentives to leave unemployment earlier in the spell as an agent anticipates the more generous benefit received later in the spell ($\varepsilon_{D_l,b_k} \geq 0$ for $l < k$). The benefit duration elasticities are weighted by the relative budget shares of the respective parts of the unemployment policy. The spillover effect of a change in one part of the policy to others parts of the unemployment policy 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.

The evolution of these moral hazard costs over the unemployment spell is key for designing the unemployment benefit profile. Seminal papers in the UI literature (Shavell and Weiss [1979] and Hopenhayn and Nicolini [1997]) have argued that benefits must be decreasing over the unemployment spell to provide the appropriate incentives to search for agents who are forward-looking . This argument can be simply linked to the gradient of the moral hazard cost over the unemployment spell in a stationary environment:

Proposition 1. *Consider a single agent in a stationary environment being borrowing-constrained (i.e., $h_{i,t}(\cdot) = \bar{h}(\cdot)$, $a_{i,1} = \bar{a}$). Assume $\beta(1+r) = 1$ and $T \to \infty$. Starting from a flat benefit profile ($b_k = \bar{b}$ for all k), the moral hazard cost of increasing benefit b_k increases during the unemployment spell,*

$$MH_k \leq MH_{k'} \text{ for } k < k'.$$

Proof. *See Appendix A.*

To provide intuition for this result, note that in a stationary environment with no savings, an agent's search behavior only depends on the continuation value of the unemployment policy. Starting from a flat profile, the impact of an increase in the unemployment benefit received after B periods of unemployment on the *remaining duration* of unemployment $\tilde{D}_B \equiv \Sigma_{t=B}^{\infty} S_t / S_B$, conditional on being unemployed at time B, is independent of B. The corresponding elasticity is thus the same for any B (including $B = 1$) and therefore coincides with $\varepsilon_{D,\bar{b}}$, the elasticity of the unemployment duration to an overall change in the flat benefit level $\bar{b}$. The latter elasticity determines the moral hazard cost of an overall change in the benefit level, $MH_{\bar{b}} = \frac{\bar{b}+\tau}{\bar{b}} \varepsilon_{D,\bar{b}}$. The moral hazard cost of the benefit increase after B periods, however, exceeds this cost. The reason is that the benefit increase reduces in addition the exit rates during the first part of the policy (causing $D_1 = \Sigma_{t=1}^{B-1} S_t$ to increase), which also increases the probability S_B to be unemployed at the start of the second part (causing a further increase in $D_2 = S_B \times \tilde{D}_B$).

In practice, the unemployment environment is not stationary. A large empirical literature has documented important non-stationary forces, which may have a substantial but ambiguous impact on how the moral hazard cost evolves over the unemployment spell. Job seekers are heterogeneous in their employability (captured by $h_{i,t}(\cdot) \neq h_{j,t}(\cdot)$) and the less employable job seekers select into longer unemployment spells. At the same time, job market opportunities for a given job seekers tend to deteriorate over the unemployment spell (captured by $h_{i,t}(\cdot) \geq h_{i,t+1}(\cdot)$).[6] While it can prove challenging to identify these non-stationary forces empirically, disentangling such forces is not necessary to evaluate local policy changes. Our framework accounts for both heterogeneity and duration-dependance and, when strong enough, these forces may well offset the forward-looking channel. As a result, the overall evolution of moral hazard over the spell is theoretically ambiguous and remains an empirical question.

Consumption Smoothing. Let us turn now to the insurance value of an increase in b_k during part k of the unemployment spell. A marginal increase in b_k will benefit the agents who are still unemployed during this stage of the spell. Due to the envelope conditions, the welfare increase is completely captured by the marginal utility of consumption for these agents,

$$\begin{aligned}\int \frac{\partial V_i(P)}{\partial b_k} di &= \int \Sigma_{t=B_{k-1}}^{B_k} S_{i,t} \frac{\partial v_i^u\left(c_{i,t}^u, s_{i,t}\right)}{\partial c_{i,t}^u} di, \\ &\equiv D_k \times E_k\left[\frac{\partial v^u(c^u, s)}{\partial c^u}\right],\end{aligned}$$

where the expectation operator E_k takes the weighted average over all individuals and all periods of the k-th part of the unemployment spell (with weights $S_{i,t}/\int \Sigma_{t=B_{k-1}}^{B_k} S_{i,t} di$). By analogy to the budgetary cost, we can write

$$\int \frac{\partial V_i(P)}{\partial b_k} di/\lambda = D_k \times [1 + CS_k], \tag{6}$$

where the consumption smoothing gain $CS_k \equiv \{E_k\left[\frac{\partial v^u(c^u,s)}{\partial c^u}\right] - \lambda\}/\lambda$. The Lagrange multiplier λ corresponds to shadow cost of the government's budget constraint and would be set equal to the average marginal value of resources at the start of this model if the government could provide a lump sum transfer to everyone. Hence, the consumption smoothing gains capture the return of a government dollar spent on the unemployed during part k of the unemployment spell, relative to the value of the lump sum transfer.[7]

The important observation here is that in spite of the presence of heterogeneity across agents,

[6] See Machin and Manning [1999] for a review on the negative duration dependence of exit rates out of unemployment. See Schmieder et al. [2012] and Kroft et al. [2013] for recent examples.

[7] In our stylized setup all individuals start unemployed, so an increase in b_1 in fact corresponds to a uniform lumpsum transfer. When this is set optimally, $\lambda = \int \partial V/\partial a_{i,1} di = \int \partial V/\partial b_1 di$. The last equation is not true more generally, for example when considering agents who start the model employed but risk to be laid off. We consider this case in the technical appendix.

of intertemporal consumption responses to the policy change and of potentially binding borrowing constraints, the wedge between the marginal value of public funds and the marginal utilities of consumption of unemployed workers at time t of the unemployment spell is sufficient to evaluate the welfare gain from an increase in benefits at time t of the unemployment spell.

The key question for designing the benefit profile is again how the consumption smoothing gains evolve during the unemployment spell. Standard models of intertemporal consumption predict that for a given individual the marginal utility of consumption is weakly increasing during the unemployment spell. For equal discount and interest rate, the Euler equation states

$$\frac{\partial v_i^u\left(c_{i,t}^u, s_{i,t}\right)}{\partial c_{i,t}^u} \geq h_{i,t}\left(s_{i,t}\right)\frac{\partial v_i^e\left(c_{i,t+1}^e\right)}{\partial c_{i,t+1}^e} + \left(1 - h_{i,t}\left(s_{i,t}\right)\right)\frac{\partial v_i^u\left(c_{i,t+1}^u, s_{i,t+1}\right)}{\partial c_{i,t+1}^u}.$$

An agent would like to use her savings or borrow against her future earnings to increase her unemployment consumption at the expense of future consumption. This is true as long as the marginal utility of consumption is higher when unemployed than when employed, which is trivially satisfied when the after-tax wage exceeds the benefit level and utility is separable in consumption and search. An unemployed agent runs down her assets and thus faces lower consumption levels the longer she is unemployed, until she becomes liquidity-constrained and starts consuming "hand-to-mouth".

Proposition 2. *Consider a single agent with separable preferences $\partial v_i^u\left(c,s\right)/\partial c = \partial v_i^e\left(c\right)/\partial c = v_i'\left(c\right)$. Assume $\beta\left(1+r\right) = 1$ and $T \to \infty$. When the agent is not borrowing constrained, the consumption smoothing gain increases during the unemployment spell,*

$$CS_k \leq CS_{k'} \text{ for } k < k'.$$

Otherwise, CS_k and b_k are inversely related.

Proof. *See Appendix A.*

Long-term unemployment implies a larger shock in resources than short-term unemployment and this larger shock requires more savings or credit to be smoothened. This implies that insurance against long-term unemployment is valued more. This force underlies the inclining optimal benefit profile in a single-agent model without search considered in Shavell and Weiss [1979].

In practice, however, heterogeneity across agents, for example in assets or in preferences, will affect the gradient of consumption smoothing gains over the spell. Dynamic selection effects may increase or decrease (and possibly revert) the gradient depending on the correlation between the marginal utility of consumption and the unemployment duration. For example, if job seekers with more assets search less, they will be unemployed for longer. As they can maintain a higher consumption level, they may value unemployment benefits less. If less employable job seekers tend to have less assets to rely on, the dynamic selection may strengthen the increase in consumption smoothing gains during the spell. Whether the consumption smoothing gains are still increasing in the presence of selection effects is therefore an empirical question.

2.3 Optimal Unemployment Policy

The optimal unemployment policy balances the provision of insurance and incentives. Following a "sufficient statistics" approach, Baily [1978] and Chetty [2006] have shown that, in the context of a simple policy giving flat UI benefits throughout the unemployment spell, the optimal balance is reached when the relative wedge in marginal utilities between employment and unemployment equals the elasticity of the unemployment duration with respect to UI benefits. Our analysis generalizes their insight for a dynamic benefit profile. We can state the following result:

Proposition 3. *Consider n-part policies $P = (b_1, .., b_n, \tau)$ with $W(P)$ strictly concave. The optimal n-part policy solves*

$$\frac{E_k\left[\frac{\partial v^u(c^u,s)}{\partial c^u}\right] - \lambda}{\lambda} = \Sigma_{l=1}^{n} \frac{D_l\,(b_l + \tau)}{D_k b_k} \times \varepsilon_{D_l, b_k} \text{ for each } k, \tag{7}$$

$$\frac{\lambda - E_\tau\left[\frac{\partial v^e(c^e)}{\partial c^e}\right]}{\lambda} = \Sigma_{l=1}^{n} \frac{D_l\,(b_l + \tau)}{(T - D)\,\tau} \times \varepsilon_{D_l, \tau}, \tag{8}$$

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

Proof. *See Appendix A.*■

The expectation operator E_k is defined as before and takes the weighted average over all individuals and all periods of part k of the unemployment spell (with weight $S_{i,t}/D_k$). Similarly, E_τ takes the weighted average over all individuals and all periods in employment (with weight $(1 - S_{i,t})/(T - D)$).

For each policy variable, the consumption smoothing gain must be equal to its moral hazard cost at the margin. Combining expressions (3), (5) and (6), we find

$$\frac{\partial W(P)}{\partial b_k} = \lambda D_k \times [CS_k - MH_k] = 0,$$

which corresponds to condition (7). As long as a change in the benefit profile can increase welfare, the policy cannot be optimal.

We can combine the optimality conditions in Proposition 3 to provide alternative characterizations of the optimal policy. For example, combining the condition for the benefit and tax level allows us to recover the well-known Baily-Chetty formula for a flat benefit profile ($b_k = \bar{b}$),[8]

$$\frac{E_{\bar{b}}[\frac{\partial v^u(c^u,s)}{\partial c^u}] - E_\tau[\frac{\partial v^e(c^e)}{\partial c^e}]}{E_\tau[\frac{\partial v^e(c^e)}{\partial c^e}]} \cong \frac{\bar{b} + \tau}{\bar{b}} \varepsilon_{D,\bar{b}}, \tag{9}$$

[8] The approximation in (9) 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.

Similarly, we can combine the conditions for different parts of the benefit profile, to characterize how the respective benefit levels should relate,

$$\frac{E_k[\frac{\partial v^u(c^u,s)}{\partial c^u}]}{E_l[\frac{\partial v^u(c^u,s)}{\partial c^u}]} = \frac{1+\Sigma_{m=1}^n \frac{D_m(b_m+\tau)}{D_k b_k} \times \varepsilon_{D_m,b_k}}{1+\Sigma_{m=1}^n \frac{D_m(b_m+\tau)}{D_l b_l} \times \varepsilon_{D_m,b_l}}. \quad (10)$$

In the spirit of the sufficient statistics approach the optimal policy characterization in Proposition 3 is robust to variations in the underlying primitives of the model.[9]

First, an unemployment policy satisfying the above conditions is optimal, regardless of the values of the primitives that map into our sufficient statistics. This confirms that, in this dynamic context, in order to determine local welfare-enhancing changes to the benefit profile, and conditional on the values of the consumption smoothing gains and moral hazard costs, it is irrelevant to separately identify the underlying sources of heterogeneity and non-stationarities. Second, one can extend the foundations of the model in various directions without substantially changing the characterization of the optimal benefit profile. Chetty [2006] has shown how the simple formula 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. This relies on the application of the envelope conditions for the different types of behavior and naturally extends for a dynamic benefit profile. Related, our stylized model focused on a single unemployment spell as is standard in the optimal UI literature, but our dynamic model can incorporate multiple unemployment spells with (endogenous) layoff rates, while the same formulae continue to apply. We explore this in more detail in the technical appendix. Assuming that the policy can only depend on the length of the ongoing spell, the relevant policy variables are still the benefit durations D_k, capturing the total expected time spent unemployed receiving benefits b_k, but now potentially spread over multiple spells. With multiple spells, the benefit duration D_k thus depends on the layoff rates in addition to the earlier survival probabilities. The layoff responses to UI policy could affect the policy-relevant elasticities ε_{D_k,b_l} if moral hazard on-the job were to be important.[10] Finally, as our characterization critically relies on the application of the envelope theorem, any externality (not internalized by our agents, but deemed relevant for welfare) would affect the optimal policy characterization. In the technical appendix, we briefly illustrate how our framework can account for the fiscal externality created by the presence of an income tax used to fund other government expenditures. Recent work has analyzed the impact of various externalities on the characterization of a static unemployment policy.[11] An interesting avenue for future research is to analyze how these externalities depend on

[9] See the review chapter by Chetty and Finkelstein [2013] for a more detailed discussion of the different issues.

[10] 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.

[11] A recent example is Nekoei and Weber [2015], who account for the fiscal impact of reservation wage responses, conditional on unemployment duration, which tend to be small relative to the duration responses themselves. Landais et al. [2010] adjust the characterization to account for general equilibrium effects and thus externalities of job seekers' behavior on firms and other job seekers. Spinnewijn [2015] accounts for "internalities" due to biased beliefs about the employment prospects.

the timing of the benefits.

2.4 Implementation

We now analyze the implementation of our approach before turning to the empirical analysis. As argued earlier, comparing the consumption smoothing gain and moral hazard cost not only allows for a characterization of the optimal policy, but also provides a simple, yet robust guide to evaluate deviations from local policy.

Our focus is on a two-part policy, paying benefit b_1 until time B and b_2 thereafter, which is the policy variation we exploit in the Swedish context. Constant benefit profiles with one or few steps are very common in practice. Panels A.I and B.I of Figure 1 illustrate such two-part schemes. Panels A.II and B.II illustrate the corresponding survival functions S_t. The areas under the survival function before and after B correspond to D_1 and D_2, the expected time spent receiving benefit b_1 and b_2 respectively.

Changes in Tilt. Our analysis demonstrates that with the appropriate data to evaluate CS_k and MH_k for either part of the unemployment policy, we can infer whether b_1 and/or b_2 are too generous or not. Moreover, we can also evaluate the welfare impact of changes in the tilt b_1/b_2 and this evaluation requires less information:

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. *See Appendix A.*■

This result follows naturally from Proposition 3 and links our analysis to a common approach to formulate policy recommendations based on "cost/benefit" ratio's (see Hendren [2013]). The ratio of the marginal consumption smoothing gains to the moral hazard cost indeed captures the marginal benefit of the policy relative to its cost, accounting for the behavioral responses. In fact, the ratio $[1 + CS_k] / [1 + MH_k]$ coincides with the so-called "marginal value of public funds", corresponding to the marginal social welfare impact of an increase in b_k per unit of government revenue spent.

When policies are multi-dimensional, but the cost/benefit ratio's are not equalized along all dimensions, one can identify simple changes in the policy that increase welfare. The main advantage of this approach for policy design is that only information on the relative cost/benefit ratio's is required. For a two-part policy, knowing the relative consumption smoothing gains CS_1/CS_2 and moral hazard costs MH_1/MH_2 is sufficient to recommend welfare-improving changes in the tilt b_1/b_2. The recommendation remains local. Clearly, the cost/benefit ratio's are interdependent and their ranking may also change when deviating from the local policy. In particular, the local recommendation regarding the tilt of the policy may not be informative about the optimal tilt of the policy.[12] One approach to overcome this limitation is to calibrate a structural model to match

[12] Of course, this can also be the case for a one-dimensional change in a benefit level b_k. Only if the policy problem

the sufficient statistics for local policy evaluation and analyze how these sufficient statistics change when moving away from the local policy. We illustrate this in Section 6.2.

Moral Hazard. While an extensive literature analyzes unemployment responses to changes in unemployment benefits, our analysis indicates that it is essential to have variation in unemployment benefits at different times during the unemployment spell. In particular, to evaluate a two-part unemployment policy, we need to estimate

$$MH_1 = \frac{b_1+\tau}{b_1}\varepsilon_{D_1,b_1} + \frac{D_2\left(b_2+\tau\right)}{D_1 b_1}\varepsilon_{D_2,b_1} \tag{11}$$

$$MH_2 = \frac{D_1\left(b_1+\tau\right)}{D_2 b_2}\varepsilon_{D_1,b_2} + \frac{b_2+\tau}{b_2}\varepsilon_{D_2,b_2}. \tag{12}$$

This estimation requires separately identifying responses to variations in short-term benefits b_1 and responses to variations in long-term benefits b_2. This requires duration-dependent policy variation as illustrated in Panel A and Panel B of Figure 1. That is, rather than having benefits change throughout the spell, which would be sufficient to evaluate a flat policy, we also need changes in benefits paid only to the short-term db_1 or to the long-term unemployed db_2.

In principle, the implementation also requires decomposing the duration responses in the first and second part of the spell. However, when evaluated at a flat benefit profile, the moral hazard cost of increasing benefits b_k simplifies to the elasticity of the average unemployment duration ε_{D,b_k}, scaled by the relative time spent receiving b_k,

$$MH_k = \frac{D}{D_k}\frac{\bar{b}+\tau}{\bar{b}}\varepsilon_{D,b_k} \text{ when } b_t = \bar{b} \text{ for all } t. \tag{13}$$

Conveniently, the ratio of moral hazard costs MH_1/MH_2 simplifies to $\left[\varepsilon_{D,b_1}/D_1\right]/\left[\varepsilon_{D,b_2}/D_2\right]$ and no longer depends on the tax distortion τ.

Consumption Smoothing. Attempts at quantifying the consumption smoothing gains of UI policies have been scarce as the estimation of differences in marginal utility levels proves very difficult in practice. Our empirical analysis follows the "consumption implementation" approach (Gruber [1997],Chetty [2006]), relating the difference in marginal utilities to the difference in consumption levels, and extend this to our dynamic setting. Using consumption wedges to actually quantify the relevant consumption smoothing gains of UI requires some assumptions, detailed below.

First, we rely on approximations of the marginal utility of consumption using Taylor expansions.

is strictly concave in all benefit levels, can we be certain that a welfare-increasing change in the benefit level b_k actually moves the policy closer to the optimal policy.

That is,

$$
\begin{aligned}
\frac{\partial v_i^u}{\partial c}\left(c_{i,t}^u, s_{i,t}\right) &= \frac{\partial v_i^u}{\partial c}\left(\tilde{c}, s_{i,t}\right) + \frac{\partial^2 v_i^u}{\partial c^2}\left(\tilde{c}, s_{i,t}\right)\left[c_{i,t}^u - c^0\right] + \frac{\partial^3 v_i^u}{\partial c^3}\left(\tilde{c}, s_{i,t}\right)\left[c_{i,t}^u - \tilde{c}\right]^2/2 + ... \\
&\cong \frac{\partial v_i^u}{\partial c}\left(\tilde{c}, s_{i,t}\right) \times \left[1 - \tilde{\gamma}_{i,t} \times \frac{\tilde{c} - c_{i,t}^u}{\tilde{c}}\right],
\end{aligned}
$$

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. This approximation relies on the third– and higher order derivatives of the utility function to be small. The expansion demonstrates how the wedge in marginal utilities relates to the relative difference in consumption levels, scaled by 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 complementarity 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. We denote this by $E_0\left[v_i'\left(c_{i,0}\right)\right]$, so that we can write

$$
CS_k = \frac{E_k\left[v_i'\left(c_{i,t}^u\right)\right] - E_0\left[v_i'\left(c_{i,0}\right)\right]}{E_0\left[v'\left(c_{i,0}\right)\right]}. \tag{14}
$$

In practice, unemployment policies not only provide insurance, but also redistribute between different types of workers, including workers who are never unemployed. Our normalization emphasizes the insurance value of the policy. To incorporate the redistributive value, the consumption smoothing gains should be expressed relative to the average marginal utility of consumption in the labor force instead. Note that the evaluation of budget-balanced changes in the benefit profile as in equation (10) are independent of this normalization.

The relation between the consumption smoothing benefits and the consumption levels (under the above assumptions) is shown mostly clearly when preferences are homogeneous, i.e., $v_i\left(c\right) = v\left(c\right)$. In this case, we can approximate

$$
CS_k \cong \frac{v'\left(\bar{c}_k^u\right) - v'\left(\bar{c}_0\right)}{v'\left(\bar{c}_0\right)} \cong -\frac{v''\left(\bar{c}_0\right)\bar{c}_0}{v'\left(\bar{c}_0\right)} \times \frac{\bar{c}_0 - \bar{c}_k^u}{\bar{c}_0}, \tag{15}
$$

where $\bar{c}_0$ and $\bar{c}_k^u$ denote the average consumption level before the onset of the spell and during part

[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 6.2.

[14]In section 5.3, we discuss in more details issues related to this assumption. One example is the substitution towards home production, which has been analyzed extensively in the context of retirement (e.g., Aguiar and Hurst [2005]).

k of the spell.[15] The first approximation relies on Taylor expansions of $v'\left(c_{i,t}^u\right)$ and $v'\left(c_{i,0}\right)$ for each individual around the averages $\bar{c}_k^u$ and $\bar{c}_0$ respectively. The second approximation relies on a Taylor expansion of $v'\left(\bar{c}_k^u\right)$ around $\bar{c}_0^u$. 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, this 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 wedges. The validity of this implementation relies on homogenous preferences, but is robust to heterogeneity in assets or employment prospects and the corresponding dynamic selection affecting the consumption profiles.

When preferences are heterogeneous, our consumption-based implementation is further complicated by preference-based selection over the unemployment spell. Using Taylor expansions of $v'\left(c_{i,t}^u\right)$ around $c_{i,0}$ for each individual, we find

$$CS_k \cong \frac{E_k\left[v_i'\left(c_{i,0}\right)\right]-E_0\left[v_i'\left(c_{i,0}\right)\right]}{E_0\left[v_i'\left(c_{i,0}\right)\right]}-\frac{E_k\left[v_i''\left(c_{i,0}\right)\left(c_{i,0}-c_{i,t}^u\right)\right]}{E_0\left[v_i'\left(c_{i,0}\right)\right]}. \tag{16}$$

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. We explore this type of selection by analyzing how consumption levels before the onset of unemployment differ depending on the ex-post duration of the unemployment spell.[16] The second part of this expression indicates that the consumption smoothing gains are reduced if households with lower risk aversion remain unemployed for longer. We explore this by analyzing selection on observables that can arguably proxy for risk aversion.[17]

A well-known challenge for the consumption-based implementation is that information on individuals' preferences is required to translate differences in consumption to welfare (see Chetty and Finkelstein [2013]). These preferences are hard to measure and the relevant preferences will depend on the type of consumption we observe empirically. It is important to note though that our recommendations on the tilt of the benefit profile, depending on CS_1/CS_2, can be robust to these implementation errors when they are uncorrelated with the unemployment duration, for example when we over- or under-estimate the overall risk aversion. We explore this further in our empirical analysis. We also note that the limitations of the consumption-based implementation

[15] Formally, $\bar{c}_k^u = \frac{1}{D_k}\int \Sigma_{t=B_{k-1}}^{B_k} S_{i,t} c_{i,t}^u di$. Empirically, it simply corresponds to the average consumption level of individuals observed in the k-th part of the policy profile.

[16] Note that our model assumes a utilitarian social welfare function. With heterogeneous Pareto weights, the dynamic selection based on these weights will matter as well.

[17] 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]).

have inspired alternative approaches relating the marginal utility gap to observable behavioral responses: Chetty [2008] 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.

3 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.

3.1 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.[18]

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. Full-time workers would get daily benefits of 80% of their pre-unemployment daily wage throughout the spell, with daily benefits capped at 580SEK a day. The cap thus applies for daily wages above 725SEK.[19] 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) 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

[18]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.

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

912.5SEK) and the cap for benefits received after the first 20 weeks was increased to 680SEK.[20]

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 4 how the 2001 and 2002 variations in the time-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.

3.2 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 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. To define unemployment spells, we use the registration date at the PES as the start date and focus on individuals with no earnings who report to be searching for a full-time work. 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.).[21]

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 portfolio's for the universe of Swedish individuals from 1999 to 2007. The register contains detailed information on all financial assets (including debt) and real assets.[22] 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.[23] The comprehensiveness and detailed nature of both the income and wealth data in Sweden is exceptional, providing a

[20]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.

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

[22]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.

[23]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.

unique opportunity to investigate what means individuals use to smooth consumption (transfers, asset rebalancing, increase in debt, etc.). We also took advantage of the richness of the income and wealth data to construct a registry-based consumption measure at the yearly level that we use to assess the robustness of our consumption analysis.[24]

Finally, information on consumption is 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. 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 about 6,500 households for which we can correlate flow measures of consumption with time since (or until) the onset of the unemployment spell.

In Table 1, we provide summary statistics on unemployment, demographics, income and wealth for the sample of unemployed individuals used in the duration analysis of section 4. For the purpose of our regression kink design strategy, we restrict our attention to individuals whose daily wage is located within 150SEK of the 725SEK-kink point. 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 duration sample has yearly gross earnings (before any tax or payroll contribution) of 190,300 SEK.[25] 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, and 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.

In Table 2, we provide similar summary statistics for unemployed individuals from the HUT sample used in our consumption analysis of section 5. 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

[24] All details on the construction of the registry-based measure of consumption are given in Appendix C and in Kolsrud et al. [2015].

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

samples. The distribution is a little bit 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.

4 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.

4.1 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_1) and after 20 weeks of unemployment (b_2). The schedule of both b_1 and b_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_1 and b_2) on unemployment durations. For spells starting after July 2001 and before July 2002, the cap for b_1 is increased, while the cap in b_2 remains unchanged. The relationship between b_1 and previous wages therefore becomes linear around the 725 SEK threshold, where the schedule of b_2 still exhibits a kink at 725 SEK. This makes it possible to identify the effect on unemployment durations of a change in b_2 only. Finally, the cap in b_2 is also increased for spells starting after July 2002, so that kinks in the schedule of both b_1 and b_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:

$$Y = y(b_1, b_2, w, \mu),$$

where Y is the duration outcome of interest, μ is (unrestricted) unobserved heterogeneity, and b_1, b_2 and w (previous daily wage) are endogenous regressors. We are interested in identifying the marginal effect of benefits given during part k of the spell on the duration outcome Y, $\alpha_k = \frac{\partial Y}{\partial b_k}$. The RK design consists in exploiting the fact that b_k is a deterministic, continuous function of the wage w, kinked at $w = \bar{w}_k$. The RK design relies on two identifying assumptions. First, the direct marginal effect of w on Y should be smooth around the kink point $\bar{w}_k$. 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 $\bar{w}_k$. This second assumption implies imperfect sorting around the threshold $\bar{w}_k$, 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, α_k can be identified as:

$$\alpha_k = \frac{\lim_{w \to \bar{w}_k^+} \frac{\partial E[Y|w]}{\partial w} - \lim_{w \to \bar{w}_k^-} \frac{\partial E[Y|w]}{\partial w}}{\lim_{w \to \bar{w}_k^+} \frac{\partial b_k}{\partial w} - \lim_{w \to \bar{w}_k^-} \frac{\partial b_k}{\partial w}}$$

In practice, we provide estimates $\hat{\alpha}_k = \frac{\hat{\delta}_k}{\nu_k}$ where $\hat{\delta}_k$ is the estimated change in slope between Y and w at $\bar{w}_k$ and ν_k is the deterministic change in slope between b_k and w at $\bar{w}_k$. We estimate the former using the following regression model:

$$E[Y|w] = \beta_0 + \beta_1(w - \bar{w}_k) + \delta_k(w - \bar{w}_k) \cdot \mathbb{1}[w \geq \bar{w}_k]. \tag{17}$$

This model is estimated for $|w - \bar{w}_k| \leq h$, 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_1 and b_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_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_1 and b_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 $\hat{\delta}_k$ for the regression model in (17). We use a linear specification and a bandwidth h = 90SEK 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.[26] In line with the evidence presented in

[26] 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.

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 $\varepsilon_{D,b_k} = \hat{\delta}_k \cdot \frac{b_k^{cap}}{D^{cap}}$, where $\hat{\delta}_k$ is the estimated marginal slope change, D^{cap} is the observed average duration at the kink and $b_k^{cap} = .8 \times 725$SEK is the benefit level at the kink. 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_1 and b_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_2 of $\varepsilon_{D,b_2} = .68$ (.13). These elasticities are large overall, and at the upper end of estimates in the literature.

The same approach can be used to estimate the effect of benefits on the survival rate in unemployment at any spell length. Figure 4 shows the RK estimates for the effect of the benefit changes on the benefit durations D_1 and D_2, where $D_1 = \sum_{t<20wks} S_t$ is the time spent receiving benefit b_1 and $D_2 = \sum_{t\geq 20wks} S_t$ is the time spent receiving benefit b_2. For both D_1 and D_2, we find that the change in slope of the relation between wage and benefit duration is significant, but substantially smaller at the kink in b_2 than at the kink in both b_1 and b_2. Table 3 summarizes all estimates from Figures 3 and 4.

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.[27]

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. 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.

[27] 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.

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 up some underlying non-linearity in the relationship between wages and unemployment duration. In Appendix Figure B.7 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].[28] 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.

4.2 Implications for Moral Hazard Costs

Our results carry important implications for the moral hazard costs of modifying the time profile of UI benefits. Table 3 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

[28] 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.

(i.e., increasing both b_1 and b_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 (*i.e.* assuming no other expenditures $\bar{G}$ in equation (1)) for the average unemployment rate during the period 1999-2007. Note that when $\bar{G} > 0$, the required tax level to balance the government's total budget will be higher. When accounting for the fiscal externality through the rest of the tax system, the implied moral hazard cost would be even higher. In subsection 3 of technical Appendix A, 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).[29]

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_k simplifies to the expression in equation (13). The moral hazard cost of increasing benefits b_2 after 20 weeks of unemployment is therefore $MH_2 = \frac{\bar{b}+\tau}{\bar{b}}\frac{D}{D_2}\varepsilon_{D,b_2} = 1.44$ (.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_1 only.[30] This then implies that $MH_1 = 1.82$ (.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_1 and MH_2 provided in Table 3 nevertheless do not allow to reject that MH_1 is equal to MH_2. From this we conclude that our evidence suggests that the moral hazard cost of increasing benefits for the first 20 weeks is equal, or, if anything, larger than that of increasing benefits after 20 weeks of unemployment.

The fact that the moral hazard cost of unemployment benefits does not increase when timed later in the spell is somewhat surprising. As discussed in Proposition 1, in a stationary environment, one would expect forward-looking incentives to play in the opposite direction. Indeed, our results unambiguously show that unemployed individuals are forward-looking. The estimated elasticity of D_1, the duration spent on the first part of the profile, with respect to benefits b_2 received in the second part of the profile, reported in Figure 4, is positive and significant ($\varepsilon_{D_1,b_2} = .60$ (.11)). Unemployed individuals are thus not fully myopic, but react early in the spell to variation in benefits paid later in the spell.

The forward-looking behavior necessitates the presence of substantial non-stationarities to ex-

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

[30]We simply use $\varepsilon_{D,b} = \varepsilon_{D,b_1}\frac{b}{b_1} + \varepsilon_{D,b_2}\frac{b}{b_2} = \varepsilon_{D,b_1} + \varepsilon_{D,b_2}$ for $b_1 = b_2$.

plain the reverse pattern of moral hazard costs. Direct evidence for such non-stationarities can be found by comparing, for different durations t, the elasticity of $\tilde{D}_t$ with respect to a flat benefit level $\bar{b}$, where $\tilde{D}_t$ is the remaining duration of unemployment at t, conditional on surviving until t. As discussed in section 2.2, these elasticities would remain constant in a stationary environment and be equal to $\varepsilon_{D,\bar{b}}$, the elasticity of the total unemployment duration.[31] Results, reported in Figure 5, however, show that the elasticity of the remaining duration strongly declines as a function of t, indicating that the responsiveness of exit rates decreases substantially over the unemployment spell. Such non-stationarity, which can be the result of dynamic selection and/or duration dependence, is large enough to offset the significant effect of forward-looking incentives, so that the incentive cost of increasing the generosity of long-term benefits is still somewhat smaller than that of increasing the generosity of short-term benefits.

In appendix section B.1 we also report the RKD estimates of the effect of UI benefits on the hazard rates out of unemployment. This allows to further investigate the non-stationary patterns in unemployment responses. Figure B.1 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. This evidence is supportive of the strong non-stationary patterns reported in Figure 5. Appendix Figure B.1 also confirms that unemployed individuals are forward-looking, as benefits b_2 received after 20 weeks do have an effect on the hazard rate in the first 10 weeks. However, the effect on hazards early in the spell of benefits received in the first 20 weeks (b_1) is almost twice as large as the effect of b_2. By reducing the survival into the second part of the unemployment spell, this also creates a large mechanical effect of b_1 on D_2, the time spent on the second part of the policy. These results are detailed and discussed in appendix section B.1.

5 Consumption Smoothing Over the Unemployment Spell

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

5.1 Consumption Over the Unemployment Spell

In order to analyze the evolution of consumption as a function of time spent unemployed, we have 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

[31] Note that the elasticity of the remaining duration $\tilde{D}_t$ with respect to the continuation policy after period t, which we considered in section 2.2 is the same as the elasticity with respect to an overall change in the benefit level in a stationary environment with a representative agent, since exit rates then only depend on the continuation policy.

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 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. Furthermore, we also provide evidence of the robustness of our results in a pure panel setting using registry data in Appendix C.

We start by providing graphical evidence for the evolution of average consumption as a function of time spent unemployed. We report in Figure 6 the estimated coefficients β_k from the regression model:

$$c_{i,t} = \sum_{k=-3}^{+4} \beta_k \cdot \mathbb{1}[t \in sem_k] + X_i'\gamma + \varepsilon_{i,t} \tag{18}$$

where $c_{i,t}$ is household consumption in the HUT survey observed at the time of the HUT interview and $\mathbb{1}[t \in sem_k]$ is an indicator for being observed in the k-th semester of the unemployment spell.[32] 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 6 represent the average consumption levels (in constant SEK) for households with an individual observed in her k-th semester of unemployment, relative to the average consumption level of households with an individual observed just one semester prior to becoming unemployed.

The graph provides evidence 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 60kSEK lower than the average consumption of households with a member before the onset of his or her spell. This represents an 18.1% drop in average household consumption. Furthermore, Figure 6 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 four semesters preceding the onset of the unemployment spell. This suggests that unemployment shocks are relatively unanticipated, or that households have little ability 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 4 to a regression setting. In column (1), we start by running the simple regression:

$$c_{i,t} = \eta_1 \cdot \mathbb{1}[0 < t <= 20 \text{ wks}] + \eta_2 \cdot \mathbb{1}[t > 20 \text{ wks}] + X_i'\gamma + \varepsilon_{i,t}, \tag{19}$$

[32] $k = 0$ is the baseline and represents consumption in the last semester 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.

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 4, 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

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

where $\mathbb{1}[L > 20 \text{ wks}]$ is a dummy variable for the total 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. 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.

To alleviate concerns regarding the small sample size and pseudo-panel nature of the HUT data, we replicate this exercise using a pure panel and a residual measure of consumption expenditures constructed from registry-based data. All details and results are reported in Appendix C. 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. While it offers the advantage of being computable for the universe of unemployed households from 1999 to 2007, the measure can only capture yearly expenditures between December of each year.

We therefore focus on individuals who become unemployed in January and investigate household consumption after exactly one year of unemployment and exactly two years of unemployment, compared to pre-unemployment consumption. We find first that pooled estimates and panel model estimates (within-estimator, and first-difference estimator) yield very similar results for the consumption profiles over the unemployment spell, which confirms the evidence from the HUT sample that the consumption drops over the spell are not driven by dynamic selection on consumption levels. Second, results confirm that there is no significant selection on consumption profile over the unemployment spell: the drop in household consumption after a year of unemployment is not statistically different for households who select into long spells and households who select into shorter spells. Finally, the magnitude of the consumption drop over the unemployment spell is very similar to that found in the HUT. After a year spent unemployed, our registry-based measure of yearly household expenditures is 50kSEK lower than at the start of the unemployment spell, which translates into a 16% drop in household consumption.

Overall, our evidence thus indicates that consumption expenditures do significantly decline over the unemployment spell, both on average and within households. This decline in consumption expenditures does not appear to be a result of differential selection on consumption levels or differential selection on consumption profiles over the unemployment spell.

5.2 Consumption Smoothing Means

We investigate the means to smooth consumption available to households and how they affect the decline in consumption expenditures.

To gauge the magnitude of the decline in consumption documented above, we first compare the drop in consumption to the drop in household income caused by job loss. Given an average replacement rate of 72% (cf. Table 1) and average individual annual earnings of 203,000 SEK prior to job loss (cf. Table 2), unemployment causes a decline in annual income of approximately 60kSEK. This amount is very close to the total drop in consumption expenditures for households where an individual has been unemployed for a year, as shown in Figure 6. This indicates that the drop in household consumption over the unemployment spell is large relative to the earnings loss generated by job displacement. After a year in unemployment, the decline in household consumption is almost similar to the decline in household income. This evidence suggests that, while most households may have means to smooth consumption early on in the spell, after a year in unemployment, they have almost entirely exhausted their means to smooth consumption.

This observation is in line with evidence attesting that unemployed households have limited means to smooth consumption beyond the unemployment benefits they receive. First, as shown in Table 2, for the vast majority of households, existing assets offer very limited ability to smooth consumption over an unemployment spell. On average, households have less than 20% of annual disposable income in liquid assets on their bank accounts at the start of an unemployment spell, and more than 50% of households have no liquid assets at all. Overall, more than 50% of households enter an unemployment spell with no positive net wealth, and more than 25% even start a spell with negative net wealth due to large household debt.

Second, evidence from registry data indicates that variations in earnings and income of other members of the households offer little consumption smoothing opportunities over the unemployment spell. In Figure 7, we report for all unemployed individuals the evolution of total disposable income of all other members of their household as a function of the time they have spent unemployed. To do so, we take advantage of the panel structure of the registry data, and estimate, using a within-estimator, fixed-effect models of the form:

$$Y_{i,t} = \sum_{k=-3}^{+8} \tilde{\beta}_k \cdot \mathbb{1}[t \in qtr_k] + X_i'\gamma + \alpha_i + \mu_{i,t} \tag{21}$$

where $Y_{i,t}$ is total yearly disposable income of all other members of the household, α_i is a household fixed-effect and $\mathbb{1}[t \in qtr_k]$ is an indicator for the unemployed being observed in the $k-$th quarter of her unemployment spell. Results, displayed in Figure 7, show that within-household changes in the disposable income of all other members of the household are not significantly different from zero throughout the unemployment spell. This suggests that in our context, the added-worker effect is not playing any significant role in increasing household consumption in response to an unemployment shock, even for long term unemployed.

Finally, evidence from registry data also shows that debt does not offer much help in smoothing consumption over the unemployment spell. A significant fraction of households enter unemployment spells with large debt, most of which is mortgage-related. Interestingly, in Figure 8, 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 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 close to hand-to-mouth, with high marginal propensity to consume.

5.3 Implications for Consumption Smoothing Gains

The empirical evidence presented above indicates that consumption declines significantly over the unemployment spell, both across and (mostly) within households. 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. We now discuss the robustness of this finding to two important issues.

Selection Our analysis in section 2.4 indicates that for a given profile of average consumption levels, when individuals have heterogeneous preferences, the evolution of the consumption smoothing gains of UI over the spell will depend on the selection of individuals at different durations. If, everything else equal, individuals with lower risk aversion select into longer unemployment spells, this would tend to decrease the consumption smoothing value of UI benefits over the spell.

Differences in preferences affect household consumption and may translate, when households have the means to do so, into different consumption levels and profiles in response to unemployment. Evidence displayed in section 5.1 shows, however, that selection on consumption levels and consumption profiles is actually quite limited over the unemployment spell in our context. Yet, even absent selection on consumption profiles, dynamic selection on risk preferences could still take place over the unemployment spell.

To further assess the potential magnitude of such selection, we investigate in Table 5 how various observable characteristics that have been shown to correlate with risk preferences are distributed across short term and long term unemployed. The patterns of selection are of small 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 absence 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 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 selection on observable characteristics does not enable us to conclude that average risk aversion would significantly increase (or decline) over the unemployment spell.

Consumption vs. expenditures Our empirical analysis considers changes in expenditures. An important question is to what extent expenditures translate into consumption and thus capture the potential welfare value of unemployment benefits. 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 6, 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.[33] 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.[34]

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.[35] 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. Results displayed in Appendix Tables C.2 and C.3 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.

Table 6 also indicates that expenditures on durable goods such as the purchase of new vehicles 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.[36] This in turn will tend to further increase 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, 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 selection and dynamic substitution across categories of expenditures. Yet, other methods exploiting comparative statics of effort choices as in Chetty [2008], or Landais [2015], could also be developed to evaluate the evolution of consumption smoothing gains over the spell. These methods could circumvent the issue, encountered in our consumption implementation approach, of having to make assumptions regarding dynamic selection on risk preferences.

[33] 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.

[34] 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]).

[35] The sign of the impact of such substitution patterns on the level 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.

[36] Note again that, although results are relatively imprecise due to the small sample size, Appendix Table C.2 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.

6 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. This allows for a transparent evaluation of local changes in the policy profile, but relies on the implementation assumptions discussed before. We then illustrate how a calibrated structural model allows us to go beyond our local policy recommendations and also to gauge their robustness.

6.1 Sufficient Statistics Approach

The welfare consequences of an increase in UI benefits are reported in Table 7. The first row examines the consequences of increasing the benefit level $\bar{b}$ throughout the unemployment spell; the second row examines the consequences of increasing the benefit b_1 during the first 20 weeks of unemployment and the third row examines the consequences of increasing the benefit b_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 3). As noted before, our estimates of the moral hazard cost are high, but lower for benefits paid later in the spell. The estimated marginal cost of making the policy more generous is 25 percent higher during the first 20 weeks of unemployment than after 20 weeks of unemployment. The second column repeats our estimates of the drops in average consumption (column (1) of Table 4), respectively over the full unemployment spell, during the first 20 weeks of unemployment and after 20 weeks of unemployment. We convert the respective consumption drops into estimates of the consumption smoothing gains CS_k, following the implementation in (15), which relies on a Taylor approximation and homogeneous preferences. As the appropriate value of risk aversion in this context is unclear, the consumption smoothing gains are reported for a range of plausible values (see Chetty [2009], Chetty and Finkelstein [2013]). Regardless of its level, 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.[37] 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, the return to the marginal krona spent on the unemployed, accounting for the duration response, is substantially lower than 1. These estimates are, however, sensitive to our implementation assumptions including the risk preference parameter γ, the tax distortion τ, our normalization of the value of public funds, etc.[38]

Comparing the benefit-costs ratio's for different parts of the policy allows us to formulate

[37] 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 3.2.

[38] Note that consumption drops are experienced at the household level, but our analysis only accounts for the consumption smoothing gains for the unemployed individual and thus underestimates the overall consumption smoothing gain of UI.

recommendations on the tilt of the benefits profile that are less sensitive to the above assumptions. The relative moral hazard costs MH_1/MH_2 are more than twice as high as the relative consumption smoothing gains CS_1/CS_2 (still assuming homogeneous preferences). The implied return to the marginal krona spent on the short-term unemployed is at least twice as low compared to the long-term unemployed. Following Corollary 1, our estimates thus indicate that welfare could be increased by introducing an inclining tilt ($b_1 < b_2$) in the flat Swedish benefit profile.

In sum, our estimates suggest that 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. A more recent reform in 2007 introduced a declining tilt in the profile for all unemployed. The replacement rate for the long-term unemployed was reduced to 70% after 40 weeks and 65% after 60 weeks, while maintaining the replacement rate at 80% for the short-term unemployed. Our analysis indicates that this reduction in generosity may well have increased welfare, but reducing benefits for the short-term unemployed instead would have increased welfare more.

6.2 Beyond Local Recommendations

To go beyond our local recommendations, we need to know how the sufficient statistics vary with the policy parameters. Both the moral hazard costs and consumption smoothing gains are endogenous to the unemployment policy, but we only have estimates evaluated at the current policy. We illustrate how this can be explored based on a calibrated, structural model.

Calibration The structural model we calibrate assumes homogeneous CRRA preferences and additive search costs, an exit rate function with heterogeneous returns to search that depreciate exponentially, and a heterogeneous asset distribution at the start of the unemployment spell with a uniform asset limit below which agents are borrowing constrained. The asset distribution is set to match the empirical distribution in our sample before the onset of the unemployment spell. We use CRRA parameter $\gamma = 2$ and calibrate the remaining parameters to target the variables underlying our sufficient statistics, including benefit duration elasticities ε_{D,b_k}, average benefit durations D_k and drops in average consumption $[\bar{c}_k - \bar{c}_0]/\bar{c}_0$. Appendix D describes functional forms, exit rate and asset distributions, and our calibration approach in more detail.

The calibrated model matches our targets reasonably well (see Appendix Table D.1). We slightly overestimate the consumption drop during the first 20 weeks of unemployment and underestimate the consumption drop after 20 weeks. The simulated drops are well within the 95% confidence intervals of our empirical estimates, but imply that we underestimate the CS_2/CS_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_2/MH_1. The local policy recommendation of the calibrated model (starting from a flat replacement rate $\bar{b} = .72$) remains unchanged: the unemployment policy is too generous overall and especially so for the short-term unemployed.

Counterfactual Analysis The main objective of this calibration exercise is to illustrate how the moral hazard costs and consumption smoothing gains change for different UI benefit levels. Figure 9 shows how the respective costs and gains for the different parts of the unemployment policy evolve when reducing the generosity of a flat profile. Two findings emerge. First, the consumption smoothing gains increase while the moral hazard costs decrease as we reduce the benefit level.[39] 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 ($CS_2 > CS_1$), while the moral hazard costs remain lower ($MH_2 < MH_1$). The introduction of an inclining tilt ($b_2 > b_1$) thus remains welfare-improving for lower replacement rates. Our calibrated model predicts that welfare would be maximized by setting $b_1 = .48$ for the short-term unemployed and $b_2 = .68$ for the long-term unemployed (while keeping the tax rate unchanged) .

The structural model allows analyzing the robustness of these findings for different model specifications. Appendix Figure D.1 illustrates this when eliminating either the depreciation or the heterogeneity in 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 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 heterogeneity continues to prescribe an inclining tilt and this remains true for lower replacement rates. The depreciation of exit rates, however, increases MH_2 relative to MH_1. As individuals exert more effort early on in the spell to avoid being stuck into long spells, the time spent on the second part of the policy D_2 decreases, while the elasticity to changes in the second part ε_{D,b_2} increases. This non-stationary force pushes towards a declining tilt and more so as the overall generosity is reduced. It is, however, more than offset by the heterogeneity in exit rates in our baseline calibration.

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 finally note that the structural model can be exploited to assess the potential importance of some implementation assumptions. Our calibrated model indicates that the Taylor approximation of the marginal CRRA utilities tends to underestimate the consumption smoothing gains, especially when the level of risk aversion is high (see Appendix Table D.2). This was already noted in Chetty [2006]. Interestingly, the approximation error on the relative consumption smoothing gains is very

[39]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.

small and thus the recommendation on the tilt unaffected, regardless of the level of risk aversion. We also use the structural model to gauge the impact of heterogeneity in preferences on the estimated consumption smoothing gains. We find that the relative drops in consumption are very similar for different levels of risk aversion (see Table D.2), but dynamic selection on marginal utility and risk preferences could in principle affect the consumption smoothing gains, as shown in section 2.4. In our calibrated model the introduction of preference heterogeneity has a limited impact on the relative consumption smoothing gains and, if anything, increases the welfare gain from introducing an inclining tilt ($b_2 > b_1$).[40]

7 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 behaviors 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. The replication of our empirical exercise for different contexts or subcategories of individuals can provide the necessary estimates in order to assess the optimality of contingent profiles such as age-dependent or business-cycle dependent profiles.

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.

[40] Note that in our model with CRRA preferences and additive search cost (see also Hopenhayn and Nicolini [1997] and Chetty [2008]), individuals with higher risk aversion have lower marginal utility of consumption (relative to the disutility of search) and are predicted to leave unemployment more slowly. We provide more detail on the simulated impact of dynamic selection on the consumption smoothing gains in Appendix D.

References

Aguiar, Mark and Erik Hurst, "Consumption versus Expenditure," *Journal of Political Economy*, 2005, *113* (5), pp. 919–948.

Andrews, Isaiah and Conrad Miller, "Optimal Social Insurance with Heterogeneity," *Mimeo, MIT*, 2013.

Baily, Martin N., "Some Aspects of Optimal Unemployment Insurance," *Journal of Public Economics*, 1978, *10* (3), 379–402.

Card, David, David S. Lee, Zhuan Pei, and Andrea Weber, "Inference on Causal Effects in a Generalized Regression Kink Design," *Econometrica*, 2015, *83* (6), 2453–2483.

Chetty, Raj, "A General Formula for the Optimal Level of Social Insurance," *Journal of Public Economics*, 2006, *90* (10-11), 1879–1901.

_ , "Moral Hazard versus Liquidity and Optimal Unemployment Insurance," *Journal of Political Economy*, 2008, *116* (2), 173–234.

_ , "Sufficient Statistics for Welfare Analysis: A Bridge Between Structural and Reduced-Form Methods," *Annual Review of Economics*, 05 2009, *1* (1), 451–488.

_ and Adam Szeidl, "Consumption Commitments and Risk Preferences," *The Quarterly Journal of Economics*, 2007, *122* (2), pp. 831–877.

_ and Amy Finkelstein, "Chapter 3 - Social Insurance: Connecting Theory to Data," in Martin Feldstein Alan J. Auerbach, Raj Chetty and Emmanuel Saez, eds., *handbook of public economics, vol. 5*, Vol. 5 of *Handbook of Public Economics*, Elsevier, 2013, pp. 111 – 193.

Cohen, Alma and Liran Einav, "Estimating Risk Preferences from Deductible Choice," *American Economic Review*, 2007, *97* (3), 745–788.

den Berg, Gerard Van, "Nonstationarity in Job Search Theory," *The Review of Economic Studies*, 1990, *57* (2), 255–277.

Eckstein, Zvi and Gerard Van den Berg, "Empirical labor search: A survey," *Journal of Econometrics*, 2007, *136* (2), 531 – 564. The interface between econometrics and economic theory.

Ganong, Peter and Simon Jaeger, "A Permutation Test and Estimation Alternatives for the Regression Kink Design," *Harvard University Discussion Paper*, 2014.

Gruber, Jonathan, "The Consumption Smoothing Benefits of Unemployment Insurance," *American Economic Review*, 1997, *87* (1), 192–205.

Hendren, Nathaniel, "The Policy Elasticity," Working Paper 19177, National Bureau of Economic Research June 2013.

Hopenhayn, Hugo A. and Juan Pablo Nicolini, "Optimal Unemployment Insurance," *Journal of Political Economy*, 1997, *105* (2), 412–438.

Koijen, Ralph, Stijn Van Nieuwerburgh, and Roine Vestman, *Judging the Quality of Survey Data by Comparison with Truth as Measured by Administrative Records: Evidence From Sweden*, University of Chicago Press, July 2014

Kolsrud, Jonas, Camille Landais, Peter Nilsson, and Johannes Spinnewijn, "The Optimal Timing of UI Benefits: Theory and Evidence from Sweden," CEPR Discussion Papers 10701, C.E.P.R. Discussion Papers July 2015.

Kreiner, Claus T., David Dreyer Lassen, and Søren Leth-Petersen, "Liquidity Constraint Tightness and Consumer Responses to Fiscal Stimulus Policy," *Mimeo, Copenhagen University*, 2014.

Kroft, Kory, Fabian Lange, and Matthew J. Notowidigdo, "Duration Dependence and Labor Market Conditions: Evidence from a Field Experiment," *The Quarterly Journal of Economics*, 2013.

Krueger, Alan B. and Bruce D. Meyer, "Labor Supply Effects of Social Insurance," Handbook of Public Economics, Elsevier 2002.

Landais, Camille, "Assessing theWelfare Effects of Unemployment Benefits Using the Regression Kink Design," *American Economic Journal: Economic Policy*, 2015.

_ , Pascal Michaillat, and Emmanuel Saez, "Optimal Unemployment Insurance over the Business Cycle," Working Paper 16526, National Bureau of Economic Research 2010.

Lentz, Rasmus and Torben Tranas, "Job Search and Savings: Wealth Effects and Duration Dependence," *Journal of Labor Economics*, July 2005, *23* (3), 467–490.

Machin, Stephen and Alan Manning, "The Causes and Consequences of Longterm Unemployment in Europe," in Orley C. Ashenfelter and David Card, eds., *Orley C. Ashenfelter and David Card, eds.*, Vol. 3, Part C of *Handbook of Labor Economics*, Elsevier, 1999, pp. 3085 – 3139.

Mas, Alexandre and Andrew Jonhston, "Potential Unemployment Insurance Duration and Labor Supply: the Individual and Market-Level Response to a Benefit Cut," *Mimeo, Princeton University*, 2015.

McCrary, Justin, "Manipulation of the Running Variable in the Regression Discontinuity Design: A Density Test," *Journal of Econometrics*, February 2008, *142* (2), 698–714.

Mogstad, Magne and Andreas Kostol, "Earnings, Income, and Consumption of Alowed and Rejected Disability Insurance Applicants," *American Economic Review, Papers and Proceedings*, 2015, *105* (5), 137–141.

Nekoei, Arash and Andrea Weber, "Does Extending Unemployment Benefits Improve Job Quality?," IZA Discussion Papers 9034, Institute for the Study of Labor (IZA) April 2015.

Pavoni, Nicola, "Optimal Unemployment Insurance, With Human Capital Depreciation, and Duration Dependence," *International Economic Review*, 2009, *50* (2), 323–362.

Pistaferri, Luigi, "Household Consumption: Research Questions, Measurement Issues, and Data Collection Strategies," *Mimeo Stanford University*, 2015.

Rothstein, Jesse, "Unemployment Insurance and Job Search in the Great Recession," Working Paper 17534, National Bureau of Economic Research October 2011.

Schmieder, Johannes F., Till von Wachter, and Stefan Bender, "The Effects of Extended Unemployment Insurance Over the Business Cycle: Evidence from Regression Discontinuity Estimates Over 20 Years," *The Quarterly Journal of Economics*, 2012, *127* (2), 701–752.

Shavell, Steven and Laurence Weiss, "The Optimal Payment of Unemployment Insurance Benefits over Time," *Journal of Political Economy*, 1979, *87* (6), 1347–1362.

Shimer, Robert and Iván Werning, "On the Optimal Timing of Benefits with Heterogeneous Workers and Human Capital Depreciation ," 2006.

_ and _ , "Reservation Wages and Unemployment Insurance," *Quarterly Journal of Economics*, 2007, *122* (3), 1145–1185.

_ and _ , "Liquidity and Insurance for the Unemployed," *American Economic Review*, 2008, *98* (5), 1922–42.

Singh, S. and G. Maddala, "A Function for the Size Distribution of Incomes," *Econometrica*, 1976, *44* (5), 963–970.

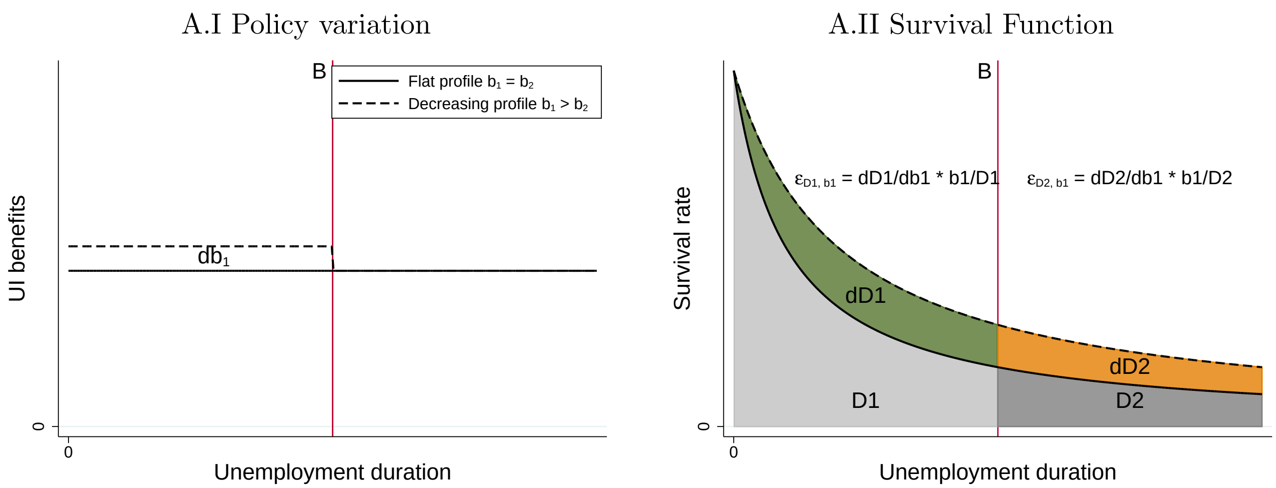
Spinnewijn, Johannes, "Unemployed but Optimistic: Optimal Insurance Design with Biased Beliefs," *Journal of the European Economic Association*, 2015, *13* (1), 130–167.

Valletta, Robert and Henry S. Farber, "Extended Unemployment Insurance and Unemployment Duration in The Great Recession: The U.S. Experience," *Mimeo, Princeton University*, 2011.

Werning, Iván, "Optimal Unemployment Insurance with Unobservable Savings," *University of Chicago and UTDT*, 2002.

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

A. Unemployment response wrt b_1



B. Unemployment response wrt b_2

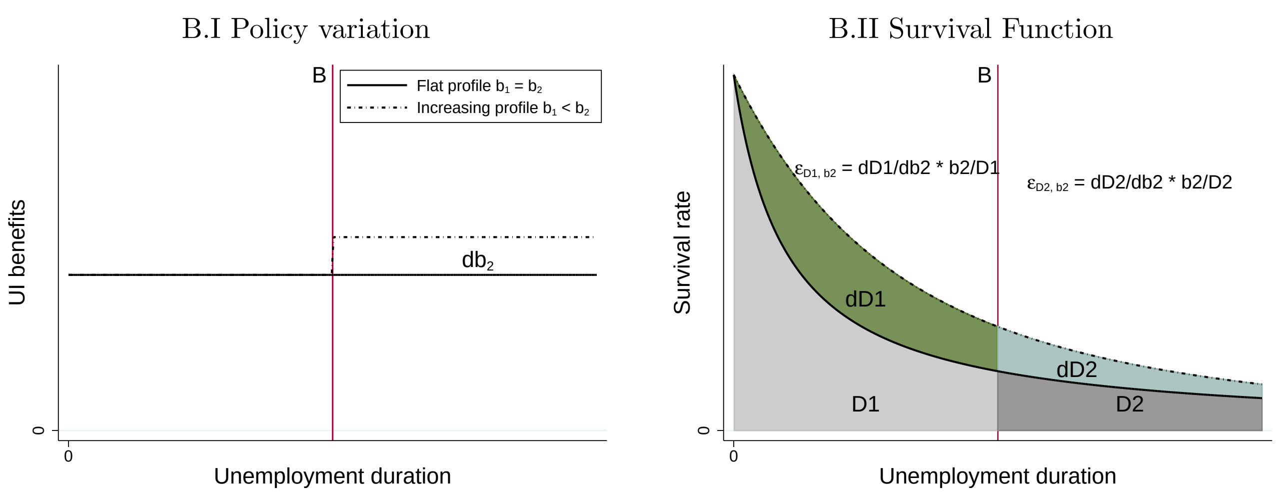
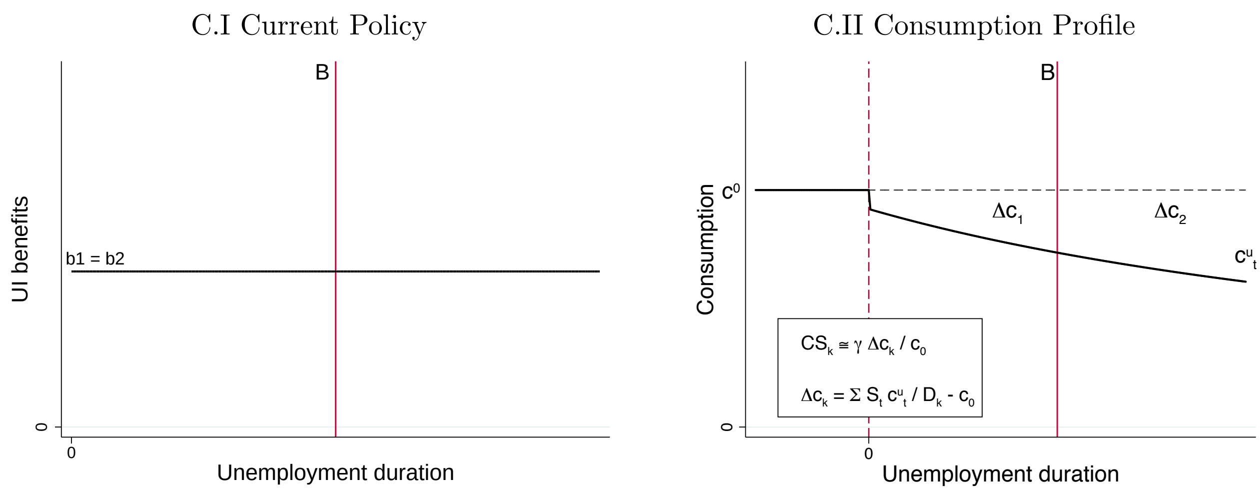


Figure 1: SUFFICIENT STATISTICS FOR WELFARE ANALYSIS OF TWO-PART POLICY (*continued)*

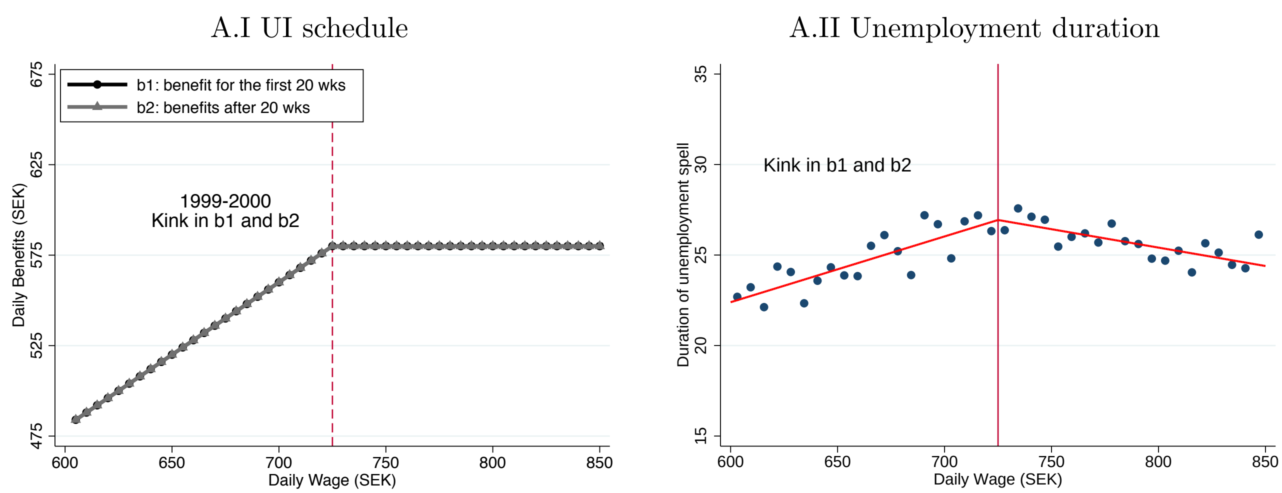
C. Consumption profile for current policy



Notes: The figure summarizes the policy variation and statistics needed to characterize an optimal two-part profile giving b_1 for the first B weeks and b_2 afterwards. In Panel A, we display policy variation db_1 in benefits given for the first B weeks that allows evaluating the moral hazard cost of a change in b_1, MH_1. The moral hazard cost depends on the responses of the duration spent in the first part of the profile D_1 and in the second part of the profile D_2, as exemplified in the right panel. These responses enable the identification of the cross-duration elasticities ε_{D_1,b_1} and ε_{D_2,b_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 (13). In Panel B, we display policy variation that allows to evaluate the moral hazard cost of a change in b_2. Like in Panel A, we start from a flat profile and display variation db_2 in benefits given after B weeks, which enables the identification of the cross-duration elasticities ε_{D_1,b_2}, ε_{D_2,b_2} entering the RHS of the corresponding dynamic Baily-Chetty formula (12). To evaluate the consumption smoothing gains of the two-part policy, following the implementation in (15), the planner requires the average drop in consumption Δc_1 for individuals in the first part of the profile receiving b_1, and the average drop in consumption Δc_2 for individuals in the second part of the profile receiving b_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



B. 2001

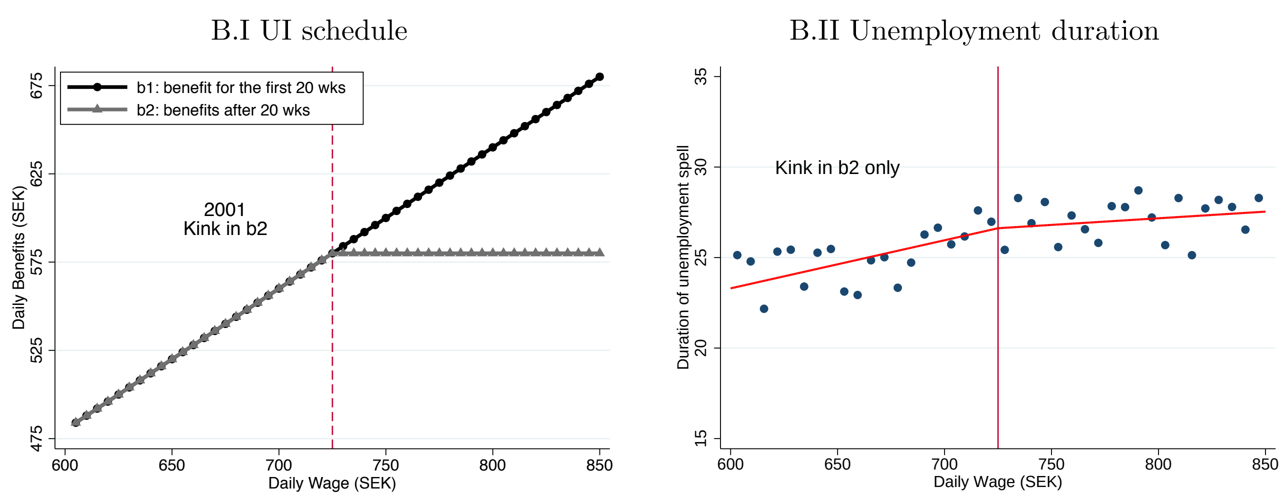
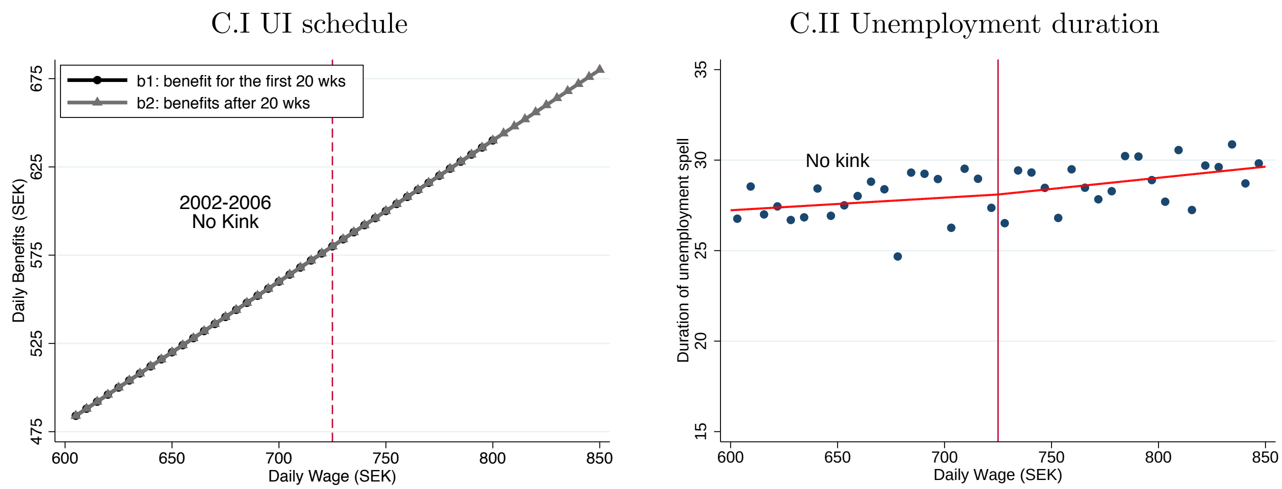


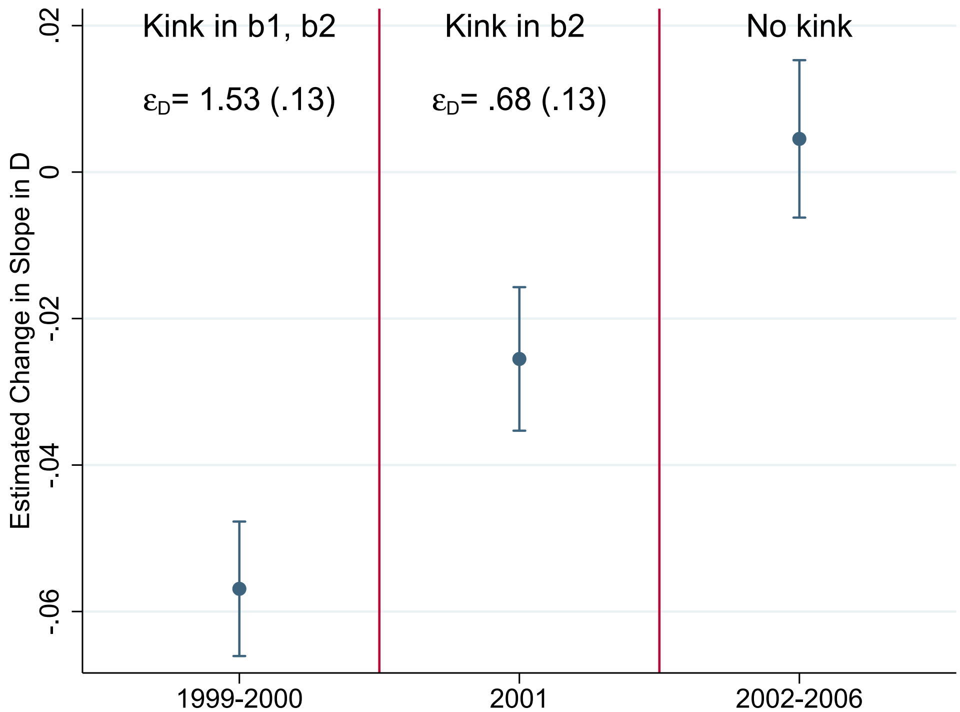
Figure 2: UI benefits and unemployment duration as a function of daily wage around the 725SEK threshold (*continued*)

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 (17) are displayed in Table 3. The red lines display predicted values of the regressions in the linear case.

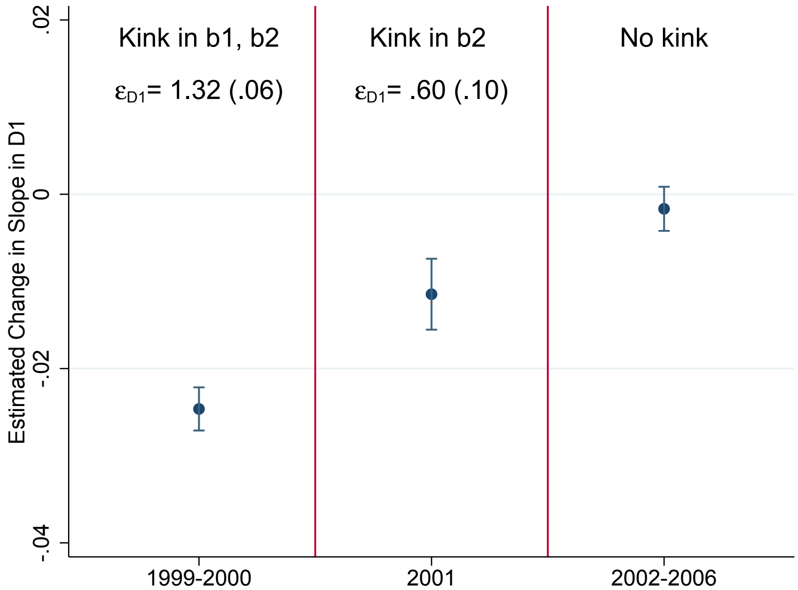
Figure 3: RKD ESTIMATES AT THE 725SEK THRESHOLD



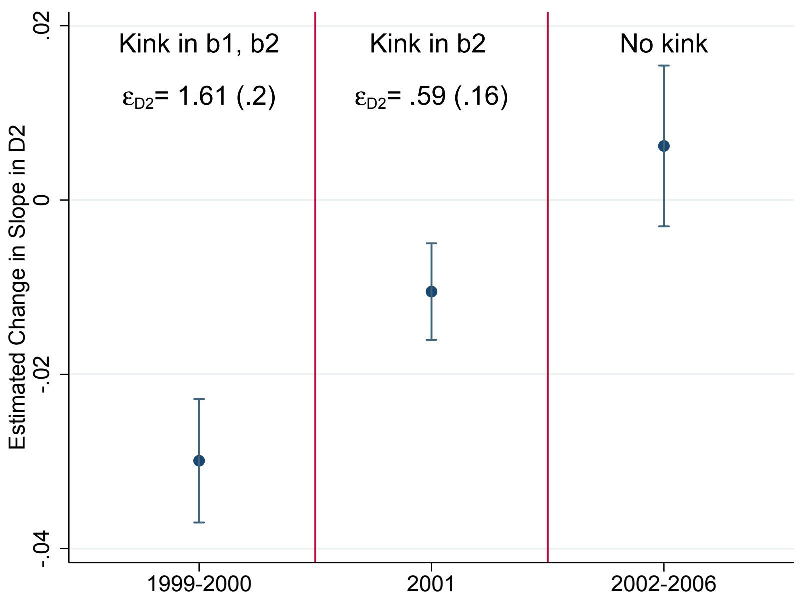
Notes: The figure reports estimates of the change in slope with 95% robust confidence interval in the relationship between unemployment duration and daily wage at the 725SEK threshold using linear regressions of the form of equation (17) with a bandwidth size h = 90SEK. These estimates are reported for 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_1 and b_2 (for period 1999-2000) and with respect to b_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 3.

Figure 4: RKD ESTIMATES AT THE 725SEK THRESHOLD

A. Time spent on first part of the profile (D_1):
unemployment duration up to 20 weeks

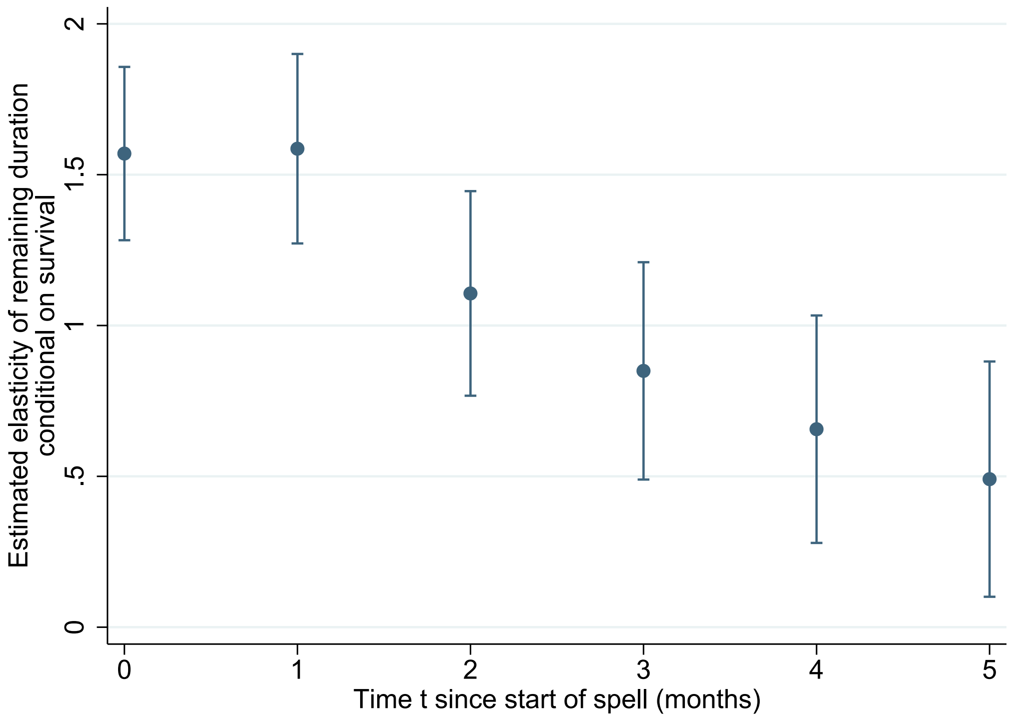


B. Time spent on second part of the profile (D_2):
unemployment duration after 20 weeks



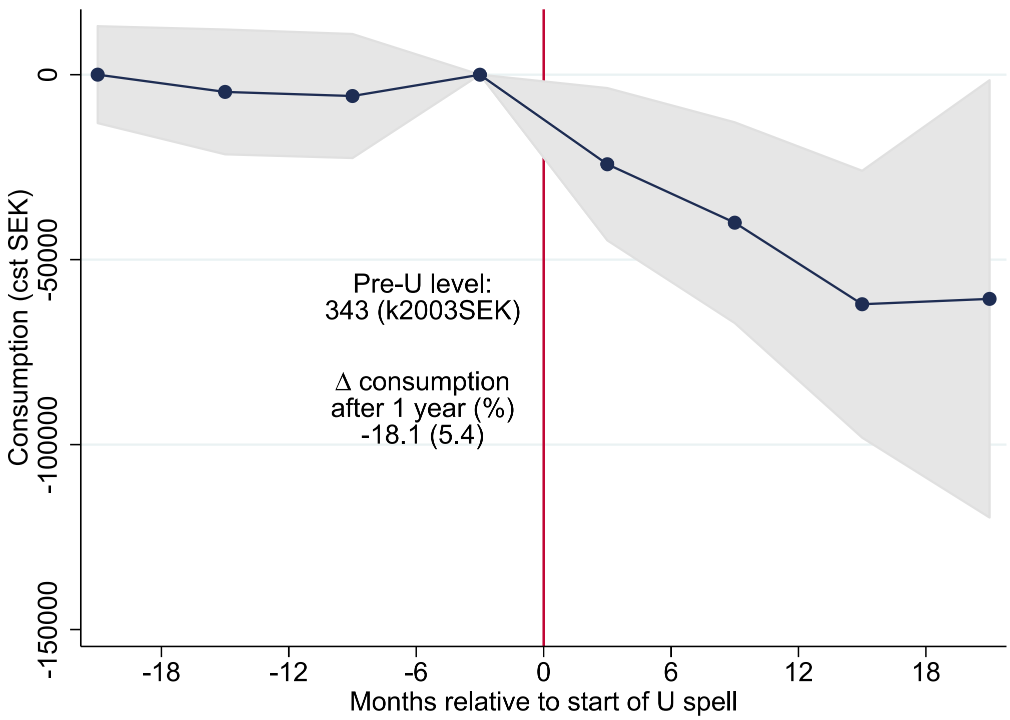
Notes: The figure reports estimates of the change in slope with 95% robust confidence interval in the relationship between daily wage and time spent in the first part of the benefit profile D_1 and time spent in the second part of the profile D_2, at the 725SEK wage threshold, using linear regressions of the form of equation (17) 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. $D_1 = \sum_{t<20wks} S_t$ 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). Total duration of unemployment is capped at 2 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_1 and b_2 (for period 1999-2000) and with respect to b_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 3.

Figure 5: TESTING FOR STATIONARITY: ELASTICITY OF THE REMAINING DURATION OF UNEMPLOYMENT, CONDITIONAL ON SURVIVING UNTIL t, WITH RESPECT TO CHANGES IN THE FLAT BENEFIT LEVEL



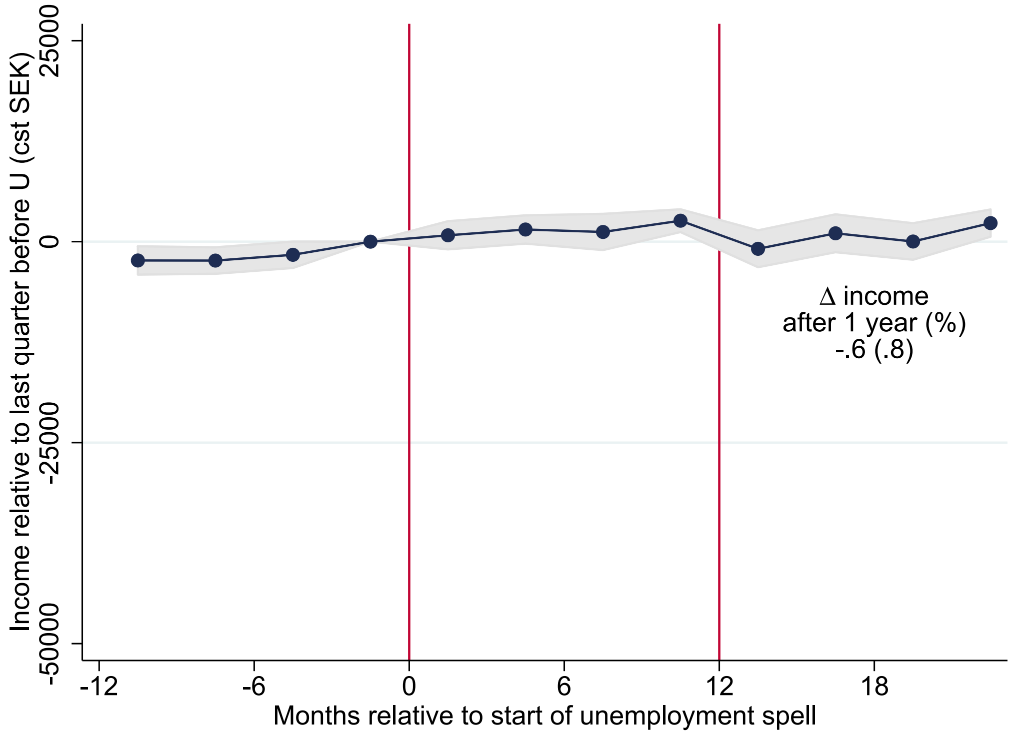
Notes: The figure reports RKD estimates (with 95% robust confidence interval) of the elasticity of the remaining duration of unemployment conditional on surviving until t with respect to changes in the flat benefit level $\bar{b}$. Estimates use the presence, for spells starting before July 2001, of 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 (17) 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 6: Household consumption relative to last semester prior to the unemployment spell



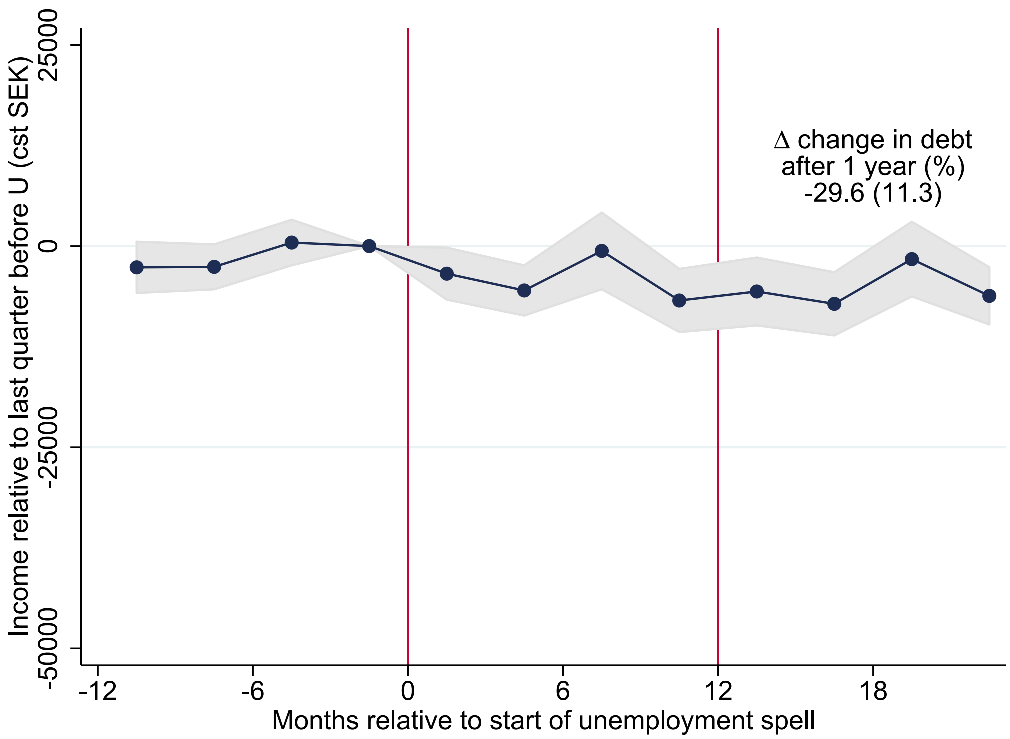
Notes: The figure correlates household consumption measured in the HUT survey with the time t since (or until) the onset of the unemployment spell observed in the administrative UI records. Bi-weekly household consumption levels measured at the moment of the HUT interview are annualized and expressed in constant SEK2003. The figure follows from regression model (18) and plots the estimated coefficients β_k for the set of indicators $\mathbb{1}[t \in sem_k]$ for being observed in the k-th semester 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 60kSEK lower than the average consumption of household with a member at the start of her spell, which represents a 18.1% drop in average household consumption.

Figure 7: Yearly income of all other members of the household as a function of unemployment duration



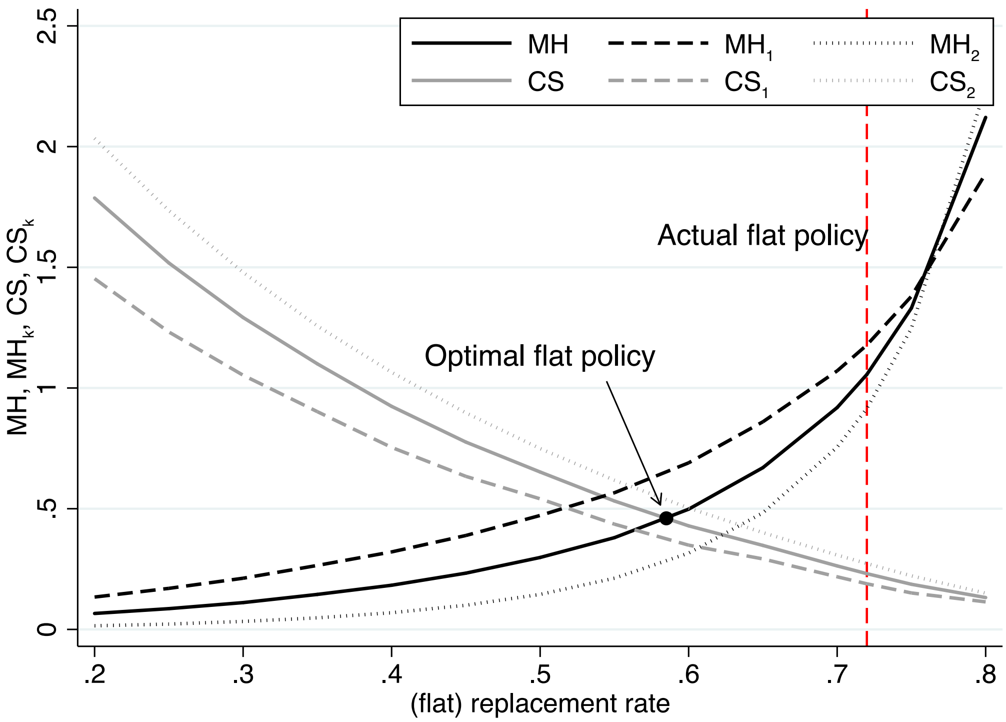
Notes: In this figure, we explore added worker effects and estimate how total yearly disposable income of all members of the household other than the unemployed individual correlates with the time t since (or until) the onset of the unemployment spell. We take advantage of the panel structure of the registry data, and estimate a model with household fixed effects using a within estimator. The figure follows from regression model (21) and plots the estimated coefficients $\tilde{\beta_k}$ for the set of indicators $\mathbb{1}[t \in qtr_k]$. These coefficients identify the average within household change in yearly disposable income of all other members of the household in the $k-$th quarter of unemployment, relative to the last quarter prior to unemployment. Results show that within-household changes in disposable income of all other members of the household are small and not significantly different from zero throughout the unemployment spell. This suggests that in our context, the added-worker effect is not playing any significant role in increasing household consumption in response to an unemployment shock.

Figure 8: Yearly change in debt for households with no real estate assets at the start of the spell as a function of unemployment duration



Notes: In this figure, we explore how the yearly consumption flow from changes in debt of households with no real estate wealth correlates with the time t since (or until) the onset of the unemployment spell. Because we cannot precisely separate mortgage debt from other types of credit in the data, restricting the sample to households with no real estate is a direct and simple way to identify how non-mortgage related debt evolves over the unemployment spell. We report estimates from regression model (18) and plot the estimated coefficients β_k for the set of indicators $\mathbb{1}[t \in qtr_k]$ for being observed, as of December of year n, in the k-th quarter since the onset of one's spell. Prior to becoming unemployed, increases in non-mortgage related debt contributes positively to yearly consumption by 8kSEK on average. After a year of unemployment ($k = 5$), unemployed individuals have decreased their consumption from changes in debt by 29.6%.

Figure 9: Calibrated 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 calibrated 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 ($CS_2 > CS_1$), while the moral hazard costs remain lower ($MH_2 < MH_1$), indicating that the introduction of an inclining tilt ($b_2 > b_1$) remains welfare improving for lower replacement rates.

Table 1: Summary statistics: RKD sample.

	Mean	P10	P50	P90
	I. Unemployment			
Duration of spell (wks)	26.78	3	16	73.43
Duration on b_1 (wks)	13.52	3	16	20
Duration on b_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 marrie	.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 sample of unemployed individuals used in the RKD 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 $\approx$ 0.11 USD2003.

Table 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 $\approx$ 0.11 USD2003.

Table 3: RKD ESTIMATES AT THE 725SEK THRESHOLD

	(1) Unemployment Duration D	(2) Duration D_1 (< 20 weeks)	(3) Duration D_2 (≥ 20 weeks)
	I. 1999-2000: Kink in b_1 and b_2		
δ_k	-.0569	-.0246	-.0299
	(.0050)	(.0012)	(.0039)
$\varepsilon_{D_k,\bar{b}}$	1.530	1.319	1.615
	(.1300)	(.0645)	(.1986)
N	187518	187518	187518
	II. 2001: Kink in b_2 only		
δ_k	-.0255	-.0115	-.0105
	(.0049)	(.0020)	(.0030)
ε_{D_k,b_2}	.6765	.6015	.5921
	(.1312)	(.1061)	(.1642)
N	65545	65545	65545
	III. 2002-. : Placebo		
δ_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)	
$MH_2 = \frac{\bar{b}+\tau}{\bar{b}} \cdot \frac{D}{D_2} \cdot \varepsilon_{D,b_2}$	1.445	(.2829)	
$MH_1 = \frac{\bar{b}+\tau}{\bar{b}} \cdot \frac{D}{D_1} \cdot (\varepsilon_{D,\bar{b}} - \varepsilon_{D,b_2})$	1.819	(.4032)	
Hypotheses testing	z-stat	p-value	
$MH_1 = MH_2$	-0.57	0.569	
$2 \times MH_1 = MH_2$	-2.11	0.035	

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 linear regressions of the form of equation (17) with a bandwidth size $h = 90$SEK. 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 $\varepsilon_{D_k,b_k} = \hat{\delta}_k \cdot \frac{b_k^{cap}}{D_b^{cap}}$, where $\hat{\delta}_k$ is the estimated marginal slope change,D_b^{cap} is the observed average duration at the kink and $b_k^{cap} = 580$ is the benefit level at the kink. In Panel IV we also report implied moral hazard costs estimates defined in equation (13). MH is the moral hazard cost of increasing benefits throughout the unemployment spell. MH_1 is the moral hazard cost of increasing benefits given for the first 20 weeks of the spell. MH_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 4: HOUSEHOLD LOG-CONSUMPTION AS A FUNCTION OF TIME SPENT UNEMPLOYED RELATIVE TO PRE-UNEMPLOYMENT CONSUMPTION: CONSUMPTION SURVEY ESTIMATES

	(1)	(2)	(3)	(4)
$\mathbb{1}[0 < t \leq 20$ weeks$]$	-0.0606*	-0.0444	-0.0408	-0.0465
	(0.0316)	(0.0311)	(0.0314)	(0.0413)
$\mathbb{1}[t > 20$ weeks$]$	-0.130***	-0.129***	-0.111***	-0.108***
	(0.0328)	(0.0332)	(0.0386)	(0.0414)
$\mathbb{1}[L > 20$ weeks$]$			-0.0294	-0.0342
			(0.0308)	(0.0378)
$\mathbb{1}[0 < t \leq 20$ months$] \times \mathbb{1}[L > 20$ weeks$]$				0.0134
				(0.0629)
Year F-E	×	×	×	×
Calendar months F-E	×	×	×	×
Marital status		×	×	×
Family size		×	×	×
Age group F-E		×	×	×
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 (19). 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 5: 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≤200k			-0.0503***	-0.0116***	-0.0122***
			(0.00234)	(0.00271)	(0.00315)
200k<Net wealth≤500k			-0.0466***	-0.0146***	-0.0114***
			(0.00324)	(0.00350)	(0.00425)
500k<Net wealth≤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.

Table 6: 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	×	×	×	×	×	×	×	×
Marital status	×	×	×	×	×	×	×	×
Family size	×	×	×	×	×	×	×	×
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 (19). 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 7: Welfare evaluation of actual profile using estimated sufficient statistics

	(1) Moral hazard hazard costs	(2) Average consumption drop	(3) Consumption smoothing gains CS_k	(4)	(5)	(6) Cost-Benefit ratio
	MH_k	Δc_k	$\gamma = 1$	$\gamma = 2$	$\gamma = 5$	CS_k/MH_k
Benefits given throughout the spell: $\bar{b}$	1.64	.09	.09	.18	.46	$\gamma \times .056$
	(.14)	(.03)	(.03)	(.05)	(.13)	
Benefits given for the first 20 weeks: b_1	1.82	.06	.06	.12	.31	$\gamma \times .034$
	(.40)	(.03)	(.03)	(.06)	(.16)	
Benefits given after first 20 weeks: b_2	1.45	.13	.13	.26	.65	$\gamma \times .090$
	(.28)	(.03)	(.03)	(.06)	(.17)	

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_1 during the first 20 weeks of the unemployment spell. The third row analyzes the welfare consequences of an increase in benefits b_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 3). The second column repeats our estimates of the average consumption drop (column 1 of Table 4). To estimate the average consumption drop over the entire spell, we run the same regression as in (19) 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_k, following the implementation in (15), 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_k/MH_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.

Appendix A: Technical Appendix

This appendix provides the technical details of our model setup and the proofs of the Propositions.

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

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

to solve

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

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_i^u, c_i^e, 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

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

which simplifies to (1) when $r=0$. For an n-part policy, we still use the short-hand notation D_k^r to refer to the discounted version of time spent on the k-th part of the policy,

$$D_k^r = \Sigma_{t=B_{k-1}}^{B_k}\left[1+r\right]^{-(t-1)}\int\int \left[1-\theta_{i,t}(\omega_{i,t})\right] dF_{i,t}(\omega_{i,t})\, di = \Sigma_{t=B_{k-1}}^{B_k}\left[1+r\right]^{t-1} S_t.$$

We also write $T^r = \Sigma_1^T\left[1+r\right]^{-(t-1)}$ and $D^r = \Sigma_{k=1}^n D_k^r$, so that

$$G(P) = [T^r - D^r]\tau - D_1^r b_1 - D_2^r b_2 - .. - D_n^r b_n.$$

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. We assume that this program is strictly concave in P.[41]

[41] Chetty (2006) provides regularity conditions such that the government's problem is strictly concave in case of

2 Local Policy Changes

Proof of Proposition 1

Starting from a flat profile ($b_k = b$)*, consider an increase in the benefit level at time* t *and time* $t+1$ *during the unemployment spell. Using notation* $S_t^r = (1+r)^{-(t-1)} S_t$*, we find*

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

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*

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

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. Since the environment is assumed to be stationary and the agent is borrowing constrained, 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. So we have* $\varepsilon_{\tilde{D}_2^r,b_{t+1}} = \varepsilon_{D^r,b_t}$ *for* $T \to \infty$*. Since the exit rate* $h_t(s_t) = h$ *is constant over time in a stationary environment, 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*

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

flat unemployment policies (i.e., $b_k = \bar{b}$).

Using the expression for $\varepsilon_{D^r,b_{t+1}}$ and the equality $D^r_2/S^r_{t+1} = D^r/S^r_t$, we can re-write

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

This implies that

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

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.■

Proof of Proposition 2

For any history ω_t, an agent's intertemporal consumption allocation satisfies the Euler condition. When unemployed at time t, 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

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

for $\beta(1+r) = 1$. This condition holds with equality if the agent's borrowing constraint is not binding at time t. With separable preferences, $\partial v^u(c,s)/\partial c = \partial v^e(c)/\partial c = v'(c)$, we have that $v'\left(c^e_{t+1}\right) < v'\left(c^u_{t+1}\right)$ as long as UI benefits are lower than the after-tax wage, since the expected life-time wealth is higher when employed than when unemployed at time $t+1$. If the borrowing constraint is not binding at time t, the Euler condition then implies $v'(c^u_t) \leq v'\left(c^u_{t+1}\right)$. Hence, 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. If the borrowing constraint is binding, the agent is hand to mouth and thus $c^u_t = b_t$. The marginal utility of consumption and thus the consumption smoothing gain of UI benefits is thus inversely related to benefit profile.■

3 Optimal Unemployment Policy

Proof of Proposition 3

The welfare impact of a change in benefit level b_k of an n-part policy P equals

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

where

$$\frac{\partial G(P)}{\partial b_k} = -D_k^r - \Sigma_{l=1}^n (b_l + \tau) \frac{\partial D_l^r}{\partial b_k} = -D_k^r \times \left[1 + \Sigma_{l=1}^n \frac{D_l^r (b_l+\tau)}{D_k^r b_k} \varepsilon_{D_l^r, b_k}\right],$$

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

$$\begin{aligned}\int \frac{\partial V_i(P)}{\partial b_k} di &= \int \{\Sigma_{t=B_{k-1}}^{B_k} \beta^{t-1} \int \frac{\partial v_i^u \left(c_{i,t}^u(\omega_{i,t}), s_{i,t}(\omega_{i,t})\right)}{\partial c_{i,t}^u} \left[1-\theta_{i,t}(\omega_{i,t})\right] dF_{i,t}(\omega_{i,t})\} di \\ &= \Sigma_{t=B_{k-1}}^{B_k} \beta^{t-1} S_t E\left(\frac{\partial v_i^u \left(c_{i,t}^u(\omega_{i,t}), s_{i,t}(\omega_{i,t})\right)}{\partial c_{i,t}^u} | t, \theta_{i,t}(\omega_{i,t}) = 0\right) \\ &= D_k^{1/\beta-1} E\left(\frac{\beta^{t-1}}{\Sigma_{t=B_{k-1}}^{B_k}\beta^{t-1}} \frac{\partial v_i^u \left(c_{i,t}^u(\omega_{i,t}), s_{i,t}(\omega_{i,t})\right)}{\partial c_{i,t}^u} | t \in [B_{k-1}, B_k], \theta_{i,t}(\omega_{i,t}) = 0\right),\end{aligned}$$

where $D_k^{1/\beta-1} = D_k^r$ for $\beta(1+r)=1$. This expression simplifies to (6) for $\beta = 1+r = 1$. The expectation operator averages over all potential states in which the agent is unemployed and receives unemployment benefit b_k. Note that since in our stylized model an agent's optimized unemployment consumption only varies with the length of the ongoing unemployment spell the weight of agent i's marginal utility at time t in calculating the average marginal utility simplifies to $S_{i,t}/D_k$.

Combining the two expressions, we find

$$\frac{\partial W(P)}{\partial b_k} = 0 \Leftrightarrow \frac{\int \frac{\partial V_i(P)}{\partial b_k} di - \lambda}{\lambda} = \Sigma_{l=1}^n \frac{D_l^r (b_l+\tau)}{D_k^r b_k} \varepsilon_{D_l^r, b_k}.$$

In the same way, we find

$$\begin{aligned}\frac{\partial G(P)}{\partial \tau} &= (T^r - D^r) \times \left[1 + \Sigma_{l=1}^n \frac{D_l^r (b_l+\tau)}{(T^r - D^r)\tau} \varepsilon_{D_l^r, \tau}\right], \\ \int \frac{\partial V_i(P)}{\partial \tau} di &= -\left(T^{1/\beta-1} - D^{1/\beta-1}\right) E\left(\frac{\beta^{t-1}}{\Sigma_{t=1}^T \beta^{t-1}} \frac{\partial v_i^e \left(c_{i,t}^e(\omega_{i,t})\right)}{\partial c_{i,t}^e} | \theta_{i,t}(\omega_{i,t}) = 1\right).\end{aligned}$$

and thus

$$\frac{\partial W(P)}{\partial \tau} = 0 \Leftrightarrow \frac{\lambda - \int \frac{\partial V_i(P)}{\partial \tau} di}{\lambda} = \Sigma_{l=1}^n \frac{D_l^r (b_l+\tau)}{(T^r - D^r)\tau} \varepsilon_{D_l^r, \tau},$$

which simplifies to expression (8) for $\beta = 1+r = 1$. The expectation operator used here is over all employment states. Compared to consumption during unemployment, employment consumption on the optimal path still depends on the unemployment history and not just on time t.

Assuming that $W(P)$ is strictly concave in the policy $P = (b_1, .., b_n, \tau)$, the $n+1$ first-order *conditions stated in the Proposition, jointly with the budget constraint, are necessary and sufficient for characterizing the optimal policy.*■

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

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

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

$$\begin{aligned} Pr\left(I_{i,t+1}^1 = 1\right) &= \int l_i(e_{i,t}(\omega_{i,t}))\theta_{i,t}(\omega_{i,t}) dF_{i,t}(\omega_{i,t}), \\ Pr\left(I_{i,t+1}^{\tilde{t}} = 1\right) &= \int (1 - h_i(s_{i,t}(\omega_{i,t}), \omega_{i,t})) I_{i,t}^{\tilde{t}-1}(\omega_{i,t}) dF_{i,t}(\omega_{i,t}). \end{aligned}$$

The budget constraint still depends on the expected time spent unemployed on each part k of the policy, D_k^r, potentially spread over multiple spells. That is,

$$D_k^r = \Sigma_{t=1}^T \Sigma_{\tilde{t}=B_{k-1}}^{B_k} [1+r]^{-(t-1)} \int\int I_{i,t}^{\tilde{t}}(\omega_{i,t}) dF_{i,t}(\omega_{i,t}) di.$$

Hence, the optimal formulae in Proposition 3 remain exactly the same, except for the potential change in the marginal utility of consumption when employed. 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.[42]

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 $z_{i,t}(x_{i,t}, \omega_{i,t})$ into the

[42] 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 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 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.

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.

Finally, 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.[43] 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

$$G(P) - \bar{G} = [T - D](\tau^u + \tau^y w) - D_1 (b_1 - \tau^y b_1) - ... - D_n (b_n - \tau^y b_n) - \bar{G},$$

where

$$\begin{aligned}
\frac{\partial G(P)}{\partial b_k} &= -D_k^r (1 - \tau^y) - \Sigma_{l=1}^n (b_l - \tau^y b_l + \tau^u + \tau^y w) \frac{\partial D_l^r}{\partial b_k} \\
&= -D_k^r (1 - \tau^y) - \Sigma_{l=1}^n (b_l + \tau^u) \frac{\partial D_l^r}{\partial b_k} - \Sigma_{l=1}^n \tau^y (w - b_l) \frac{\partial D_l^r}{\partial b_k} \\
&= -D_k^r \left[(1 - \tau^y) + \Sigma_{l=1}^n \frac{D_l}{D_k} \frac{b_l + \tau^u}{b_k} \varepsilon_{D_l^r, b_k} + \Sigma_{l=1}^n \frac{D_l}{D_k} \frac{\tau^y (w - b_l)}{b_k} \varepsilon_{D_l^r, b_k} \right].
\end{aligned}$$

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.

4 Implementation

Proof of Corollary 1

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

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

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

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

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

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

$$\frac{db_2}{db_1}|_2 = -\frac{D_1 MH_1}{D_2 MH_2}$$

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.$\blacksquare$

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.

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.

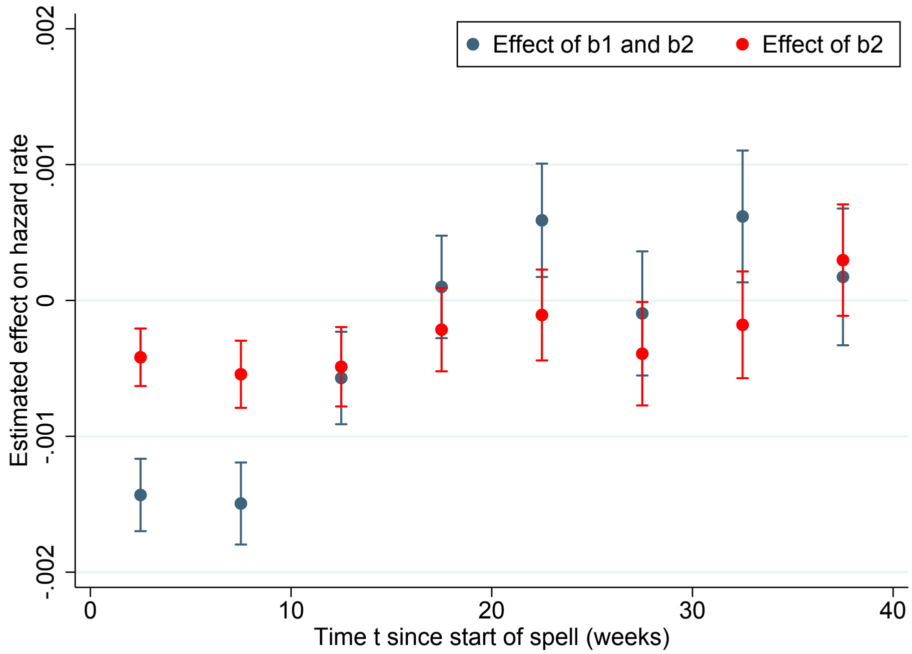
Second, the graph shows that b_2 (benefits received after 20 weeks) do have an effect on the hazard rate in the first 10 weeks. This confirms that unemployed individuals are forward-looking. b_2 does have a somewhat negative effect on contemporaneous hazard rates (after 20 weeks), but this effect is 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 of b_1 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

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_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. All estimates are from linear specifications using the changes in the UI schedule at the 725SEK kink with a 90SEK bandwidth. 95% confidence interval around the point estimates, from robust standard errors, are displayed.

weeks are too small and insignificant to undo, in the MH costs estimates, the effects on hazard rates early in the spell.

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 (17) are displayed in Table 3 and Figure 4. The red lines display predicted values of the regressions in the linear case.

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.

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:

$$Y = y(b_1, b_2, w, \mu),$$

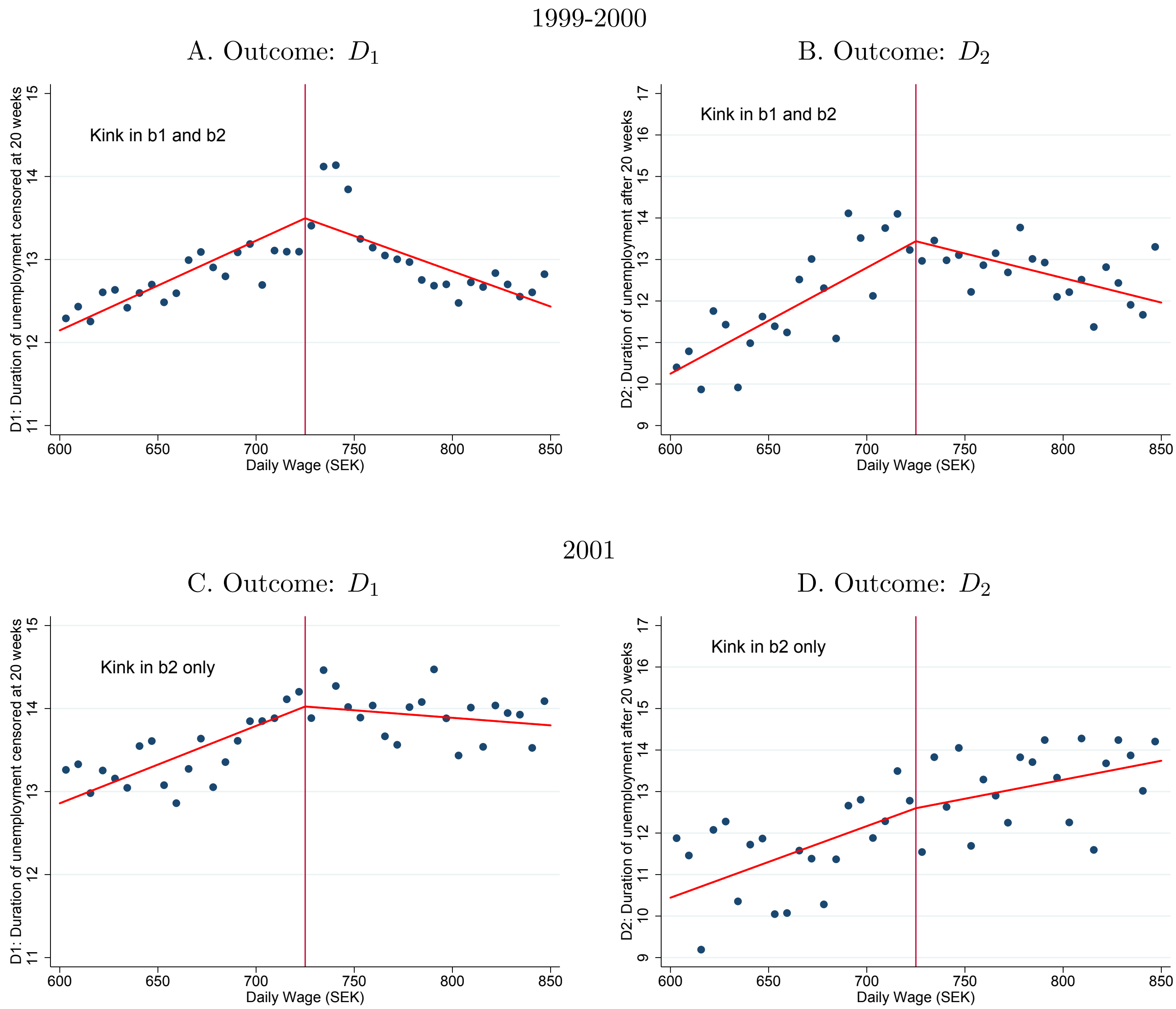
We are interested in identifying the marginal effect of benefits $b_k, k = 1, 2$ on the duration outcome Y, $\alpha_k = \frac{\partial Y}{\partial b_k}$. b_k is a deterministic, continuous function of the wage w, kinked at $w = \bar{w}_k$. Identification of α_k in the RK design relies on two assumptions.

Assumption 1: the direct marginal effect of the assignment variable w on Y is assumed to be smooth around the kink point $\bar{w}_k$. This means that $\frac{\partial y(b_1,b_2,w,\mu)}{\partial w}$ 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.

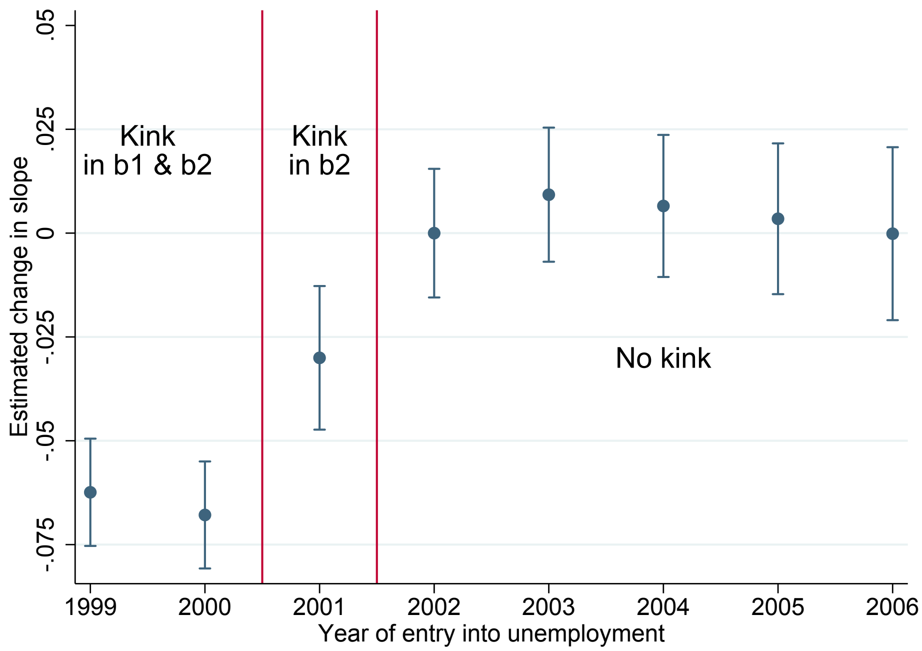
These identifying assumptions are, by definition, untestable. Yet, we can use the various "experiment arms" of our quasi-experimental setting as well as sensitivity analysis to try to detect

Figure B.2: RKD 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 (17) are displayed in Table 3 and Figure 4. 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_1 and b_2. Spells starting in 2001 are subject to a kink in b_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 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.

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.

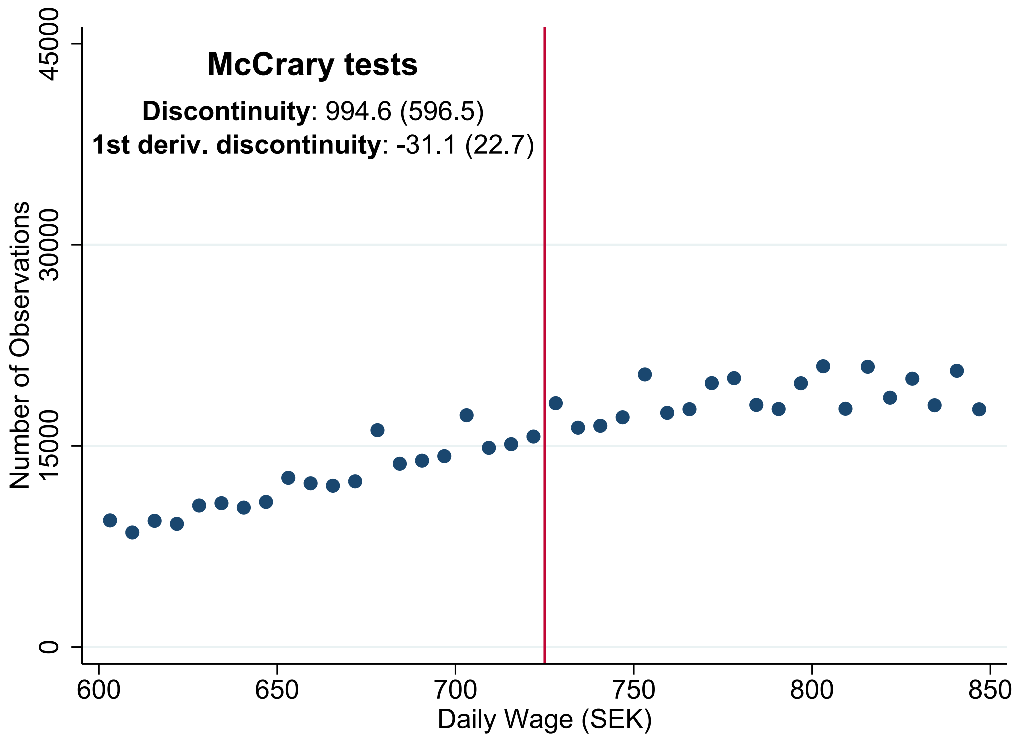
Testing for clear violations of Assumption 2: manipulation The most obvious violation of the assumption of smooth distribution of heterogeneity at the kink arises if individuals are able to locate their daily wage strategically around the kink point. A few tests can help assess the robustness of this 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 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_1 and b_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

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.

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.

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 $h = 60$SEK.

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.7 shows the R-squared when we run the RKD regression in (17) 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.

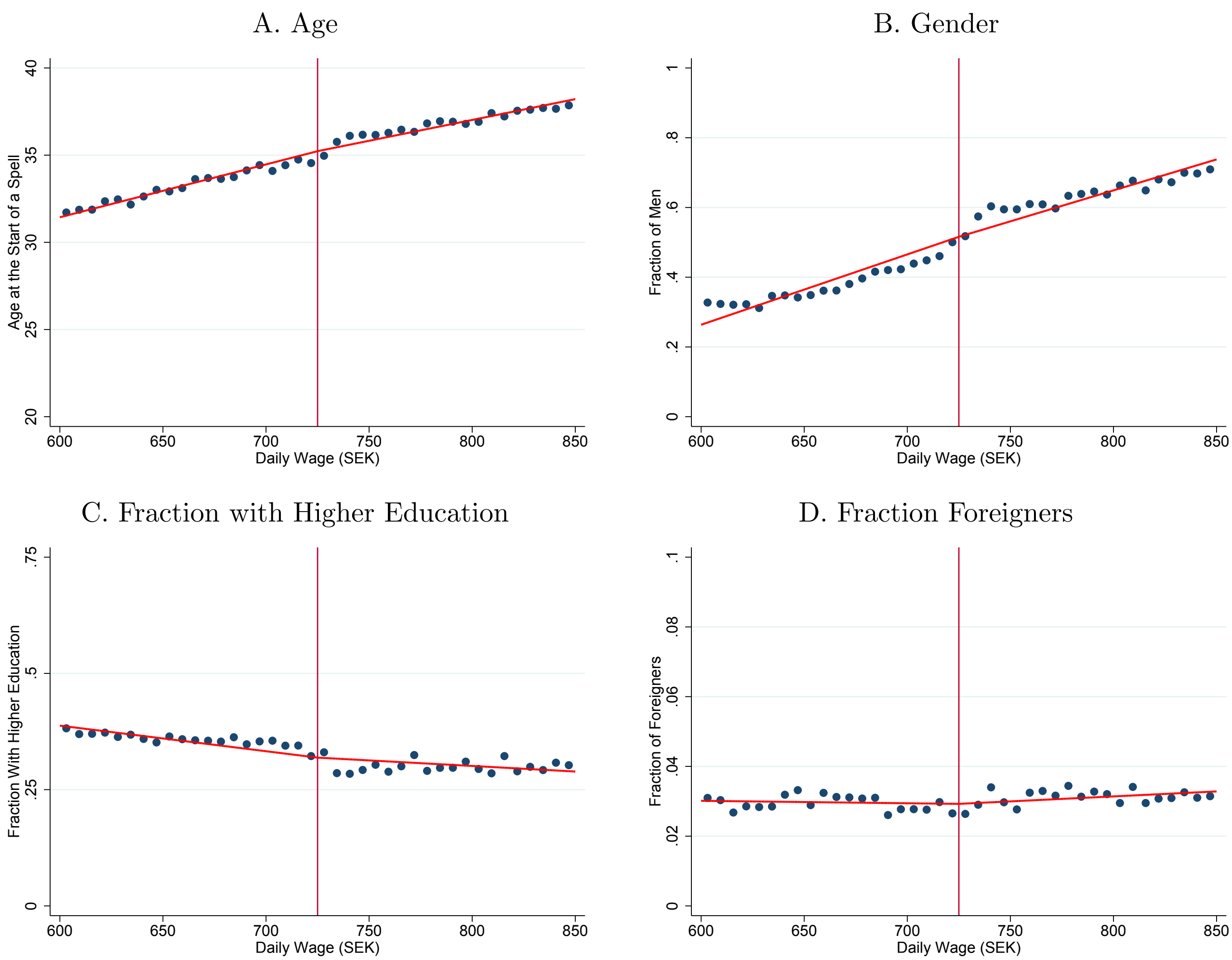
Polynomial order Table B.3 shows estimates of the change in slope at the kink for linear, quadratic and cubic specifications, assessing the model fit for these different specifications.

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 D1 and D2 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.

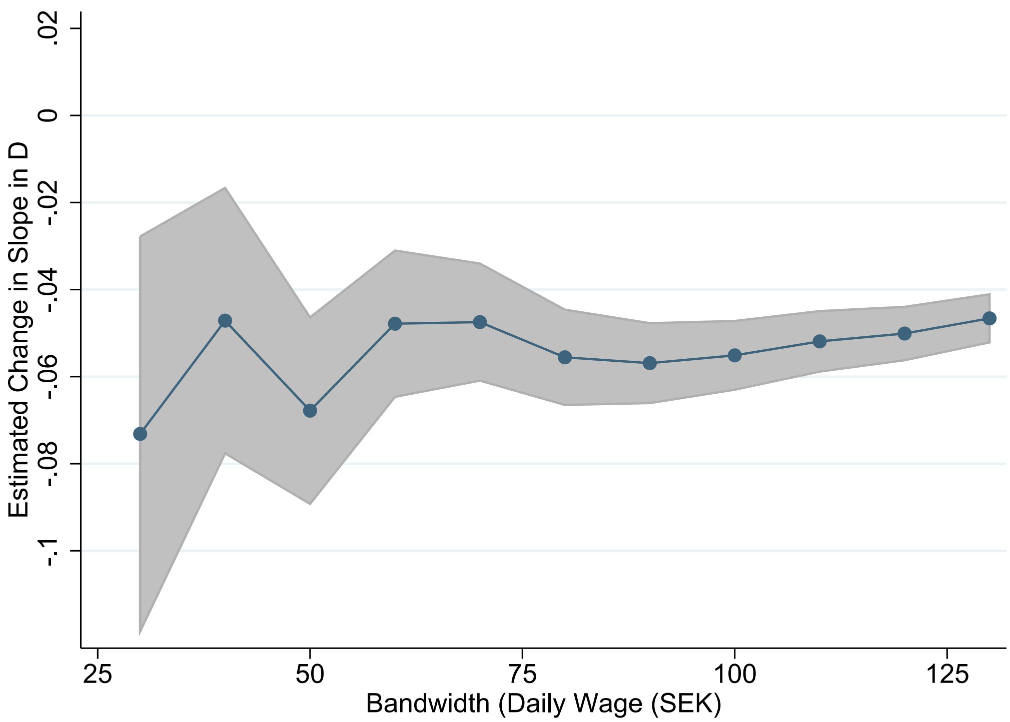
Figure B.5: Robustness of the RK design: Covariates



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 (17) 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**

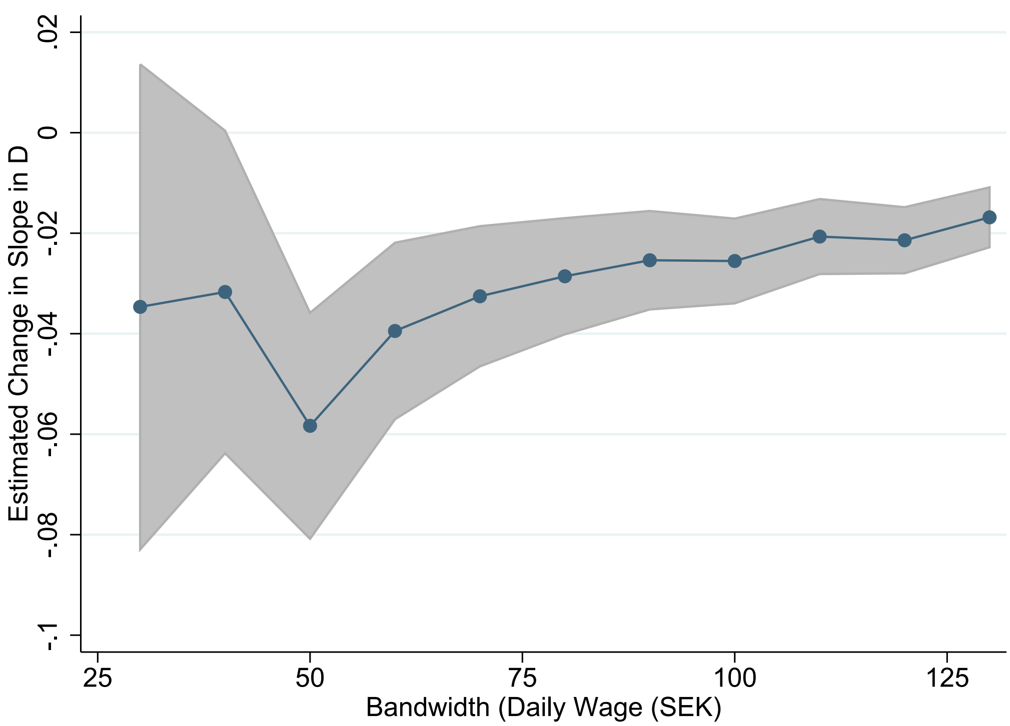
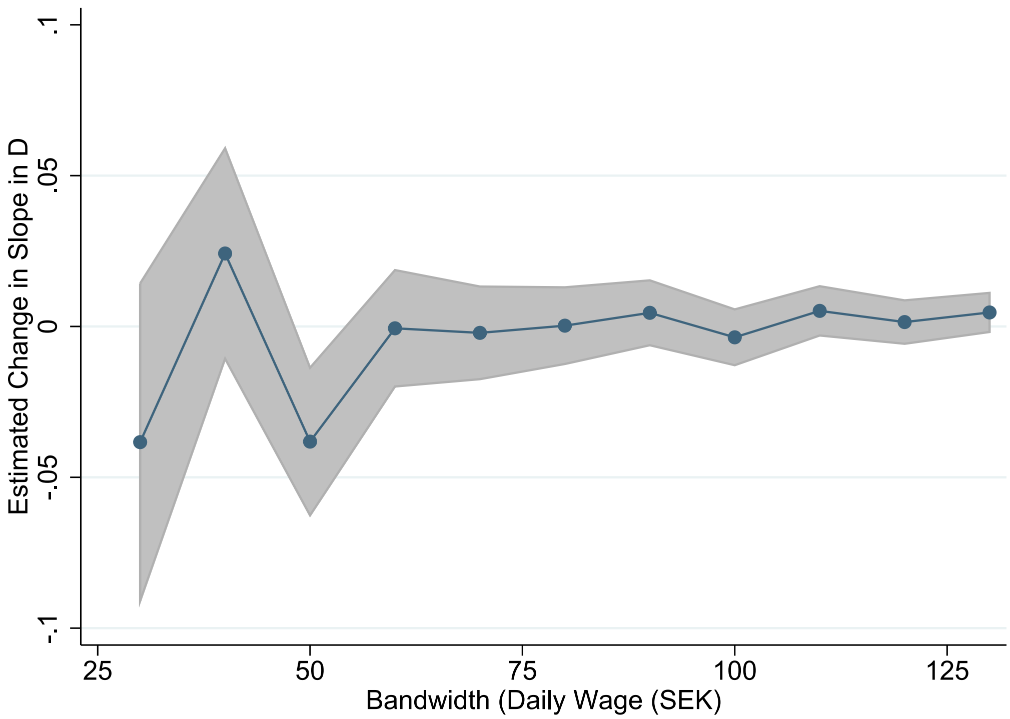


Figure B.6: RKD estimates as a function of bandwidth size (*continued)*

C. **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 (17) 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.

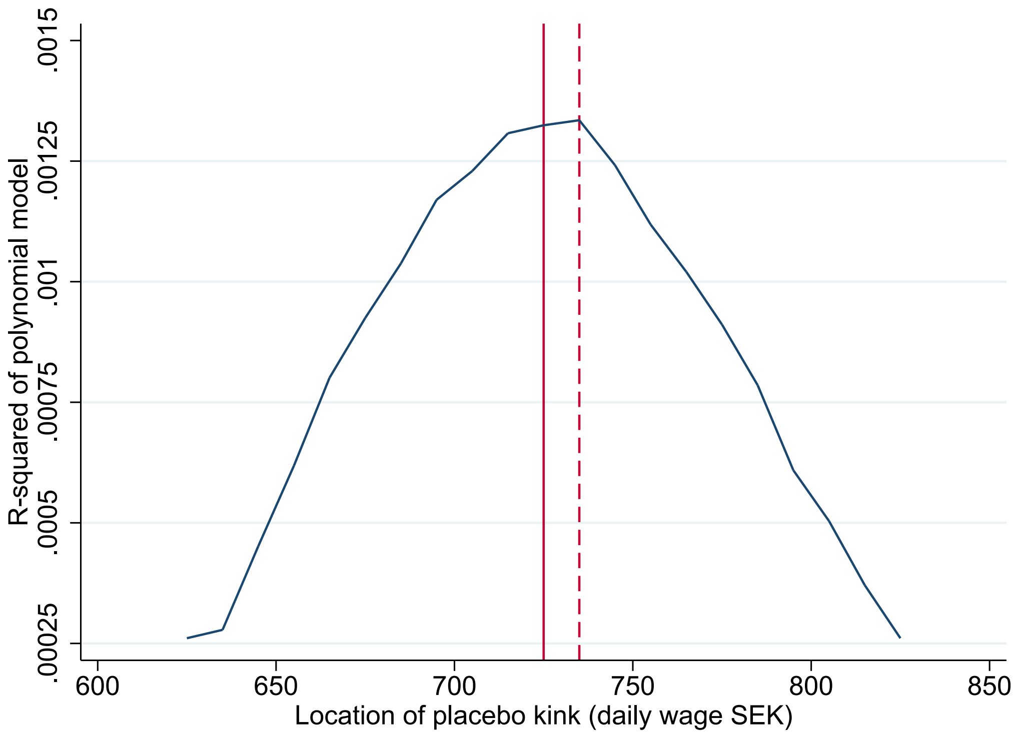
Table B.2: RKD ESTIMATES AT THE 725SEK THRESHOLD: INFERENCE

	(1) Unemployment Duration D	(2) Duration D_1 (< 20 weeks)	(3) Duration D_2 (≥ 20 weeks)
	I. 1999-2000: Kink in b_1 and b_2		
Linear - δ_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 - δ_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 -δ_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 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)). $D_1 = \sum_{t<20wks} S_t$ 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 (17) 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 δ_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].

Figure B.7: Non-parametric detection of kink location

A. 1999 - 2000



B. 2001

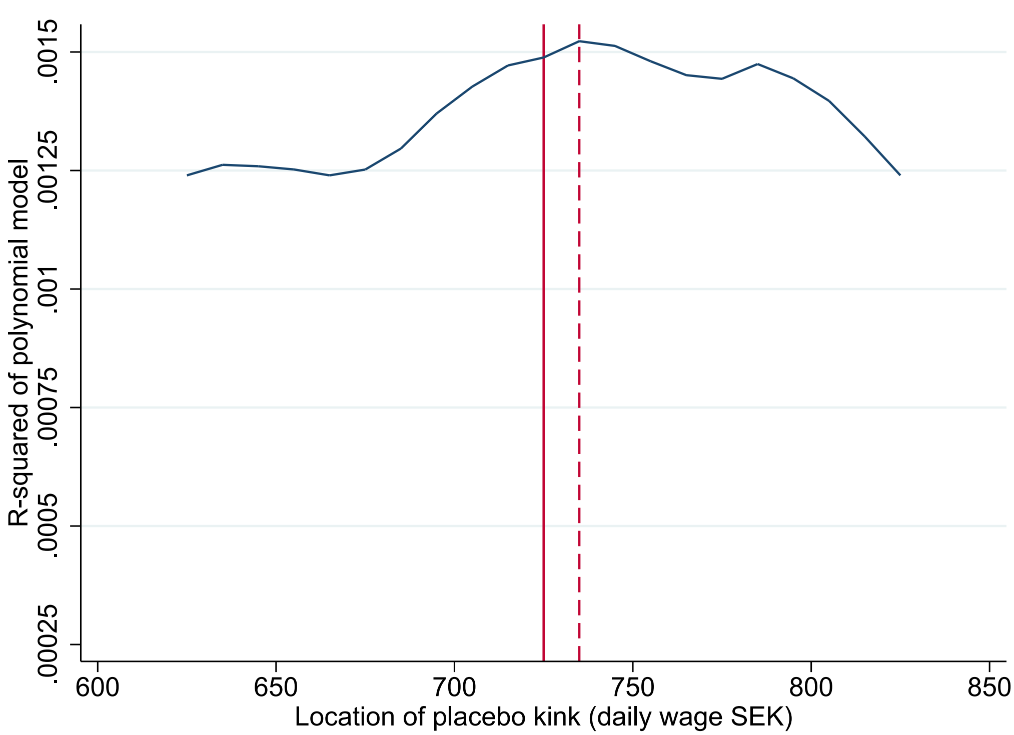
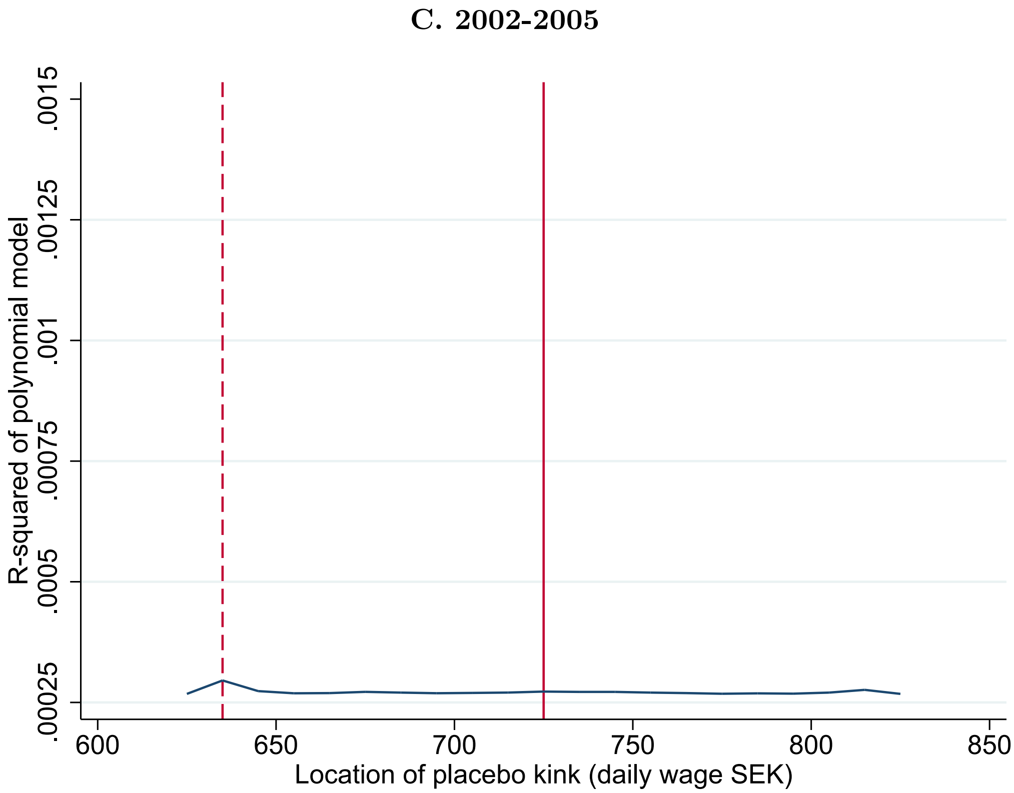


Figure B.7: Non-parametric detection of kink location (*continued)*



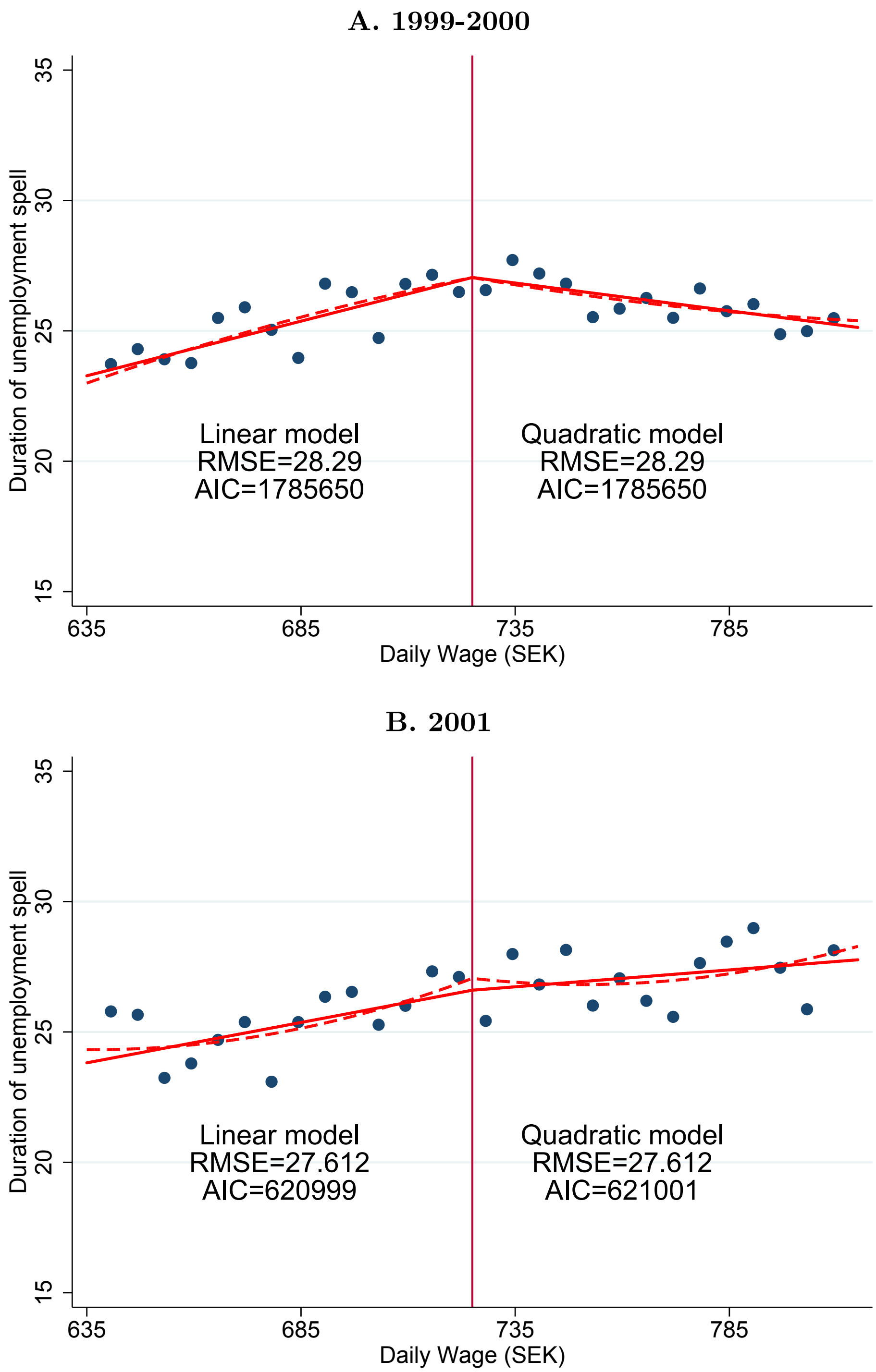
Notes: The figure reports the R-squared of polynomial regressions of the form of equation (17) for alternative (placebo) locations of the kink point $\bar{w}_k$ for all observations with wages between 625SEK and 825SEK. The red line indicates the location of the true kink in the schedule. The dashed 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.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 (≥ 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

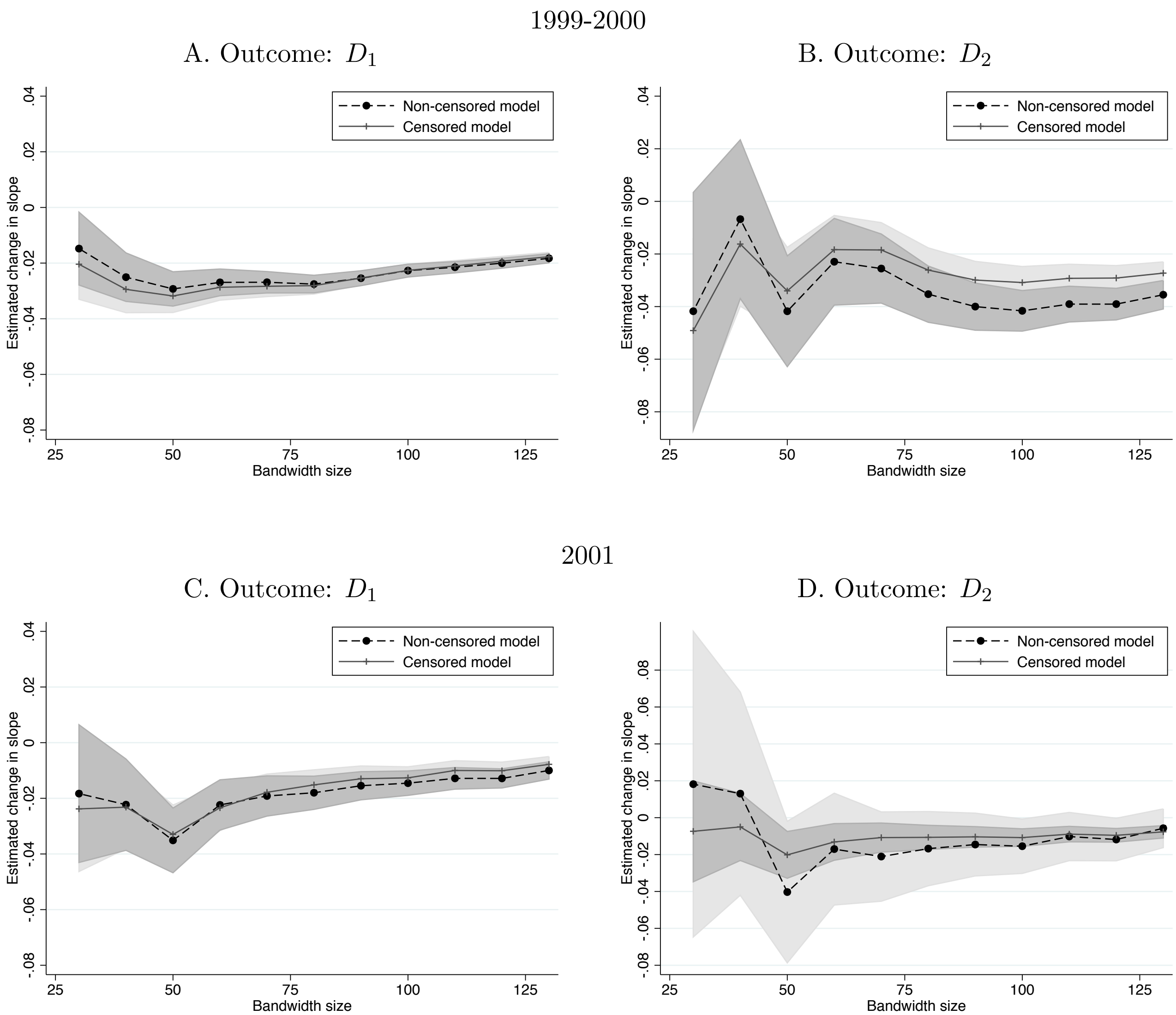
Notes: The table reports estimates of the change in slope δ_k, at the 725SEK threshold, in the relationship between daily wage and the total duration of unemployment D (col. (1)), the time D_1 spent on the first part of the Swedish UI profile (col. (2)) and the time D_2 spent on the second part of the Swedish UI profile (col. (3)). Estimates are obtained from polynomial regressions of the form of equation (17) 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.

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



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 (17) 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 (17) 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).

Appendix C: Additional results and robustness of consumption analysis

This appendix presents additional results on consumption patterns over the unemployment spell and dynamic selection using a residual measure of consumption expenditures based on registry data.

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. While it offers the advantage of being computable for the universe of unemployed households from 1999 to 2007, the measure can only capture yearly expenditures between December of each year. All details on the procedure to construct this measure are given in Kolsrud et al. [2015].

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 is closely related to 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. 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.
- T_n accounts for all income taxes and transfers, including unemployment insurance, disability insurance, sick pay, housing and parental benefits, etc.
- $\tilde{C}_n^b = y_n^b - \Delta b_n$ 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.
- $\tilde{C}_n^h = y_n^h - \Delta h_n$ 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.

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

Yearly consumption profiles over the spell using registry-based measures of consumption

To analyze the effect of unemployment duration on consumption using the registry-based measures of consumption, we restrict the sample to all individuals unemployed as of December of year n, or who will become unemployed in year $n+1$, for all years n from 1999 up to 2007. We then correlate our yearly household consumption measure, which records total yearly expenditures between December of year $n-1$ and December of year n, with time t spent unemployed since the onset of the unemployment spell. Negative values for t represent the time remaining until the start of the unemployment spell.[44] In practice, we run the following regression:

$$C_{i,t} = \sum_{k=-3}^{+8} \tilde{\beta}_k \cdot \mathbb{1}\left[t \in qtr_k\right] + X_i'\gamma + \varepsilon_{i,t} \tag{C.1}$$

where $\mathbb{1}\left[t \in qtr_k\right]$ is an indicator for being observed in the k-th quarter of the ongoing unemployment spell. We include in this regression a set of controls X, which consists of year dummies, age

[44]To ensure that yearly consumption observed in quarters $k=-3$ to $k=0$ only aggregates consumption flows while employed, we impose the further restriction that individuals must not have experienced any unemployment spell in the past two years prior to becoming unemployed.

dummies, a dummy for being just out of school in year n, and a set of dummies for family status.

We report in Figure C.1 the estimated coefficients $\tilde{\beta}_k$ from regression model (C.1). These coefficients represent the average yearly consumption levels in constant SEK for individuals observed in their k-th quarter of unemployment, relative to the yearly average consumption level of individuals observed just one quarter prior to becoming unemployed.

The graph confirms the evidence from the HUT data that unemployment is correlated with a substantial drop in consumption. The average yearly consumption of households where an individual has been unemployed for a full year ($k = 5$ quarters) is almost 50kSEK lower (15.8%) than the average consumption of households at the start of the unemployment spell. Interestingly, the magnitude of the drop in household consumption expenditures after a year of unemployment is very similar to the drop in consumption measured in the HUT data.

Exploring selection on consumption profiles using the registry-based measure of consumption

We can use our registry-based measure of consumption to assess the robustness of the results on selection patterns from the HUT. The advantages of the registry-based measures of consumption are that it is available for the universe of unemployed households from 1999 to 2007 (which gives a greater statistical power compared to the HUT), and that it has a panel structure, as opposed to the pseudo-panel structure of the HUT merged with UI records. However, the measure can only capture yearly expenditures between December of each year. We therefore focus on individuals who become unemployed in January and investigate household consumption after one year of unemployment and after two years of unemployment, compared to pre-unemployment consumption.

We estimate the following model, which is similar to that estimated on the HUT data:

$$c_{i,t} = \alpha_i + \eta_1 \cdot \mathbb{1}[t = 1 \text{ year}] + \eta_2 \cdot \mathbb{1}[t = 2 \text{ years}] + X_i'\gamma + \varepsilon_{i,t}, \tag{C.2}$$

where c_{it} is log yearly consumption of the household and η_1 (resp. η_2) captures the effect of having a member of the household unemployed for a year (resp. for 2 years) at the time of the interview, relative to household pre-unemployment consumption. α_i is a household fixed-effect.

Results are reported in Table C.1 and convey two important takeaways.

First, we find that pooled estimates (column (2)) and panel model estimates (within-estimator in column (3), and first-difference estimator in column (4)), controlling for household fixed effects in consumption levels, yield very similar results for the consumption profiles over the unemployment spell. This indicates limited selection on consumption levels and confirms the evidence from the HUT sample that the consumption drops over the spell are not driven by dynamic selection on consumption levels.

Second, in columns (5) and (6), we analyze the role of selection on consumption profiles. We estimate differences in consumption profiles between households who select into spells longer than a year versus households who select into spells shorter than a year. To do so, we interact $\mathbb{1}[L =$

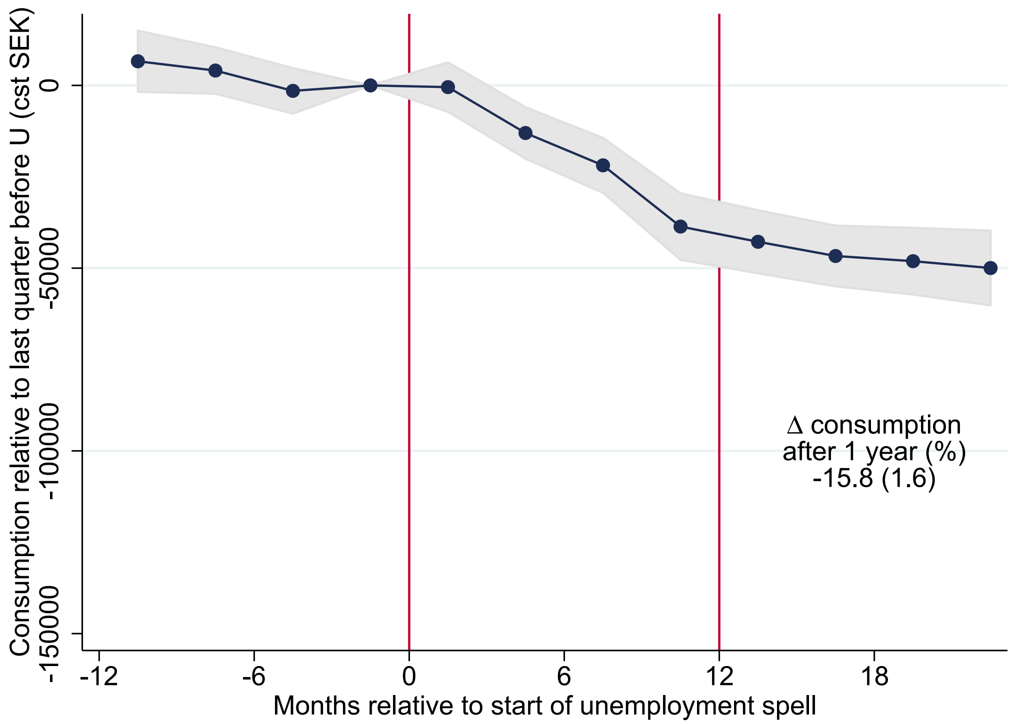
2 years], a dummy variable for the total length of the unemployment spell being longer than a year, with the event time dummies to estimate two separate consumption profiles over the unemployment spell: one for households with long spells and one for households with short spells. Results reported in columns (5) and (6) indicate that the consumption drop in the first year for households who select into long spells is 1% smaller but not statistically different from the consumption drop in the first year of households who select into short spells. This evidence is in line with the evidence from the HUT sample and suggests that differences in consumption profiles are relatively small and insignificant, so that selection on consumption profiles are not significantly driving the observed patterns of average household consumption over the unemployment spell.

Selection on different categories of expenditures in the HUT data

Tables C.2 and C.3 replicate the analysis of, respectively, column 3 and 4 of Table 4 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 5.3 . 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 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.[45]

[45] 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.

Figure C.1: Yearly household consumption as a function of unemployment duration: registry-based consumption estimates



Notes: The figure correlates yearly household consumption with the time t since (or until) the onset of the unemployment spell. To compute household consumption, we compute an individual residual measure of consumption for all members aged above 18 of the household in which the unemployed individual lives, and then sum these measures at the household level. The figure follows from regression model (C.1) and plots the estimated coefficients $\tilde{\beta}_k$ for the set of indicators $\mathbb{1}\left[t \in qtr_k\right]$ for being observed, as of December of year n, in the k-th quarter since the onset of one's spell. Regression includes year dummies, age dummies, a dummy for being just out of school in year n, and a set of dummies for family status. We also plot the 95% confidence interval from robust standard errors.

Table C.1: Household consumption as a function of time spent unemployed relative to pre-unemployment consumption: registry-based consumption estimates

	(1) Pooled	(2) Pooled	(3) Within	(4) First-diff	(5) Within	(6) First-diff
$\mathbb{1}[t = 1 \text{ year}]$	-0.118***	-0.137***	-0.140***	-0.123***	-0.152***	-0.134***
	(0.00862)	(0.00933)	(0.0181)	(0.0116)	(0.0464)	(0.0378)
$\mathbb{1}[t = 2 \text{ year}]$	-0.161***	-0.167***	-0.198***	-0.176***	-0.198***	-0.176**
	(0.0121)	(0.0126)	(0.0322)	(0.0222)	(0.0322)	0.0223)
$\mathbb{1}[t = 1 \text{ year}] \times \mathbb{1}[L = 2 \text{ years}]$					0.0130	0.0119
					(0.0448)	(0.0376)
Year F-E	×	×	×	×	×	×
Calendar months F-E	×	×	×	×	×	×
Spell length F-E		×	×	×	×	×
Marital status		×	×	×	×	×
Family size		×	×	×	×	×
Age group F-E		×	×	×	×	×
R^2	0.00724	0.261	0.0917	0.00391	0.0917	0.00391
N	51369	34458	34466	29372	34466	29372

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 using a registry-based measure of consumption. All details on the construction of this measure are given in Kolsrud et al. [2015]. The consumption drops are estimated following the model of equation (C.2). $\mathbb{1}[t = 1 \text{ year}]$ is an indicator for having a member of the household unemployed for a year at the moment yearly consumption is measured. $\mathbb{1}[t = 2 \text{ year}]$ is an indicator for having a member of the household unemployed for two years at the moment yearly consumption is measured. Columns (1) and (2) present pooled estimates, i.e. without household fixed effects. Columns (3) and (4) present panel model estimates, controlling for household fixed effects in consumption levels. Column (3) uses a within-estimator, while column (4) uses a first-difference estimator. Columns (5) and (6) estimate differences in consumption profiles between households who select into spells longer than a year versus households who select into spells shorter than a year, by interacting the event time dummies with $\mathbb{1}[L = 2 \text{ years}]$, i.e. a dummy variable for the total length of the unemployment spell being longer than a year.

Table C.2: 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 (19). 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.3: 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
$\mathbb{1}[0 < t \leq 20$ 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)
$\mathbb{1}[t > 20$ 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)
$\mathbb{1}[L > 20$ 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)
$\mathbb{1}[L > 20$ weeks] $\times$ $\mathbb{1}[0 < t \leq 20$ 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 (19). 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.

Appendix D: Calibration Details and Simulation Results

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

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

$$v_{i,t}^{u}\left(c_{i,t}, s_{i,t}\right) = c_{i,t}^{1-\gamma} / \left(1-\gamma\right) - \psi_0 \left(s_{i,t}\right)^{\psi_1},$$
$$v_{i,t}^{e}\left(c_{i,t}\right) = c_{i,t}^{1-\gamma} / \left(1-\gamma\right).$$

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

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

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.

Model Fit The first column in Table D.1 lists the targets of the calibration, which are the variables underlying our sufficient statistics. The targets include the benefit duration elasticities ε_{D,b_k}, average benefit durations D_k and the average consumption drops $[\bar{c}_k - \bar{c}_0]/\bar{c}_0$. 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. We minimize the relative difference between the targeted moments and the simulated values of these moments generated by our model, putting a higher weight on the ratios. As discussed in the main text, the calibrated model slightly overestimates the consumption drop in the first 20 weeks and underestimates the consumption drop thereafter and thus underestimates the relative consumption smoothing gains overall. 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.

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, $\lambda \equiv E_0(v'(c_{i,0}))$, where $c_{i,0}$ is the agent's level of consumption when she would have been employed from the start of the model. Figure 9 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 $CS_1 = MH_1$ and $CS_2 = MH_2$ jointly hold for the unique pair $(b_1, b_2) = (.48, .68)$.

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.

Finally, 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. To analyze the impact of preference heterogeneity, 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. Note that in our model with CRRA preferences and additive search cost (a standard specification also used in Hopenhayn and Nicolini [1997] and Chetty [2008]), 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 (16) in section 2.4, 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_1 and CS_2 are lower relative to the baseline scenario. The ratio CS_1/CS_2 decreases as well, which further increases the value of introducing an inclining tilt ($b_2 > b_1$). We note though that with heterogeneous preferences, the calculation of CS_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 $\left[1+CS_1\right]/\left[1+CS_2\right] = E_1\left(v_i'\left(c_{i,t}\right)\right)/E_2\left(v_i'\left(c_{i,t}\right)\right)$, relevant for evaluating a budget-balanced change in the tilt as in condition (10). This ratio changes from .934 in the baseline calibration to .935 in the simulation with heterogeneity (cf. Table D.2), which indicates that incorporating preference heterogeneity in our calibrated model is unlikely to overturn our policy recommendation on the tilt.

Table D.1: Model Calibration: Parameters and Targets

Model inputs

Parameters	Description	Calibrated value
h_0	baseline job finding rate at zero effort	0.007
h_1	baseline return to search	1.063
a	parameter of $\beta(a, b)$ for indiv. return to search	1.862
b	parameter of $\beta(a, b)$ for indiv. return to search	2.309
θ	depreciation rate	0.109
ψ_0	search costs (scalar)	1.920
ψ_1	search costs (curvature)	1.821
$\bar{a}$	asset limit	0.208
α_{SM}	parameter asset distribution	1.170
c_{SM}	parameter asset distribution	1.389
k_{SM}	parameter asset distribution	1.500

Model moments: 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
Δc_1	consumption drop ST unemployed	0.060	0.072
$\Delta c_1/\Delta c_2$	ratio of consumption drops	0.462	0.658
D_1	ST benefit duration	12.600	6.675
D_1/D_2	ratio of benefit duration	0.940	0.991
$\varepsilon_{D_{b1}}$	elasticity wrt ST benefits	0.854	0.549
$\varepsilon_{D_{b1}}/\varepsilon_{D_{b2}}$	ratio of elasticities	1.262	1.271

Notes: The top panel lists the parameters we calibrate in our structural model, while the bottom panel lists the variables we target, which are the moments underlying our sufficient statistics. 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 parameter values are obtained by minimizing the relative distance between the targeted moments and the simulated values of these moments generated by our model. The exception are the parameters of the asset distribution which 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). For the calibration, we also set the parameter of CRRA $\gamma = 2$, discount factor $\beta = 1$ and interest rate $r = 0$.

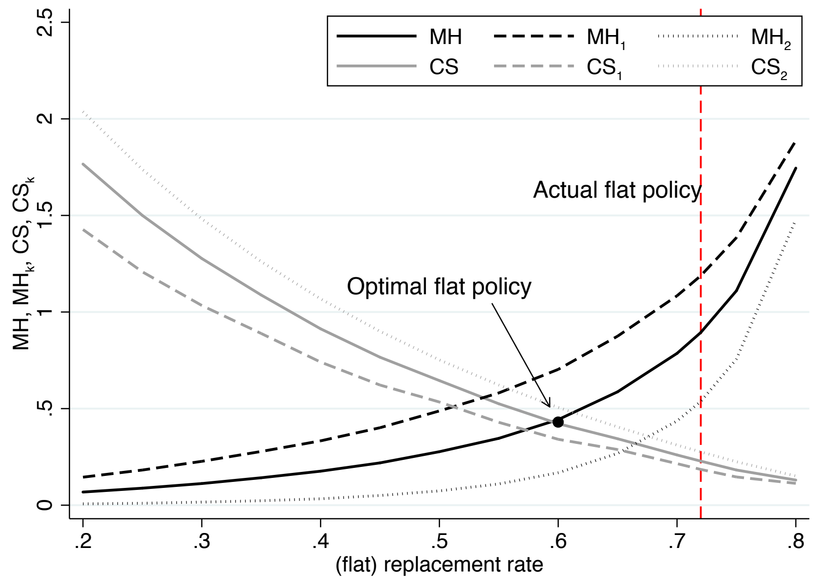
Table D.2: Implementation of consumption smoothing gains

	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

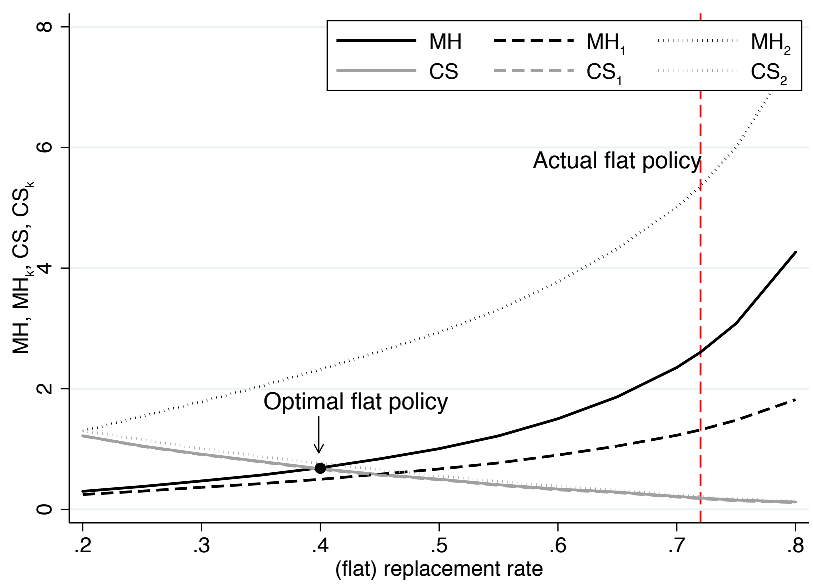
Notes: This Table illustrates alternative implementations of the consumption smoothing gains for different underlying risk preferences. The top panel reports the simulated consumption drops and consumption smoothing gains for the baseline model with homogeneous CRRA γ=2. The third and fourth panel report 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_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_k following the implementation in (15), 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.

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

A. Only heterogeneity in returns to search



B. Only depreciation of returns to search



Notes: The figure analyzes the robustness of the findings in Figure 9 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.