REFERENCE-DEPENDENT JOB SEARCH: EVIDENCE FROM HUNGARY*

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December 2016

Abstract

We propose a model of job search with reference-dependent preferences, with loss aversion relative to recent income (the reference point). In this model, newly unemployed individuals search hard since consumption is below their reference point. Over time, though, they get used to lower income, and thus reduce their search effort. In anticipation of a benefit cut their search effort rises again, then declines once they get accustomed to the lower post-cut benefit level. The model fits the typical pattern of exit from unemployment, even with no unobserved heterogeneity. To distinguish between this and other models, we use a unique reform in the UI benefit path. In 2005, Hungary switched from a single-step UI system to a two-step system, with overall generosity unchanged. The system generated increased hazard rates in anticipation of, and especially following, benefit cuts in ways the standard model has a hard time explaining. We estimate a model with optimal consumption, endogenous search effort, and unobserved heterogeneity. The reference-dependent model fits the hazard rates substantially better than plausible versions of the standard model, including habit formation. Our estimates indicate a slow-adjusting reference point and substantial impatience, likely reflecting present-bias.

*We thank Patrick Arni, Saurabh Bhargava, Lajos Bódis, Moshe Buchinsky, David Card, Henry Farber, Sebastian Findeisen, Larry Katz, Gábor Kézdi, Patrick Kline, Matthew Rabin, Hedvig Horváth, Erzo Luttmer, Edward O'Donoghue, Daniele Paserman, Alex Rees-Jones, Emmanuel Saez, Mihály Szoboszlai, Owen Zidar and audiences at Boston University, Columbia, Cornell, Harvard, IZA, Mannheim (Department and ZEW), Tinbergen, UCLA, UCL, University of Regensburg, University of Zurich, the UCSB 2014 Behavioral Conference, the 2014 BEAM conference, the 2014 NBER Labor Studies Summer Institute, the Brookings conference on "Policy lessons from behavioral economics", the 2016 ASSA meetings, and the Hungarian Society of Economists Annual Conference for very helpful comments. We are grateful to János Köllő, Kitti Varadovics, Mónika Bálint, Dániel Biró for giving us access to the administrative data and providing continuous help throughout the project, and to Nagy Csaba and Pelek Zsolt for explaining the practical implementation of the reform. We also thank Tristan Gagnon-Bartsch, Matthew Gudgeon, Avner Shlain, Jessica Shui, Alex Steiny, and Ferenc Szucs for excellent research assistance and Gautam Rao for sharing his code. Financial support from the Center for Equitable Growth at UC Berkeley, from the Alfred P. Sloan Foundation, from the National Science Foundation and from the Firms, Strategy and Performance Lendület Grant of the Hungarian Academy of Sciences is gratefully acknowledged.

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I Introduction

Unemployment insurance programs in most Western countries follow a common design. The benefits are set at a constant replacement rate for a fixed period, typically followed by lower benefits under unemployment assistance. In such systems, the hazard rate from unemployment typically declines from an initial peak the longer workers are unemployed, surges at unemployment exhaustion, and declines thereafter.[1]

It is well-known that a basic job search model a la Mortensen (1986) and van den Berg (1990) is unable to match this pattern. This model predicts an increasing exit hazard up until benefit expiration, with a constant exit rate thereafter. To match the time path, job search models add unobserved heterogeneity among workers. More productive workers are more likely to find a job initially, leading to a decrease in the hazard over time as the workers still unemployed are predominantly of the less productive type.

In this paper, we propose, and test, a behavioral model of job search which can account for this time path of unemployment even in the absence of unobserved heterogeneity. Namely, we incorporate one of the best established facts in psychology, that people's perceptions and decisions are influenced by relative comparisons. We assume that workers have reference-dependent preferences over their utility from consumption. As in prospect theory (Kahneman and Tversky, 1979), workers are loss-averse with respect to consumption below a reference point. Further, we assume that this reference point is given by recent earnings.

To fix ideas, consider a reference-dependent worker who was just laid off and assume, for now, hand-to-mouth consumption. At the time of job loss, the reference point of the unemployed individual is the previous wage, which is significantly higher than the unemployment benefit, the new consumption level. The unemployed worker, therefore, finds the new state of unemployment particularly painful given the loss relative to the reference point, and so she searches hard at the beginning of a UI spell. Over the weeks of unemployment, however, the reference point shifts as the individual adapts to the lower benefit level, and the loss is thus mitigated. Hence, the worker's search effort decreases. As the end of UI benefits draws near, the worker, if still unemployed, anticipates the loss in consumption due to the exhaustion of the benefits, and searches harder. This force is at work also in the standard model, but it is heightened by the anticipation of the future loss aversion. If the worker does not find a job by the UI expiration, the worker once again slowly adjusts to the new, lower benefit level.

The hazard from unemployment for this reference-dependent worker decreases from the initial peak, increases at exhaustion, then decreases again. Hence, the predicted hazard matches

[1]This has been shown in a variety of settings, such as in the United States (Katz and Meyer, 1990), Hungary (Micklewright and Nagy, 1999), Austria (Card et al., 2007a), Slovenia (van Ours and Vodopivec, 2008), Germany (Schmieder et al., 2012a), and France (Le Barbanchon, 2016).

the patterns documented in the literature, even in absence of unobserved heterogeneity.

How would one distinguish the standard job search model from a reference-dependent model? Consider two UI systems, the first offering a constant benefit path until period T, with the second offering higher initial benefits up to period $T_1 < T$ but lower benefits between T_1 and T (Figure Ia). The two systems have the same welfare benefit level after period T. The standard model with no heterogeneity predicts that, starting from period T, the hazard rate in the two systems would be the same, as the future payoffs are identical (Figure Ib). Furthermore, the hazard rate in the periods right before period T will be higher in the system with two-step benefits given the lower benefits at that point.

The reference-dependent model makes three different predictions (Figure Ic). First, right after period T the hazard in the one-step system would be higher because of the higher loss in consumption compared to the recent benefits. Second, this difference would attenuate over time and ultimately disappear as the reference point adjusts to the lower benefit level. Third, the hazard rate in the first UI system increases already *in advance of* period T, in anticipation of the future loss aversion. Notice that, while these predictions are developed in the absence of heterogeneity to highlight the intuition, we fully integrate heterogeneity in our estimates.

We evaluate a change in the Hungarian unemployment insurance system which is ideally suited for a test of the above predictions. Before November 2005, the Hungarian system featured a constant replacement rate for 270 days, followed by lower unemployment assistance benefits. After November 2005, the system changed to a two-step unemployment system: benefits are higher in the first 90 days, but lower between days 90 and 270, compared to the pre-period (Figure IIa)). There was no major change in the unemployment assistance system taking place after 270 days. As such, this UI set-up corresponds to the hypothetical case outlined above with period T corresponding to 270 days.

An important feature of the Hungarian reform is that the total benefits paid out until day 270 remain about the same after the reform.[2] Hence, differences in savings and in selection in the pre- and post- period are likely to be small, allowing for a more straightforward comparison.

We evaluate the reform by comparing the hazard rates into employment in the year before and after the reform. The evidence is well in line with the predictions of the reference-dependent model. In the weeks immediately preceding the 270-day exhaustion of benefits, the pre-reform hazard rates rise above the post-reform hazard rates. In the months following the exhaustion, the pre-reform hazard rates remain higher, and they ultimately converge to the post-reform level only after a couple months. The observed pattern around the exhaustion is consistent with the anticipation of, and then the direct effect of the higher loss in consumption for individuals in the pre-reform. The ultimate convergence between the two hazards indicates,

[2]Notice that this does not imply that the reform is revenue neutral, as individuals unemployed for fewer than 270 days receive higher benefits after the reform.

in this interpretation, the timing of the reference point adjustment.

We present several robustness checks. Controlling for a broad set of controls and alternative definitions of our sample barely affects the estimated hazards. Also, an interrupted time series analysis shows that the break in the hazards occurs immediately in the quarter of introduction of the reform, and does not appear to reflect previous trends.

While the evidence is qualitatively consistent with predictions of the reference-dependent model, it is important to compare the quantitative fit of the behavioral model with the fit of the standard model allowing for unobserved heterogeneity. To do so, we estimate a model with both an optimal search effort decision and an optimal consumption-saving decision. The standard model allows for unobserved heterogeneity in the form of three types with different search costs. The reference-dependent model has two extra behavioral parameters (loss-aversion and updating horizon for the reference point) but assumes no unobserved heterogeneity and thus has two fewer parameters overall. We estimate the model with a minimum-distance estimator, with the empirical hazard rates in the pre- and post-reform as moments.

The preferred estimate for the standard model does a relatively good job of fitting the hazards in the first 200 days. More specifically, the presence of heterogeneous types allows the model to qualitatively match the spike in the hazard at 90 days post-reform. The standard model, however, is unable to capture the observed behavior leading up to, and following, the exhaustion of benefits. In particular, the hazard rates from day 270 on in the pre- and post-period are predicted to be almost identical, contrary to the empirical findings.

The reference-dependent model captures the spike at 90 days and the subsequent decrease in the exit hazard, similar to the standard model (and with a closer fit). This behavioral model also captures key features of the data which the standard model does not fit: the increase in hazard in the month prior to the expiration of benefits in the pre-period, the spike at 270 days, the decrease thereafter, and the ultimate convergence of the hazard between the pre- and post-period after a few months. The fit of the model is by no means perfect: the model underpredicts the spike at 270 days and the difference in hazards in the following two months. Still, it captures most of the qualitative features which the standard model does not fit at all. Importantly, it does so without assuming any unobserved heterogeneity.

What parameters characterize the best-fitting reference-dependent model? The estimated loss aversion is in the range of the previous literature and the estimated reference point horizon extends back about six months. A key role is played by the time discounting parameters, which indicate high impatience. The estimated discount factor is arguably implausibly small at 0.9 for a 15-day period, leading to an annual discount factor of 0.08. (The estimated discounting for the standard model is similar). Allowing for present-biased time preferences (Laibson, 1997; O'Donoghue and Rabin, 1999), with an additional discount factor β between

the current period and the future, implies much more reasonable discounting. We estimate a present-bias parameter β=0.58, well within the range of estimates in the literature, for an implied annual discount factor of 0.52 for the first year and 0.88 for later years.

Thus, the results point to an important interaction between reference dependence and impatience. The reference-dependent model does not provide a good fit for the data if workers are patient: these workers would build precautionary savings to smooth the upcoming loss utility due to a benefit decrease, eliminating the elevated hazards at benefit exhaustion. If workers are instead impatient, as estimated, consumers essentially go hand-to-mouth, and respond to the loss utility associated with the benefit declines by increasing search effort.

We highlight two key components of the reference-dependent model: loss aversion and a backward-looking, adaptive reference point. We show that estimates with reference points fixed to the pre-unemployment wage, or with forward-looking expectations a la Koszegi-Rabin (2006) do not fit the data. Is our reference-dependence model distinct from habit formation? Models a la Constantinides (1990) and Campbell and Cochrane (1999), like the reference-dependent model, induce a temporarily higher marginal utility of income following a benefit cut as consumption gets closer to the habit. We highlight a key difference. In the reference-dependent model, the impact of the loss on search effort is approximately proportional to the size of the loss. Instead, in the habit-formation model larger decreases in consumption have disproportionate effects. Given this, the habit-formation model underpredicts the spike in hazard at 270 days, since the benefit step-down at 90 days is proportionally larger.

Could alternative versions of the standard model fit the data as well as the reference-dependent model? We allow for, among other assumptions, time-varying search costs and a delay between search effort and the job start. We also estimate the model using probability of exit instead of hazard and excluding the spikes from the moments. None of these changes sizeably affect the fit of the standard model, or of the reference-dependent model.

We then examine alternative forms of unobserved heterogeneity. While so far we assumed heterogeneity in search cost in the spirit of Paserman (2008), we allow for more cost types, for heterogeneity in re-employment wage, or in the search elasticity. The first two forms of heterogeneity do not close the gap with the reference-dependent model, but the model with heterogeneous search elasticity fits better, even outperforming the reference-dependent model. This model explains the spikes by allowing for a type with such high search elasticity (over 50) that she only searches once benefits fall below a threshold. Is this model plausible? This model fits the data well only for very high elasticities, making the unlikely prediction that, if welfare benefits were increased by just 10 percent, individuals on welfare would stop searching.

To provide further evidence on this model with heterogeneous elasticities, as well as on the standard and reference-dependent model, we make two out-of-sample predictions. We consider

an earlier UI reform which lengthens the duration of unemployment assistance, and a sample of workers with lower pre-unemployment earnings, and thus a different benefit structure. In both cases, the reference-dependent model provides the best out-of-sample fit, fitting the qualitative patterns well. The standard model with cost heterogeneity does not do as well fit-wise, and the model with heterogeneous elasticity does worst.

As a final piece of evidence, we also compare the dynamic selection implied by the models to the selection on observables in the data. The selection implied by the standard model differs both qualitatively and quantitatively from the observed selection; instead, the implied selection is close to the observed selection for the reference-dependent model.

Finally, we briefly discuss other job search models which we do not estimate but which could potentially explain some of the findings. A model of storable offers (as in Boone and van Ours, 2012) could explain the spike in hazard at benefit exhaustion, but not the pattern of the hazards in the following months. A model of workers who are temporarily laid off and later recalled to the same employer (as in Katz, 1986 and Katz and Meyer, 1990) can explain a declining hazard rate early in the spell and a spike at the exhaustion point. However, recalls are quite rare in our context and the hazard rates are very similar when we exclude workers who are recalled. A model of skill depreciation or screening (e.g. Schmieder et al., 2016) can explain decreasing hazards over the spell, but such decreases would plausibly be the same pre- and post-reform. Two relevant models are worker learning and consumption commitments (Chetty, 2003 and Chetty and Szeidl, 2016). A worker may learn over time that finding jobs is harder than expected, and this learning may take place later in the pre-reform period, given the different benefit structure. A worker with committed consumption would increase search effort to avoid paying a fixed cost of adjustment; if despite this, she does not find a job soon enough, she will pay the cost and then decrease search. These dynamics could generate some of the hazard patterns after day 270. While both models have intuitive features, neither is tractable enough to estimate on our sample. For tractability reasons, we also do not estimate models with reservation wage choice and optimal consumption-savings.[3]

To sum up, reference dependence, in combination with impatience, can help explain patterns in job search that are hard to rationalize with most alternative models, even allowing for unobserved heterogeneity. These results have implications for potential redesigns of unemployment insurance policies, a point to which we return briefly in the conclusions.

The paper relates to the literature on unemployment insurance design (e.g. Chetty, 2008; Kroft and Notowidigdo, forthcoming; Schmieder et al., 2012a). Within this literature, we

[3]The consumption commitment model requires one to keep track of a fixed cost decision, making the model cumbersome to estimate. To address this issue, Chetty (2003) makes the timing of fixed cost payment exogenous. A consumption commitment model with exogenous consumption readjustment would not explain our findings. We present in the appendix estimates with reservation wage choice for the hand-to-mouth case. The results should be considered only suggestive, as endogenizing consumption is very important.

evaluate a unique reform: changing the time path of the benefit schedule, keeping the overall payments approximately constant. The paper is consistent with recent evidence of sharp consumption drops at unemployment entry and UI exhaustion for unemployed workers (Ganong and Noel, 2015; Kolsrud et al., 2015), suggesting approximate hand-to-mouth behavior.

The paper also contributes to a literature on behavioral job search (DellaVigna and Paserman, 2005 and Spinnewijn, 2013)[4] It also adds to the field evidence on reference dependence. (Sydnor, 2010; Barseghyan et al., 2013; Fehr and Goette, 2007; Farber, 2015; Card and Dahl, 2011; Barberis and Huang, 2001; Allen et al., forthcoming; Simonsohn and Loewenstein, 2006; Engström et al., 2015; Rees-Jones, 2013). We provide evidence on the speed of updating of a backward-looking reference point as in Bowman et al. (1999) and Post et al. (2008). This paper is also part of a literature on structural behavioral economics (Laibson et al., 2007; Conlin et al., 2007; DellaVigna et al., 2012).

The paper proceeds as follows. In Section II, we present the model. In Section III we present the institutional details for the unemployment insurance reform, which we evaluate in Section IV. In Section V we present the structural estimates, and we conclude in Section VI.

II Model

In this section we present a discrete-time model of job search with reference-dependent preferences and present-biased preferences. We build on the job search intensity model presented in Card et al. (2007a) and in Lentz and Tranaes (2005) by adding a reference dependent utility function in consumption with a backward looking reference point.

Model Setup. As in Card et al. (2007a) we make two simplifying assumptions. First, jobs last indefinitely once found. Second, wages are fixed, eliminating reservation-wage choices. In each period, a job seeker decides search effort $s_t \in [0,1]$, representing the probability of receiving a job offer at the end of period t and thus of being employed in period $t+1$. Search costs are given by the function $c(s_t)$, which we assume to be time-separable, twice continuously differentiable, increasing, and convex, with $c(0) = 0$ and $c'(0) = 0$.

In each period individuals receive income y_t, either UI benefits b_t or wage w_t, and consume c_t. Consumers can accumulate (or run down) assets A_t with a borrowing constraint $A_t \geq -L$. Assets earn a return R so that consumers face a budget constraint $\frac{A_{t+1}}{1+R} = A_t + y_t - c_t$. We also consider a simplified model with hand-to-mouth consumption, $c_t = y_t$.

The utility from consumption in period t is $v(c_t)$, where $v(.)$ is an increasing and concave

[4]Koenig et al. (2016) model a reference-dependent reservation wage choice in that the wage offers with some probability equal a previous wage (the reference). Their paper focuses on job matches and reservation wages, as opposed to the dynamics of exit from unemployment.

function. Following the functional form of Kőszegi and Rabin (2006), flow utility is

$$u\left(c_t|r_t\right) = \begin{array}{ll} v\left(c_t\right) + \eta\left[v\left(c_t\right) - v\left(r_t\right)\right] & \text{if} \quad c_t \geq r_t \\ v\left(c_t\right) + \eta\lambda\left[v\left(c_t\right) - v\left(r_t\right)\right] & \text{if} \quad c_t < r_t \end{array} \tag{1}$$

where r_t denotes the reference point in period t. The utility consists of consumption utility $v\left(c_t\right)$ and gain-loss utility $v\left(c_t\right) - v\left(r_t\right)$. When consumption is above the reference point ($c_t \geq r_t$), the individual derives gain utility $v\left(c_t\right) - v\left(r_t\right) > 0$, which receives weight η. When consumption is below the reference point ($c_t < r_t$), the individual derives loss utility $v\left(c_t\right) - v\left(r_t\right) < 0$, with weight $\lambda\eta$. The parameter $\lambda \geq 1$ captures loss aversion: the marginal utility is higher for losses than for gains. This utility function builds on prospect theory of Kahneman and Tversky (1979) without, for simplicity, modeling diminishing sensitivity or probability weighting. The standard model is a special case with $\eta = 0$.

The second key assumption regards the reference point r_t. Unlike in the literature on forward-looking reference points (Kőszegi and Rabin, 2006), but in the spirit of papers on backward-looking reference points (Bowman et al., 1999) and on habit formation (Constantinides, 1990), the reference point is the average income over the $N \geq 1$ previous periods:

$$r_t = \frac{1}{N}\sum_{k=t-N}^{t-1} y_k.$$

Note that an alternative backward-looking reference point could be past consumption. This alternative reference point, however, is much less tractable: when setting current consumption, an individual has to take into account the impact on future preferences through the effect on future reference points. In our formulation, instead, the reference point path is exogenous. In addition, psychologically, recent paychecks are a plausibly salient comparison point.

To gain perspective on the impact of reference dependence, consider an individual in steady state with consumption, income, and reference point equal to y. Then in period T, consider a small, permanent decrease in income from y to $y - \Delta y < y$, and an identical decrease in consumption from $c = y$ to $y - \Delta y$.[5] In period T, utility changes to $v\left(y - \Delta y\right) + \eta\lambda\left[v\left(y - \Delta y\right) - v\left(y\right)\right]$. The short-term change in utility equals, up to a linear approximation, $-\left(1 + \eta\lambda\right)\Delta y * v'\left(y\right)$. Over time, however, the reference point adjusts to $y - \Delta y$ so that the utility after N periods is $v\left(y - \Delta y\right)$. Hence, the long-term change in utility equals $-\Delta y v'\left(y\right)$ and $\eta\lambda$ captures the weight on additional short-term utility in response to an income loss.

Value Functions. The unemployed choose search effort s_t and consumption c_t in each

[5] A sudden permanent drop in consumption could occur, for example, if the individual were a hand-to-mouth consumer and benefits suddenly dropped.

period and face the following value function, where δ is the discount factor:

$$\begin{aligned} V_t^U(A_t) &= \max_{s_t \in [0,1]; A_{t+1}} u\left(c_t | r_t\right) - c\left(s_t\right) + \delta\left[s_t V_{t+1|t+1}^E\left(A_{t+1}\right) + \left(1 - s_t\right) V_{t+1}^U\left(A_{t+1}\right)\right] \quad (2) \\ &\text{subject to: } c_t = A_t + y_t - \frac{A_{t+1}}{1+R}. \end{aligned}$$

For the unemployed, the value function depends only on assets A_t, since the reference point is fully determined by t and thus is not an explicit state variable: $V_t^U(A_t)$.

For the employed, the value function can be written as $V_{t|j}^E(A_t)$ for an individual who is employed in period t and who found a job in period j, where the combination of t and j determines the reference point:

$$V_{t|j}^E\left(A_t\right) = \max_{c_t > 0} u\left(c_t | r_t\right) + \delta V_{t+1|j}^E\left(A_{t+1}\right). \quad (3)$$

Given Equation (2) the first order condition for the optimal level of search effort s_t^* in the case of an interior solution can be written as:

$$c'\left(s_t^*\left(A_{t+1}\right)\right) = \delta\left[V_{t+1|t+1}^E\left(A_{t+1}\right) - V_{t+1}^U\left(A_{t+1}\right)\right]. \quad (4)$$

Thus we can rewrite the value of unemployment as:

$$\begin{aligned} V_t^U\left(A_t\right) &= \max_{A_{t+1}} u\left(A_t + y_t - \frac{A_{t+1}}{1+R} \middle| r_t\right) - c\left(s_t^*\left(A_{t+1}\right)\right) \\ &\quad + \delta\left[s_t^*\left(A_{t+1}\right) V_{t+1|t+1}^E\left(A_{t+1}\right) + \left(1 - s_t^*\left(A_{t+1}\right)\right) V_{t+1}^U\left(A_{t+1}\right)\right] \end{aligned}$$

We solve the model by backwards induction, deriving first the steady-state consumption in the employed state. This allows us to solve for the consumption path for each asset level and to obtain the value functions $V_{t|t}^E\left(A_t\right)$ for each t and each asset level A_t. Then we solve the problem for the unemployed, moving backwards from the steady state, solving for the optimal consumption path and search effort path for each starting value of the asset vector.

Front-Loading The Benefit Path. To highlight the implications of reference dependence, we consider a hypothetical unemployment insurance reform that closely corresponds to our empirical setting. To build intuition and for tractability, we consider in detail the case of hand-to-mouth consumers with no heterogeneity and then briefly discuss the extension to the general case. In the case of hand-to-mouth consumers, assets are not a control variable and thus $s_t^* = \mathcal{C}\left(\delta\left[V_{t+1|t+1}^E - V_{t+1}^U\right]\right)$,where $\mathcal{C}(.) = c'^{-1}(.)$.

Consider a UI system with benefits b_1 for the first T_1 periods benefits and benefits b_2 from period T_1+1 until T. After period T, there is a lower second tier (such as social assistance) with

benefits $\underline{b}$. A single-step UI system, like the one in the US, is captured by $b_1 = b_2 = b_{constant}$ and is illustrated by the blue solid line in Figure Ia).

Consider a reform that front-loads the benefit path by raising benefits b_1 in the first T_1 periods and reducing benefits b_2 in periods T_1 to T, while leaving second-tier benefits $\underline{b}$ unchanged, as illustrated by the red dashed line in Figure Ia). Furthermore, the reform leaves untouched the total amount of benefits paid to an individual unemployed for T periods:

$$b_1 T_1 + b_2(T - T_1) = b_{constant} T. \tag{5}$$

Equation (5) implies $\frac{\partial b_2}{\partial b_1} = -\frac{T_1}{T-T_1}$. We now partially characterize how search s_t^* is affected by an increase in b_1 subject to constraint (5). We express the results in terms of $\frac{ds_t^*}{db_1} = \frac{\partial s_t^*}{\partial b_1} - \frac{T_1}{T-T_1}\frac{\partial s_t^*}{\partial b_2}$, where the total derivative takes the implied adjustment of b_2 into account.[6]

Proposition 1. *Assume a hand-to-mouth unemployed job-seeker and consider a shift in the benefit path that front-loads the benefits b_1 keeping the overall benefits paid constant.*

a) In the standard model ($\eta = 0$), the search effort in all periods after benefit expiration at T is unaffected: $\frac{ds_{T+i}^*}{db_1} = 0$, for $i = 0, 1, ...$

b) In the reference-dependent model ($\eta > 0$ and $\lambda \geq 1$) search effort (weakly) decreases in the first N periods after T, and remains constant in later periods: $\frac{ds_{T+i}^*}{db_1} \leq 0$, for $i = 0, 1, ...N-1$ and $\frac{ds_{T+i}^*}{db_1} = 0$, for $i = N, N+1, ...$ Furthermore, if the adjustment speed N of the reference point is shorter than T, then the inequality is strict: $\frac{ds_{T+i}^*}{db_1} < 0$, for $i = 0, 1, ..., N-1$.

These predictions are illustrated in Figures Ib) and c). In the standard model (Figure Ib), optimal search effort increases under both regimes up until period T, and then plateaus after period T at a level that is unaffected by the front-loading of benefits (Proposition 1a). Generally, the hazard rate for the front-loaded regime (the dotted red line) will be higher than the one for constant benefits in the periods right before period T, given the moral hazard.

In contrast, under reference dependence (Figure Ic), search effort in period T is substantially higher under the constant-benefit regime (continuous blue line). Individuals in this regime experience a sharper drop in consumption and thus (for $N < T$) significant loss utility due to the higher reference point (Proposition 1b). The difference in hazards persists but shrinks for N periods, at which point the reference point has fully adapted to the lower benefits under either regime, and thus search effort converges. Another implication (not captured in the Proposition) is that in the last periods before period T, for sufficiently large loss aversion λ, the hazard is higher under the constant-benefit regime: the anticipation of larger future losses counters the moral hazard effect of more generous benefits.

[6]Note that search effort in period t is not affected by UI benefits in period t, since the individual will only start a job found in period t in period $t+1$. Thus search effort s_t corresponds to the exit hazard from unemployment in period $t+1$: $s_t = h_{t+1}$.

Proposition 1 does not hold with either heterogeneity or optimal consumption. With heterogeneous types, differences in the path of benefits up to period T may lead to a different composition of types surviving at period T, and thus differences in the hazard even in the standard model, violating Proposition 1a). With endogenous savings, individuals may have different savings at period T, thus creating differences in hazards. However, given that the total benefit payout is constant, differences in type composition and in savings are likely to be small. We address both heterogeneity and savings when we estimate the structural model.

Present Bias. We extend the model to allow for present-bias (Laibson, 1997; O'Donoghue and Rabin, 1999), with an additional discount factor $\beta \leq 1$ between the current period and the future, as in DellaVigna and Paserman (2005). We assume naivete': the workers (wrongly) assume that in the future they will make decisions based on regular discounting δ. This assumption simplifies the problem, since we can use the value functions of the exponential agent (given that the naive worker believes she will be exponential from next period). In addition, the evidence on present bias is largely consistent with naivete' (DellaVigna, 2009).

The naive present-biased individual solves the following value functions:

$$\begin{aligned} V_t^{U,n}(A_t) &= \max_{s_t \in [0,1]; A_{t+1}} u\left(c_t | r_t\right) - c\left(s_t\right) + \beta\delta\left[s_t V_{t+1|t+1}^E\left(A_{t+1}\right) + \left(1-s_t\right) V_{t+1}^U\left(A_{t+1}\right)\right] \quad (6) \\ &\text{subject to: } c_t = A_t + y_t - \frac{A_{t+1}}{1+R}, \end{aligned}$$

where the functions V_{t+1}^U and $V_{t+1|t+1}^E$ are given by equations (2) and (3) above for the exponential discounters. We thus first solve for all possible values of V_{t+1}^U and $V_{t+1|t+1}^E$ and then we solve for consumption and search paths given $V_{t+1}^{U,n}$.

III Data and Institutions

III.A Unemployment Insurance in Hungary

Hungary had a generous two-tier unemployment insurance system up to 2005. In the first tier, potential duration and benefit amount depended on past UI contribution.[7] The maximum potential duration, obtained after around 4 years of contribution, was 270 days,[8] while the benefit level was based on the earnings in the previous year. After the exhaustion of first-tier benefits, unemployment assistance (UA) benefits could be claimed in the second tier. The UA benefit amount was the same for everybody, with the potential duration depending on age.

[7]Every worker in the formal sector must pay a UI contribution. In 2005, employers contributed 3% to the UI fund, while employees 1%. There is no experience rating of UI benefits in Hungary.

[8]More specifically, potential benefit in the first tier was calculated as UI contribution days in the last 4 years divided by 5, but at most 270 days.

On May 30th, 2005 the Hungarian government announced a comprehensive reform of the UI system[9], with the goal of speeding up transition from unemployment to employment. The government changed the benefit calculations formula in the first tier, but did not alter the way potential duration and the earnings base were calculated. Before the reform, the benefit in the first tier was constant with a replacement rate of 65% and with minimum and maximum benefit caps. The reform introduced a two-step benefit system, as summarized in Web Appendix Figure A-1. For the first step, the length was half of the potential duration in the first tier, and at most 91 days, and the replacement was lowered to 60%, but with increased minimum and maximum benefit caps. In the second tier everybody received the new minimum benefit amount, reducing benefits for most claimants compared to before..

The most prominent change occurred for those with 270 days of eligibility (four years of UI contributions before lay-off) and base year earnings above the new benefit cap (earnings above 114,000 HUF ($570) per month in 2005). As Figure IIa) shows, for this group benefit duration in the first tier remains 270 days, but with higher benefits in the first 90 days, and lower benefits afterwards. Importantly, the increase in weekly benefits in the first 90 days is about twice as large as the decrease in weekly benefits between 90 and 270 days, keeping the total benefit pay-out for individuals unemployed for 270 days the same.

Even though the main element of the reform was the new benefit formula, other changes were introduced, including a reemployment bonus scheme equal to 50% of the remaining total first-tier benefits. However, claiming the bonus was not without costs, as the entitlement for the unused benefit days was nulled. Also, the bonus could only be claimed after the first-tier benefits were exhausted, and UI claimants had to show up and fill out the paper work in the local UI office. Given these costs, it is not surprising that the take-up rate was only 6%, making it unlikely that it had substantial effects. As a robustness check, we show that the results are not sensitive to dropping the reemployment bonus users from the sample.[10]

In addition to the introduction of the reemployment bonus, there were two other minor relevant changes. First, those who claimed UI benefit before February 5th, 2005 faced a longer, but somewhat lower, benefit in the second tier.[11] To avoid this complication, we only focus on those who claimed their benefits after February 5th, 2005. Second, there were minor changes

[9]The reform was part of a wider government program called "100 steps". Policies related to the labor market and unemployment insurance (such as reemployment bonus and training policies) are discussed later. In addition, VAT and corporate income tax were decreased from January 1st 2006.

[10]Lindner and Reizer (2015) further show that the bonus does not affect the unemployment duration.

[11]The change in the duration and benefit level in the second tier was introduced at November 1st, 2005 at the same time as other changes. However, it affected everybody who claimed second tier (UA) benefits after November 1st, 2005. A UI claimant who claimed her benefits after February 5th, 2005 and had 270 days potential eligibility could only claim second tier benefits (UA) after November 1st, 2005. Therefore, claimants between February 5th, 2005 and November 1st, 2005 are under the old benefit system for the first tier, but face the same second tier (UA) insurance scheme, see Figure 2b).

in financing training programs.[12] However, participation in training programs was very low (less than 5%) in our sample and our results are robust to dropping these claimants.

Those who exhausted benefits in both tiers (UI and UA) and were still unemployed could claim means-tested social assistance. The duration of social assistance is indefinite, while the amount depends on family size, family income, and wealth. In most cases social assistance benefits are lower than the second tier UI benefit level.[13]

III.B Data

We use administrative data[14] on social security contributions for roughly 4 million individuals between January 2002 and December 2008. The sample consists of a 50% de facto random sample of Hungarian citizens older than 14 and younger than 75 in 2002.[15] The data contains information on UI claims from February 2004 to December 2008 as well as basic information used by the National Employment Service, like the start and end date of the UI benefit spells and the earnings base. This data allows us to calculate non-employment durations, that is the time between claiming UI benefit and the starting date of the next job.

In this paper we only focus on UI claimants who are eligible for the maximum potential duration (270 days) in the first tier. In addition, we restrict our sample to those who are older than 25 years and younger than 49 years, since specific rules apply close to retirement. Moreover, we identify as our main sample UI claimants with high earnings base, since our goal is to explore the variation in Figure IIa). To construct a consistent sample over time, we focus on the unemployed with earnings base above the 70th percentile among the UI claimants in a given year. In 2005, a UI claimant at the 70th percentile earned 100,800 HUF ($504).[16]

To evaluate the reform, we construct two comparison groups of workers who entered UI just before or just after the reform, since the claiming date determined the relevant regime. Due to the change in unemployment assistance in February 2005, we use all UI claimants between February 5th, 2005 and October 15, 2005 (to avoid getting too close to the reform) as our pre-reform group. For the post-reform group, we take UI entrants in the same date range

[12]Unemployed participating in training programs received the so-called income substituting benefit. Before November 1st, 2005 this amount was 22,200HUF ($111) or 44,400HUF ($222), depending on household characteristics and type of training. This benefit was payed in excess of the UI. After November 1st, the benefit was 34,200HUF ($171) for everybody, but the UI benefit was suspended during training. Although we only observe training participation after November 1st, 2005, aggregate data show that the probability of participation in training programs remained constant throughout this period Frey (2009).

[13]For large families, social assistance can be more generous than UI. However, social assistance cannot be claimed before all other benefits have been exhausted in the UI system.

[14]The data set is provided by the Centre for Economic and Regional Studies - Hungarian Academy of Sciences.

[15]The sample is composed of everybody born on the 1st of January, 1927, and every second day thereafter.

[16]Our results are robust to alternative earnings thresholds. For example, we obtain very similar results for individuals with (real) earnings base above 114,000 HUF ($570).

(February 5 to October 15) in 2006 so as to match possible seasonal patterns. Figure IIb) shows the timing of the two comparison groups and the range for which our data is available. For robustness checks, we later show results using data in the earlier and later ranges as well.

The basic demographic characteristics, such as age at time of claiming, education and log earnings in the years 2002 - 2004, are similar before and after the reform.[17] The take-up rates of the reemployment bonus scheme, which was introduced in 2005, are low.

IV Reduced Form Results

IV.A Estimation of Hazard Into Employment

In this section, we evaluate the impact of the reform on the exit rates from unemployment into employment. We estimate the hazard rates with a linear probability model separately for each 15-day period[18], indexed by t, after entering unemployment insurance:

$$I(t_i^* = t | t_i^* \geq t) = \beta_{0,t} + \beta_{1,t} POST_i + X_i\gamma + \epsilon_{it}, \tag{7}$$

where i indexes individuals and t_i^* represents the duration of non-employment of individual i. The left hand side is an indicator for individual i finding a job in period t, conditional on still being unemployed at the beginning of the period. The variable $POST_i$ is an indicator for individual i claiming benefits in the post-reform period, while X_i is a matrix of control variables. The equation is estimated separately for each period t on the sample of individuals who are still unemployed at time t (that is conditional on $t_i^* \geq t$). In our baseline estimates we do not control for any observables X_i, and show results controlling for X_i as robustness.

Figure IIIa) shows the estimates of equation (7) for each t with no controls. The blue line represents the coefficient estimates of $\beta_{0,t}$, the estimated hazard in the before period, while the red line represents the estimated $\beta_{0,t} + \beta_{1,t}$, the after period hazard. Vertical lines between the two series indicate differences that are statistically significant at the 5% level.

The exit rate from unemployment in the pre-reform period shows a familiar pattern for a one-step unemployment system. The exit hazard falls in the first months after entering UI, then increases as it approaches the exhaustion point of UI benefits (at 270 days). After this exhaustion point, it falls and spikes again as people exhaust the second tier benefits, unemployment assistance, at 360 days. The hazard rate then decreases monotonically, as unemployed people are only eligible for welfare programs.

[17] See Web Appendix Table A-1. Furthermore Web Appendix Figure A-3 shows that the unemployment rate was quite stable at around 7 percent, and the GDP growth was also stable during the sample periods, only slowing down at the beginning of 2007.

[18] We choose a 15-day period so that the benefit path steps after 90 days and 270 days occur at integer values of these periods. The results are very similar with hazards computed at 7 days or 30 days.

The exit hazard changes substantially after the reform. The hazard rate increases at 90 days, at the end of the higher UI benefit, and remains elevated compared to the pre-reform period for the following 2.5 months. By 180 days, the pre- and post-reform hazards have converged, and both hazards increase at the exhaustion of UI benefits at 270 days. Importantly though, the post-reform hazard increases significantly less, and the pre-reform hazard remains significantly higher for three months following UI exhaustion. Finally, by 360 days, the end of unemployment assistance, the two hazards once again converge.

The most striking difference occurs around day 270, when in the pre-reform period the hazard remains significantly higher after the UI exhaustion point (270 days), relative to the post-reform period. This difference in hazards fits nicely with the reference-dependent model: the workers in the pre-reform period experience a larger drop-off in benefits around day 270, inducing a spike in loss utility and thus an increase in the value of search. The persistence for three months of the higher hazard indicates slow adjustment of the reference point. Furthermore, the increase in hazard in the pre-period happens already in anticipation of benefit expiration at day 270, consistent again with reference dependence.

While we focus mainly on the hazard rate around day 270 because it leads to the most distinct predictions, other patterns are also consistent with reference dependence. The spike in the hazard at 90 days in the post-period, corresponding to the first step down in benefits, disappears after 3-4 months, consistent once again with loss utility and a slowly-adjusting reference point. However, this spike could also be explained by the standard model with unobserved heterogeneity, as we show below.

Figure IIIb) shows the estimated survival function for the two groups. For these estimates we use a variant of equation (7), including the whole sample and taking $I(t_i^* \geq t)$ as the outcome. The survival functions diverge after 90 days, with lower survival in the after group than in the before group. This difference persists until around 300 days, after which the lines converge. Since the expected duration is the integral over the survival function from 0 onward, the expected unemployment duration is significantly reduced in the after period.

IV.B Robustness Checks

The results presented so far do not control for demographic characteristics. Even though the differences in demographics between the pre- and the post-period are quite small (Web Appendix Table A-1), they could potentially explain differences in the hazard patterns if the demographic impacts on the hazard rates are large. Thus, we re-estimate equation (7) controlling for a rich set of characteristics, where we allow these characteristics to have arbitrary effects on the hazard function at each point, the only restriction being that the effect is the same in the before and after period. As Figure IVa) shows, controlling for observables has

virtually no effect on the hazard rates, implying that they cannot explain our findings.[19]

A second concern regards the introduction of a reemployment bonus in November 1st, 2005. While the take-up rate of the bonus was just 6% in our sample, it may still affect the hazard rate in the post-reform period, especially in the first 90 days. As a check, we re-estimate our baseline specification dropping all individuals that received a reemployment bonus; the results are virtually unchanged (Figure IVb)).

A third issue is the potential role of recalls. Recalls may be timed to UI benefits contributing to the spike in the exit hazard from unemployment at the exhaustion point (Katz (1986), Katz and Meyer (1990) and Nekoei and Weber (2015)). Furthermore, learning about the probability of recall may affect search effort throughout the unemployment spell.

In our main dataset we do not observe employer identifiers and therefore cannot distinguish recalls to the same employer from exits to new jobs. We therefore investigated the role of recalls using the CERS-HAS Linked-Employer Employee Dataset, where we observe whether workers are recalled, that is return to the same employer, at the end of a nonemployment spell. A limitation of this data set is that we only measure nonemployment durations on the monthly level and we do not observe UI receipt. Nevertheless, the exit hazards from nonemployment spells in this data look quite similar to our baseline results and in particular show the crossing of the hazard rates at 270 days (Figure A-4a) in the Web Appendix).[20] Importantly, the exit hazards are almost identical when we drop recalls (Figure A-4b)). This is consistent with Web Appendix Figure A-5, which shows that the recall hazard in our sample is small and does not vary much over the unemployment path, or between the pre- and post-reform periods. Overall, recalls and recall expectations do not explain the exit patterns with the reform.

Next, we consider the concern that the differences in the hazard rates may be due to a time trend. We estimate two placebo tests for whether there are differences in the two years before the reform and the one year before the reform. Web Appendix Figure A-7a) shows that the hazard rates are virtually unchanged between 2004 and 2005. There is a small difference right after the 270 line, which is expected due to the reduction in unemployment assistance in February 2005. Similarly Web Appendix Figure A-7b) shows that there are virtually no differences between the hazards 1 and 2 years after the reform, again indicating that the differences between our before- and after-period line up nicely with the reform.

We explore the timing further by plotting time-series graphs of the hazards. Figure Va) shows the evolution over time of the hazard between 30 and 90 days (red line) and between 90 and 150 days (black line). Each dot indicates the average hazard for a 3-month period between 2004 and 2007, with quarter 1 indicating the first 3-month period after the reform. Prior to

[19] Alternatively we also used propensity score reweighting to estimate the hazards in the pre- and post-period, holding the observables constant over time and obtained almost identical results (not shown).

[20] We mimic the baseline sample restrictions as much as possible.

the reform, the hazard at 90-150 days is smaller than the hazard at 30-90 days, consistent with the patterns in Figure IIIa). Subsequent to the reform introducing a step down of benefits after 90 days, the pattern abruptly changes. Already in the first quarter after the reform, the hazard at 90-150 days increases sizeably, becoming similar to the hazard at 30-90 days, a pattern that persists over the next 6 quarters. The figure provides little evidence of previous trends, suggesting that the changes in hazards are indeed a causal effect of the reform.

Figure Vb) provides parallel evidence for the hazard at 210-270 days versus at 270-330 days. In the quarters pre-reform, the hazard at 270-330 days is significantly higher than the hazard at 210-270 days, a pattern that changes abruptly with the first quarter following the reform. The time-series plots again indicate a change that is coincidental with the reform and not due underlying trends or changes in the macroeconomic environment.

V Structural Estimation

The evidence on the exit hazards is consistent with the predictions of the reference dependent model, but it is difficult to say to what extent the standard job search model with unobserved heterogeneity could explain the same patterns. We therefore turn to structural estimation as a more precise test, allowing flexibly for unobserved heterogeneity in the standard model.

Note that this is somewhat different from the typical goal of structural estimation of conducting welfare analysis or policy predictions. In our setting, the structural estimation is necessary to test between models, which is impossible to fully do in reduced form. It also allows us to compare the estimated parameters, such as the job search elasticity, the loss aversion, and impatience, to estimates in other settings, and thus check their plausibility.

V.A Set-up and Estimation

We use the model of section II, imposing six additional assumptions, some of which we relax later. First, we assume a search cost function of power form as in Paserman (2008) and Chetty (2003): $c(s) = ks^{1+\gamma}/(1+\gamma)$. The parameter γ is the inverse of the elasticity of search effort with respect to the net value of employment.[21] Second, we assume log utility, $v(b) = \ln(b)$. Third, similar to Bloemen (2005), Paserman (2008), Fougere et al. (2009), and van den Berg and van der Klaauw (2015) we model heterogeneity in the cost of search k.

Fourth, to avoid modeling on-the-job search, we start the worker problem in the first period of unemployment, and thus fit the hazard from the second period on.[22] Fifth, we set past wages equal to the median earnings in our sample, which is 135,000 HUF ($675), and assume

[21] The first-order condition of search effort (equation 4) is $c'(s^*) = v$, where we denote with v the net value of employment. This yields $s^* = (v/k)^{1/\gamma}$, and the elasticity of s^* with respect to v is $\eta_{s,v} = (ds/dv)\, v/s = 1/\gamma$.

[22] Recall that a successful job search in period t yields a job in period $t+1$.

that reemployment wages are constant over the UI spell and equal to past wages. Sixth, we assume that individuals start with zero assets, that they cannot borrow against their future income, and that they earn no interest on saved assets.

The vector of parameters ξ for the standard model are: (i) the three levels of search cost k_{high}, k_{med}, and k_{low}, with $k_{high} \geq k_{med} \geq k_{low}$, and two probability weights p_{low} and p_{med}; (ii) the search cost curvature γ; (iii) the time preference parameters δ and β. For the reference-dependent model, we estimate in addition: (iv) the loss aversion parameter λ; and (v) the number of (15-day) periods N over which the backward-looking reference point is formed.[23] For the the reference-dependent model we assume no heterogeneity, thus removing parameters k_{high}, k_{med}, p_{low}, and p_{med}. Notice that the weight η on gain-loss utility is set to 1 rather than being estimated; thus, the loss-aversion parameter λ can be interpreted also as the overall weight on the losses. Over the course of the unemployment spell the individual is always on the loss side, hence it is difficult to estimate a separate weight on gain utility and loss utility.[24]

Estimation. Denote by $m(\xi)$ the vector of moments predicted by the theory as a function of the parameters ξ, and by $\hat{m}$ the vector of observed moments. The moments $m(\xi)$ are the 15-day hazard rates from day 15 to day 540 for both the pre-reform and post-reform period, for a total of 35*2=70 moments. The estimator chooses the parameters $\hat{\xi}$ that minimize the distance $(m(\xi) - \hat{m})' W (m(\xi) - \hat{m})$. As weighting matrix W, we use a diagonal matrix that has as diagonal elements the inverse of the variance of each moment. To calculate the theoretical moments, we use backward induction. First we numerically compute the steady-state search and value of unemployment. Then we solve for the optimal search and consumption path in each period as a function of the asset level. Finally, we use the initial asset level as a starting value to determine the actual consumption path and search intensity in each period.

Under standard conditions, the minimum-distance estimator using weighting matrix W achieves asymptotic normality, with estimated variance $(\hat{G}'W\hat{G})^{-1}(\hat{G}'W\hat{\Lambda}W\hat{G})(\hat{G}'W\hat{G})^{-1}/N$, where $\hat{G} \equiv N^{-1}\sum_{i=1}^{N} \nabla_{\xi} m_i(\hat{\xi})$ and $\hat{\Lambda} \equiv Var[m(\hat{\xi})]$ (Wooldridge, 2010).

Identification. While the parameters are identified jointly, it is possible to address the main sources of identification of individual parameters. The cost of effort parameters k_j are identified from both the level of search intensity and the path of the hazards over time. This is clearest in the standard model, where the heterogeneity in the parameters is needed, for example, to explain the decay in the hazard after day 360. The search cost curvature parameter, γ, is identified by the responsiveness of the hazard rate to changes in benefits since $1/\gamma$ is the elasticity of search effort with respect to the (net) value of finding a job.

[23] In the estimations tables we report the speed of adjustment in days, which is just N*15.

[24] In principle, the weight on gain utility η can be separately identified as we show in a robustness section, since gain utility affects the utility of reemployment, but the reemployment utility does not allow for very precise identification of η.

The time preference parameters are identified by the presence of spikes around benefit cuts, among other moments. If the unemployed workers are patient, they save in advance of benefit decreases so as to smooth consumption. Impatient workers, instead, save little if at all and thus experience a sharp decrease in consumption around the benefit change. This consumption drop then induces a sharp increase in search effort as the benefits decrease.

Turning to the reference-dependence parameters, a major component to identification of the loss utility λ is the extent to which the hazard for the pre-period is higher both before and after day 270, in response to a larger loss. By contrast, the standard model implies essentially identical hazards from day 270 onwards. The loss parameter is also identified by the response to other benefit changes, such as at 90 days in the post-period. The parameter N, which indicates the speed at which the losses are reabsorbed into the reference point, is identified by the speed of convergence of the pre- and post-reform hazards after the benefit decreases.

V.B Benchmark Estimates

Figure VIa) presents the fit for the standard model with 3-type heterogeneity. The model fits quite well the surge in hazard around day 90 in the post-period, and the decreasing path of the hazard in the first 200 days. The fit is also reasonably good for the period from day 400 on. However, the fit between days 250 and 400 is poor. As discussed above, the standard model predicts that the hazard rates for the pre- and post-period should be almost exactly the same after day 270. As such, the model misses both the sharp difference in hazard between day 260 and day 360, as well as the spikes at both 260 and 360 days.

In comparison, Figure VIb) displays the fit of the reference-dependent model with no heterogeneity (and thus two fewer parameters). The fit in the first 250 days is very good, though it was quite good also for the standard model. But, as anticipated, the model does much better for longer durations, where the standard model fits poorly. The model fits better the surge in the hazard rate in the pre-period in anticipation of the benefit cut after 270 days (which is larger in the pre period than in the post-period), as well as the elevated level for the following three months, compared to the pre-period. Then the model tracks quite well the period following the exhaustion of unemployment assistance (after 360 days).

The fit of the reference-dependent model, while superior to the standard model, is certainly not perfect. The model does not capture the large spike on day 270 for the pre-period; storable offers may play a role in this case. In addition, the reference-dependent model under-fits the difference in hazards between the pre- and post-period after day 270.

As Table I shows, the goodness of fit (GOF) measure $(m(\xi) - \hat{m})' W (m(\xi) - \hat{m})$ is substantially better for the reference-dependent model, despite having fewer parameters. The estimates for the standard model (Column (1)) indicate substantial heterogeneity in cost k

and a relatively high elasticity of search effort ($1/\gamma$) to incentives. The estimates for the reference-dependent model (Column (2)) indicate a substantial weight on loss utility, $\hat{\lambda} = 4.9$ (s.e. 0.2), and slow adjustment of the reference point, $\hat{N} = 188$ (s.e. 11) days.

The most striking result, though, is the estimated degree of impatience: 15-day discount factors of $\delta = 0.89$ for the reference-dependent model and $\delta = 0.93$ for the standard model, implying annual discount factors of 0.17 or lower. Web Appendix Figure A-9a) provides evidence on the identification of the discount factor, indicating the goodness of fit of the best-fitting estimate for a particular (15-day) discount factor. For patient individuals ($\delta = 0.995$ or higher), the reference-dependent model does poorly: loss-averse workers with a high degree of patience would build a buffer stock, thus smoothing the loss utility. As individuals become more impatient, already for $\delta = 0.95$ the reference dependent model has a good fit (and better than the standard model), with the best fit for a lower discount factor. The fit of the standard model also improves as the discount factor decreases, though less steeply.

Thus, to fit the data the models require a degree of impatience which is hard to reconcile with other estimates in the literature. Yet, this high degree of impatience may be due to a mis-specification of the discounting function. In Columns (3) and (4) we assume beta-delta preferences (Laibson (1997) and O'Donoghue and Rabin (1999)), allowing for an additional discount factor β between the present and the next period to capture the present bias. To keep the number of parameters constant, we set the long-term discount factor δ to 0.995. The fit is better than in the models with delta discounting for the reference-dependent model with much more plausible discounting: the estimated present-bias parameter is $\beta = 0.58$, implying a discount factor of 0.46 for the first year and of 0.88 for subsequent years. This indicates a substantial degree of impatience, but in line with estimates in the literature.[25]

In light of both the higher plausibility and the better fit, we adopt the reference-dependent model with beta-delta discounting as the benchmark behavioral model in the rest of the paper. For the standard model, especially given the small difference in fit between the two discounting functions, we use the more standard delta discounting.[26]

How do the two models achieve their fit? In Web Appendix Figures A-10 and A-11 we report plots for key model components, focusing on the high-cost type for the standard model. In the standard model, the flow utility follows the step down in benefits, with the size of the later steps accentuated by the curvature of the utility function. In the reference-dependent model, the flow utility captures also the intensity of the loss relative to the reference point. The value of unemployment decreases over time in the standard model, while in the reference-

[25]For example, Paserman (2008), building on DellaVigna and Paserman (2005), using job search data obtains estimates for β ranging between 0.40 and 0.89, depending on the sample. Laibson et al. (2007), based on life-cycle consumption choices, estimates a β between 0.51 and 0.82.

[26]The results for the reference-dependent models with either delta or beta-delta discounting are qualitatively similar in all the subsequent specifications.

dependent model it actually increases over most of the range, reflecting the importance of reference point adaptation (and fitting the observed decrease in search effort over time). The value of unemployment declines sharply in correspondence to the benefit drop. The reference point is decreasing over time and is higher in the pre-reform group from around day 250, generating higher loss utility and thus a larger increase in search effort near benefit expiration.

The value of employment, which is almost constant in the standard model, increases monotonically over time for the reference-dependent model, as getting a job is associated with a larger gain utility as the reference point declines. This latter force does not account for much of the results, as we illustrate later when we turn off gain utility. Consumption tracks quite closely the per-period earnings, especially in the reference-dependent model, and assets are close to zero over the spell, broadly consistent with Ganong and Noel (e.g. 2015).

V.C Reference-Dependence Variants and Habit Formation

In Table II, we consider alternative formulations of reference dependence along two dimensions. First, we keep the same utility function but consider alternative reference points, including status quo and forward-looking expectations. Second, we take as given a backward-looking reference point, but examine alternative utility functions, including habit formation models.

We assumed a backward-looking reference point with a memory of N periods. Does it matter how one models the backward-looking averaging? In Column (2) we assume an updating rule with longer "memory" and smoother adjustment, an AR(1) process: $r_t = \rho r_{t-1} + (1-\rho)y_{t-1} = (1-\rho)\sum_{i=1}^{\infty}\rho^i y_{t-i}$. The fit, as Figure 7a shows, is quite similar to our standard assumption, with a shorter estimated memory for the reference point.[27]

While a backward-looking reference point has parallels in the literature (e.g., Simonsohn and Loewenstein 2006), other reference points are also common, such as the status quo. In endowment effect experiments (Kahneman et al. (1991)), this is the initial allocation, and in asset pricing papers (Barberis and Huang, 2001), it is the purchase price of the asset. In our context, we take the reference point to be the last wage before the start of the unemployment spell. This reference point is still backward-looking, but with no adaptation. This specification does poorly (Web Appendix Table A-3): the adaptation over time is critical to reproducing the initial surge in hazard, the decline, and then second surge at benefit exhaustion.

A second common class are forward-looking reference points a' la Kőszegi and Rabin (2006): individuals use forward-looking rational expectations formed in the recent past as reference points. We thus take as reference point for period t the expected earnings for period

[27]When we implement this estimate we assume that the memory of the AR(1) update goes back to 1050 days (or 70 15-day periods). We adopt as benchmark the N-period reference point for computational reasons, since the long memory of the AR(1) model makes the estimates more time-consuming.

t, as expected in period $t-1$.[28] As Web Appendix Table A-3 documents, this reference point also provides a poor fit to the data. Importantly, forward-looking reference points cannot reproduce the large difference in hazards past day 270.

Thus, a backward-looking, adaptive reference point is critical. Does it matter how it is incorporated in the utility function? In our benchmark model, gain utility gets weight η while loss utility gets weight $\eta\lambda$, and for estimation we set $\eta = 1$. Assuming no gain utility when workers get a job, but still estimating the loss utility weight $\eta\lambda$, results in a fit similar to the benchmark one (Web Appendix Table A-3). Conversely, assuming no loss utility but estimating gain utility with weight η, instead, leads to a worse fit than the standard model, indicating the key role played by loss utility (Column (3) of Table II). To further focus on loss aversion, we show that assuming no loss aversion ($\lambda = 1$) and estimating η leads to a fit that is not quite as good as that of the benchmark model. When estimating both η and λ, we obtain imprecise estimates for η (1.63, s.e. 2.92). As such, in the rest of the paper we hold η fixed at 1. Finally, if the loss utility is expressed entirely in terms of income, that is, $\eta\lambda\left[v\left(y_t\right) - v\left(r_t\right)\right]$, the results are essentially identical (Web Appendix Table A-3).

Importantly, can we distinguish our model from habit formation models, which also share an adaptive reference point component? Models a la Constantinides (1990) and Campbell and Cochrane (1999) assume utility $u(c - zr)$, where r is the habit formed from past consumption and u is a concave function. Habit formation, like reference dependence, induces a temporarily high marginal utility following a benefit cut, as consumption c gets closer to the habit zr. Thus, it could also plausibly fit the patterns in the data.

We estimate a habit formation model replacing the utility in (1) with $v(c_t, r_t) = \log(c_t - zr_t)$, where z captures the responsiveness to changes in the habit and r_t is calculated as before, but reinterpreted as a measure of habit stock.[29] The estimates, allowing for two unobserved types, are in Columns (6) and (7) (for the habit formed with an AR1 process). The fit is significantly worse than the fit of the reference-dependent model (see also Figures VII c-d).

This may appear surprising given the similar intuition behind the two models. The models however differ in a key aspect. In the reference-dependent model, the impact of the loss, $\lambda(u(c) - u(r))$, on search effort is approximately proportional to the size of the loss. Instead, in the habit-formation model larger decreases in consumption have disproportionate effect, as c gets closer to zr. Given this, the habit-formation model fits the data less well, since it predicts a larger spike at the 90-day (post reform) benefit decrease, and a much smaller spike for the later (proportionally smaller) benefit decreases.[30]

[28] We provide further details in the Web Appendix.

[29] Observe that this function is not defined whenever $c_t < zr_t$, complicating the estimation. To avoid this problem, Campbell and Cochrane (1999) made z a non-linear function of $y_t - r_t$. We treat z as a parameter instead and check in the optimum that our utility function is defined for the relevant y_t and r_t.

[30] We provide further detail in Web Appendix Figure A-13. The habit-formation model is also computation-

V.D Robustness and Unobserved Heterogeneity

In Table III we consider the robustness of the standard and reference-dependent models to alternative specifications, including alternative assumptions on heterogeneity.

We attempt to estimate both patience parameters, β and δ, but the parameters are quite collinear (Column (1)). We allow for a linear time trend in the baseline cost to capture skill depreciation (Column (2)), and we explore the role played by the spikes at days 270 and 360, dropping these moments (Column (3)). The qualitative findings are unaltered.

We also consider the importance of timing in the model. We assume that jobs start one period after the offer is received, but what if the hiring process takes longer? Assuming a 2-period (that is, one-month) delay between the job offer and the first wage payment leaves the results unaltered and does not shift the spike (Figure VIIIa)). Job-seekers start their search earlier, taking into account the longer delay.

In Web Appendix Tables A-5 and A-6, we document that the results are robust to a series of statistical checks: using equal-weighted moments, controlling for observables, or using the (unconditional) probability of exiting unemployment in each 15-day period as moments. (The latter procedure allows us to use the full variance-covariance matrix for weights.) The estimates are also similar if we use 30-day or 7-day hazards. Turning to modeling assumptions, allowing for background consumption and for alternative assumption on welfare receipt does not significantly affect the results. Under alternative assumptions for the reemployment wage and the initial assets, the fit of the models is somewhat worse, but the results are similar.

Heterogeneity. So far, we have modeled a single form of heterogeneity, in search costs, and fixed the number of types at 3. We now relax both assumptions.

Allowing for 4 types improves the fit of the standard model, though the fit remains significantly worse than in the reference-dependent model (Figure VIIIb); estimates with 5 types have trouble converging. For the reference-dependent model, there is an improvement in fit going to 2 types, with no improvement thereafter.

Next, we consider alternative forms of unobserved heterogeneity, such as in the reemployment wage. We take the 10th, 50th, and 90th percentile of the reemployment wage from the data, fix the proportion of each type at 20, 60, and 20 percent respectively, and estimate three cost parameters k_j. This specification improves somewhat the fit of the standard model (Web Appendix Table A-7), but the fit of the reference-dependent model is still significantly better.

Allowing for heterogeneity in the curvature γ, instead, improves the fit dramatically. The estimates fit the spikes at 270 and at 360 days, as well as the difference between the pre- and post-reform period (Figure VIIIc). How does this model attain a fit superior to the reference-dependent model? The fit relies on vast heterogeneity in the elasticity of search $1/\gamma$. Initially,

ally trickier to estimate, as the estimated habit parameter γ has to always satisfy the condition $c > \gamma r$.

most exits are of the low-elasticity types ($\gamma_{high} = 1.04$), but the spike at day 270 is driven by the medium-elasticity types ($\gamma_{med} = 0.20$), whose search intensity surges with the lower benefits. The spike at 360 days is due to the high-elasticity types (with elasticities over 50, i.e. $\gamma_{low} = 0.017$), that start searching once benefits hit the welfare level.

The high-elasticity types are critical to the fit, as estimates with elasticity capped at 5 (Column (8) and Figure VIIId)) do not match the fit of the reference-dependent model any more. To assess the implications of such elasticities, consider a hypothetical 10 percent increase in the level of welfare benefits paid out after 360 days. In response to such a small benefit change, the search effort of reference-dependent workers would barely change, but in the gamma-heterogeneity model, instead, the hazard rates would plummet (Web Appendix Figure A-18). We find this implied response unrealistic.

V.E Out-of-Sample Predictions

To further compare the different models, we consider two sets of out-of-sample predictions: a reform of the unemployment assistance system two years prior to our main sample, and the response to our main reform for individuals with a lower earnings basis, who experienced a different change in benefits. We predict the response at the estimated parameters using the reference-dependent model (Column (4) in Table I), the standard cost-heterogeneity model (Column (1) in Table I) and the gamma-heterogeneity model (Column (7) in Table III).

Turning to the first case, one year before our 'pre' period (see Figure 2b), the duration of the unemployment assistance was 180 days, compared to 90 days for the individuals in our sample.[31] As Figure IXa) shows, the reference-dependent model fits this earlier period well, with an out-of-sample GOF of 53.3. The gamma heterogeneity model, instead, fits quite poorly the period of the lengthened unemployment assistance (between 300 and 450 days), with an out-of-sample GOF of 111.3. The out-of-sample fit of the standard model with heterogeneity in cost levels is better (GOF of 81.1), but does not reach the reference-dependent model.

Second, we consider individuals in our main sample period, but with lower pre-unemployment income: a low-wage sample and a medium-wage sample (Web Appendix Figure A-1). Both groups experience less generous benefits post reform in the first 90 days, compared to our main sample. Figure IXb) displays the hazards for the low-wage sample in the pre-reform period, and the out-of-sample predictions according to the three models. The reference-dependent model captures quite well the patterns, while the model with heterogeneous elasticity provides the worst fit. Web Appendix Figure A-19 provides similar evidence for this sample in

[31] This change applied to individuals aged 45 and lower. Since our sample includes only individuals up to age 50, the reform applies to the large majority of the sample. Excluding the 46-50 year olds barely affects the estimates.

the post-reform period, as well as for the medium-wage sample. The reference-dependent model does consistently better out of sample than the other models.

V.F Dynamic Selection throughout the UI spell

The standard model reasonably captures the exit hazards in the first 270 days as well as some of the trend after that, especially in the gamma-heterogeneity version. To achieve this fit, changes in the unobserved types over time play a key role. How plausible then is the amount of heterogeneity that the standard model requires? While we cannot measure the time-changing unobserved heterogeneity, we propose that a useful metric is the time-varying selection on observables of the unemployed, under the assumption that unobservable factors that influence job search correlate with these observable characteristics.

To document the dynamic selection along observables, we regress at the individual level the realized unemployment duration (censored at 540 days) on a rich set of observables.[32] These variables are reliable predictors of non-employment duration, with an R^2 of 0.05-0.06 (Web Appendix Table A-10) and the predicted duration based on these estimates for the pre-period varies between 230 days (5th percentile) and 370 days (95th percentile), a good amount of variation.The dotted lines with crosses in Figure Xa)-b) show the predicted duration for individuals who exit unemployment in a given 15-day period. While predicted unemployment increases (unsurprisingly) throughout the spell, the relationship is quite flat and barely affected by the benefit path, with fairly parallel lines for the pre- and post-reform periods.[33] Selection on observables thus plays only a limited role in the data.

The selection over time in predicted duration has a counterpart in the models. For each type, we compute the expected unemployment duration in the pre-reform period, and calculate the average expected duration for unemployed individuals who leave in a given period according to the estimated models. The reference-dependent model (solid lines in Figure Xa-b), which predicts no type shift, is similar to what we observe empirically. The standard model with cost heterogeneity (Figure Xa)) instead displays a large swing until 200 days, with no corresponding selection on observables. The gamma-heterogeneity model (Figure Xb)) is even more at odds with the data, with an initial swing, and then a second swing between 300 and 360 days, corresponding to the transition from the medium- to the high-elasticity type.

Thus, the patterns of dynamic selection implied by two versions of the standard model appear at odds with the much more muted and monotonic selection in the data. Of course,

[32]The observables are education, age groups, gender, waiting period (the number of days between job lost and UI claimed), log past earnings, indicators for county of residence, day of the month UI was claimed, and occupation (1 digit) of the last job.

[33]The fact that dynamic selection seems to be small and not much affected by the UI regime is similar to the finding in Schmieder et al. (2016).

it is not surprising that the selection on observables is more muted than the selection implied by the model, given that we only observe part of the selection. However, the extent of the difference is quite striking, given the rich set of variables used in calculating dynamic selection. Moreover, it is puzzling for the standard model that the observed selection does not display any of the trends in the model predictions, even on a more muted scale.

V.G Reservation Wages

So far the reemployment wage is fixed so the unemployed accept every job offer. While this is consistent with a literature documenting a small role of reservation wages for job search dynamics (e.g. Card et al., 2007a; Schmieder et al., 2016; Krueger and Mueller, 2016), a natural question is whether introducing a reservation wage would change our conclusions. We estimated standard and reference-dependent models that incorporate reservation wages, though for tractability assuming hand-to-mouth consumers as well as some other simplifying assumptions (see the Web Appendix for details of model and estimation). These results should be considered only suggestive, as endogenizing consumption is important, but they continue to support the importance of reference dependence: Web Appendix Table A-11 shows that the reference-dependent model has a better fit than the standard model (GOF of 273 versus 300), largely due to the reference-dependent model providing a better fit for the hazard moments (see also Web Appendix Figures A-16). Notice however that the estimates have relatively patient unemployed workers, at odds with the maintained assumption of hand-to-mouth consumption.

VI Conclusion

We provided evidence that a model with reference-dependent preferences can explain qualitative features of the hazards which plausible versions of the standard model have a hard time fitting. The model itself builds on one of the most robust behavioral models, reference dependence, and uses a natural candidate for a backward-looking reference point. We also find that job seekers are substantially impatient, likely in the form of present-bias preferences.

This evidence has policy implications. Reference-dependent job-seekers respond strongly in their search effort to front-loaded benefits. Lindner and Reizer (2015) show that the Hungarian UI reform examined here did not just speed up exits to employment, but it was revenue-neutral from the perspective of the government. This evidence suggests that multiple-step unemployment insurance systems could prove advantageous.

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Table I: Benchmark Estimates of the Standard and Reference-Dependent Model

	δ-discounting		$\beta\delta$-discounting	
	Standard 3 type	Ref. Dep. 1 type	Standard 3 type	Ref. Dep. 1 type
	(1)	(2)	(3)	(4)
Parameters of Utility Function				
Loss aversion λ		4.91		4.54
		(0.23)		(0.25)
Adjustment speed of reference point N in days		188.4		167.4
		(11.0)		(9.25)
Discount factor (15 days) δ	0.93	0.89	0.995	0.995
	(0.01)	(0.01)		
Discount factor β	1	1	0.91	0.58
			(0.01)	(0.03)
Parameters of Search Cost Function				
Curvature of search cost γ	0.35	0.81	0.07	0.37
	(0.04)	(0.16)	(0.01)	(0.02)
Search cost for high cost type k_{high}	173.1	359.8	964.4	108.4
	(114.1)	(33.3)	(.)	(14.5)
Search cost for medium cost type k_{med}	49.6		105.2	
	(2.0)		(9.0)	
Search cost for low cost type k_{low}	12.9		13.6	
	(1.1)		(1.7)	
Share of high cost UI claimant	0.21		0.17	
	(0.07)		(0.01)	
Share of medium cost UI claimant	0.63		0.73	
	(0.06)		(0.01)	
Model Fit				
Number of moments used	70	70	70	70
Number of estimated parameters	7	5	7	5
Goodness of Fit	227.5	193.3	229.1	183.7

Notes: The table shows parameter estimates for the standard and the reference-dependent search models. Estimation is based on minimum distance estimation, using the hazard rates in the pre- and post-reform periods as the moments.
Standard errors for estimated parameters in parentheses.
(.) indicates that the parameter is not well identified, i.e. the Hessian cannot be inverted close to the reported values and therefore we do not provide standard errors. The other standard errors are calculated by inverting the Hessian matrix after dropping the parameter from the matrix.

Table II: Alternative Specifications for Reference-Dependent Model

	Benchmark RD 1-type	AR(1) Updating	No Loss Utility	Estimate η Fix $\lambda = 1$	Estimate λ and η	Habit Model a la Constantinides (1990) $log(c_t - zr_t)$	
	1-type	1-type	1-type	1-type	1-type	2-type	2-type, AR(1)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Parameters of Utility Function							
Loss aversion λ	4.54	16.9	0	1	3.13		
	(0.25)	(4.08)			(3.75)		
Gain utility η	1	1	0.002	39.3	1.63		
			(.)†	(18.9)	(2.92)		
Habit formation parameter z						0.29	0.26
						(.)	(.)
Adjustment speed of reference point N in days	167.4		584.4	188.0	168.3	211.1	
	(9.25)		(734493.2)	(21.9)	(12.9)	(20.7)	
AR(1) parameter		0.74					0.93
		(0.06)					(0.01)
Implied half life of AR(1) process		34.9					150.0
		(9.87)					(30.6)
Discount factor (15 days) δ	0.995	0.995	0.995	0.995	0.995	0.927	0.904
						(0.002)	(0.018)
Discount factor β	0.58	0.05	0.72	0.08	0.55	1	1
	(0.03)	(.)†	(2.15)	(0.09)	(0.35)		
Curvature of search cost γ	0.37	1.39	1.70	1.33	0.40	0.33	0.52
	(0.02)	(0.50)	(2.70)	(0.48)	(0.37)	(0.01)	(0.01)
Model Fit							
Number of moments used	70	70	70	70	70	70	70
Number of estimated parameters	5	5	5	5	6	7	7
Goodness of Fit	183.7	175.9	477.1	197.3	183.6	228.4	247.4

Notes: The table shows parameter estimates for the reference-dependent search model. Estimation is based on minimum distance estimation, using the hazard rates in the pre- and post-reform periods as the moments. Standard errors for estimated parameters in parentheses.

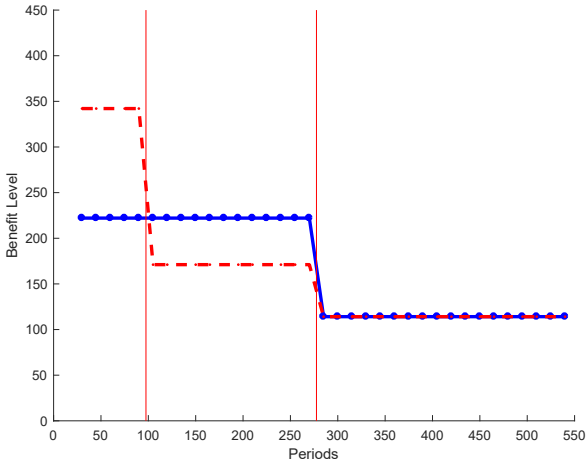
† The AR(1) and no loss utility estimations do not converge to an interior solution within our parameter space and therefore standard errors are not reported. Column (4) also corresponds to the habit formation model in Abel (1990) with utility function $log(c_t) - \eta log(r_t)$. The gain utility parameter η in this model is close to the boundary we set for the estimation. In columns (6) and (7), the parameter z is at the upper bound of possible values for z (for higher values the gain loss part would be the log of a negative number) and we therefore do not provide standard errors.

Table III: Robustness to Alternative Specifications of Utility Function, Estimation Methods and Heterogeneity

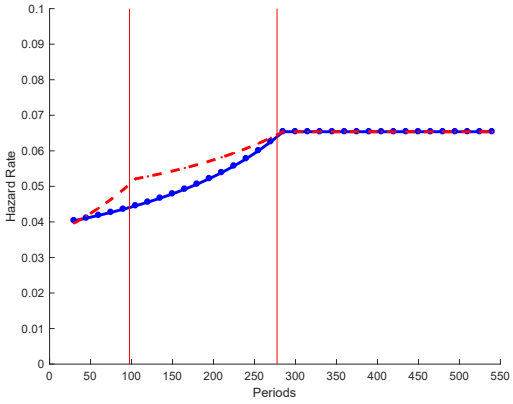
	Estimate β and δ	Time-varying search cost	Estimation without Spikes	Delayed Job Start Date	2 cost types	4 cost types	Heterogeneity search cost curvature γ	Heterogeneity search cost curvature $\gamma \geq 0.2$
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Standard Model								
Discount factor (15 days) δ	0.937	0.913	0.930	0.927	0.898	0.918	0.889	0.858
	(0.239)	(0.017)	(0.053)	(0.038)	(0.019)	(0.006)	(0.013)	(0.35)
Discount factor β	0.617							
	(0.103)							
Time varying cost		-0.016						
		(0.002)						
Curvature of search cost γ_{high}					0.49	0.31	1.04	1.92
					(0.02)	(0.04)	(0.002)	(0.016)
Curvature of search cost γ_{med}							0.20	0.49
							(0.004)	(0.01)
Curvature of search cost γ_{low}							0.017	0.20
							(0.164)	(0.006)
Number of moments used	70	70	66	68	70	70	70	70
Number of estimated parameters	8	8	7	7	5	9	7	7
Goodness of fit (SSE)	222.0	225.3	152.0*	213.6*	296.9	222.9	155.6	209.7
Reference Dependent Model								
Loss aversion λ	4.60	13.25	4.97	3.98	4.06		3.94	
	(1.18)	(5.24)	(0.85)	(0.25)	(0.26)		(0.04)	
Adjustment speed of reference point N	169.8	198.8	174.9	184.9	189.6		193.7	
	(13.5)	(11.1)	(17.4)	(11.5)	(13.3)		(265.2)	
Discount factor (15 days) δ	0.981	0.995	0.995	0.995	0.995		0.995	
	(0.437)							
Discount factor β	0.52	0.55	0.52	0.70	0.58		0.58	
	(0.10)	(0.12)	(0.26)	(0.03)	(0.03)		(0.03)	
Time varying cost		-0.020						
		(0.004)						
Curvature of search cost γ_{high}	0.47	0.58	0.47	0.25	0.37		13.4	
	(0.83)	(0.17)	(0.30)	(0.03)	(0.02)		(0.20)	
Curvature of search cost γ_{low}							0.364	
							(1.03)	
Number of moments used	70	70	66	68	70		70	
Number of estimated parameters	6	6	5	5	7		7	
Goodness of fit (SSE)	183.8	175.0	132.5*	177.5*	175.7		175.4	

Notes: The table shows parameter estimates for the standard and the reference-dependent search model. Estimation is based on minimum distance estimation, using the hazard rates in the pre- and post-reform periods as the moments. Standard errors for estimated parameters in parentheses.
In column (2) we parameterize the search cost function as: $c(s) = k(1 + \xi t)(s^{1+\gamma}/(1+\gamma))$ and the reported parameter is ξ: the change in the level of search cost per period. The 4-type reference-dependent model (col 6) does not converge.
* These are the SSE with the alternative moments and they are not directly comparable to the goodness of fit statistics in the other columns.

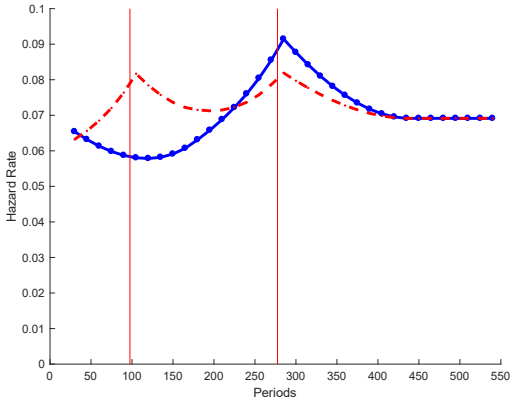
Figure I: Model Simulations of the Standard and the Reference-Dependent model



(a) Benefits



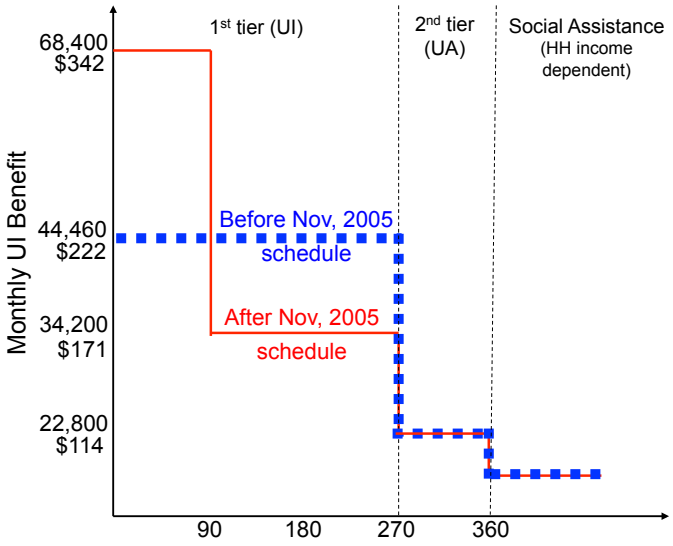
(b) Standard Model

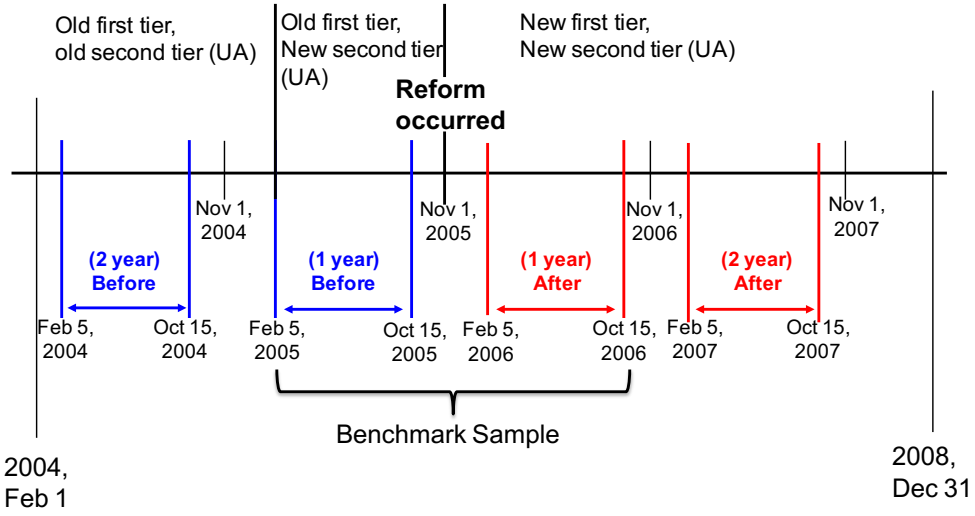


(c) Reference-Dependent Model

Notes: Panel (a) shows two benefit regimes, both of them having a step-down benefit system. After the first step benefits are higher in the regime represented by the circled blue line than in the regime represented by the red dashed line. After the second step benefits drop to the same level. Panel (b) shows the hazard rates predicted by the standard model (with $k = 130$, $\gamma = 0.2$, $w = 555$, $\delta = 0.99$) while Panel (c) the prediction of the reference-dependent model (with $k = 130$, $\gamma = 0.2$, $w = 555$, $\delta = 0.99$, $\lambda = 2$, $N = 10$ (150 days)).

Figure II: Institutional Setting: Change in Benefit Path and Sample Periods

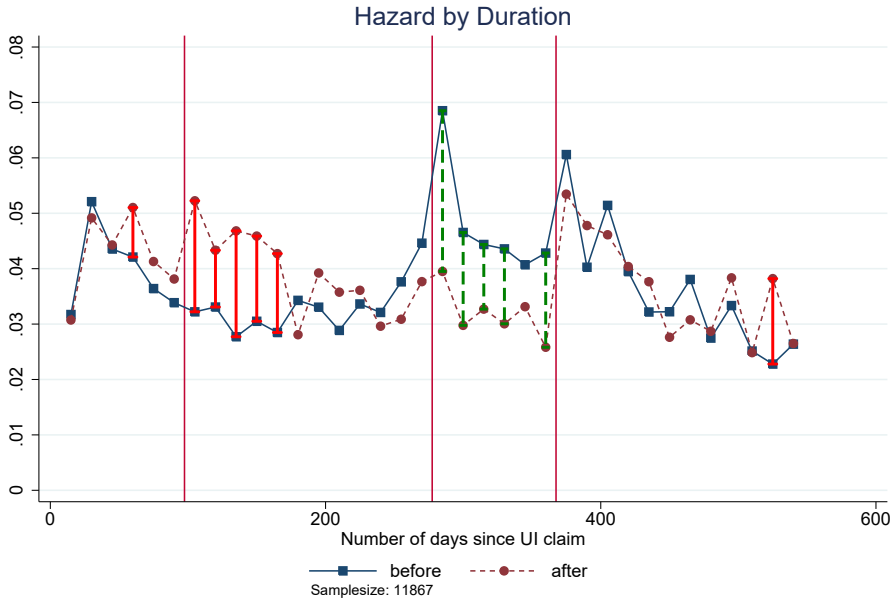


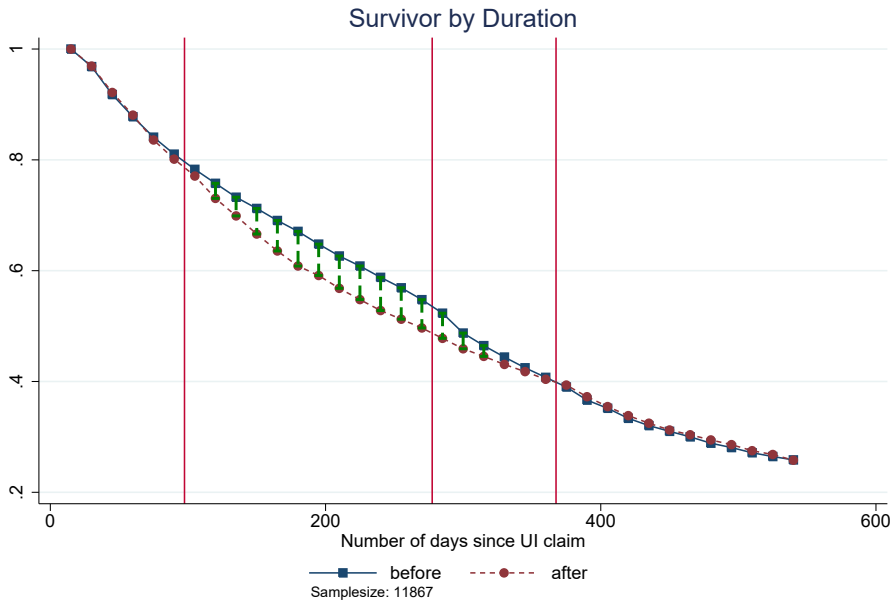


(b) Before-After Comparison Groups

Notes: Panel a) shows the benefit schedule if UI is claimed on October 31, 2005 (old benefit schedule, dashed blue line) and benefit schedule if UI is claimed on November 1st, 2005 (new benefit schedule, solid red line) for individuals who had 270 days potential duration in the first-tier, were less than 50 years old and earned more than 114,000 HUF ($570) prior to entering UI. Benefits levels in social assistance are approximate as they depended on family income, household size and wealth.
Panel b) shows the time frame for which we have access to administrative data on unemployment insurance records, the time of the reform and how we define the before and after periods that we use for our before-after comparison.

Figure III: Empirical Hazard and Survival Rates under the Old and the New Benefit Schedule

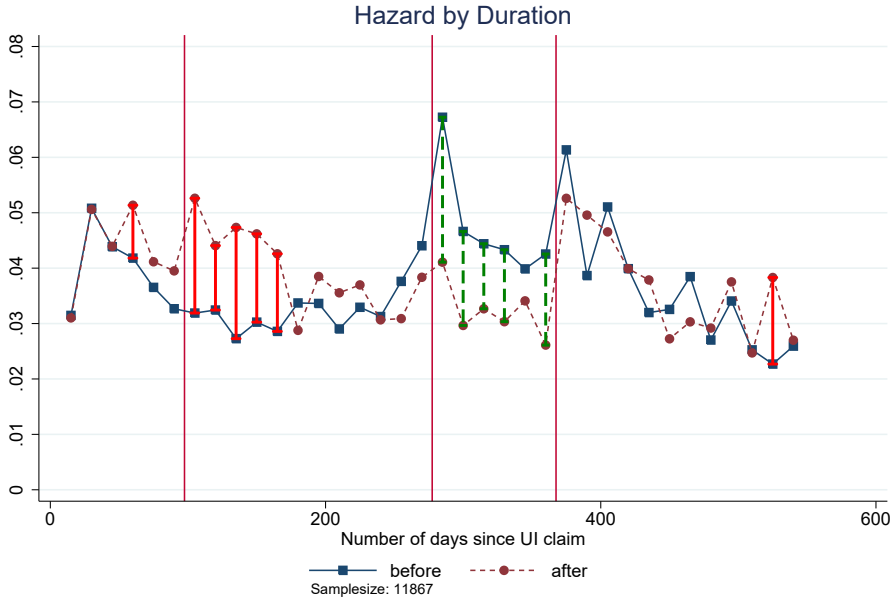


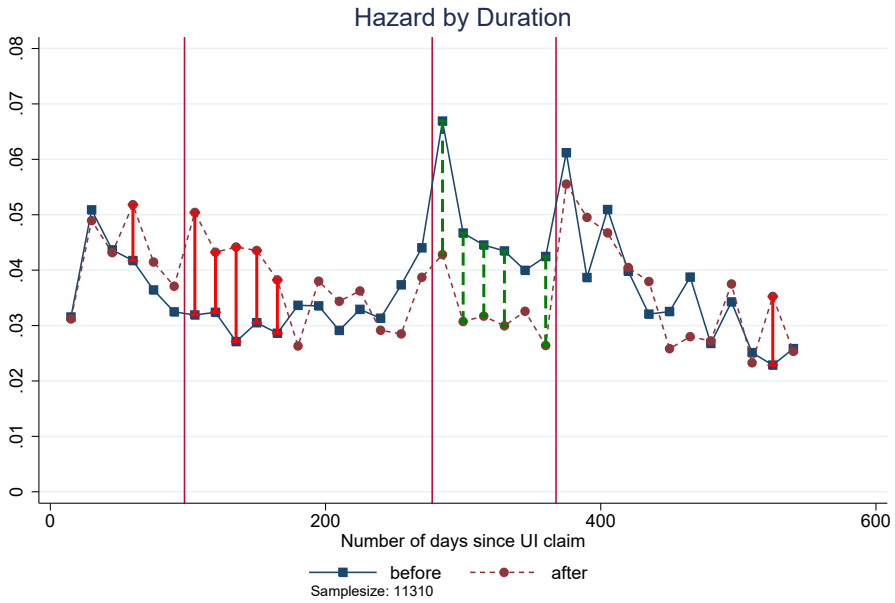


(b) Empirical survival rates

Notes: The figure shows point wise estimates for the empirical hazards, Panel (a), and for the empirical survival rates, Panel (b), before and after the reform. The differences between the two periods are estimated point-wise at each point of support and differences which are statistically significant ($p \leq 0.05$) are indicated with a vertical bar (green dashed if pre-period hazard is above post period hazard, red solid otherwise). The three major (red) vertical lines indicate periods when benefits change in the new system. The sample consists of unemployed workers claiming UI between February 5th, 2005 and October 15th, 2005 (before sample) and February 5th, 2006 and October 15th, 2006 (after sample), who had 270 days of potential duration, were 25-49 years old, and were above the 70th percentile of the earnings base distribution of the UI claimants in the given year.

Figure IV: Robustness Checks for change of Hazard rates before and after the reform

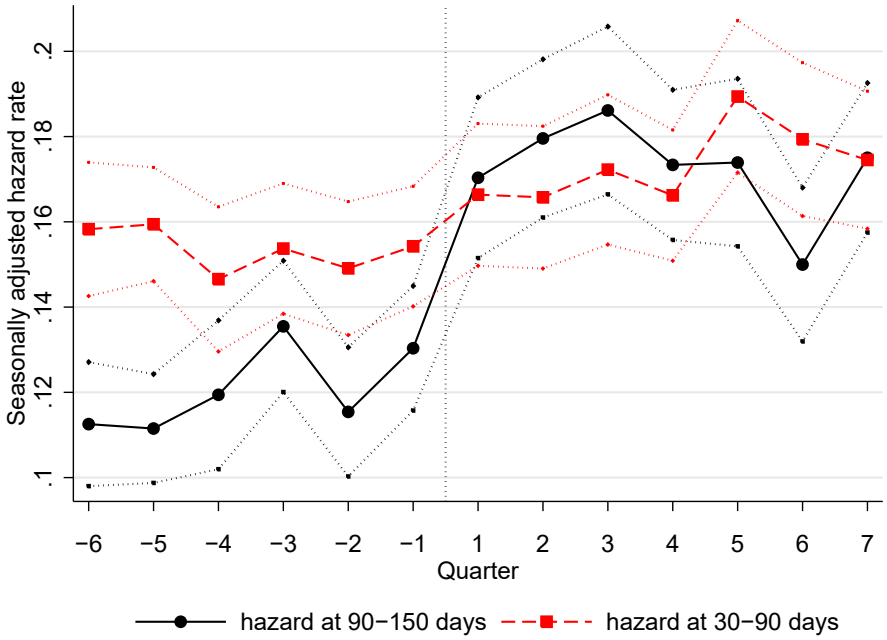


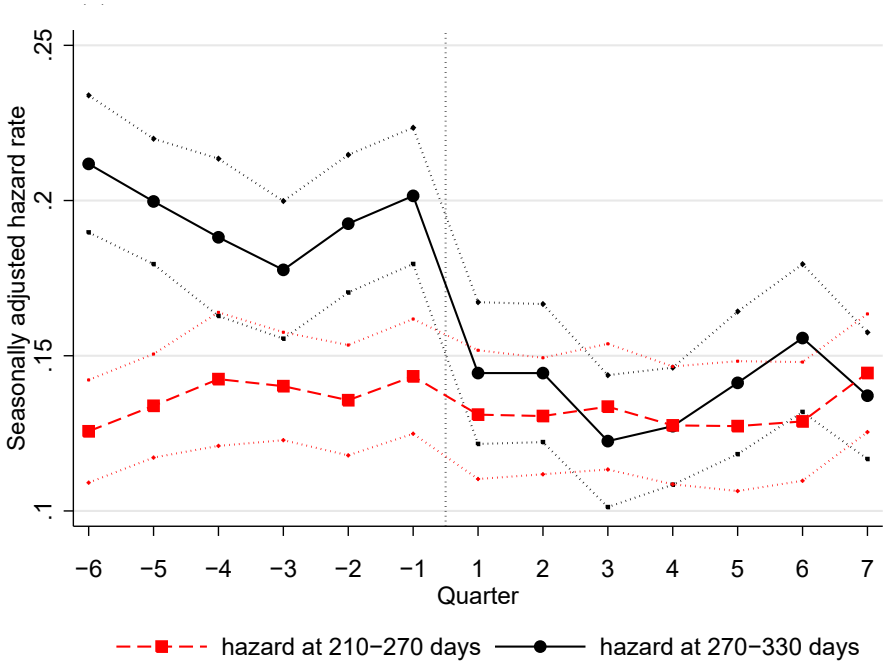


(b) Restricted sample

Notes: The figure shows point wise estimates for the empirical hazards before and after the reform. The differences between the two periods are estimated point-wise at each point of support and differences which are statistically significant are indicated with a vertical bar (green dashed if pre-period hazard is above post period hazard, red solid otherwise). The three major (red) vertical lines indicate periods when benefits change in the new system. In Panel (a) we added demeaned control variables for sex, age, age square, waiting period (the number of days between job lost and UI claimed), the county of residence, day of the month UI claimed, education, occupation (1 digit) of the last job, and log earnings in 2002 and 2003. In Panel (b) in addition to controlling for these control variables we dropped reemployment bonus claimants and those participating in training programs (after the reform), see text for the details. The sample is otherwise the same as in Figure III.

Figure V: Interrupted Time Series Analysis of Exit Hazards

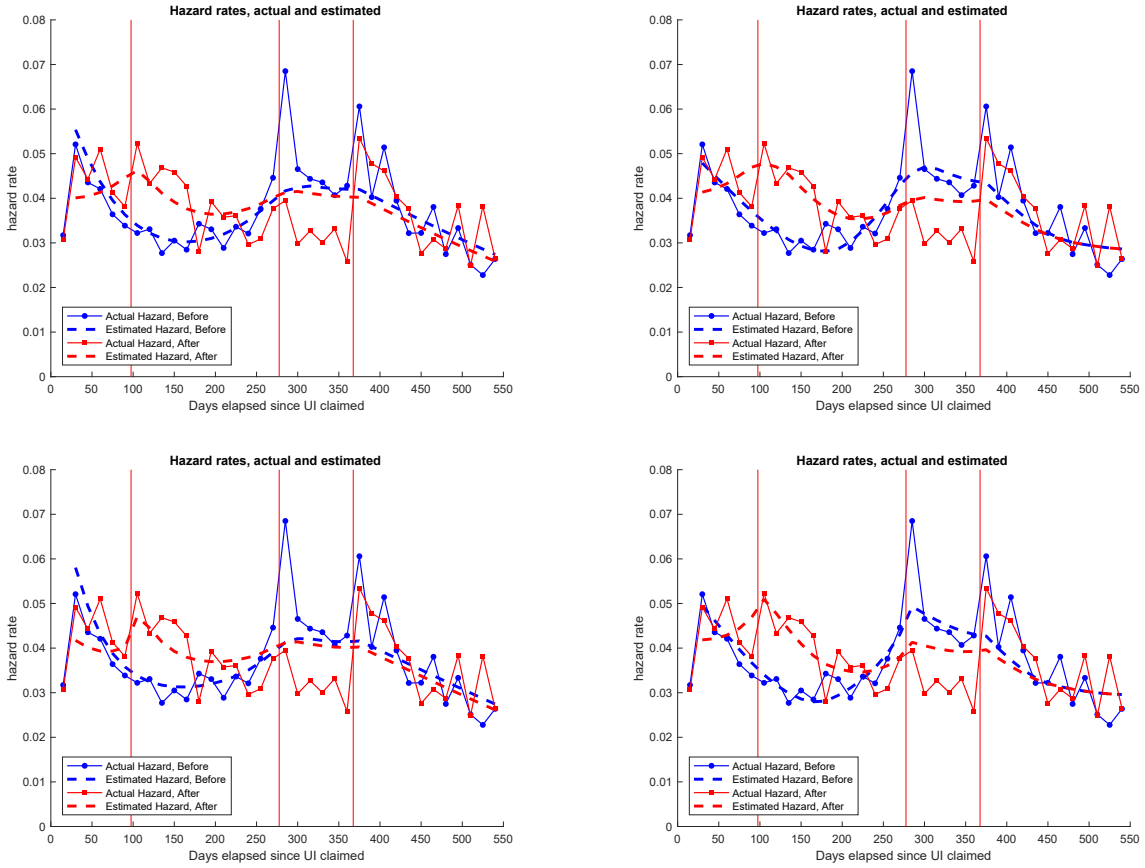




(b) The evolution of the hazard rates between 210 and 330 days

Notes: The figure shows the level of the most important hazard rates 6 quarters before and 7 quarters after the reform. Panel (a) shows the seasonally adjusted hazard rates between 30 and 150 days, while Panel (b) shows the seasonally adjusted hazard rates between 210 and 330 days. The monthly seasonal adjustment of hazard rates takes into consideration the level shift present in the data in November, 2005. The figures highlight that the shift in the hazard plots documented earlier corresponds to the precise timing of the reform. Other sample restrictions are the same as in Figure III.

Figure VI: Predicted Hazards of the Benchmark Standard and Reference-Dependent Models

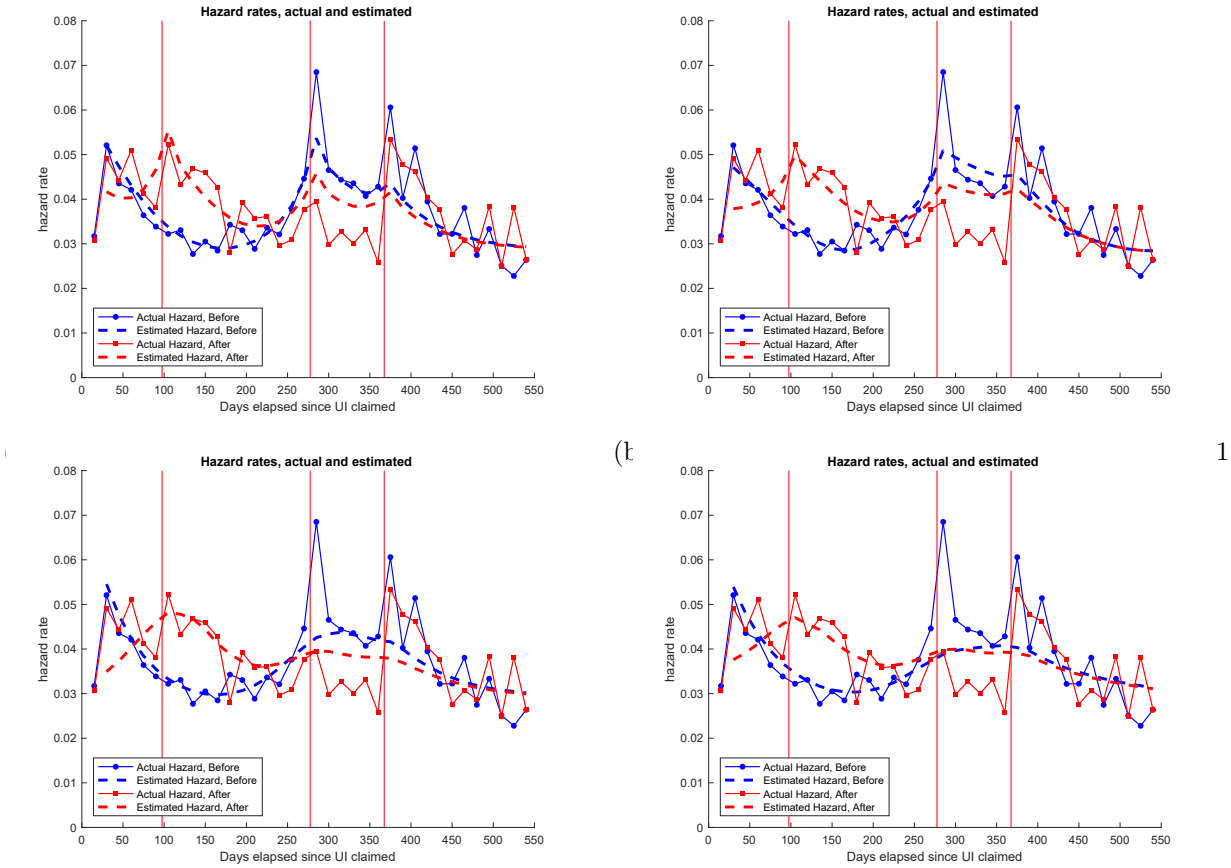


(c) Standard Model, $\beta\delta$-discounting (present bias)

(d) Ref.-Dep. Model, $\beta\delta$-discounting (present bias)

Notes: The figure shows the empirical hazards and the predicted hazards of the standard and the reference-dependent models with endogenous savings shown in Table I. Panel (a) corresponds to the standard model with 3 cost types and δ-discounting. Panel (b) corresponds to the reference-dependent model with 1 cost type and δ-discounting. Panel (c) shows the standard model (3 cost types) with $\beta\delta$-discounting (present bias) and Panel (d) the corresponding reference-dependent model. The three major (red) vertical lines indicate periods when benefits change in the new system.

Figure VII: Alternative Models with Reference Dependence and Habit Formation

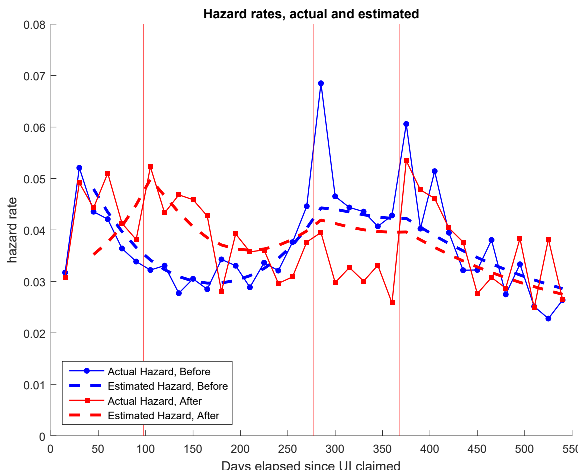


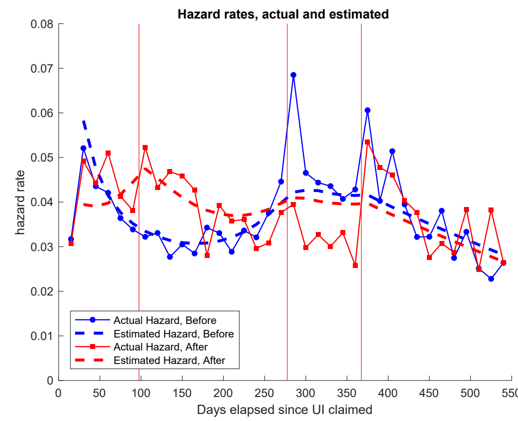
(c) Habit Model with 2 types (Constantinides 1990)

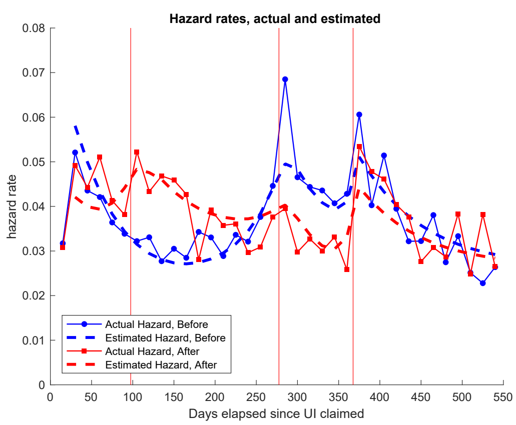
(d) Habit Model with 2 types (Constantinides 1990), AR(1)

Notes: The figure shows the empirical hazards and the predicted hazards of estimates of alternative versions of the reference-dependent model. Panel (a) shows the reference-dependent model where the reference point is updated using a AR(1) process (Table II column 2). Panel (b) shows the reference-dependent model with η estimated and λ set to 1 (Table II column 4). Panel (c) shows the habit formation model of Constantinides (1990) with 2 types and the same reference point as our baseline model (Table II column 6). Panel (d) shows the habit formation model of Constantinides (1990) with 2 types and AR(1) updating of the reference point (Table II column 7).

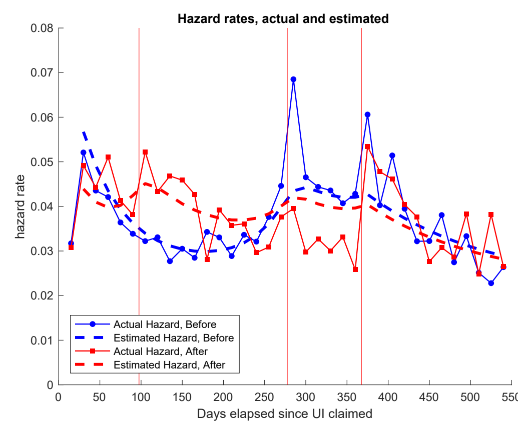
ıtes of the Standard Model







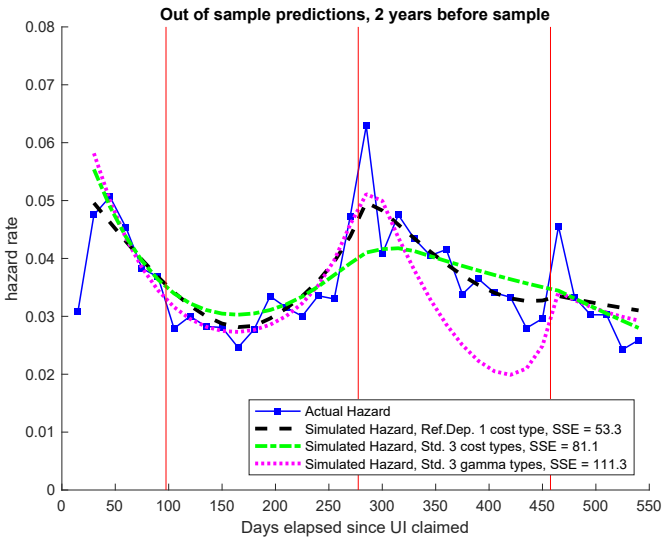
(c) Heterogeneity in Search Cost Curvature



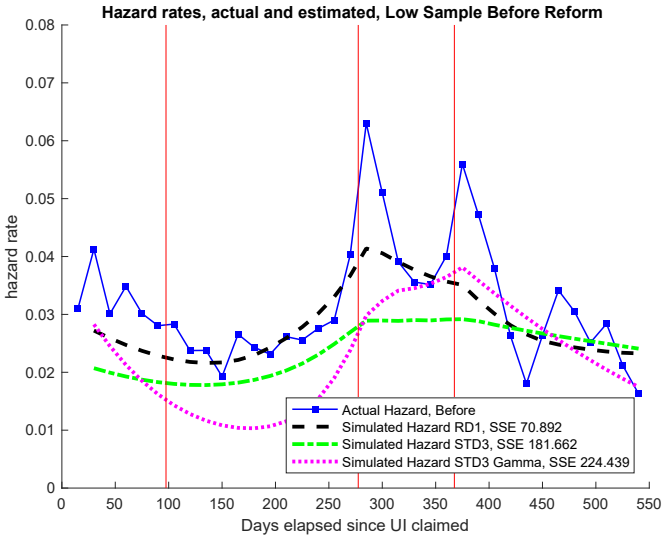
(d) Heterogeneity in Search Cost Curvature (γ restricted to ≥ 0.2)

Notes: The figure shows the empirical hazards and the predicted hazards for alternative estimates of the standard model (See Table III). Panel (a) allows the job starting date to be delayed by one period. Panel (b) allows for 4 cost types. Panel (c) allows for three different types in the elasticity of the cost of job search γ and Panel (d) is the same as Panel (c) but restricting the γ to be larger or equal than 0.2, which would imply an elasticity of search effort with respect to the returns to job search of less than 5.

Figure IX: Out-of-sample Performance of Models



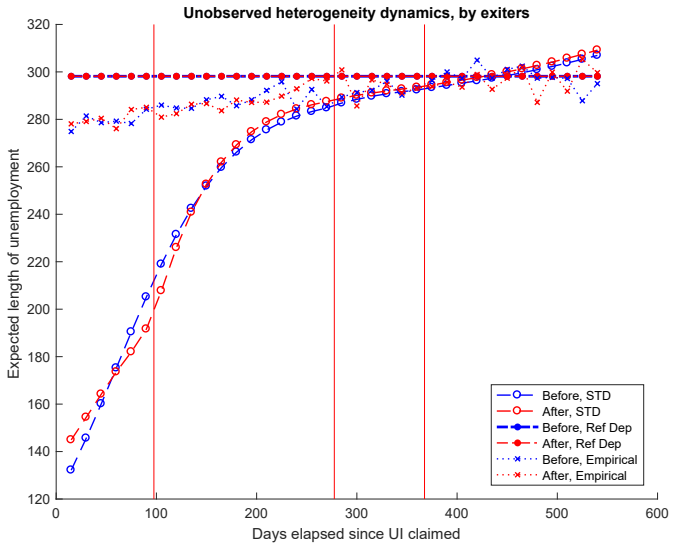
(a) Out-of-sample predictions of models for unemployment system 2 ye



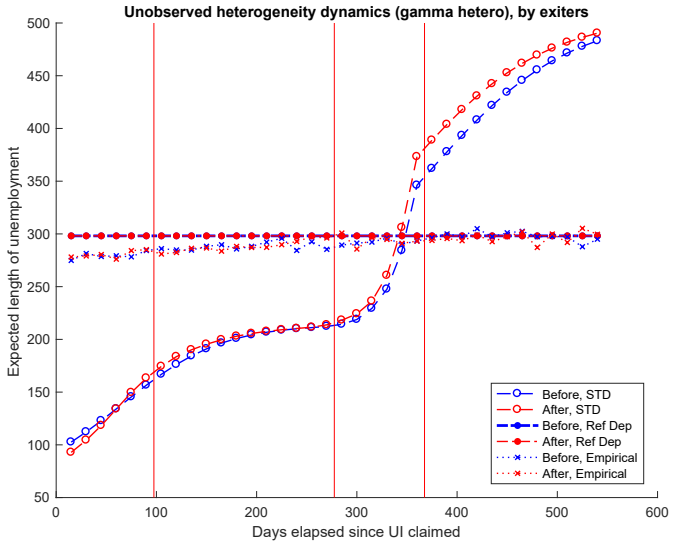
(b) Out-of-sample predictions of models for low earnings sample, pre-reform period

Notes: The figure shows the out of sample fit of the estimated the reference-dependent model with 1 cost type, the standard model with 3 search cost types and the standard model with three γ-types. Panel a) shows the empirical and simulated hazard rates for the period 2 to 1 years before the reform when unemployment assistance could be claimed until 460 days. Panel b) shows the empirical and simulated hazard rates for individuals who had lower pre-unemployment earnings and thus faced a different benefit path (lower benefits during UI), see Web Appendix Figure A-1.

Figure X: Changes in Heterogeneity throughout the Unemployment Spell: Empirical Heterogeneity vs. Model Predictions



(a) Predicted total unemployment duration of individuals exiting



(b) Predicted total unemployment duration of individuals exiting at time t: Heterogeneity in search cost curvature γ

Notes: The figure shows estimates of the expected nonemployment duration of individuals who left unemployment in each time period, contrasting the empirically observed selection with the predicted selection from the estimated standard and reference-dependent models. The empirical expected nonemployment duration (lines with x's) for each individual is calculated as the predicted values from a regression of nonemployment duration on observable characteristics at the time of entering unemployment (see Web Appendix Table A-10). The expected nonemployment durations are calculated for the standard 3 cost type (Table I, Column 1), the standard 3 gamma type (Table III Column 7) and the reference dependent (Table I, Column 4) model.