

Should Unemployment Insurance Vary With the Unemployment Rate? Theory and Evidence

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April 2011*

Abstract

We study theoretically and empirically how optimal unemployment insurance (UI) benefits vary over the business cycle. Theoretically, we characterize how the moral hazard cost and the consumption smoothing benefit of UI vary over the business cycle in a standard job search model. This analysis motivates our empirical strategy which tests whether the effect of UI on unemployment durations and the consumption drop at unemployment varies with the unemployment rate. In our preferred specification, we find that a one standard deviation increase in the state unemployment rate reduces the magnitude of the duration elasticity by 42%. By contrast, using consumption changes upon unemployment, we do not find evidence that the consumption smoothing benefit of UI varies with the unemployment rate. Combining these empirical estimates to calibrate the optimal level of UI benefits implied by our model, we find that a one standard deviation increase in the unemployment rate leads to a roughly 13 to 27 percentage point increase in the optimal wage replacement rate. (JEL H5, J64, J65)

*E-mail: kory.kroft@yale.edu; noto@chicagobooth.edu. We thank Joe Altonji, Judy Chevalier, Jonathan Guryan, Erzo Luttmer, Andrew Metrick, Giuseppe Moscarini, Fiona Scott Morton, and seminar participants at Chicago Booth, Federal Reserve Bank of Philadelphia, Harvard, LSE, MIT, NBER, University of Texas at Austin, University of Toronto, Wisconsin, and Yale for providing excellent comments. We thank Jesse Burkhardt and Christian Goldammer for providing outstanding research assistance. Notowidigdo gratefully acknowledges the National Institute of Aging (NIA grant number T32-AG000186) for financial support.

1 Introduction

It is commonly accepted that higher unemployment benefits prolong unemployment durations (Hamermesh 1977, Moffitt 1985, Meyer 1990, Chetty 2008). Most of the evidence for this “moral hazard effect” comes from empirical studies that do not distinguish between changes in benefits when labor market conditions are good and changes in benefits when labor market conditions are poor. If the moral hazard cost of unemployment insurance (UI) depends on labor market conditions, this may imply that optimal UI benefits should respond to shifts in labor demand. However, many of the studies that conduct a welfare analysis of UI do not consider whether and to what extent UI benefits should vary with labor market conditions (Baily 1978, Hopenhayn and Nicolini 1997, Chetty 2006, 2008, Shimer and Werning 2007, Kroft 2008, Lentz 2009).¹ As Alan Krueger and Bruce Meyer (2002, p64-65) remark:

[F]or some programs, such as UI, it is quite likely that the adverse incentive effects vary over the business cycle. For example, there is probably less of an efficiency loss from reduced search effort by the unemployed during a recession than during a boom. As a consequence, it may be optimal to expand the generosity of UI during economic downturns ... Unfortunately, this is an area in which little empirical research is currently available to guide policymakers.

Similarly, the Congressional Budget Office writes that the availability of long-term unemployment benefits “could dampen people’s efforts to look for work, [but that concern] is less of a factor when employment opportunities are expected to be limited for some time.”²

In this paper, we investigate how the optimal UI benefit level varies over the business cycle. Combining our theoretical model with reduced form empirical evidence, we conclude that a one standard deviation increase in the unemployment rate leads to a roughly 13 to 27 percentage point increase in the optimal UI benefit level. We briefly summarize how we reach this conclusion at the end of this section.

Our main theoretical contribution is to exploit the structure of a standard job search model to derive analytical conditions for how the consumption smoothing benefit of UI (insurance value) and the elasticity of unemployment duration with respect to the UI benefit level (moral hazard cost) vary with the job offer arrival rate. Several studies in the job search

¹Nicholson and Needels (2006) discuss how worsening labor market conditions in the U.S. in the 1970s and 1980s triggered large, policy-driven, increases in benefit payments.

²The CBO quote is available from the following URL: http://www.washingtonpost.com/wp-dyn/content/article/2010/03/08/AR2010030804927_pf.html.

literature derive conditions under which the duration of unemployment varies with the job offer arrival rate (Burdett 1981, Flinn and Heckman 1983, Burdett and Ondrich 1985, Jensen and Vishwanath 1985, Vroman 1985, van den Berg 1994); however, this is the first paper to our knowledge that uses a standard job search model to derive conditions under which the *behavioral responses* to UI vary with the job offer arrival rate. These conditions allow us to explain, in a clear and rigorous way, the economic forces that shape the cyclicity of these effects. Moreover, since our model nests other models in the literature that are used to evaluate optimal UI over the cycle, we are able to sharply delineate the key differences in results across models.³

Our key theoretical finding is that the cyclicity of the unemployment duration elasticity and the cyclicity of the consumption smoothing benefit of UI are both theoretically ambiguous. Each effect may be either procyclical or countercyclical. To see part of the intuition, consider the cyclicity of the duration elasticity. We show that there are two opposing forces in the standard search model that shape how this varies over the cycle. On the one hand, in a downturn, the job offer arrival rate or labor demand is less responsive to an increase in labor supply or search effort. This causes the duration elasticity to be smaller in a recession and is related to the speculation of Krueger and Meyer (2002) above. On the other hand, when labor demand is low, a worker values an increase in the benefit level more, since she expects to collect UI for some time. This acts to increase the duration elasticity in a recession. Which effect dominates depends on the assumptions placed on the structural parameters, as we discuss more fully below. The theoretical ambiguity highlighted by the job search model indicates that how the moral hazard cost and the consumption smoothing benefit of UI vary with labor market conditions is ultimately an empirical question. This motivates our two-part empirical strategy, which directly estimates each of these two terms.

The first part of our empirical contribution is to examine how the elasticity of unemployment duration with respect to the UI benefit level varies with labor market conditions. We do this by estimating a hazard model where the effect of the UI benefit level on unemployment durations is allowed to depend on the state unemployment rate. We find that the elasticity of unemployment durations with respect to the level of unemployment benefits is 0.61 at the average state unemployment rate. This average duration elasticity estimate is similar to the elasticity estimates reported in the literature. Our new empirical result is

³For example, our paper is the only paper in the literature on optimal UI over the cycle that includes reservation wages as a choice variable, in addition to search intensity. Additionally, since we use a dynamic job search model, we can compare and contrast our results to results that are based on a static job search model (Andersen and Svarer 2009).

that the duration elasticity varies with local labor market conditions; specifically, we find that the duration elasticity is statistically significantly lower when the state unemployment rate is relatively high. Furthermore, the magnitude of this interaction effect is economically large: in our preferred specification, a one standard deviation increase in the unemployment rate (an increase of 1.3 percentage points from a base of 6.2%) reduces the magnitude of the duration elasticity from 0.61 to 0.35 (a decline in magnitude of 42%).

The second part of our empirical contribution is to estimate how the consumption smoothing benefit of UI varies with the unemployment rate. We do this by estimating a model where the effect of UI on the consumption change upon unemployment is allowed to depend on the state unemployment rate. In contrast to our duration elasticity results, we do not find evidence that the consumption smoothing benefit of UI varies with the unemployment rate. Our consumption smoothing estimates are both economically and statistically insignificant, and though our statistical power is somewhat limited, we can rule out large effects. Putting these two pieces together, they imply that the moral hazard cost of UI is procyclical while the consumption smoothing benefit of UI is acyclical. These findings form the basis of our conclusion that optimal benefits are decreasing in the unemployment rate.

The identification of both empirical models comes from exploiting variation in UI benefits within states over time interacted with within- and between-state variation in the unemployment rate. We pursue this time-series, cross-sectional research design using state unemployment rates rather than a purely time-series design using the national unemployment rate in order to have sufficient variation in UI benefit levels across a wide range of labor market conditions. While this has the benefit of allowing us to more precisely identify how the duration elasticity varies with the unemployment rate, there are several caveats to extrapolating this estimate to national business cycles which we discuss in detail in Section 4.1 below.

An immediate concern with our empirical strategy is that when the state unemployment rate is high, benefits may (endogenously) become more generous. When benefits respond partially to observed and unobserved labor market conditions, we will consistently estimate our interaction term as long as the correlation between UI benefits and labor market conditions does not vary with the unemployment rate. When this condition is violated, estimates of the interaction term of interest will suffer from endogeneity bias. We pursue three strategies to address this concern.

First, we always measure the local unemployment rate *relative* to the national unemployment rate, and we control for this relative local unemployment rate directly in all specifica-

tions. The use of relative unemployment rates alleviates the concern that UI benefit levels may partially respond to national business cycles. Additionally, if states raise UI benefits in recessions, but do not systematically adjust benefits in good times, then this strategy is preferable to using absolute unemployment rates with year fixed effects, as we discuss below. By controlling for the local unemployment rate in all specifications, we address the concern that benefits may respond endogenously to local labor market conditions.

Second, we directly investigate the association between the state unemployment rate and UI benefits. We find weak but suggestive evidence that the local unemployment rate and the maximum UI benefit level are positively correlated. Moreover, our evidence weakly suggests that the positive association is stronger when the unemployment rate is relatively high. This implies that benefits may respond more strongly to local labor market conditions during bad times. We illustrate with a simple model that this type of differential correlation between the unemployment rate and UI benefits works against our findings. In particular, any omitted variables bias due to this type of policy endogeneity will make the duration elasticity artificially larger during times of high unemployment; however, we find the opposite result: the duration elasticity is *lower* when the local unemployment rate is relatively high.

Third, we investigate several alternative identifying assumptions to gauge the magnitude of omitted variables bias, and we find, if anything, that our results become stronger. First, we include a flexible polynomial in the state unemployment rate, which addresses the concern that benefits may vary non-linearly with the unemployment rate. Second, we include as additional controls the interaction of the state unemployment rate with state fixed effects and year fixed effects. This allows for a more flexible correlation between observable local labor demand shocks and UI benefits. In particular, it captures the possibility that in certain states and/or years, UI benefits may be unusually responsive to changes in labor market conditions. Third, we investigate alternative specifications which allow for unobserved trends across states within a region and within states over time, and we find, if anything, slightly stronger results. Lastly, we find stronger (though less precise) results when we define local labor markets as metropolitan areas (MSAs) rather than states and exploit purely across-MSA, within-state variation in unemployment rates, holding UI benefit levels constant. We therefore interpret our preferred specification as a conservative estimate of the magnitude of the relationship between the duration elasticity and the local unemployment rate. We also investigate a wide variety of alternative explanations for this finding, and we find no consistent evidence that the interaction effect we estimate is primarily determined by compositional changes, endogenous takeup, or bias from using both between-state

and within-state variation in state unemployment rates. Therefore, our interpretation of the duration elasticity results is that they are most consistent with a negative relationship between the moral hazard of cost of UI and the local unemployment rate.

Our third contribution in this paper is to calculate how the optimal UI benefit level varies over the business cycle using our reduced form empirical evidence. Specifically, we derive the Baily-Chetty formula for the marginal welfare gain of UI that illustrates the standard trade-off between the consumption smoothing benefit of UI and the moral hazard cost of UI. Unlike previous work, we explicitly allow for the dependence of each term on the unemployment rate, and we numerically implement this formula using our empirical results. Since we state our formula in terms of our reduced form empirical estimates, our welfare analysis is in the spirit of the “sufficient statistics” approach (Chetty 2009). The primary advantage of this approach is that it does not place any restrictions on the model primitives and is therefore valid for a wide range of underlying mechanisms which cause the duration elasticity and the consumption smoothing benefit to vary with the unemployment rate.⁴ For example, our welfare analysis does not require assuming a specific functional form for how effort is translated into a job offer arrival rate, which – as we show below – can affect how the duration elasticity varies over the cycle. We also do not need to separately identify how reservation wages and search effort respond to a change in benefits and how these behavioral responses vary over the cycle, nor do we need to make specific assumptions about liquidity effects.

Combining our reduced form empirical estimates to calibrate the optimal UI benefit level implied by our model, we find that a one standard deviation increase in the local unemployment rate leads to a roughly 13 to 27 percentage point increase in the optimal replacement rate, depending on the coefficient of relative risk aversion used in the calibration.⁵ To give a sense of the magnitude of this change, it is roughly equivalent to the change in the optimal UI benefit level stemming from a one unit change in the coefficient of relative risk aversion (e.g., from $\gamma = 3$ to $\gamma = 4$), holding constant the duration elasticity and the effect of UI on the consumption drop at unemployment.

Our results suggest a countercyclical UI policy. This is broadly consistent with the observed UI policy in the U.S., which is based on extending the number of weeks for which an unemployed worker can claim benefits – typically 26 weeks in the United States. We

⁴Chetty (2009) describes the advantages and disadvantages of the sufficient statistics approach in more detail.

⁵We report results across a range of CRRA values from $\gamma = 2$ to $\gamma = 4$.

show how one can use our elasticity estimates to shed light on the optimality of current UI policy in the US, and we also provide an illustration of how one may use our empirical results to shed light on how extending benefits in a recession affects the aggregate unemployment rate.

Several papers in the literature have explored optimal UI over the business cycle theoretically. Kiley (2003) and Sanchez (2008) impose particular functional forms on the job finding probability to ensure search effort and a variable affecting the job offer arrival rate are highly complementary.⁶ Under these functional forms, UI benefits are more distortionary in good times than bad times, and – as a consequence – optimal UI benefits are unambiguously countercyclical. Andersen and Svarer (2009) consider a static model and impose a similar functional form assumption on the job finding probability. They incorporate UI financing requirements and show that if the government budget must balance in each state, benefits could be procyclical due to a “budget effect.” Our contribution relative to these papers is to consider a more general dynamic search model with stochastic wage offers, as in Shimer and Werning (2007). This framework allows us to shed light on several new dimensions of the optimal UI problem. First, the model permits us to characterize the cyclical behavior of the behavioral responses of both search effort and reservation wages. This is important if in practice individuals set a reservation wage in addition to choosing search intensity. Second, since our model nests other search models used in the literature, we can use the model to zoom in on the distinctions between them. For example, our results highlight that the response of search effort to UI benefits over the cycle is pinned down by three factors – (1) a static effect, (2) a dynamic effect and (3) a reservation wage effect – and we show that these effects may go in opposite directions. Previous studies have not highlighted this distinction since they considered special cases of our model (typically a static model without reservation wages).

Another strand of the literature has begun to explore optimal UI over the cycle in a general equilibrium framework. Andersen and Svarer (2010) consider a stylized general equilibrium model, and they demonstrate that allowing for changes in the business cycle situation changes how the distortion to effort created by UI varies over the cycle, since search effort depends on anticipated changes in the labor market. Another general equilibrium approach is Landais, Michaillat, and Saez (2011), who consider a matching model with search effort and focus on characterizing the optimal benefit level over the cycle. The

⁶The dynamic, discrete-time, search effort models in these papers build on the work of Hopenhayn and Nicolini (1997).

primary innovation in this paper is the introduction of endogenous job rationing coming through the combination of diminishing marginal returns to production and wage rigidity. They derive a version of the Baily-Chetty formula for optimal UI in terms of a “micro” and “macro” elasticity, the latter capturing the direct effect of a change in UI benefits on search and the indirect effect that arises via changes in the aggregate job finding rate. Though our model is a partial equilibrium job search model, it can be reinterpreted as a general equilibrium model following Rogerson, Shimer, and Wright (2005). More specifically, one can interpret our model as the Landais et al. model with the addition of reservation wages and the elimination of job rationing (which would be obtained by assuming constant returns to scale in production, for example).

Finally, while we focus on the optimal *level* of UI benefits over the business cycle, contemporaneous research by Schmieder, von Wachter and Bender (2011) explore theoretically and empirically the optimal *potential duration* of UI benefits over the cycle using unique administrative data from Germany. We provide a detailed discussion of the differences between the findings in this paper and our findings in Section 3.1.5 below. Overall, we view our work which focuses on the optimal benefit level as highly complementary to work which focuses on optimal potential duration, and an important task in future work will be to investigate the problem of jointly choosing the optimal benefit level and potential duration over the business cycle.

In summary, to our knowledge this is the first paper to estimate how the optimal UI benefit level varies over the business cycle by calibrating an optimal UI model using reduced form empirical estimates of how the consumption smoothing benefit and moral hazard cost of UI varies with the unemployment rate.

The remainder of the paper proceeds as follows. The next section develops the search model and describes both the agent and planner problems. Section 3 presents our empirical analysis which estimates how the duration elasticity and consumption smoothing benefit of UI vary with the unemployment rate. Section 4 considers the welfare implications of our empirical findings. Section 5 concludes.

2 Theory

In this section, we describe the setup of a standard continuous-time, infinite-time horizon, job search model. The model closely follows Shimer and Werning (2007). The objective of the theory is primarily to provide economic intuition for how UI affects job search in-

centives and how these incentives vary over the business cycle. While previous studies have examined the behavioral responses to UI (Mortensen 1977), few have considered how these behavioral responses vary over the cycle, and our study is novel in this respect. While the welfare analysis that follows is less model-dependent, we nevertheless feel that it is a useful and important exercise to derive and study the comparative statics for how the behavioral responses vary over the cycle.

2.1 Assumptions

We make several assumptions. First, we focus on benefit level, not potential benefit duration, although the latter is clearly also an important policy parameter.⁷ Second, workers consume hand-to-mouth. The only way an unemployed worker can smooth consumption across states is through publicly provided unemployment insurance. We make this assumption since it allows us to characterize analytically how the behavioral responses to UI vary over the cycle. Importantly, our sufficient statistics welfare analysis is valid even when workers have access to liquidity, a point we discuss below. Third, there is no value from leisure time during an unemployment spell.⁸ Fourth, workers are homogeneous. Finally, we work in a partial equilibrium setting focusing on the worker’s problem.

2.2 Agent’s Problem

Consider a single worker with flow utility $U(c)$, where $U' > 0$, $U'' < 0$. The worker’s subjective discount rate is $\rho \geq 0$. The worker maximizes the expected present value of utility from consumption

$$E_0 \int_0^{\infty} e^{-\rho t} U(c(t)) dt \quad (1)$$

If the worker is unemployed, she receives unemployment benefit b and samples wage offers from a known distribution function, $F(w)$. Wage offers arrive randomly and individuals affect the offer arrival rate through costly search effort, e . Following Andersen and Svarer (2009), we assume a linear, separable cost of search, denoted by $\psi(e)$. The offer arrival rate $\lambda(e, \alpha)$ has the following properties: $\lambda_1 \geq 0$, $\lambda_{11} \leq 0$, $\lambda_{12} \geq 0$ and $\lambda_2 \geq 0$. We

⁷Shimer and Werning (2008) find that socially optimal UI policy is infinite duration, constant benefits in a model with free access to savings and lending and CARA preferences.

⁸We relax this assumption in Extension 1 in the Appendix.

characterize business cycles as shifts in labor demand. Variation in labor demand comes via the parameter α , which proxies for productivity.⁹

Workers who accept a wage offer commence employment immediately. When the worker is employed, she earns a wage w and pays taxes τ which are used to finance unemployment benefit payments. Consumption when employed is her net wage, $w - \tau$. Employment is assumed to end exogenously with separation rate s .

Finally, the structural parameters s , $F(w)$, b , τ and ρ are independent of time. All duration dependence in the model comes from shifts in α , which we assume are unanticipated by the individual; that is, she always thinks that it will remain constant at its current value. As Van Den Berg (2001) notes, this is likely to be a reasonable assumption in the case of a random macroeconomic shock.¹⁰ For ease of notation, we suppress the dependence of α on time, t . Empirically, we do not directly observe α . Instead, in our empirical specifications, we allow α to vary across individuals and over time (but not within spells) by using variation in the unemployment rate at the start of an individual's unemployment spell. We describe this in more detail in our empirical section below. It is a well-known result that workers adopt a reservation wage strategy accepting wage offers above the reservation wage, $\bar{w} = \bar{w}(b, \tau)$. We refer the reader to the Appendix for the characterization of the agent's reservation wage and the agent's optimal search intensity.

2.3 Elasticity Concepts

For ease of notation, we define $\theta(\bar{w}) \equiv \frac{f(\bar{w})}{F(\bar{w})}$, the hazard rate (or failure rate) of the wage offer distribution, evaluated at the reservation wage, where $\bar{F}(w) \equiv 1 - F(w)$. Next, let $\delta(e, \alpha) \equiv \frac{d \log \lambda(e, \alpha)}{de}$ represent the *percentage change* in the job offer arrival rate from an additional unit of search, evaluated at the optimal search effort. In this model, expected duration is the inverse of the job finding probability, $D \equiv 1/[\lambda(e, \alpha)\bar{F}(\bar{w})]$. Define the total elasticity of expected unemployment duration with respect to the UI benefit level as $\varepsilon \equiv \frac{d \log D}{d \log b}$. We can conveniently express the duration elasticity as:

$$\varepsilon = \varepsilon_{\bar{w}} + \varepsilon_e \tag{2}$$

⁹In Extension 3, we consider business cycles driven by changes in $F(w)$.

¹⁰An alternative formulation is the case where α is permitted to vary in a deterministic way over an individual's unemployment spell. Individuals correctly anticipate a change in α . See Van Den Berg (1990) for an example.

The first term in (2), $\varepsilon_{\bar{w}} \equiv \theta(\bar{w}) \times \bar{w} \times \frac{d \log \bar{w}}{d \log b}$, is the duration elasticity in a reservation model with exogenous job offer arrivals (Shimer and Werning 2007). The second term in (2), $\varepsilon_e \equiv -\delta(e, \alpha) \times e \times \frac{d \log e}{d \log b}$, is the duration elasticity in a search effort model with a fixed wage (Chetty 2008). This decomposition is useful in characterizing how the duration elasticity varies over the cycle.

2.4 Planner's Problem

In this section, we consider the optimal unemployment insurance problem. Our approach is to solve for the optimal level of UI in a given labor market state. We then focus on the question of how optimal UI varies over the cycle. A social planner maximizes an unemployed worker's expected lifetime utility, V_u . We restrict the class of feasible policies to those where the unemployment benefit level, b , and the employment tax, τ , are constant. The planner's policy must satisfy a balanced-budget requirement which means that expected benefits paid out equals expected taxes collected. We assume throughout that $r = \rho$, which is a standard assumption in the literature. The social planner's problem is stated formally as:

$$\begin{aligned} & \max_{b, \tau} V_u(b, \tau) \\ & \text{s.t. } D(b, \tau(b))b = \frac{\tau}{r + s} \end{aligned}$$

The following proposition characterizes the money-metric marginal welfare gain of increasing benefits by \$1.¹¹

Proposition 1 *With $r = \rho = 0$, the money-metric welfare gain of raising b is given by*

$$\frac{dW}{db} = \frac{u}{1 - u} \left\{ \frac{U'(b) - E[U'(w - \tau) | w \geq \bar{w}]}{E[U'(w - \tau) | w \geq \bar{w}]} - \varepsilon \right\} \quad (3)$$

At the optimum,

$$\frac{U'(b) - E[U'(w - \tau) | w \geq \bar{w}]}{E[U'(w - \tau) | w \geq \bar{w}]} = \varepsilon \quad (4)$$

Proof. See Appendix. ■

This is the standard ‘‘Baily-Chetty condition’’ of optimal unemployment insurance (Baily 1978, Chetty 2006). If at current levels of UI benefits, $dW/db > 0$, raising the UI benefit

¹¹We solve the planner's problem by substituting the budget constraint $\tau(b)$ into the objective function and solve the unconstrained problem.

level increases welfare. Thus, dW/db can be viewed as a “test” for whether current benefit levels are optimal. The test for the optimality of UI benefits illustrates the standard trade-off between the insurance role of UI benefits against the disincentive effect. Moral hazard arises in the second-best world, since agents do not internalize the planner’s balanced-budget constraint. Thus, they impose an externality on the planner’s budget, which is captured by the elasticity of expected duration with respect to UI benefits, ε .¹² Importantly, Chetty (2006) shows that equations (3) and (4) are robust to variation in the financial environment. In particular, it does not matter if agents consume hand-to-mouth or have access to liquidity. This result carries over to our setting; as a result, our welfare analysis will be valid, even with savings and borrowing.

2.5 Business Cycle Dependent UI

We now study how the optimal benefit level, characterized in equation (4) varies over the business cycle. First, we consider how the moral hazard cost of UI varies over the cycle.

2.5.1 Duration Elasticity Over the Cycle

Equation (4) demonstrates that the moral hazard cost of UI is determined by the magnitude of the total duration elasticity, ε . To see how ε varies over the cycle, we start by differentiating the agent’s reservation wage and optimal search intensity with respect to b , holding taxes constant, to obtain the following lemma¹³:

Lemma 1 *The marginal effects with endogenous search intensity satisfy*

$$\frac{\partial \bar{w}}{\partial b} = \frac{U'(b)}{U'(\bar{w} - \tau)} u > 0 \quad (5)$$

where $u \equiv \frac{\rho+s}{\rho+s+\lambda(e,\alpha)\bar{F}(\bar{w})}$.

$$\frac{\partial e}{\partial b} = \left[\frac{d \log \lambda_1(e, \alpha)}{de} \right]^{-1} \frac{U'(b)}{E[U(w - \tau) | w \geq \bar{w}] - U(b) + \psi(e)} < 0 \quad (6)$$

where $\lambda_1(e, \alpha) \equiv \frac{\partial \lambda}{\partial e}$.

¹²We show in Extension 1 in the Appendix that the basic structure of formula (3) remains intact if unemployment has leisure value and/or there are monetary costs of search that reduce consumption.

¹³We assume that $\frac{\partial \bar{w}}{\partial \tau} = \frac{\partial e}{\partial \tau} = 0$ and focus on the partial derivatives of a benefit change. We do this for analytical tractability and we have verified in simulations that this has a negligible quantitative impact on the results.

Proof. See Appendix. ■

Corollary 1 *In a fixed wage, **dynamic** search effort model, the marginal effect for search effort is*

$$\frac{\partial e}{\partial b} = \left[\frac{d \log \lambda_1(e, \alpha)}{de} \right]^{-1} \frac{U'(b)}{U(w - \tau) - U(b) + \psi(e)} < 0 \quad (7)$$

Corollary 2 *In a fixed wage, **static** search effort model, the marginal effect for search effort is*

$$\frac{\partial e}{\partial b} = \left[\frac{d \log \lambda_1(e, \alpha)}{de} \right]^{-1} \frac{U'(b)}{U(w - \tau) - U(b)} < 0 \quad (8)$$

There are several things worth noting about equations (5) and (6). First, higher UI benefits raise reservation wages. Second, since a higher reservation wage is associated with a lower search intensity, we see that higher benefits reduce search effort. Third, for $\rho \approx 0$, the behavioral responses depend on the *predetermined* unemployment rate, u , that prevails at the time of the benefit increase; consequently, we use the predetermined unemployment rate in our empirical framework below. In addition to the theoretical motivation, this is advantageous from an econometrics standpoint, since it avoids the “reflection problem”, a simultaneity bias that is caused by the fact that an increase in UI benefits increases unemployment duration and hence, the unemployment rate. Fourth, equation (5) shows that when $U'' \approx 0$, $\frac{\partial \bar{w}}{\partial b} \approx u$, we can identify the responsiveness of the reservation wage to changes in benefits simply by measuring the unemployment rate that prevails at the time of the benefit increase.¹⁴ Lastly, one can easily see how the search effort response to benefits generalizes from a fixed wage, static search effort model (equation (8)), to a fixed wage, dynamic search effort model (equation (7)) to a dynamic reservation wage model (equation (6)). This will help in characterizing the cyclical behavior of $\partial e / \partial b$ and relating this behavior to earlier results in the literature. In particular, note that equation (8) is exactly the expression in the static search effort model of Andersen and Svarer (2009).

¹⁴For the U.S. over the period 1999-2009, $u \in [3.8\%, 10.1\%]$. On the other hand, Feldstein and Poterba (1984) empirically estimate, using variation in UI benefits and reservation wages, $\frac{\partial \bar{w}}{\partial b} \in [13\%, 42\%]$. Therefore, in order to reconcile this range of estimates with the model-based expression for $\frac{\partial \bar{w}}{\partial b}$, it must be the case that risk aversion is relevant at existing UI benefit levels. While others have derived similar expressions for $\frac{\partial \bar{w}}{\partial b}$ (e.g., Chesher and Lancaster 1983) we believe that this is a novel point that has not been highlighted in the literature on job search.

The next proposition characterizes the cyclical behavior of $\frac{\partial \bar{w}}{\partial b}$ and $\frac{\partial e}{\partial b}$.¹⁵ Denote the coefficient of relative risk aversion evaluated at the after-tax reservation wage by $\gamma(\bar{w} - \tau)$. For arbitrary x and y , let $\varepsilon_{x,y} \equiv \frac{d \log x}{d \log y}$.

Proposition 2 *Cyclical Behavior of Marginal Effects.*

$$(i) \quad \frac{d}{d\alpha} \left[\frac{\partial \bar{w}}{\partial b} \right] > 0 \iff \varepsilon_{u,\alpha} + \gamma(\bar{w} - \tau) \varepsilon_{\bar{w},\alpha} > 0 \quad (9)$$

$$(ii) \quad \frac{d}{d\alpha} \left[\left[\frac{\partial e}{\partial b} \right] \right] > 0 \iff \varepsilon_{\frac{d \log \lambda_1}{de}, \alpha} + \varepsilon_{\rho+s+\lambda \bar{F}, e} \times \varepsilon_{e,\alpha} + \frac{u}{1-u} \times \varepsilon_{\rho+s+\lambda \bar{F}, \bar{w}} \times \varepsilon_{\bar{w},\alpha} < 0 \quad (10)$$

Proof. See Appendix. ■

2.5.2 Interpretation and Discussion

Condition (9) illustrates that two effects determine the cyclicity of the responsiveness of reservation wages to UI benefits:

Discount Rate Effect ($\varepsilon_{u,\alpha} < 0$) – The key thing to note is that u effectively represents how the agent weights or discounts an increase in the benefit level b in all periods. An increase in λ (via α) effectively boosts the agent’s discount rate, making her less responsive to variation in her future income stream. Intuitively, when u is low, the agent does not expect to be unemployed in the future and thus, attaches little weight to future UI benefits. It is therefore not surprising that if an individual is less likely to be unemployed in the future, an increase in b has a smaller effect on her current reservation wage. This expression also explains why it does not matter if we model the business cycle as coming from shifts in s or λ , as it is variation in u that is important.

Risk Aversion Effect ($\gamma(\bar{w} - \tau) \varepsilon_{\bar{w},\alpha} > 0$) – An increase in λ raises the option value of search and therefore increases \bar{w} , as agents become more choosy about which jobs to accept when times are good. If the agent’s utility function is highly curved at the existing reservation wage, an increase in \bar{w} can substantially lower the marginal utility of consumption

¹⁵For expositional purposes, we focus on exogenous variation in α , but the conditions we provide are essentially identical if we consider exogenous variation in the separation rate, s . Intuitively, this is because increases in α and decreases in s have similar effects on the agent’s behavior. In Extension 3 in the Appendix, we derive analogous results for a change in the mean of the wage offer distribution. We show that the conditions are very similar to the conditions we derive in this section.

for the employed relative to the unemployed. This exacerbates the reservation wage response to UI benefits.

Thus, whether reservation wages respond more or less to benefits in a downturn depends on the strength of the “discount rate effect” to the “risk aversion effect.” Interestingly, the model predicts a *larger* behavioral response during recessions for conservative levels of risk aversion, in contrast to the hypothesis of Krueger and Meyer discussed in the introduction.

Condition (10) again illustrates a tension between several opposing economic forces in shaping how search effort varies with UI benefits over the cycle:

Static Effect ($\varepsilon_{\frac{d \log \lambda_1}{de}, \alpha} < 0$) – This term reflects the degree of complementarity between e and α . This expression shows that if search effort complementarities are sufficiently strong ($\lambda_{12} \gg 0$), then benefits are more distortionary in good times, consistent with the Krueger and Meyer intuition. Intuitively, a boost to labor demand raises the marginal efficiency of search and the behavioral response to UI benefits.¹⁶ Andersen and Svarer (2009) consider a static, fixed wage model of search with a linear search cost and show that the cyclicity of $\partial e / \partial b$ depends entirely on the sign of this effect; thus, we label it a “static effect.”

Dynamic Effect ($\varepsilon_{\rho+s+\lambda, e} \times \varepsilon_{e, \alpha} > 0$) – This term arises in a fixed wage, dynamic search effort model. In a dynamic model, a permanent increase in benefits raises the value of unemployment in all future periods. The agent’s behavioral response in a dynamic model is pinned down by the discounted value of this increase. A positive and permanent labor demand shock lowers the probability that the agent is unemployed in all future periods, since it raises search effort. This acts to effectively reduce the agent’s discount factor and attenuates the behavioral response to UI benefits.

Reservation Wage Effect ($\frac{u}{1-u} \times \varepsilon_{\rho+s+\lambda \bar{F}, \bar{w}} \times \varepsilon_{\bar{w}, \alpha} > 0$) – Finally, in a model with stochastic wages, one also needs to additionally account for the effect of benefits on reservation wages. When the arrival rate of job offers is high, agents search with a higher reservation wage and this acts to reduce the discount factor and attenuate their behavioral response.

The main result of this section is to show that how the marginal effects vary over the cycle depends on the precise specification and structural parameters of the search model. We now study the consequences for the elasticity of expected duration with respect to the benefit level. Let $\frac{d\varepsilon}{d\alpha} = \frac{d\varepsilon_{\bar{w}}}{d\alpha} + \frac{d\varepsilon_e}{d\alpha}$. We consider each of these in turn.

¹⁶Card & Levine (2000) speculate that in good times, the opportunity cost of cutting back on job search activities may be too great for many and that this would tend to *attenuate* the behavioral response to UI. Our results suggest that although search effort is high in good times, the behavioral response is *exacerbated*.

Proposition 3 *Cyclical Behavior of $\varepsilon_{\bar{w}}$ and ε_e .*

(i)

$$\frac{d\varepsilon_{\bar{w}}}{d\alpha} > 0 \iff \varepsilon_{u,\alpha} + (\gamma(\bar{w} - \tau) + \varepsilon_{\theta,\bar{w}}) \varepsilon_{\bar{w},\alpha} > 0 \quad (11)$$

(ii)

$$\frac{d|\varepsilon_e|}{d\alpha} > 0 \iff \varepsilon_{\frac{d \log \lambda_1}{de},\alpha} + \varepsilon_{\rho+s+\lambda\bar{F},e} \times \varepsilon_{e,\alpha} + \frac{u}{1-u} \times \varepsilon_{\rho+s+\lambda\bar{F},\bar{w}} \times \varepsilon_{\bar{w},\alpha} - \varepsilon_{\delta,e} \times \varepsilon_{e,\alpha} < 0 \quad (12)$$

Proof. See Appendix. ■

Comparing condition (11) to condition (9), we see that there is an additional term $\varepsilon_{\theta,\bar{w}}$ which reflects how the hazard rate of the wage offer distribution varies with the wage, evaluated at the reservation wage. According to Van den Berg (1994), most of the distributions used in structural job search analysis have hazards that are decreasing in the wage, $\varepsilon_{\theta,\bar{w}} < 0$. This would tend to attenuate the “risk aversion effect” and therefore make $\varepsilon_{\bar{w}}$ more likely to be countercyclical.

Next, comparing condition (12) to condition (10), we see that the condition is augmented by the factor $\varepsilon_{\delta,e} \times \varepsilon_{e,\alpha}$. Under the stated assumptions on $\lambda(e, \alpha)$, $\varepsilon_{\delta,e} < 0$. This acts to make ε_e more likely to be countercyclical.

In summary, we see that taking into account how the technical aspects of $\varepsilon_{\bar{w}}$ and ε_e vary with the cycle (e.g., how $\theta(\bar{w})$ and $\delta(e)$ vary with α) make *both* of them more likely to be countercyclical. It follows that the total duration elasticity, ε , is also likely to be more countercyclical. Ultimately, whether ε is procyclical or countercyclical depends on the precise specification of the primitives and functional forms of the model. Figures 1 and 2 illustrate this ambiguity by calibrating two versions of the job search model in this section. Figure 1 shows that the duration elasticity is increasing in the unemployment rate in a model with an exogenous arrival rate and stochastic wage offers, while figure 2 shows that the duration elasticity is decreasing in the unemployment rate in a model with a fixed wage and an endogenous arrival rate that depends on search effort. The figures suggest that whether the duration elasticity increases or decreases with the unemployment rate ultimately depends on the parameterization of the search model. We next analyze how the insurance value of UI varies over the cycle.

2.5.3 Consumption Smoothing Over the Cycle

Define $\bar{g} = \frac{U'(b)}{E[U'(w-\tau)]_{w \geq \bar{w}}}$ as the money-metric amount such that, the government is indifferent between giving \$1 to someone who is unemployed and \bar{g} to someone who is employed.

Proposition 4 *Cyclicalities of Insurance Value of UI.*

$$\frac{d\bar{g}}{d\alpha} = \frac{d\bar{g}}{d\bar{w}} \frac{d\bar{w}}{d\alpha} + \frac{d\bar{g}}{d\tau} \frac{d\tau}{d\alpha} > 0$$

Proof. See Appendix. ■

Reservation Wage Effect ($\frac{d\bar{g}}{d\bar{w}} \frac{d\bar{w}}{d\alpha} > 0$) – The first term comes from the fact that the reservation wage varies over the business cycle. Intuitively, an agent searching for a job with a higher reservation wage expects a higher wage during employment and thus values insurance more.¹⁷ This effect is positive and therefore calls for procyclical UI benefits. Clearly, in a search effort model with a fixed wage, this term will not appear. Thus, relative to Andersen and Svarer (2009), our results suggest a new way that the insurance effect can vary over the cycle.

Budget Effect ($\frac{d\bar{g}}{d\tau} \frac{d\tau}{d\alpha} > 0$) – The second term represents an effect that operates through the balanced-budget constraint. Intuitively, when α is high, fewer taxes need to be raised to finance a given level of benefits. This increases the marginal utility of consumption for the employed relative to marginal utility of consumption for the unemployed since the net wage is $w - \frac{u}{1-u}b$; in order to restore optimality, benefits need to be increased. This effect is positive and ceteris paribus, implies that benefits should be procyclical. In Extension 2 in the Appendix, we relax the budget balance condition and allow the planner to run deficits in bad times and surpluses in good times. We show that this eliminates the budget effect and the insurance value of UI only depends on the reservation wage effect. In summary, we see that both the reservation wage effect and the budget effect imply that the insurance value of UI varies positively with α in this model.

Liquidity Effect – As discussed above and in relation to Chetty (2008), UI can be very valuable if it affects agents’ search behavior primarily through a non-distortionary “liquidity effect”, rather than through a substitution effect. It is possible that in a bad labor market,

¹⁷Note that this does not say that individuals with higher reservation wages value publicly provided insurance more. In practice, workers with access to liquidity will tend to have higher reservation wages and better means to smooth consumption. Rather, this statement applies to two similar individuals, with one having a higher reservation wage solely because his job opportunities are more plentiful.

workers have fewer means to smooth consumption; for instance, one’s spouse may also be out of work, and so a secondary source of income cannot be relied on. In this case, UI would be more valuable in a downturn since it has a larger liquidity effect. We develop this intuition more formally in Extension 4 where we consider a fully credit constrained model where individuals set consumption equal to income in each period.¹⁸ We assume individuals have access to exogenous, non-labor income A each period, in addition to the UI benefit or the net wage. We show that the consumption smoothing benefit of UI – the left-hand side of equation (4) – is identified by the ratio of the liquidity effect in search effort to the substitution effect in search effort. This shows that how the consumption smoothing benefit of UI in our model varies over the cycle additionally depends on how A varies with α . We explore this question empirically below and consider the implications for our welfare analysis.

In summary, both the reservation wage effect and the budget effect imply that the insurance value of UI varies positively with α in this model, while the liquidity effect is theoretically ambiguous. This implies that the insurance value of UI is theoretically ambiguous and also implies that the optimal UI benefit level over the cycle is theoretically ambiguous. Figures 1 and 2 illustrate this by showing the optimal UI benefit level as a function of the unemployment rate. As above, whether the optimal benefit level increases or decreases with the unemployment rate depends on the precise specification and specific parameters of the model. The figures also show that the magnitude of the relationship between the optimal UI benefit level and the unemployment rate can be quantitatively large.

We conclude that for realistic parameter values, the duration elasticity and the optimal benefit level can vary substantially over the cycle. Our calibration results also suggest that, in contrast with some claims in the literature, the reservation wage model and the fixed-wage, endogenous search effort may have very different normative implications when considering how UI should optimally vary over the cycle.¹⁹ This theoretical ambiguity motivates the two-part empirical analysis in this paper, which estimates how the duration elasticity and the consumption smoothing benefit vary with labor market conditions.

¹⁸Results from Card, Chetty and Weber (2007) suggest this model more accurately describes job search behavior than a perfect consumption smoothing model.

¹⁹For example, Lentz and Traenes (2005) write that “We do not believe that it is crucial whether the problem is formulated as a choice of reservation wage given a fixed search intensity or (as here) as a choice of search intensity given a fixed wage.” While there are many settings where this is true, our calibration results in this section suggest that when studying the *interaction* between optimal UI and the unemployment rate, this modeling choice is not innocuous.

3 Empirical Analysis

The theoretical model above predicts that the unemployment duration elasticity (ε) and the insurance effect (\bar{g}) vary with labor market conditions (α), but the sign and magnitude of these comparative statics are theoretically ambiguous. To take the model to the data, we make three important assumptions.

First, we assume that the *predetermined* unemployment rate (u) at the start of an unemployment spell is a valid proxy for α . Under this assumption, the relationship between (ε, \bar{g}) and u provides information on the relationship between (ε, \bar{g}) and α . Moreover, using the predetermined unemployment rate – as opposed to the actual unemployment rate at a given time during an unemployment spell – alleviates the concern that the unemployment rate is endogenous to the UI benefit level, as discussed above.

Second, we assume that α is constant within an unemployment spell. This assumption is motivated by the fact that virtually all of the variation in unemployment rates is across-spell variation, with negligible within-spell variation.²⁰ Under this assumption, we identify how ε and \bar{g} vary with α by estimating how ε and \bar{g} vary with the unemployment rate. This assumption also implies that individuals do not need to anticipate how α will evolve during an unemployment spell.

Lastly, we rely on variation in unemployment rates between and within states, which implicitly assumes that the relevant local labor market conditions are proxied by the state-level unemployment rate. As discussed above, we pursue this time-series, cross-sectional research design in order to have sufficient variation in UI benefit levels across a wide range of labor market conditions. The primary benefit of this strategy is that we are able to more precisely identify how ε and \bar{g} vary with labor market conditions. We discuss several caveats to extrapolating our estimates over national business cycles in the next section.

3.1 Part 1: Duration Elasticity

The first part of the empirical analysis is to estimate how the duration elasticity varies with the unemployment rate. We present two pieces of evidence: (1) graphical evidence and nonparametric tests of survival curves and (2) semi-parametric estimates of proportional hazard models (Cox models). The empirical strategy closely follows Chetty (2008). Both strategies focus on estimating the effect of UI benefit levels on expected unemployment

²⁰A variance decomposition of monthly local unemployment rates reveals that 98% of the variance is between-spell and 2% is within-spell.

duration. All empirical specifications will focus on the benefit level rather than the maximum potential duration that an individual can receive UI benefits, although the latter is clearly also an important policy parameter. Our theory focuses on the optimal benefit level in a model with infinite potential duration, and our results are robust to controlling for variation in maximum potential duration that may be correlated with changes in UI benefit levels.²¹

We use unemployment spell data from the Survey of Income and Program Participation (SIPP) spanning 1985-2000. We impose the same sample restrictions as in Chetty (2008): we focus on prime-age males who (1) report searching for a job, (2) are not on temporary layoff, (3) have at least three months of work history, and (4) took up UI benefits. We also follow Chetty (2008) and censor unemployment spells at 50 weeks. We use several alternative proxies for an individual’s actual UI benefits: (1) average benefits for each state-year pair, (2) the maximum weekly benefit amount, (3) the average replacement rate, and (4) a simulated UI benefit variable constructed by using a fixed national sample for each state-year in the sample (Currie and Gruber 1996). All proxies (and all nominal dollar values in the data) are adjusted to real dollars using the 2000 CPI-U series.

3.1.1 Graphical evidence and nonparametric tests

We begin by providing graphical evidence on the effect of unemployment benefits on durations. We split the sample into two sub-samples according to whether individuals began their unemployment spell in states with above-median unemployment rates or in states with below-median unemployment rates, where each year we define the median unemployment rate across states that year. We then assign monthly state unemployment rates to unemployment spells based on the unemployment rate in the state that the individual resided in when his spell began. Lastly, we categorize unemployment spells based on whether the prevailing UI benefit level at the start of the spell in a given state and year is above or below the median UI benefit level across the sample.

Figures 3 and 4 show the effect of UI benefits on the probability of unemployment for individuals in above-average and below-average unemployment state-years, respectively. In each figure, we plot Kaplan-Meier survival curves for individuals in low-benefit and high-benefit states.²² The results in figure 3 show that the curves are fairly similar in both

²¹Specifically, we estimate models where the duration elasticity is only identified off of the weeks before benefits expire, and we find very similar results.

²²Following Chetty (2008), the plotted curves are adjusted for the “seam effect” in the SIPP panel data, but the test that the survival curves are identical is fully nonparametric and does not make this adjustment.

low-benefit and high-benefit states when the unemployment rate in a state-year is above the median unemployment rate. The curve in high-benefit states is slightly higher, indicating that UI benefits may marginally increase benefits, but a nonparametric test that the curves are identical does not reject at conventional levels ($p = 0.932$).²³ By contrast, in figure 4 the curves are noticeably different; in particular, durations are significantly longer in high-benefit states, and the difference between the survival curves is strongly statistically significant ($p = 0.007$).²⁴

These figures suggest that the moral hazard cost of UI benefits depends crucially on whether unemployment is high or low. In particular, our findings suggest that the effect of UI benefits on durations is not statistically significant when the unemployment rate is high but is strongly statistically significant when the unemployment rate is low.²⁵ These effects are based on simple comparisons across spells. It is possible, however, that the characteristics of individuals vary with unemployment rate in a way that would bias these effects. To investigate this issue and other potential biases, as well as to quantify the magnitude of this interaction effect, the next subsection reports results from the estimation of semi-parametric proportional hazard models that include a rich set of individual-level controls. Overall, we find that the results from the hazard models are broadly consistent with the results based on these figures.

3.1.2 Semiparametric Hazard Models

We investigate the robustness of the graphical results by estimating a set of Cox proportional hazard models. Each table reports results with alternative sets of control variables in the

²³Across all the figures, we report p-values of log-rank tests of equality across the two survival curves. This is the appropriate test to use when data are censored (as is the case in our data). Results using Wilcoxon rank sum test, as are reported in Chetty (2008), are generally very similar.

²⁴While the survival curves are statistically significantly different in figure 4 but not in figure 3, one might ask whether the difference-in-difference (DD) across the two figures is statistically significant. To answer this question, we construct a semiparametric test by estimating a Cox proportional hazard model with separate nonparametric baseline hazard estimates for above-median and below-median unemployment state-years. We include two covariates in this Cox model, an indicator for above-median benefits and a DD term which is 1 for above-median benefits in above-median unemployment state-years and 0 otherwise. The p-value on the estimated DD coefficient is 0.050.

²⁵We have also looked at the subsample of workers with above-median liquid wealth, and we find broadly similar results (see Appendix Figures A2 and A3). These results suggest that liquidity effects are not primarily accounting for the differential duration elasticity between high and low unemployment, which is broadly consistent with our results in Table 11, described below.

columns. The baseline estimating equation is the following:²⁶

$$\log \lambda_{i,s,t} = \alpha_t + \alpha_s + \beta_1 \log(b_{s,t}) + \beta_2(\log(b_{s,t}) \times u_{s,t}) + \beta_3 u_{s,t} + \mathbf{X}_{i,s,t} \Gamma + e_{i,s,t} \quad (13)$$

where $\lambda_{i,s,t}$ is the hazard rate of exit out of unemployment for individual i in state s at time t , α_t and α_s represent year and state fixed effects, $b_{s,t}$ is the unemployment benefit for individual i at the start of the spell based on the state the individual resided in at the start of the spell, and $\mathbf{X}_{i,s,t}$ is a set of (possibly time-varying) control variables. Our primary proxy for local labor market conditions, $u_{s,t}$, is the log state unemployment rate at the start of the spell relative to the log national unemployment rate. The decision to use log unemployment rates follows Bertrand (2004), and we find similar results with the unemployment rate in levels as shown below. We discuss the decision to use relative rather than absolute unemployment rates in detail in Section 3.1.3 below. All variables are de-measured so that $-\beta_1$ represents the elasticity of unemployment durations with respect to the UI benefit level at the average state unemployment rate.²⁷ The coefficient on the interaction term ($-\beta_2$) is the incremental change in the duration elasticity for a one log point change in the state unemployment rate.

The identifying assumption that allows us to interpret β_2 as a test of whether the duration elasticity varies with the unemployment rate is the following: conditional on the average UI weekly benefit amount, state unemployment rate, state fixed effects, year fixed effects, and control variables, there are no omitted determinants of the duration of an unemployment spell that vary with the *interaction* of the average UI weekly benefit amount and the state unemployment rate. This assumption is considerably more plausible with the inclusion of state and year fixed effects, though there remains the concern that benefits respond endogenously to both observed and unobserved local labor market conditions. In Section 3.1.3, we discuss this (and many other) threats to validity in more detail.

Before turning to our regression results, we present descriptive statistics in Panel A of Table 1. The table presents summary statistics for the overall sample and the two sub-samples used to create figures 3 and 4. The two sub-samples are broadly similar, though

²⁶The notation of the estimating equation is a simplified presentation of the true model. The (latent) hazard rate is not actually observed in the data, and there is a flexible (nonparametric) baseline hazard rate which is also estimated when fitting the Cox proportional hazard model. Also, following Chetty (2008), we fit a separate baseline hazard rate for each quartile of net liquid wealth, although our results are very similar when a single nonparametric baseline hazard rate is estimated instead (see Appendix Table A1).

²⁷We will use this approximation throughout for the expected unemployment duration $\log(D) \approx \log(1/\lambda) = -\log(\lambda)$, so that the duration elasticity and other marginal effects of interest are given by the negative of the coefficient in the hazard model.

unemployed individuals are slightly older in states with high unemployment.²⁸

The main results are reported in Table 2. Column (1) of Table 2 reports results of a specification broadly similar to the previous literature (Moffitt (1985), Meyer (1990), Chetty (2008)). This specification controls for age, marital status, years of education, a full set of state, year, industry and occupation fixed effects, and a 10-knot linear spline in log annual wage income. Column (2) reports estimates of equation (13). This column includes the same set of controls in column (1) and estimates the same hazard model; the only difference is the addition of an interaction term between the UI benefit level and the log state unemployment rate. The results indicate that the elasticity of unemployment durations with respect to the UI benefit level is 0.605 (s.e. 0.295) at the average unemployment rate, and this estimate is statistically significant at conventional levels ($p = 0.040$). The coefficient on the interaction term ($-\beta_2$) represents the change in the duration elasticity for a one log point increase in the log state unemployment rate, holding the national unemployment rate constant. The results in column (2) show an estimate of $-\beta_2$ of -1.256 (s.e. 0.433). The bottom two rows of Table 2 show an alternative way to interpret the interaction term. These rows report the duration elasticity when the state unemployment rate is one standard deviation above and below the mean unemployment rate. At one standard deviation above the mean, the duration elasticity is 0.348 (s.e. 0.294), while at one standard deviation below the mean the duration elasticity is 0.863 (s.e. 0.321). These results imply that the magnitude of the duration elasticity decreases with the unemployment rate and suggest that the moral hazard cost of unemployment insurance is lower when the unemployment rate is relatively high.

We can also interpret these results using our job search model. Our empirical finding that the duration elasticity is decreasing with the unemployment rate is consistent with a parameterization of our model where search effort (e) and labor demand conditions (α) are strongly complementary. Unlike a simple static model, complementarity between e and α is not sufficient to generate our empirical results, as the dynamic effect may dominate the static effect if the complementarity is weak. By contrast, if the complementarity is sufficiently strong, then the static effect will dominate and the duration elasticity will be decreasing in the unemployment rate. The simulation results in Figure 2 are qualitatively similar to the duration elasticity results in Table 2 and assume strong complementarity between e and α by using the following functional form: $\lambda(e, \alpha) = \Lambda e^\alpha$.

²⁸In Table 8 below, we control for the small compositional changes in the sample of unemployed individuals across labor market conditions, and we find extremely similar results.

3.1.3 What If UI Benefits Respond to Labor Market Conditions?

An immediate concern with our identification strategy is that UI benefits may be correlated with labor market conditions. We pursue several strategies to address this concern. While the sign of the bias due to the endogeneity of UI benefits is not clear a priori, the collection of evidence in this section suggest that our baseline result is likely a conservative estimate of how the duration elasticity varies with the unemployment rate.

As a first step, we directly examine the association between UI benefits and the unemployment rate. Table 3 reports OLS estimates from a regression of log UI benefits on the unemployment rate with two sets of controls. In Panel A, the controls are state fixed effects and year fixed effects, while in Panel B the controls include state-specific linear time trends in addition to the fixed effects in Panel A. In both panels, the results in column (1) show a small positive coefficient estimate. In Panel A, the point estimate in column (1) implies that a one standard deviation increase in the relative unemployment rate increases UI benefits by 1.64% (s.e. 1.03%). While these results suggest that benefits appear to be somewhat responsive to local labor market conditions, the association is not statistically significant at conventional levels in either panel.

A concern with this simple test is that it may fail to capture the fact that UI benefits are more responsive to the unemployment rate in bad times than in good times. The pattern of benefit changes that we observe in the data suggest this relationship.²⁹ This policy endogeneity would bias estimates of β_1 and β_2 in equation (13). We therefore explore alternative assumptions about how the unemployment rate affects benefits. Column (2) in both panels report results which add a quadratic in the state unemployment rate. In both panels, the quadratic term is positive, suggesting that benefits are more responsive to the unemployment rate when the unemployment rate is high. Columns (3) and (4) report results which split the sample depending on whether the state unemployment rate is above or below the median, and the results are similar to column (2). While the results are often imprecise, the results in this table provide weak but suggestive evidence that benefits are more responsive to labor market conditions in bad times than in good times.

Motivated by this analysis, we report results in Table 4 which control more flexibly for the local unemployment rate in equation (13). We assess the possible bias from such policy endogeneity through several alternative specifications. Column (1) reports our baseline

²⁹We find that there are 1374 instances where the change in the maximum weekly benefit amount is weakly positive and only 12 instances where the maximum weekly benefit amount was reduced.

specification for comparison. Columns (2) through (4) include various polynomial functions of the local unemployment rate and the UI benefit level. These tests address the concern that UI benefits respond non-linearly to the local unemployment rate. Additionally, to the extent that the flexible polynomial in the unemployment rate more thoroughly controls for unobserved local labor market conditions, this specification can be used to gauge the extent of the bias due to policy endogeneity. Though the results are somewhat less precise, the results in these columns suggest that, if anything, the magnitude of our interaction term is larger with these more flexible controls.³⁰ Columns (5) through (7) include specifications which include some combination of interactions between the state unemployment rate and state fixed effects and interactions between the state unemployment rate and year fixed effects. These specifications capture the possibility that in certain states and/or certain years, UI benefits may be unusually responsive to changes in local labor market conditions. Again, the results suggest that, if anything, the magnitude of our interaction term is larger with these more flexible controls.

The previous discussion ignored the possibility of unobserved factors that determine both UI benefits and unemployment durations. Our second strategy now explores this possibility. As our empirical strategy is essentially equivalent to estimating a difference-in-differences regression, we consider the case where the unemployment rate $u_{s,t}$ can only take two values: u_H and u_L , with $u_H > u_L$.³¹ In this case, estimating equation (13) is equivalent to estimating the following two equations:

$$\log(D_{i,H,t}) = \beta_H \log(b_{s,t}) + v_H + \alpha_t + \alpha_s + e_{i,H,t} \quad \text{if } u_{s,t} = u_H \quad (14)$$

and

$$\log(D_{i,L,t}) = \beta_L \log(b_{s,t}) + v_L + \alpha_t + \alpha_s + e_{i,L,t} \quad \text{if } u_{s,t} = u_L \quad (15)$$

The coefficient β_2 on the interaction term $\log(b_{s,t}) \times u_{s,t}$ in equation (13) is given by the difference between β_H and β_L . Each of these two equations is subject to an identification problem. In particular, each equation is a reduced form equation from a system of two equations: an equation determining durations and an equation determining UI benefits. Thus, the standard endogeneity problem due to simultaneous equations arises. Consider

³⁰We have also investigated robustness to flexible, non-linear effects of UI benefit level in addition to the local unemployment rate. These results are in Appendix Table A1, where we re-estimate our baseline specification with various non-linear (polynomial) functions of the unemployment rate and the UI benefit level. The results are similar to our baseline specification.

³¹This discussion extends Bertrand (2004).

the following simplified two equation system:

$$\begin{aligned} D_q &= \beta_q b + \xi_q \\ b &= \lambda_q \xi_q + \eta_q \end{aligned}$$

where $q = \{H, L\}$. The first equation describes the duration equation and the second equation describes the UI benefit equation. The variable ξ_q represents unobserved labor demand shocks, and η_q represents unobserved factors which shift UI benefits and are orthogonal to local labor market conditions. We can now ask what happens if one estimates (14) and (15) ignoring the endogeneity of benefits? It is straightforward to show that $\widehat{\beta}_q$ will be given by:

$$\widehat{\beta}_q = \beta_q + \lambda_q$$

This illustrates the well-known identification problem that $\widehat{\beta}_q \neq \beta_q$ when $\lambda_q \neq 0$. Under the assumption that $\lambda_q = 0$, it is easy to see that β_H and β_L (and therefore $\beta_H - \beta_L$) can be consistently estimated. This assumption requires that all variation in benefits be driven by shocks that are uncorrelated with unobserved labor demand shocks.

By contrast, if $\lambda_q \neq 0$, we need stronger assumptions for identification. Under the strong assumption that $\lambda_H = \lambda_L$, then $\widehat{\beta}_H - \widehat{\beta}_L = \beta_H - \beta_L$. Thus, while we cannot identify the main effect when $\lambda_H = \lambda_L \neq 0$, we will be able to consistently estimate the interaction term of interest.

A key remaining challenge arises when λ_q depends on q . For ease of exposition, consider the case where benefits are exogenous in good times, but are endogenous to local labor demand conditions in bad times (e.g., $\lambda_L = 0$ and $\lambda_H > 0$). In this case, $\widehat{\beta}_H - \widehat{\beta}_L = \beta_H - \beta_L + \lambda_H$. This illustrates that this particular type of policy endogeneity works against the findings in our baseline specification. Intuitively, if variation in benefits is plausibly exogenous during good times, then we will consistently estimate the duration elasticity in good times; however, if variation in benefits is correlated with unobserved labor market conditions during bad times, then this will cause upward bias in the magnitude of the duration elasticity during bad times (e.g., $\widehat{\beta}_H = \beta_H + \lambda_H > \beta_H$). Since we find that the magnitude of the duration elasticity is significantly *smaller* during bad times, this suggests that this policy endogeneity likely causes us to understate the magnitude of the interaction term.

The last threat to identification we discuss comes from the implicit assumption that UI

benefits respond symmetrically to both local and national labor market shocks. If benefits respond to observable and unobservable national labor market conditions, then including year fixed effects, as we do in our baseline specification, addresses the problem. The concern is that when the national labor market is bad, benefits are more correlated with labor market conditions than when national labor market is good. In this case, year fixed effects will not capture the fact that the correlation between benefits and unobservable labor market conditions depends on the year. One strategy for dealing with this is interacting year fixed effects with UI benefits and including this as a control. This means that any variation in benefits that is correlated with the (unobservables in the) year fixed effects is not variation we use to identify the interaction term. Alternatively, one may use relative unemployment rates to address the concern. In regression results not reported, we find that UI benefits are larger when national unemployment rate is high. We also find results that suggest when the national unemployment rate is high, state benefits are more responsive to relative state unemployment rates. Motivated by this concern, we use the state unemployment rate relative to the national unemployment rate in all specifications.³²

Finally, Table 5 presents results from modifications of our baseline specification which focus on alternative assumptions regarding contemporaneous trends across states within a region and within states over time. To the extent that such smoothly-varying unobserved trends are correlated with the interaction of local labor market conditions and UI benefit levels, this may cause our baseline specification to be biased. The results in columns (2) through (4) show that results are similar when we flexibly control for unobserved trends.

Overall, given the results in Tables 3 through 5, we interpret our preferred estimate as a conservative estimate of the magnitude of the relationship between the duration elasticity and the local unemployment rate. We find no evidence that a simple policy endogeneity story is primarily responsible for our findings. The next section explores additional threats to validity and alternative explanations for our findings.

3.1.4 Alternative Explanations and Robustness Tests

Decomposing Variation in the State Unemployment Rate

In our baseline specification, identification of the interaction term of interest comes from both across-state and within-state variation in unemployment rates. In figures 5 and 6,

³²We find similar results if we include year fixed effects interacted with UI benefits as a control, and we also find similar results when we include the national unemployment rate interacted with UI benefits as a control.

we report survival curves analogous to figures 3 and 4 using only within-state variation in unemployment rates. We do this by subtracting the unemployment rate in each state-year by the average unemployment rate in the state over the sample period. Figures 5 and 6 show that the same pattern emerges when using only within-state variation in the unemployment rate. These figures show that there is a statistically significant difference between the high- and low-benefit survival curves when the unemployment rate is relatively low, but not when it is relatively high. Figures 7 and 8 show similar results using only cross-state variation in the state unemployment. To construct these figures, we compute the average state unemployment rate over the sample period and divide states based on whether they are above or below the median.

We quantify these patterns in Table 6, where we report results from a specification where we decompose the variation in the state unemployment rate into across-state variation and within-state variation. The first row reports results from our baseline specification for comparison, and the second and third rows report results which interact the UI benefit level with (1) the average state unemployment rate (calculated separately for each state over the sample period) and (2) the state unemployment rate minus the average state unemployment rate. This allows us to see separately how across-state and within-state variation in the state unemployment rate affects the duration elasticity. We also find that both interaction terms are the same sign as the interaction term in the baseline specification, with a larger interaction term using solely across-state variation in unemployment rates. Most importantly, the magnitude of the interaction term using purely within-state variation is very similar to the baseline results.

Exploiting Variation Across Metropolitan Areas Within States

Thus far, we have defined local labor markets as states. As an additional robustness test, we instead use metropolitan areas (MSAs) to define local labor markets.³³ In Table 7, we report results from using the MSA unemployment rate rather than the state unemployment rate, which allows us to exploit within-state variation in local labor market conditions. Overall, the results in this table are similar to the baseline specification. In the final column of Table 7, we report results which include a full set of state-by-year fixed effects, so that the only variation used to estimate the interaction term is within-state-year, across-MSA variation in the unemployment rate, holding the state-year UI benefit level constant. The interaction term remains statistically significant ($p = 0.027$), and the magnitude is somewhat

³³To preserve the sample size, we assign the state unemployment rate to all unemployed individuals who do not have an MSA code, which is roughly 50% of the sample.

larger than in the baseline specification.

Composition Bias, Selection on Observables, and Endogenous Take-up

As the local unemployment rate fluctuates, there may be compositional changes in the pool of unemployed workers receiving UI benefits. For example, if there is heterogeneity in moral hazard across demographic groups, and the distribution of demographics of the unemployed varies with the level of unemployment, then this compositional change could generate an observed change in the average duration elasticity. We first note that the appropriate measure for the welfare calibrations below is how the average duration elasticity varies with the unemployment rate, and that this is true whether or not the change in the average duration elasticity is primarily due to compositional changes or individual-level changes in moral hazard. Nevertheless, we investigate the extent to which compositional changes can account for our findings, as understanding this may be important for other economic problems.

In Table 8 we report estimates of an augmented version of our baseline specification where we add interactions between UI benefits and the demographic controls in our baseline specification. If the estimates of the baseline interaction term is primarily due to compositional changes among demographic groups with different duration elasticities, then we would expect to see a reduction in the magnitude of the coefficient as we include additional interactions between UI benefits and demographic controls. The results in Table 8 show that our main result is very robust to including such controls – looking across all the columns, we see that adding interactions between demographics and UI benefits has a negligible effect on our main coefficient of interest.³⁴

The final column investigates a related source of compositional bias, which is selection bias due to endogenous take-up. As shown in column (8), we find that the effect of UI benefits on take-up varies with the unemployment rate. We also find that the unemployment rate itself is a strong predictor of take-up. These results raise concerns about possible selection bias, though the results in the rest of the columns in Table 8 suggest negligible effects of selection on observables. While the duration elasticity could also vary with unobservable characteristics of individuals, the robustness to selection on observables suggests that it is unlikely that our interaction term is primarily due to selection on unobservables, though of course we cannot test this directly.³⁵

³⁴The results in this table also reveal suggestive evidence that the duration elasticity varies with years of education. Though accounting for this interaction has no effect on our interaction term of interest, it nevertheless suggests that moral hazard cost may vary with the observable level of human capital.

³⁵We find that our results are similar if we estimate the following two-step estimator. In the first step,

Alternative Measures of the Interaction Term

In our baseline specification, the interaction term of interest is formed by interacting the log of the average weekly UI benefit in the state with the log of the state unemployment rate (relative to the log of the national unemployment rate). Both of these variables are proxies for the underlying variables of interest (an individual’s actual UI benefit level and the latent local labor demand conditions). Table 9 explores several alternative measures of the interaction term, and we find that our results are generally robust to these alternatives. Each row of Table 9 reports results from estimating our baseline specification with an alternative measure of our interaction term. The first row reproduces our baseline estimates for comparison. The second row replaces the state unemployment rate with a dummy variable for whether or not the unemployment rate is greater than the median state unemployment rate in that year. This specification corresponds more closely to the nonparametric results presented above. The third row reports results using the state unemployment rate in levels (rather than logs). In both cases, the results are similar to the baseline specification. The fourth row replaces the average weekly UI benefit amount with the maximum weekly benefit amount. The maximum weekly benefit amount corresponds more closely to a specific policy parameter set by states. While this alternative measure may be more plausibly exogenous, the primary drawback is that it does not exploit all of the variation in UI benefit generosity. The results indicate that the magnitude of our interaction term is very similar with this alternative measure. The fifth row shows that results are similar using the average replacement rate rather than the average weekly UI benefit amount. The sixth row reports reduced form results using a “simulated instrumental variable” following Currie and Gruber (1996). By construction, the variation in this UI benefit variable is only due to changes in program parameters, holding sample composition constant.³⁶ Reassuringly, we find similar results to our baseline specification.³⁷ Finally, the last row uses an alternative proxy for local labor demand instead of the local unemployment rate. One concern with the unemployment rate is that it reflects both labor demand and labor supply shocks. We construct variation in state unemployment rates that is driven by plausibly exogenous shifts in local

we estimate a probit model of UI receipt on interaction term using the same set of controls used in the baseline proportional hazard model using the expanded sample which includes eligibles who do not receive UI benefits. In the second step, we estimate the baseline hazard specification including as an additional control the inverse Mills ratio evaluated at the fitted values.

³⁶This variable is constructed by using a fixed 20% 1993 (national) sample and computing the average weekly UI benefit in this fixed sample for every state-year combination in the data set.

³⁷Appendix Table A1 reports similar results when this simulated instrument is used as an instrumental variable for the average UI WBA. This is implemented using a two-step control function approach.

labor demand using a well-established procedure developed in Bartik (1991).³⁸ Figures 9 and 10 plot survival curves comparing the effect of UI benefits across high and low predicted employment-to-population ratios. Consistent with figures 3 and 4, this nonparametric evidence indicates that the behavioral effect of UI benefits is largest during periods of high predicted employment. Row 7 of Table 9 reports hazard model estimates, where the magnitude of the interaction term is similar to our baseline specifications, but our estimates are imprecise.³⁹

To summarize, across all the specifications in this section, we find no evidence that our baseline results are primarily due to composition bias, liquidity effects, mismeasurement, or nonlinear effects. We therefore conclude that the most likely explanation for our findings is that the disincentive effect of UI benefits decreases with the unemployment rate.

3.1.5 Comparison to Schmieder et al. (2011) and Landais et al. (2011)

While we focus on how the optimal UI benefit level varies over the cycle, contemporaneous work by Schmieder et al. (2011) considers how the optimal potential duration of UI benefits varies over the business cycle. They consider a fixed wage, search effort model and derive a formula for the marginal welfare gain of extending the maximum length of time an individual is eligible to receive UI benefits. Similar to our work, they pursue a sufficient statistics approach and empirically estimate how several of the inputs to their formula vary over the cycle. Specifically, they implement a regression discontinuity design using administrative data from Germany to identify the elasticity of non-employment and actual benefit duration with respect to the potential benefit duration. Their research design allows them to estimate an elasticity for each year in their sample (1987-2004). They show that their annual estimate of the non-employment duration elasticity does not correlate significantly with the annual unemployment rate over this time period. On the other hand, the actual benefit duration elasticity is significantly positively correlated with the unemployment rate. While they do not fully evaluate their sufficient statistics formula (because they do not estimate how the insurance value or liquidity effect of UI varies over the cycle), their empirical findings also indicate that the welfare gain of an increase UI benefit generosity (in this case, through UI

³⁸We closely following the implementation of the Bartik (1991) procedure in Autor and Duggan (2003). We predict the employment to population ratio by interacting initial cross-sectional distribution of state-level employment shares with national industry employment trends.

³⁹Lastly, Appendix Table A1 reports similar results using alternative sets of control variables, alternative specifications allowing for nonlinear direct effects, and controls for extended benefits.

benefit extensions) is higher during recessions.⁴⁰

There are several differences between the studies to highlight: institutional setting (United States versus Germany), research design (fixed effects panel versus regression discontinuity), definition of the labor market (local versus national), UI benefit variation (benefit level versus potential duration). Additionally, the marginal worker affected by the UI variation in their setting is an experienced worker older than age 40, while the UI variation in our setting affects a majority of the workers eligible for UI. All of these differences make the empirical estimates not strictly comparable.

The recent work of Landaï et al. (2011) also suggests an economic explanation for the differences in results. While we study changes in UI benefit levels which affect a majority of unemployed workers, Schmieder et al. study changes in UI benefit generosity across age thresholds, holding program parameters constant. Therefore, one may interpret the Schmieder et al. estimates as partial equilibrium (PE) estimates and our estimates as general equilibrium (GE) estimates. The logic is that since they analyze the effect of UI by comparing across two groups experiencing the *same* labor market conditions, it is possible that some of the GE effects of UI are “differenced out”. By contrast, our estimation strategy primarily relies on comparing observably similar individuals receiving the same benefits in *different* labor markets. Under this interpretation, the combined evidence across the two papers suggests that the PE duration elasticity does not vary over the cycle while the GE duration elasticity is strongly decreasing in the unemployment rate. This is precisely the predictions of the job rationing model in Landaï et al. (2011), providing a parsimonious explanation for the empirical findings in the two papers. One problem with this interpretation is that the PE estimates in Schmieder et al. are lower than the estimates in this paper, while the Landaï et al. paper predicts a larger PE elasticity than GE elasticity. We speculate that this difference may simply be due to different behavioral responses to UI benefit *levels* versus UI benefit *extensions*, although we leave a rigorous investigation of this to future work.

3.2 Part 2: Consumption Smoothing

The second part of the empirical analysis replicates and extends previous work on the consumption smoothing benefit of UI (Gruber 1997). Specifically, we estimate how the effect of

⁴⁰They also calibrate their search model and consider two comparative static exercises: (1) an increase in the agent’s search cost and (2) a decline in re-employment wages. They find these lead to opposite effects on search behavior and the marginal welfare gain of extensions.

UI on the consumption drop upon unemployment varies with the state unemployment rate. The empirical strategy closely follows Gruber (1997), which uses the UI replacement rate rather than the UI benefit level and uses the change in total food consumption as a proxy for the change in total consumption.

We use data from the Panel Study of Income Dynamics (PSID) between 1968 and 1987. We impose the same sample restrictions as in Gruber (1997): we focus on all heads of household who are employed at interview date $t - 1$ and unemployed at date t , and we define individuals as unemployed if they are looking for a new job and are not on temporary layoff; furthermore, observations are excluded if any element of food consumption is imputed or there is more than a threefold change in total food consumption.

The baseline specification is the following:

$$\Delta \log C_{i,t} = \alpha_t + \alpha_s + \delta_1 b_{i,s,t} + \delta_2 (b_{i,s,t} \times u_{s,t}) + \delta_3 u_{s,t} + \mathbf{X}_{i,s,t} \Gamma + e_{i,s,t} \quad (16)$$

where $\Delta \log C_{i,t}$ is the difference in log total food consumption for individual i between date $t - 1$ and date t , $b_{i,s,t}$ is the UI replacement rate, $u_{s,t}$ is the state unemployment rate, α_t and α_s are year and state fixed effects, and $\mathbf{X}_{i,s,t}$ is the same set of control variables used in Gruber (1997).⁴¹

Table 10 reports results of estimating equation (16). Column (1) reproduces column (4) in Table 1 of Gruber (1997) and column (2) reports our replication effort. Column (2) shows the average consumption smoothing benefit of UI using our replication sample, and we find that a ten percentage point increase in the UI replacement rate reduces the consumption drop upon unemployment by 2.6%.⁴² Column (3) reports our preferred specification which includes the interaction between the replacement rate and the unemployment. The interaction term is not statistically significant at conventional levels ($p = 0.985$), and the magnitude is about 23% of the magnitude of the interaction term in the duration elasticity results above. Moreover, there is enough precision to rule out large magnitudes, though not modest magnitudes. The remaining columns of Table 10 report a variety of alternative of specifications of equation (16), and none of the results provide significant evidence that

⁴¹As in the first part of the empirical analysis, in our preferred specification we use the log difference between the state unemployment rate and the national unemployment rate. Also, we use the previous year's unemployment rate because we do not observe individuals at the start of their spell, and we want to ensure that the unemployment rate is predetermined, for reasons discussed above.

⁴²The coefficient estimates from our replication sample match the results in Gruber (1997) fairly closely. We have been unable to account for the remaining differences. We are grateful to Jonathan Gruber for sharing his UI benefit calculator, which greatly improved our original replication effort.

the consumption smoothing benefit of UI varies with the state unemployment rate. We view this conclusion as tentative, however, as the consumption-based results lack enough statistical power to rule out modest effects.⁴³

In addition to the results based on consumption data, we modify the duration elasticity specifications from above to study the insurance value of UI. Recent work by Chetty (2008) raises a concern with interpreting the duration elasticity as a pure moral hazard effect; he presents evidence that a component of the observed duration elasticity represents an income effect (or “liquidity effect”). This suggests that the interaction term which we estimate in our baseline specification in the first part of our empirical analysis could plausibly represent a liquidity effect which varies systematically with local labor market conditions. In principle, the consumption smoothing results rule out this possibility as they incorporate liquidity effects. Nevertheless, we present these additional results as a complementary consumption smoothing test (specifically, whether our interaction term in the duration elasticity specification is plausibly driven by liquidity effects). An important caveat to these duration-based results is that – as in Chetty (2008) – these specifications do not exploit plausibly exogenous variation in liquidity constraints, but rather compare duration elasticities across individuals who differ in net liquid wealth.

Table 11 reports these alternative specifications which investigate whether liquidity effects affect the interaction term. Column (1) reports our baseline specification for comparison. Column (2) reports results for the subsample of workers in the 3rd and 4th quartiles of net liquid wealth, where liquidity effects are likely to be less important. The coefficient on the interaction is slightly larger than in the baseline specification.⁴⁴ Columns (3) and (4) report results which include a full set of liquid wealth quartile dummy variables interacted with a combination of occupation fixed effects, industry fixed effects, unemployment duration, and the UI benefit level, and the results are, if anything, slightly stronger. Lastly, column (5) verifies that the interaction term does not vary with liquid wealth. The results consistently support the interpretation that the moral hazard cost of UI decreases with the unemployment rate, and that our results are not primarily due to liquidity effects varying with local labor market conditions. More broadly, the results in this table do not provide any evidence that

⁴³We have also tested whether the consumption level varies with the unemployment rate, and we have found no statistically or economically significant evidence of such an interaction effect. As with the consumption smoothing results, however, we are unable to rule out modest effects with available statistical precision.

⁴⁴The interaction term when using the subsample of unemployed workers without a mortgage (another proxy for individuals that are not liquidity constrained used in Chetty (2008)) is even larger in magnitude than columns (1) and (2), again providing no evidence that liquidity effects are primarily responsible for the interaction term in our baseline specification.

the consumption smoothing benefit (or insurance value) of UI varies with the unemployment rate, which is consistent with the consumption-based tests above.

Overall, we conclude that the consumption smoothing benefit of UI is approximately constant across labor market states. We next discuss the implications of our empirical results for national business cycles, conduct a welfare analysis using our baseline empirical results, compare optimal UI policy to current policy in the U.S., and discuss the aggregate implications of extending benefits in recessions.

4 Interpretation and Implications

4.1 Extrapolating Estimates to National Business Cycles

Our estimate of how the duration elasticity varies with labor market conditions is identified off of variation in local unemployment rates. To study the impact of national business cycles on the duration elasticity, ideally one would want to conduct a full general equilibrium analysis that explicitly accounts for all margins that could potentially be distorted. While such an analysis is outside the scope of this paper, the aim of this subsection is to shed light on this question by drawing inferences from our empirical estimates.

First, the nature of a national shock compared to a local shock may differ. In particular, one may argue that the persistence of national shocks and local shocks differs, and that the behavioral response to UI, conditional on a given level of persistence, may differ. One piece of evidence against differential responses by persistence comes from the results in Table 6. If we associate the between-state variation in unemployment rates with a measure of permanent shocks and the within-state variation with a measure of temporary shocks, we find that with available statistical precision, we cannot reject the hypothesis that the correlation between the duration elasticity and the unemployment rate is the same across both sources of variation.

Second, a local unemployment shock may differ from a national unemployment shock since it changes mobility incentives. If only New York gets hit with a labor market shock, out-migration may increase as individuals move to neighboring states to search for a job. However, the relevant question is whether mobility patterns induced by such a shock vary across states with different levels of UI benefits. Current U.S. law mandates that individuals collect UI benefits in the state of their previous employer. Given this restriction, we believe the mobility response to UI benefits is likely to be low.

Finally, a national shock may have different budgetary implications than a local shock, though it is not clear which direction this will cause our estimates to be biased. In the case of a local shock that persists for some time, states could use the federal government as a buffer to finance UI expenditures; however, a particularly acute local shock may force states to raise taxes to finance the additional UI expenditures. Since the federal government has the ability to run budget deficits over a longer time horizon, this budget effect may have additional welfare implications.

4.2 Calibrating the Welfare Implications

Our empirical results suggest that moral hazard decreases with the unemployment rate. To see what this finding implies for optimal policy, we now calibrate the optimal UI level implied by our model, following the spirit of the “sufficient statistic” approach to welfare analysis. To review, this method requires using the reduced form empirical estimates as inputs into the optimal UI formula.

One can think of $\varepsilon_{D,b} = h(\log(u))$, where $h(\cdot)$ is a non-linear function. In order to exploit our empirical estimates, we assume that $h(\cdot)$ can be locally approximated by a linear function of $\log(u)$. A first-order approximation of $h(\log(u))$ around $\log(u) = \log(\bar{u})$ yields:

$$\varepsilon_{D,b}(\log(u)) = \varepsilon_{D,b}(\log(\bar{u})) + \frac{d\varepsilon_{D,b}(\log(\bar{u}))}{d\log(u)} \times (\log(u) - \log(\bar{u}))$$

This can also be derived directly from our reduced form estimating equation (13):

$$-\log(D) \approx \log(\lambda) = \alpha + \beta_1 \log(b) + \beta_2 \log(b) \times (\log(u) - \log(\bar{u})) + e \quad (17)$$

With this specification,

$$\varepsilon_{D,b}(\log(u)) = \frac{d\log(D)}{d\log(b)} = -\beta_1 - \beta_2 \times (\log(u) - \log(\bar{u}))$$

Thus, $-\beta_1 = \varepsilon_{D,b}(\log(\bar{u}))$ and $-\beta_2 = \frac{d\varepsilon_{D,b}(\log(\bar{u}))}{d\log(u)}$. Our empirical results imply that $-\hat{\beta}_1 = 0.605$ and $-\hat{\beta}_2 = -1.256$.⁴⁵ To analyze the welfare implications, we will assume that

⁴⁵The optimal UI formula in this section delivers the optimal UI replacement rate, which implies that the relevant duration elasticity is the elasticity of unemployment durations with respect to the UI replacement rate, rather than the UI benefit level. For simplicity, we use the results from the baseline specification using the level of benefits, since the duration elasticity results using replacement rate are similar (see Appendix Table A1).

the budget effect can be ignored. This is motivated by current US policy where states' UI trust funds run deficits in recessions and surpluses in booms.

Recall, the consumption smoothing benefit of UI is the following:

$$\frac{U'(b) - E[U'(w - \tau)|w \geq \bar{w}]}{E[U'(w - \tau)|w \geq \bar{w}]}$$

Following Chetty (2006), we can show that the consumption smoothing benefit can be approximated as

$$\left[\gamma \frac{\Delta c}{c}(b) \left(1 + \frac{1}{2} \delta \frac{\Delta c}{c}(b) \right) + 1 \right] F - 1$$

where $\frac{\Delta c}{c}(b)$ is the consumption drop at unemployment as a function of UI benefits, δ is the coefficient of relative prudence, $F = \frac{1}{1 + \frac{1}{2} \gamma \delta s_e^2}$ is a correction factor that accounts for the volatility in consumption when employed. It is defined in terms of s_e , the coefficient of variation of consumption when employed. For ease of computation, we assume a fixed wage, search effort model, so that $F = 1$. In this case, the consumption smoothing benefit of UI becomes

$$\gamma \frac{\Delta c}{c}(b) \left(1 + \frac{1}{2} \delta \frac{\Delta c}{c}(b) \right)$$

We calibrate the consumption smoothing benefit using the estimate for the consumption drop based on our reduced form empirical results.⁴⁶ Given the considerable uncertainty over the value of risk aversion, we present calibration results for a range of values $\gamma = 2, 3, 4$ ($\delta = 3, 4, 5$). Table 12 presents results from the numerical implementation of expression (4). At $\bar{u} = 6.2\%$ and $\gamma = 3$, the optimal replacement rate is 38.9%. At an unemployment rate of 7.5% (roughly one standard deviation above the mean unemployment rate), the formula implies an optimal replacement rate of 57.1%. Thus, we see that variation in the unemployment rate can substantially affect replacement rates. The basic lesson to emerge from the table is that plausible variation in the unemployment rate generates wide variation in the optimal level of UI. To give a sense of the quantitative importance of this variation, the magnitude is roughly equivalent to a one unit change in the coefficient of relative risk aversion in the model (e.g., from $\gamma = 3$ to $\gamma = 4$), holding the unemployment rate constant. While the previous literature has emphasized the sensitivity of the optimal UI benefit level to the level of risk aversion, our results suggest that the optimal UI benefit level is equally

⁴⁶The results are similar if we instead use the results from Gruber (1997) across all labor market conditions, which is as expected given the economically insignificant results in Table 10 above.

sensitive to labor market conditions. This sensitivity highlights the value of future work which produces more precise estimates of how the duration elasticity and effect of UI on consumption drop at unemployment vary with the unemployment rate.⁴⁷

4.3 Comparing Optimal UI Policy to Extended Benefits Policy in the U.S.

In terms of existing UI policy in the U.S., the potential duration of UI benefits has typically been adjusted in response to slackness in the labor market, rather than the level of UI benefits. Historically, benefits have been extended by 13 weeks when a state’s insured unemployment rate exceeded some threshold (Card and Levine, 2001). To shed light on the optimality of such a policy, we consider the case where the unemployment rate increases from 4.9% to 6.2%.⁴⁸ The question we seek to address is whether the expected payout from an increase in the optimal benefit level comes close to the expected payout from an extension of benefits in line with what we observe with existing UI policy. We emphasize that these stylized calculations are primarily meant to be illustrative and that considerable caution should be exercised in interpreting the quantitative results.

Table 12 shows, at $\gamma = 4$, the weekly optimal benefit level increases from \$137 to \$187. Given a mean unemployment duration of roughly 18 weeks based on Table 1, this implies total expected benefits paid out to a given individual increase by \$900 ($= 18 \times (\$187 - \$137)$), assuming no behavioral response. A duration elasticity with respect to the benefit level of 0.605 increases the total expected payout to \$1099 ($= (\$187 - \$137) \times (18 + 0.605 \times 18 \times \frac{\$187 - \$137}{\$137})$).

Next, consider instead a benefit extension of 13 weeks. The expected increase in payments to an individual, assuming a potential duration of 26 weeks for the regular UI program, is given by

$$\$137 \sum_{t=1}^{13} t \times \Pr\{duration = 26 + t\} + \$137 \times 13 \times \Pr\{duration > 39\}$$

⁴⁷In particular, the results using the consumption drop at unemployment are not precise enough to rule out modest interaction effects. If we take the upper and lower bound of the 95% confidence interval on the estimated consumption smoothing interaction term and assume $\gamma = 3$, then we calculate that a one standard deviation increase in the unemployment rate increases the optimal replacement rate by between 5 and 45 percentage points.

⁴⁸Schwartz (2010) notes that extensions are triggered if the insured unemployment rate exceeds 5% and has increased by 20% over the two prior years.

For simplicity, let's assume that $\Pr\{duration = 26 + t\} = \Pr\{duration = 26\}$ for all t . In our SIPP sample, we calculated that the average frequency that a spell lasts exactly t weeks where $t \in [27, 39]$ is roughly 0.009 and $\Pr\{duration > 39\} = .12$. Roughly 76% of the unemployment spells in our data end at 26 weeks or less. Therefore, the expected increase in total payments paid to an individual are \$314 ($= 0.009 \times 91 \times \$137 + 0.12 \times 13 \times \137), assuming no behavioral response. The results from Card and Levine (2000) suggest that an extra 13 weeks of benefits raise durations by 1 week. If we assume that this one week increase would result in receipt of a full week's worth of benefits, the total expected payout incorporating behavioral responses is \$451. Thus, based on this simple illustration, the actual UI policy appears to be less generous in terms of expected payouts as the optimal policy from adjusting the replacement rate would imply.

4.4 Implications of Extending Benefits During Recessions

An important policy question is how extended benefits affect the unemployment rate in a recession. Ideally, credible empirical evidence on how durations respond to benefit extensions and how this effect varies with labor market conditions would shed light on this question. Unfortunately, such evidence does not exist for the U.S.. The key challenge is that benefits are typically extended during recessions, making them endogenous and unreliable for empirical analysis. For illustrative purposes, we show how one can use our reduced form empirical results to shed light on how extended benefits affect the unemployment rate in a recession.

To fix ideas, we focus on the recent extension in potential duration from 26 weeks to 104 weeks in the U.S. Card and Levine (2000) estimate during normal economic times that a 13-week extension increases unemployment duration by 1 week. Extrapolating from their estimate, a 78 week increase in potential duration is expected to prolong durations on average by 6 weeks in normal times. We now compute how much the UI benefit *level* has to increase to raise durations by 1 week, based on our estimates. We know that unemployment duration is 18 weeks at the mean unemployment rate in our sample of 6.2%. Further, the duration elasticity is 0.605. Thus, 1 extra week of unemployment duration would require a percent increase in benefits equal to $.092 = (1/18)/.605$.

Now suppose that the unemployment rate has increased to 8.8%. Our estimates imply that a 9.2% increase in benefits would increase durations by roughly 1.5% or $.015 \times 18 = 0.3$ weeks. Using our estimates, and assuming that the potential duration elasticity is also lower

in bad times, a 13-week extension would raise durations by 0.3 weeks in bad times, compared to the 1 week from Card and Levine. Thus, a 78 week increase in potential duration leads to a 1.8 week increase in durations, rather than the 6 week increase one would obtain based on Card and Levine's estimate.

To see what this implies for the aggregate unemployment rate, we make use of the steady-state relationship, $u = s/(p + s)$, where s is the separation rate and p is the job finding rate. For the monthly job separation rate, we set $s = .0125$, which is similar to the estimate in Shimer (2007). At this separation rate, $p = 0.19$ gives you an unemployment duration of roughly 22.8 weeks and an unemployment rate of 6.2%. After the benefit extension, durations rise to $24.6 = 22.8 + 1.8$ weeks or 5.7 months, giving a new unemployment rate of $u = .0125/ (.0125 + 1/4.95) = 6.65\%$. Thus, our calculations imply that extended benefits can account for roughly 17% of the observed increase in the aggregate unemployment rate during the recession in the U.S..

5 Conclusions

In this paper, we have considered a standard job search model and have shown that the relationship between both the moral hazard cost and the consumption smoothing benefit of UI and the unemployment rate is theoretically ambiguous. This motivated our two-part empirical strategy which (1) estimated how the elasticity of unemployment duration with respect to the UI benefit level varies with the unemployment rate and (2) estimated how the effect of UI on the consumption drop upon unemployment varies with the unemployment rate.

Our empirical findings indicate that moral hazard is lower when unemployment is high, consistent with the speculation of Krueger and Meyer (2002) who claimed that there is likely less of an efficiency loss from reduced search effort by the unemployed when local labor market conditions are poor. On the other hand, we do not find evidence that the consumption smoothing benefit of UI varies with the unemployment rate. We have also shown how one can use the empirical relationship between the duration elasticity and the unemployment rate to calibrate a simple optimal UI formula.

We view the concept that the moral hazard cost of social policies may vary with local labor market conditions as quite general, extending beyond the application of unemployment insurance considered in this paper. It is plausible that the disincentive effects of other government policies may also be lower in times of high unemployment. For example, if the

labor supply response to tax changes is lower during recessions, it may be more efficient to redistribute during recessions. In the case of disability insurance and workers compensation, the adverse incentive effect of such programs may be influenced by the severity of the health or income shock. It would be interesting to study whether moral hazard varies with the size of the shock that triggers these programs, particularly since, as we showed here, benefits can be conditioned on the size of the shock. To the extent that the moral hazard cost of social insurance programs varies with labor market conditions, one should draw caution in comparing elasticity estimates across studies to the extent that there are different labor market conditions that underlie these estimates.

Lastly, while we focused on the UI benefit level as the policy parameter, in practice, the potential benefit duration is typically extended during times of high unemployment. In ongoing work, we are studying theoretically how governments should optimally set the time path of UI benefits and how this varies with labor market conditions. We hope that this analysis will shed light on UI benefit extension programs in the U.S. and other developed countries.

References

- [1] **Andersen, T. M. and M. Svarer** (2009), "State Dependent Unemployment Benefits", *Journal of Risk and Insurance*, forthcoming.
- [2] **Andersen, T. M. and M. Svarer** (2010), "Business Cycle Dependent Unemployment Insurance," IZA DP No. 5196.
- [3] **Autor, D. and M. Duggan** (2003), "The Rise in the Disability Rolls and the Decline in Unemployment", *Quarterly Journal of Economics*, 118(1) (February): 157-206.
- [4] **Baily, M. N.** (1978), "Some Aspects of Optimal Unemployment Insurance", *Journal of Public Economics*, 10 (December): 379-402.
- [5] **Bartik, T. J.** (1991), "Who Benefits From State and Local Economic Development Policies", W.E. Upjohn Institute for Employment Research: Kalamazoo, Mich.
- [6] **Bertrand, Marianne** (2004), "From the Invisible Handshake to the Invisible Hand? How Import Competition Changes the Employment Relationship", *Journal of Labor Economics*, 22(4): 723-766.
- [7] **Burdett, K.** (1978), "A Theory of Employee Job Search and Quit Rates", *American Economic Review*, 68(1): 212-20.
- [8] **Burdett, K.** (1981), "A Useful Restriction on the Offer Distribution in Job Search Models", in G. Eliasson, B. Holmlund, and F.P. Stafford, eds., *Studies in Labor Market Behavior: Sweden and the United States*. Stockholm: Industrial Institute for Economic and Social Research.
- [9] **Burdett, K. and J. I. Ondrich** (1985), "How Changes in Labor Demand Affect Unemployed Workers", *Journal of Labor Economics*, 3(1): 1-10.
- [10] **Card, D., Chetty R., and A. Weber** (2007), "Cash-on-Hand and Competing Models of Intertemporal Behavior: New Evidence from the Labor Market", *Quarterly Journal of Economics* 122(4): 1511-1560.
- [11] **Card, D. and P.B. Levine** (2000), "Extended benefits and the duration of UI spells: evidence from the New Jersey extended benefit program", *Journal of Public Economics*, (78): 107-138.
- [12] **Chesher, A. and T. Lancaster** (1983), "An Econometric Analysis of Reservation Wages", *Econometrica*, 51(6): 1661-1676.
- [13] **Chetty, R.** (2004), "Optimal Unemployment Insurance When Income Effects are Large", NBER working paper #10500.
- [14] **Chetty, R.** (2006), "A General Formula for the Optimal Level of Social Insurance", *Journal of Public Economics* 90 (November): 1879-1901.

- [15] **Chetty, R.** (2008), "Moral Hazard versus Liquidity and Optimal Unemployment Insurance," *Journal of Political Economy*, 116(2): 173-234.
- [16] **Chetty, R.** (2009), "Sufficient Statistics for Welfare Analysis: A Bridge Between Structural and reduced form Methods", *Annual Review of Economics* 1: 451-488.
- [17] **Cohen, A. and L. Einav** (2007), "Estimating Risk Preferences from Deductible Choice", *American Economic Review* 97(3) (June): 745-788.
- [18] **Currie, J. and J. Gruber** (1996), "Saving babies: The efficacy and cost of recent changes in the Medicaid eligibility of pregnant women", *The Journal of Political Economy* 104(6): 1263–1296.
- [19] **Feldstein, M. and J. Poterba** (1984), "Unemployment Insurance and Reservation Wages", *Journal of Public Economics* 23 (February-March): 141-67.
- [20] **Gruber, J.** (1997), "The Consumption Smoothing Benefits of Unemployment Insurance", *American Economic Review* 87 (March): 192-205.
- [21] **Hamermesh, D.** (1977), "Jobless pay, and the economy (Johns Hopkins Press, Baltimore, MD).
- [22] **Han, A. and J. Hausman** (1990), "Flexible Parametric Estimation of Duration and Competing Risk Models," *Journal of Applied Econometrics*.
- [23] **Hopenhayn, H. A. and J. Nicolini** (1997), "Optimal unemployment insurance", *Journal of Political Economy* 105: 412–438.
- [24] **Kiley, M. T.** (2003), "How Should Unemployment Benefits Respond to the Business Cycle?", *The B.E. Journal of Economic Analysis and Policy*. Topics, Vol 3, Issue 1.
- [25] **Kroft, K.** (2008), "Takeup, Social Multipliers and Optimal Social Insurance", *Journal of Public Economics*, 92: 722-737.
- [26] **Krueger, A. and B. Meyer** (2002), "Labor Supply Effects of Social Insurance" in A. Auerbach and M. Feldstein, eds. *Handbook of Public Economics* Vol 4. North-Holland: Amsterdam.
- [27] **Lentz, R.** (2009), "Optimal Unemployment Insurance in an Estimated Job Search Model with Savings", *Review of Economic Dynamics*. Vol 12(1): 37-57.
- [28] **Landais, C., P. Michailat and E. Saez** (2011). "Optimal Unemployment Insurance over the Business Cycle," NBER Working Paper No. 16526.
- [29] **Liu, H.** (2009), "Life Cycle Human Capital Formation, Search Intensity, and Wage Dynamics", Working Paper.
- [30] **Manski, C. F.** (1993), "Identification of Endogenous Social Effects: The Reflection Problem", *Review of Economic Studies*. Vol 60(3): 531-42.

- [31] **Meyer, B.** (1990), "Unemployment Insurance and Unemployment Spells". *Econometrica* 58 (July): 757-782.
- [32] **Moffitt, R.** (1985), "Unemployment Insurance and the Distribution of Unemployment Spells", *Journal of Econometrics*, 28, 85-101.
- [33] **Mortensen, D.** (1977), "Unemployment Insurance and Job Search Decisions". *Industrial and Labor Relations Review*, ILR Review, ILR School, Cornell University, 30(4), (July): 505-517.
- [34] **Mortensen, D. and C. Pissarides** (1999), "New developments in models of search in the labor market", O. Ashenfelter & D. Card (ed.) *Handbook of Labor Economics*, chapter 39: 2567-2627.
- [35] **Nicholson, W. and K. Needels** (2006), "Unemployment Insurance: Strengthening the Relationship between Theory and Policy", *Journal of Economic Perspectives*, 20(3) (Summer): 47-70.
- [36] **Redding, S. J. and D. M. Sturm** (2008), "The Costs of Remoteness: Evidence from German Division and Reunification", *American Economic Review* 98(5): 1766-1797.
- [37] **Røgerson, R., R. Shimer and R. Wright** (2005), "Search-Theoretic Models of the Labor Market: A Survey", *Journal of Economic Literature*, 43(4), 959-988.
- [38] **Saez, E.** (2001), "Using Elasticities to Derive Optimal Income Tax Rates", *Review of Economic Studies*, 68: 205-229.
- [39] **Sanchez, J.** (2008), "Optimal State-Contingent Unemployment Insurance", *Economics Letters*, 98(3): 348-357.
- [40] **Schmieder, J. F., von Wachter, T. and S. Bender** (2011), "The Effects of Extended Unemployment Insurance Over the Business Cycle: Evidence from Regression Discontinuity Estimates over Twenty Years", Working Paper.
- [41] **Schwartz, J.** (2010), "Do Temporary Extensions to Unemployment Insurance Benefits Alter Search Behavior? The Effects of the Standby Extended Benefit Program in the United States", Working Paper.
- [42] **Shimer, R. and I. Werning** (2007), "Reservation Wages and Unemployment Insurance", *Quarterly Journal of Economics* 122(3): 1145-1185.
- [43] **Shimer, R.** (2007), "Reassessing the Ins and Outs of Unemployment", Working Paper.
- [44] **Shimer, R. and I. Werning** (2008), "Liquidity and Insurance for the Unemployed", *American Economic Review*, 98(5): 1922-1942.
- [45] **Van den Berg, G. J.** (1990), "Nonstationarity in Job Search Theory", *Review of Economic Studies*, 57: 255-277.

- [46] **Van den Berg, G. J.** (1994), "The Effects of Changes of the Job Offer Arrival Rate on the Duration of Unemployment", *Journal of Labor Economics*, 12(3): 478-498.
- [47] **Van den Berg, G. J.** (2001), "Duration Models: Specification, Identification and Multiple Durations", *Handbook of Econometrics*, Volume 5, Chapter 55, Edited by J.J. Heckman and E. Leamer

Table 1
Descriptive Statistics

<u>Panel A: Duration elasticity sample (SIPP)</u>							
	Full Sample		State Unemp. Rate < Median		State Unemp. Rate ≥ Median		p-value of difference in means
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	
Annual Income (\$000's)	20.925	13.570	20.932	13.554	20.921	13.581	0.979
Age	37.165	11.066	36.590	11.112	37.483	11.030	0.011
Years of Education	12.171	2.877	12.119	2.869	12.200	2.882	0.372
1 {Married}	0.616	0.486	0.609	0.488	0.620	0.486	0.501
Average UI Weekly Benefit Amount (\$'s)	163.33	26.80	163.08	26.07	163.46	27.21	0.660
1 {Net liquid wealth in 1st quartile}	0.259	0.438	0.260	0.439	0.258	0.438	0.914
2 {Net liquid wealth in 2nd quartile}	0.238	0.426	0.232	0.422	0.240	0.427	0.544
3 {Net liquid wealth in 3rd quartile}	0.271	0.444	0.273	0.446	0.269	0.444	0.775
4 {Net liquid wealth in 4th quartile}	0.233	0.423	0.235	0.424	0.232	0.422	0.842
Unemployment duration (weeks)	18.510	14.351	16.950	13.605	19.373	14.678	0.000
Number of Spells	4307		2774		1533		
<u>Panel B: Consumption change sample (PSID)</u>							
	Full Sample		State Unemp. Rate < Median		State Unemp. Rate ≥ Median		p-value of difference in means
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	
After-tax UI Replacement Rate	0.606	0.155	0.599	0.153	0.613	0.156	0.083
Age	34.583	11.441	34.377	11.211	34.766	11.644	0.498
1 {Female}	0.233	0.423	0.250	0.433	0.217	0.413	0.129
1 {Married}	0.571	0.495	0.565	0.496	0.576	0.495	0.661
1 {White}	0.496	0.500	0.495	0.500	0.496	0.500	0.964
1 {Black}	0.460	0.499	0.478	0.500	0.444	0.497	0.180
Change in log food needs	0.002	0.255	0.002	0.254	0.003	0.256	0.910
Number of Children Under 18	1.240	1.422	1.223	1.391	1.255	1.449	0.650
Change in log consumption upon unemployment	-0.070	0.415	-0.061	0.402	-0.079	0.427	0.396
Number of Observations	1595		749		846		

Notes: In Panel A, data are individual-level unemployment spells from 1985-2000 SIPP. Average UI Weekly Benefit Amount is 2000 CPI-U adjusted dollars. In Panel B, data are individual-level observations from 1968-1987 PSID. See text for details.

Table 2
How Does Duration Elasticity Vary With the
Unemployment Rate?

	(1)	(2)
log(Average UI WBA) (A)	-0.503 (0.291) [0.083]	-0.605 (0.295) [0.040]
log(Average UI WBA) × log(State Unemp. Rate / National Unemp. Rate) (B)		1.256 (0.433) [0.004]
log(State Unemp. Rate / National Unemp. Rate)	-0.002 (0.128) [0.989]	-0.014 (0.119) [0.904]
log(Average UI WBA) × Unemployment Duration	0.003 (0.009) [0.704]	0.003 (0.009) [0.746]
Age	-0.017 (0.002) [0.000]	-0.017 (0.002) [0.000]
Marital Dummy	0.207 (0.040) [0.000]	0.207 (0.040) [0.000]
Years of Education	0.004 (0.006) [0.500]	0.004 (0.006) [0.500]
Number of Spells	4307	4307
Post-estimation: (A) + σ × (B)		-0.348 (0.294) [0.236]
Post-estimation: (A) - σ × (B)		-0.863 (0.321) [0.007]

Notes: All columns report semiparametric (Cox proportional) hazard model results from estimating equation (13). Data are individual-level unemployment spells from 1985-2000 SIPP. All specifications include state, year, industry and occupation fixed effects, 10-knot linear spline in log annual wage income and a control for being on the seam between interviews to adjust for the "seam effect." The Average UI WBA is the average weekly benefit amount paid to individuals claiming unemployment insurance in a given state. All columns estimate nonparametric baseline hazards stratified by quartile of net liquid wealth. The final two rows report linear combinations of parameter estimates to produce the duration elasticity when the state unemployment rate is one standard deviation above/below the mean. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix for each state over time, are in parentheses and p-values are in brackets.

Table 3
How Do UI Benefits Vary With the Unemployment Rate?

Dependent variable: Log of Maximum UI Weekly Benefit Amount				
	Sample Restrictions:			
	Full Sample		Unemp. Rate \geq Median	Unemp. Rate $<$ Median
	(1)	(2)	(3)	(4)
<u>Panel A: State FEs + Year FEs</u>				
log(State Unemp. Rate / National Unemp. Rate)	0.080	0.103	0.117	0.008
	(0.050)	(0.062)	(0.137)	(0.088)
	[0.122]	[0.104]	[0.399]	[0.924]
(log(State Unemp. Rate / National Unemp. Rate)) ²		0.100		
		(0.107)		
		[0.355]		
N	672	672	336	336
<u>Panel B: State FEs + Year FEs + State-specific linear trends</u>				
log(State Unemp. Rate / National Unemp. Rate)	0.015	0.044	0.109	-0.010
	(0.032)	(0.025)	(0.058)	(0.046)
	[0.641]	[0.078]	[0.068]	[0.822]
(log(State Unemp. Rate / National Unemp. Rate)) ²		0.091		
		(0.049)		
		[0.071]		
N	672	672	336	336

Notes: All columns report OLS regressions with the log of the statutory maximum weekly UI benefit in the state as the dependent variable. Data set is state-level panel of the 42 states used in the baseline SIPP sample between 1985 and 2000. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix for each state over time, are in parentheses and p-values are in brackets.

Table 4
Robustness to Allowing UI Benefits to Respond Flexibly to the Unemployment Rate

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
log(Average UI WBA) (A)	-0.605 (0.295) [0.040]	-0.615 (0.331) [0.063]	-0.741 (0.373) [0.047]	-0.693 (0.351) [0.048]	-0.781 (0.331) [0.018]	-0.631 (0.341) [0.064]	-0.575 (0.369) [0.119]
log(Average UI WBA) × log(State Unemp. Rate / National Unemp. Rate) (B)	1.256 (0.433) [0.004]	1.264 (0.463) [0.006]	1.599 (0.442) [0.000]	1.737 (0.472) [0.000]	2.576 (1.205) [0.033]	1.320 (0.388) [0.001]	2.529 (0.901) [0.005]
Quadratic in State Unemployment Rate	N	Y	N	N	N	N	N
Cubic in State Unemployment Rate	N	N	Y	N	N	N	N
Quartic in State Unemployment Rate	N	N	N	Y	N	N	N
State FEs x State Unemployment Rate	N	N	N	N	Y	N	Y
Year FEs x State Unemployment Rate	N	N	N	N	N	Y	Y
Post-estimation: (A) + σ × (B)	-0.348 (0.294) [0.236]	-0.348 (0.294) [0.236]	-0.356 (0.316) [0.259]	-0.414 (0.348) [0.234]	-0.253 (0.329) [0.442]	-0.361 (0.326) [0.269]	-0.056 (0.362) [0.876]
Post-estimation: (A) - σ × (B)	-0.863 (0.321) [0.007]	-0.863 (0.321) [0.007]	-0.874 (0.371) [0.018]	-1.069 (0.417) [0.010]	-1.309 (0.483) [0.007]	-0.902 (0.372) [0.015]	-1.093 (0.458) [0.017]

Notes: All columns report semiparametric (Cox proportional) hazard model results from estimating equation (13). Data are individual-level unemployment spells from 1985-2000 SIPP. Number of spells = 4307. See Table 2 for more details on the baseline specification. The final two rows reports linear combinations of parameter estimates to produce the duration elasticity when the state unemployment rate is one standard deviation above/below the mean. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix for each state over time, are in parentheses and p-values are in brackets.

Table 5
Robustness to Flexibly Controlling for Unobserved Trends

	(1)	(2)	(3)	(4)
log(Average UI WBA) (A)	-0.605	-0.765	-0.625	-0.809
	(0.295)	(0.379)	(0.456)	(0.467)
	[0.040]	[0.044]	[0.171]	[0.083]
log(Average UI WBA) × (B)	1.256	1.430	1.737	1.236
log(State Unemp. Rate / National Unemp. Rate)	(0.433)	(0.514)	(0.841)	(0.547)
	[0.004]	[0.005]	[0.039]	[0.024]
State FEs + Year FEs	Y	Y	Y	Y
Region-specific linear time trends	N	Y	N	N
Region × Year FEs	N	N	Y	N
State-specific linear time trends	N	N	N	Y
Post-estimation: (A) + σ × (B)	-0.348	-0.472	-0.269	-0.556
	(0.294)	(0.388)	(0.460)	(0.493)
	[0.236]	[0.224]	[0.559]	[0.259]
Post-estimation: (A) - σ × (B)	-0.863	-1.058	-0.981	-1.063
	(0.321)	(0.399)	(0.514)	(0.467)
	[0.007]	[0.008]	[0.057]	[0.023]

Notes: All columns report semiparametric (Cox proportional) hazard model results from estimating equation (13). Number of spells = 4307. Data are individual-level unemployment spells from 1985-2000 SIPP. See Table 2 for more details on the baseline specification. The final two rows reports linear combinations of parameter estimates to produce the duration elasticity when the state unemployment rate is one standard deviation above/below average. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix for each state over time, are in parentheses and p-values are in brackets.

Table 6
Decomposing Variation in State Unemployment Rate

	Hazard Model Results			Post-estimation	
	(A)	(A) × (B)	(B)	(A) + σ × (B)	(A) - σ × (B)
(1) (A) log(Average UI WBA) ×	-0.605	1.256	-0.014	-0.348	-0.863
(B) log(State Unemp. Rate / National Unemp. Rate)	(0.295)	(0.433)	(0.119)	(0.294)	(0.321)
	[0.040]	[0.004]	[0.904]	[0.236]	[0.007]
(2) (A) log(Average UI WBA) ×		2.830	-1.467	-0.220	-1.100
(B) Average of log(State Unemp. Rate / National Unemp. Rate), 1985-2000		(2.122)	(0.518)	(0.363)	(0.580)
	-0.660	[0.182]	[0.005]	[0.545]	[0.058]
(A') log(Average UI WBA) ×	(0.354)	1.059	-0.026	-0.511	-0.808
(B') (log(State Unemp. Rate / National Unemp. Rate) - (B))	[0.062]	(0.465)	(0.121)	(0.368)	(0.352)
		[0.023]	[0.826]	[0.165]	[0.022]

p-value of test (B) = (B'): 0.440

Notes: All rows report semiparametric (Cox proportional) hazard model results from estimating equation (13); each column reports separate parameter estimate. Data are individual-level unemployment spells from 1985-2000 SIPP. Number of spells = 4307. See Table 2 for more details on the baseline specification. The final two columns report linear combinations of parameter estimates to produce the duration elasticity when the state unemployment rate is one standard deviation above/below the mean. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix for each state over time, are in parentheses and p-values are in brackets.

Table 7
Exploiting Variation Across Metropolitan Areas Within States

	(1)	(2)	(3)	(4)
log(Average UI WBA) (A)	-0.430 (0.277) [0.121]	-0.397 (0.393) [0.312]	-0.894 (0.490) [0.068]	
log(Average UI WBA) × (B) log(Metropolitan Unemp. Rate / National Unemp. Rate)	0.986 (0.396) [0.013]	1.145 (0.636) [0.072]	1.284 (0.483) [0.008]	2.013 (0.970) [0.038]
log(Metropolitan Unemp. Rate / National Unemp. Rate)	-0.065 (0.094) [0.494]	0.066 (0.157) [0.675]	0.009 (0.096) [0.922]	0.078 (0.157) [0.620]
MSA FEs	Y	Y	Y	Y
Year FEs	Y	Y	Y	Y
Region × Year FEs	N	Y	N	N
State-specific linear time trends	N	N	Y	N
State × Year FEs	N	N	N	Y
Post-estimation: (A) + σ × (B)	-0.149 (0.294) [0.613]	-0.071 (0.428) [0.868]	-0.528 (0.523) [0.312]	
Post-estimation: (A) - σ × (B)	-0.710 (0.304) [0.020]	-0.723 (0.437) [0.098]	-1.259 (0.494) [0.011]	

Notes: All columns report semiparametric (Cox proportional) hazard model results from estimating equation (13). Data are individual-level unemployment spells from 1985-2000 SIPP. Number of spells = 4307. See Table 2 for more details on the baseline specification. To preserve sample size, observations without MSA codes are grouped together within a state and assigned the state unemployment rate. The final two rows reports linear combinations of parameter estimates to produce the duration elasticity when the state unemployment rate is one standard deviation above/below average. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix for each state over time, are in parentheses and p-values are in brackets.

Table 8
How Much Do Demographics Explain Why Moral Hazard Varies
with the State Unemployment Rate?

Dependent variable:	Unemployment Duration							Take-up dummy
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
log(Average UI WBA) (A)	-0.605 (0.295) [0.040]	-0.583 (0.291) [0.045]	-0.606 (0.295) [0.040]	-0.584 (0.290) [0.044]	-0.504 (0.302) [0.094]	-0.488 (0.322) [0.130]	-0.447 (0.309) [0.147]	0.105 (0.079) [0.192]
log(Average UI WBA) × (B) log(State Unemp. Rate / National Unemp. Rate)	1.256 (0.433) [0.004]	1.244 (0.434) [0.004]	1.255 (0.427) [0.003]	1.249 (0.437) [0.004]	1.268 (0.442) [0.004]	1.211 (0.413) [0.003]	1.240 (0.413) [0.003]	-0.313 (0.134) [0.024]
log(State Unemp. Rate / National Unemp. Rate)	-0.014 (0.119) [0.904]	-0.015 (0.119) [0.900]	-0.014 (0.119) [0.904]	-0.014 (0.118) [0.906]	-0.022 (0.117) [0.854]	-0.021 (0.116) [0.859]	-0.018 (0.115) [0.878]	0.117 (0.023) [0.000]
log(Average UI WBA) × Age		0.009 (0.008) [0.284]					0.010 (0.010) [0.281]	
log(Average UI WBA) × 1{Married}			0.005 (0.175) [0.979]				-0.078 (0.202) [0.701]	
log(Average UI WBA) × Years of Education				0.050 (0.026) [0.053]			0.051 (0.031) [0.101]	
Number of Spells	4307	4307	4307	4307	4307	4307	4307	16322
log(Average UI WBA) × Occupation FEs	N	N	N	N	Y	N	Y	Y
log(Average UI WBA) × Industry FEs	N	N	N	N	N	Y	Y	Y
Post-estimation: (A) + σ × (B)	-0.348 (0.294) [0.236]	-0.328 (0.292) [0.262]	-0.348 (0.293) [0.235]	-0.328 (0.289) [0.255]	-0.245 (0.297) [0.410]	-0.240 (0.313) [0.444]	-0.193 (0.299) [0.519]	0.036 (0.064) [0.579]
Post-estimation: (A) - σ × (B)	-0.863 (0.321) [0.007]	-0.838 (0.316) [0.008]	-0.863 (0.321) [0.007]	-0.840 (0.317) [0.008]	-0.764 (0.332) [0.021]	-0.736 (0.352) [0.036]	-0.701 (0.339) [0.039]	0.174 (0.100) [0.083]

Notes: Columns (1) through (7) report semiparametric (Cox proportional) hazard model results from estimating equation (13) using individual-level unemployment spells from 1985-2000 SIPP. See Table 2 for more details on the baseline specification. Column (8) reports OLS estimates using individual-level data from 1968-1987 PSID. The final two rows reports linear combinations of parameter estimates to produce the marginal effects when the state unemployment rate is one standard deviation above/below average. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix for each state over time, are in parentheses and p-values are in brackets.

Table 9
Alternative Measures of the Interaction Term

	Hazard Model Results			Post-estimation	
	(A)	(A) × (B)	(B)	(A) + $\sigma \times$ (B)	(A) - $\sigma \times$ (B)
(1) (A) $\log(\text{Average UI WBA}) \times$ (B) $\log(\text{State Unemp. Rate} / \text{National Unemp. Rate})$	-0.605 (0.295) [0.040]	1.256 (0.433) [0.004]	-0.014 (0.119) [0.904]	-0.348 (0.294) [0.236]	-0.863 (0.321) [0.007]
(2) (A) $\log(\text{Average UI WBA}) \times$ (B) $\mathbf{1}\{\text{State Unemp. Rate} \geq \text{Median}\}$	-1.236 (0.355) [0.001]	0.966 (0.199) [0.000]	-0.016 (0.044) [0.718]	-0.270 (0.314) [0.390]	
(3) (A) $\log(\text{Average UI WBA}) \times$ (B) $(\text{State Unemp. Rate} - \text{National Unemp. Rate})$	-0.564 (0.318) [0.076]	0.136 (0.069) [0.048]	-0.001 (0.018) [0.965]	-0.383 (0.304) [0.207]	-0.745 (0.356) [0.036]
(4) (A) $\log(\text{Statutory Maximum UI WBA}) \times$ (B) $\log(\text{State Unemp. Rate} / \text{National Unemp. Rate})$	-0.228 (0.282) [0.417]	1.219 (0.470) [0.009]	0.010 (0.130) [0.940]	0.021 (0.264) [0.936]	-0.478 (0.328) [0.145]
(5) (A) $\log(\text{Average UI Replacement Rate}) \times$ (B) $\log(\text{State Unemp. Rate} / \text{National Unemp. Rate})$	-0.425 (0.252) [0.092]	1.425 (0.502) [0.005]	-0.010 (0.111) [0.926]	-0.133 (0.274) [0.628]	-0.717 (0.270) [0.008]
(6) (A) $\log(\text{Simulated Average UI WBA}) \times$ (B) $\log(\text{State Unemp. Rate} / \text{National Unemp. Rate})$	-0.925 (0.363) [0.011]	1.220 (0.772) [0.113]	0.009 (0.120) [0.937]	-0.674 (0.350) [0.054]	-1.175 (0.437) [0.007]
(7) (A) $\log(\text{Average UI WBA}) \times$ (B) $-1 * \log(\text{Predicted Employment-to-Pop Ratio})$	-0.736 (0.427) [0.084]	0.867 (1.043) [0.406]	0.687 (1.992) [0.730]	-0.559 (0.326) [0.087]	-0.914 (0.591) [0.122]

Notes: All rows report semiparametric (Cox proportional) hazard model results from estimating equation (13); each column reports separate parameter estimate. Data are individual-level unemployment spells from 1985-2000 SIPP. Number of spells = 4307. See Table 2 for more details on the baseline specification. In row (2), the median unemployment rate across all states in sample is calculated separately each year. In row (3), the Average UI WBA is the average weekly benefit amount paid to individuals claiming unemployment insurance. In row (4), the Statutory Max UI WBA is the maximum weekly benefit amount an individual can receive in a state-year. In row (5), the Average UI Replacement Rate is the Average UI WBA divided by the average weekly wages in a given state-year for prime-age males (computed using the CPS). In row (6), the Simulated Average UI WBA is constructed following the simulated instrumental variables procedure in Currie and Gruber (1996), isolating variation in generosity due to changes in program parameters, holding composition of unemployed constant. In row (7), the Predicted Employment to Population Ratio is computed following the "shift share" procedure of Bartik (1991); see text for details. The final two columns report linear combinations of parameter estimates to produce the duration elasticity when the state unemployment rate is one standard deviation above/below the mean. In row (2) we set $\sigma = 1.0$ because the interaction term includes a dummy variable rather than a continuous measure. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix for each state over time, are in parentheses and p-values are in brackets.

Table 10
How Does Effect of UI on Consumption Change Upon Unemployment Vary with the
Unemployment Rate?

	Gruber (1997)	Replication Sample				
	(1)	(2)	(3)	(4)	(5)	(6)
UI replacement rate	0.280 (0.105)	0.255 (0.116) [0.034]	0.257 (0.115) [0.031]	0.245 (0.112) [0.034]	0.224 (0.128) [0.087]	0.244 (0.116) [0.042]
Implied consumption change at replacement rate of 0	-0.231	-0.253	-0.254	-0.247	-0.237	-0.243
UI replacement rate × log(State Unemp. Rate / National Unemp. Rate)			0.004 (0.213) [0.985]	0.069 (0.221) [0.756]	0.027 (0.223) [0.903]	0.065 (0.232) [0.781]
log(State Unemp. Rate / National Unemp. Rate)			-0.033 (0.155) [0.833]	-0.080 (0.178) [0.654]	0.108 (0.207) [0.605]	-0.095 (0.189) [0.617]
N	1604	1595	1595	1595	1595	1595
R ²		0.099	0.099	0.104	0.164	0.114
State and Year FEs	Y	Y	Y	Y	Y	Y
Region-specific linear trends	N	N	N	Y	N	N
Region × Year FEs	N	N	N	N	Y	N
State-specific linear trends	N	N	N	N	N	Y

Notes: Column (1) reproduces the results from column (4) in Gruber (1997), Table 1. Remainder of columns report OLS results on a replication sample. Data are individual-level observations from 1968-1987 PSID. See text for more details on the baseline specification. The implied consumption drop is computed as the average fitted value across the sample when the replacement rate is set to 0 for all observations. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix for each state over time, are in parentheses and p-values are in brackets.

Table 11
Moral Hazard and Net Liquid Wealth

	(1)	(2)	(3)	(4)	(5)
log(Average UI WBA) (A)	-0.605 (0.295) [0.040]	-0.467 (0.451) [0.300]	-0.563 (0.289) [0.052]		
log(Average UI WBA) × (B)	1.256	1.553	1.427	1.463	1.619
log(State Unemp. Rate / National Unemp. Rate)	(0.433)	(0.734)	(0.466)	(0.481)	(0.594)
	[0.004]	[0.034]	[0.002]	[0.002]	[0.006]
log(Average UI WBA) × log(State Unemp. Rate / National Unemp. Rate) × 1 {1st and 2nd liquid wealth quartiles}					-0.390 (0.698) [0.576]
Number of Spells	4307	2170	4307	4307	4307
3rd and 4th liquid wealth quartiles only	N	Y	N	N	N
Occupation FEs × Liquid wealth quartile	N	N	Y	Y	Y
Industry FEs × Liquid wealth quartile	N	N	Y	Y	Y
Unemployment duration × Liquid wealth quartile	N	N	N	Y	Y
log(Average UI WBA) × Liquid wealth quartile	N	N	N	Y	Y
Post-estimation: (A) + σ × (B)	-0.348 (0.294) [0.236]	-0.149 (0.469) [0.750]	-0.270 (0.295) [0.359]		
Post-estimation: (A) - σ × (B)	-0.863 (0.321) [0.007]	-0.786 (0.482) [0.103]	-0.855 (0.315) [0.007]		

Notes: All columns report semiparametric (Cox proportional) hazard model results from estimating equation (13). Data are individual-level unemployment spells from 1985-2000 SIPP. See Table 2 for more details on the baseline specification. The final two rows reports linear combinations of parameter estimates to produce the duration elasticity when the state unemployment rate is one standard deviation above/below average. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix for each state over time, are in parentheses and p-values are in brackets.

Table 12
Sufficient Statistics Calibrations:
Optimal UI and the Unemployment Rate

u	3.6%	4.9%	6.2%	7.5%	8.8%
$\varepsilon_{D,b}$	1.288	0.901	0.605	0.366	0.165
<u>Panel A: Coefficient of Relative Risk Aversion, $\gamma = 2$</u>					
r^*	0.0%	0.0%	10.9%	38.3%	65.8%
b^*	\$0	\$0	\$38	\$134	\$230
<u>Panel B: Coefficient of Relative Risk Aversion, $\gamma = 3$</u>					
r^*	0.0%	19.7%	38.9%	57.1%	75.3%
b^*	\$0	\$69	\$136	\$200	\$263
<u>Panel C: Coefficient of Relative Risk Aversion, $\gamma = 4$</u>					
r^*	23.6%	39.2%	53.3%	66.7%	80.1%
b^*	\$83	\$137	\$187	\$234	\$280

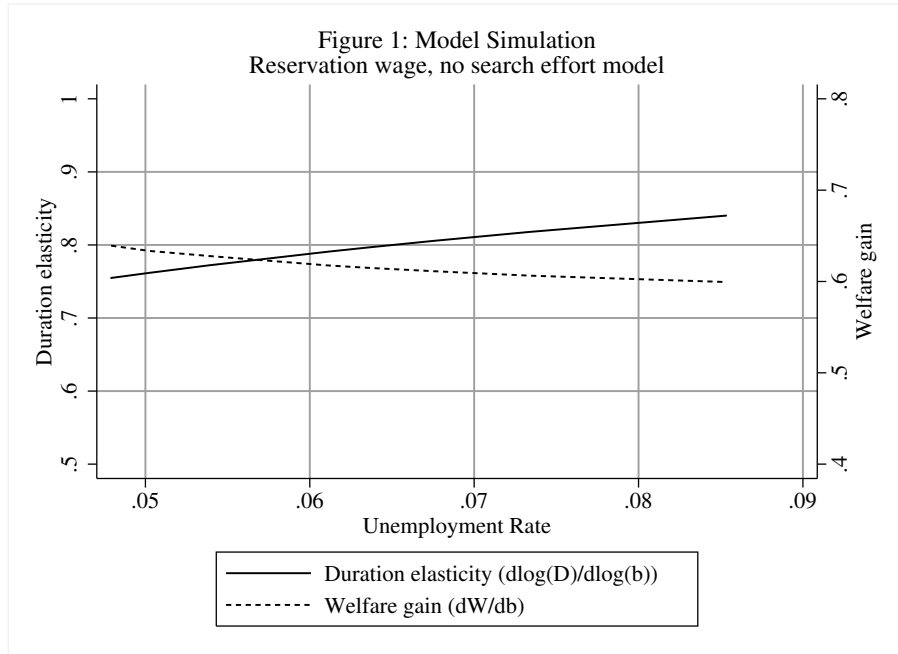
Notes: All columns report optimal UI benefit levels at various levels of the unemployment rate. Subsequent rows report the elasticity of unemployment duration with respect to UI benefit level, the optimal UI benefit level (b^*) and the optimal UI replacement rate (r^*). Both b^* and r^* are censored at 0. The optimal replacement rate is computed by dividing UI benefit level by the average wage. See Section 4.2 for more details on the computations. The optimal benefit level is computed assuming a weekly wage of \$350.

Appendix Table A1

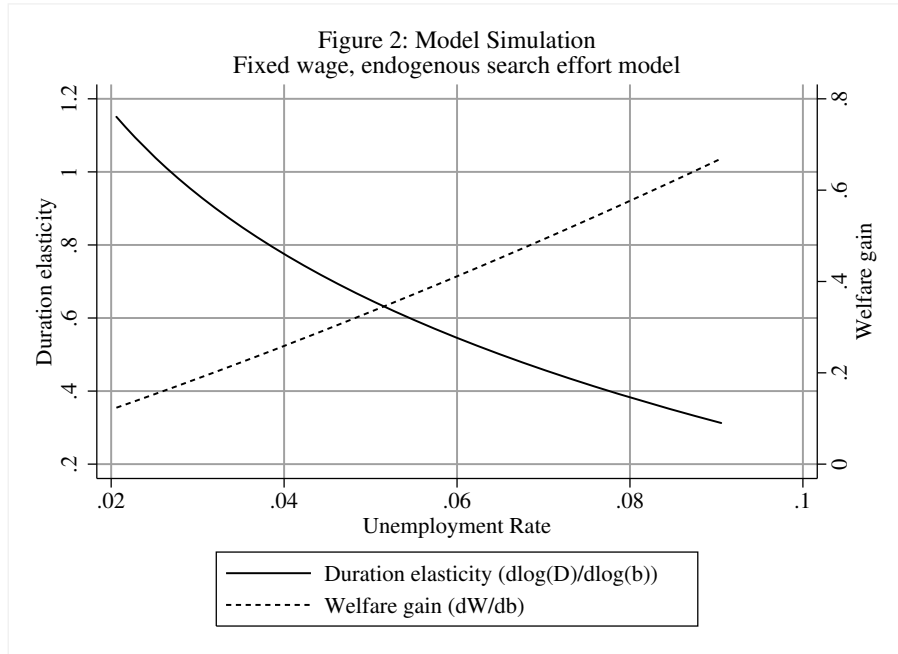
Other Robustness Tests: Alternative Controls, Nonlinear Direct Effects, Extended Benefits, Control Function

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
log(Average UI WBA) (A)	-0.605 (0.295) [0.040]	-0.584 (0.297) [0.049]	-0.438 (0.193) [0.023]	-0.408 (0.190) [0.032]	-0.534 (0.426) [0.210]	-0.262 (0.516) [0.611]	-0.551 (0.244) [0.024]	-0.819 (0.430) [0.057]
log(Average UI WBA) × log(State Unemp. Rate / National Unemp. Rate) (B)	1.256 (0.433) [0.004]	1.243 (0.446) [0.005]	0.896 (0.537) [0.095]	0.759 (0.529) [0.151]	1.267 (0.463) [0.006]	1.707 (0.437) [0.000]	1.419 (0.413) [0.001]	1.074 (0.613) [0.080]
Stratified baseline hazard	Y	N	N	N	N	Y	Y	Y
State, Occupation, Industry FEs	Y	Y	N	N	Y	Y	Y	Y
Age, Marital Dummy, Education, Wage Spline	Y	Y	Y	N	Y	Y	Y	Y
Quadratic in State Unemp. Rate, Average UI WBA	N	N	N	N	Y	Y	N	N
Cubic in State Unemp. Rate, Average UI WBA	N	N	N	N	N	Y	N	N
Controls for potential duration and extended benefits	N	N	N	N	N	N	Y	N
Instrument log(Avg. UI WBA) with log(Simulated Avg. UI WBA)	N	N	N	N	N	N	N	Y
Post-estimation: (A) + σ × (B)	-0.348 (0.294) [0.236]	-0.329 (0.296) [0.266]	-0.254 (0.191) [0.184]	-0.253 (0.198) [0.203]	-0.274 (0.409) [0.503]	0.088 (0.520) [0.866]	-0.261 (0.261) [0.317]	-0.599 (0.435) [0.169]
Post-estimation: (A) - σ × (B)	-0.863 (0.321) [0.007]	-0.839 (0.324) [0.010]	-0.621 (0.249) [0.013]	-0.564 (0.237) [0.017]	-0.793 (0.462) [0.086]	-0.612 (0.526) [0.245]	-0.842 (0.256) [0.001]	-1.039 (0.460) [0.024]

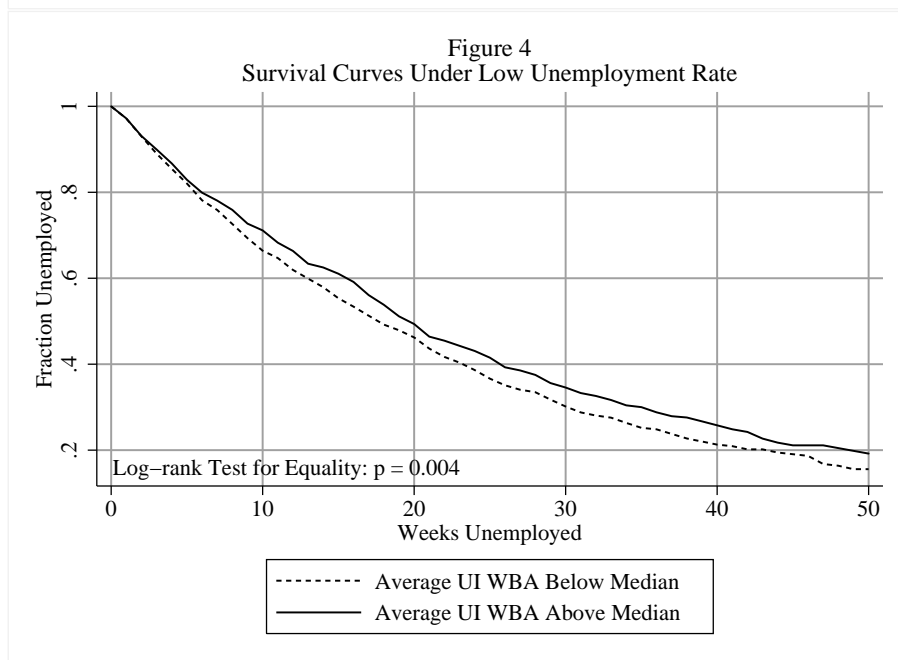
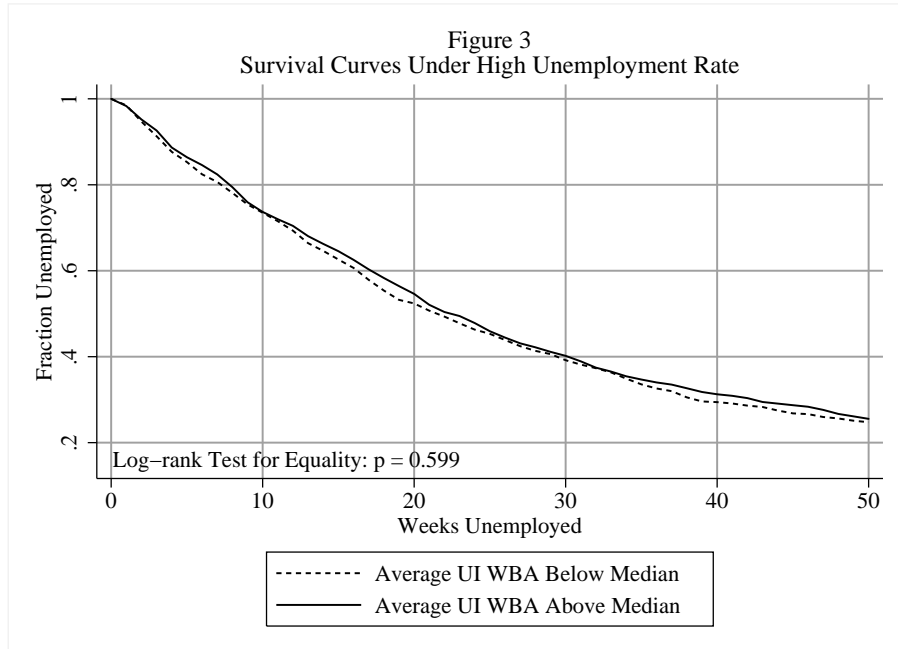
Notes: All columns report semiparametric (Cox proportional) hazard model results from estimating equation (13). Data are individual-level unemployment spells from 1985-2000 SIPP. Number of spells = 4307. See Table 2 for more details on the baseline specification. Column (7) controls for maximum potential duration of benefits by setting log(Average UI WBA) to 0 for all weeks beyond this maximum number of weeks (where the maximum accounts for extended benefits programs). Column (8) reports results from a two-stage instrumental variables specification using log(Simulated Average UI WBA) as an instrument for log(Average UI WBA), where in the second stage a fifth-order polynomial in the first stage residuals is used as a control function. See text for details on the simulated instrument. The final two rows report linear combinations of parameter estimates to produce the duration elasticity when the state unemployment rate is one standard deviation above/below the mean. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix for each state over time, are in parentheses and p-values are in brackets. In column (8), the standard errors are bootstrapped standard errors based on 1000 replications, sampling states with replacement.



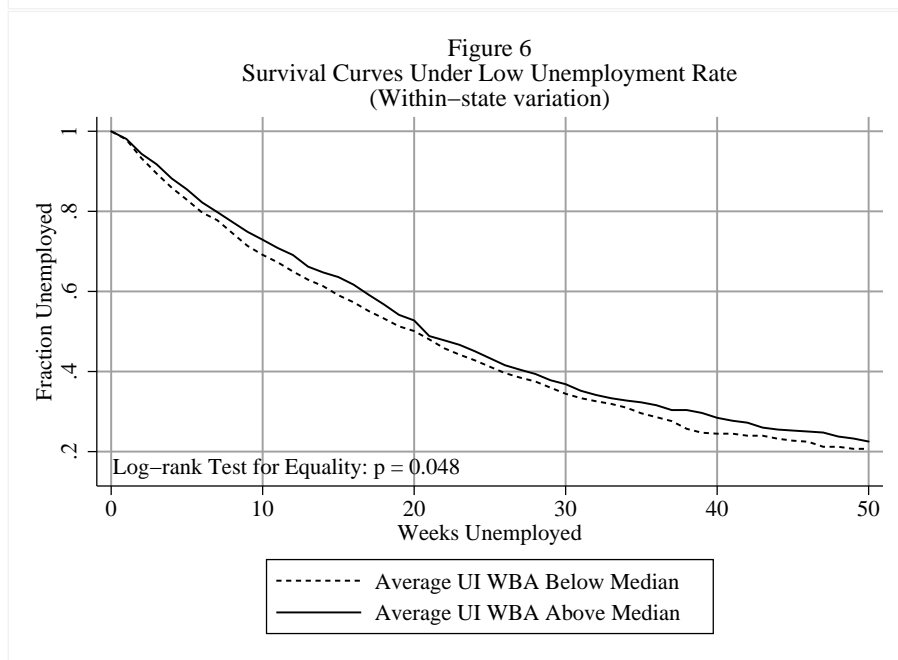
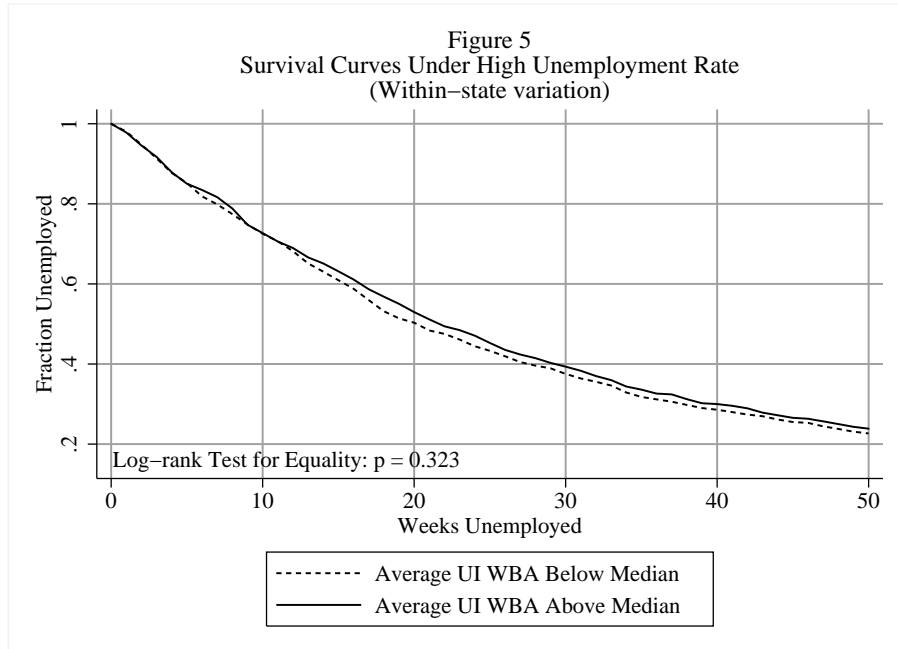
Notes: This figure is generated by calibrating the job search model in the main text with the following parameters. The wage distribution is assumed to be log-normal with mean 0.1 and standard deviation of 0.3. The benefit level is set to 0.067. There is no discounting and interest rate is set to 0 (i.e., $r = \rho = 0$). The job offer arrival rate is $\lambda(e, \alpha) = \alpha$; i.e., there is no search effort decision. The job separation rate is $s = 0.002$.



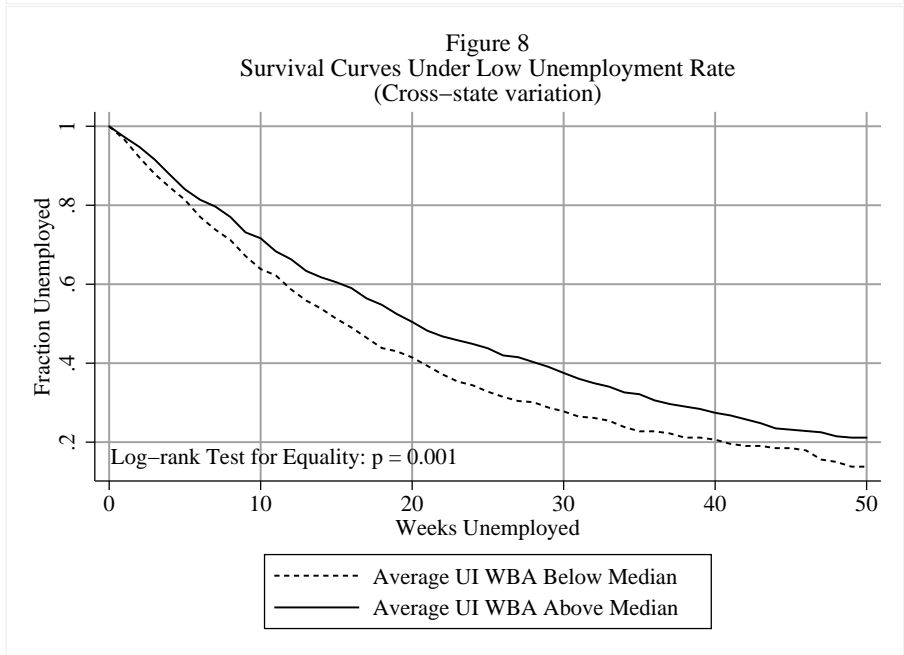
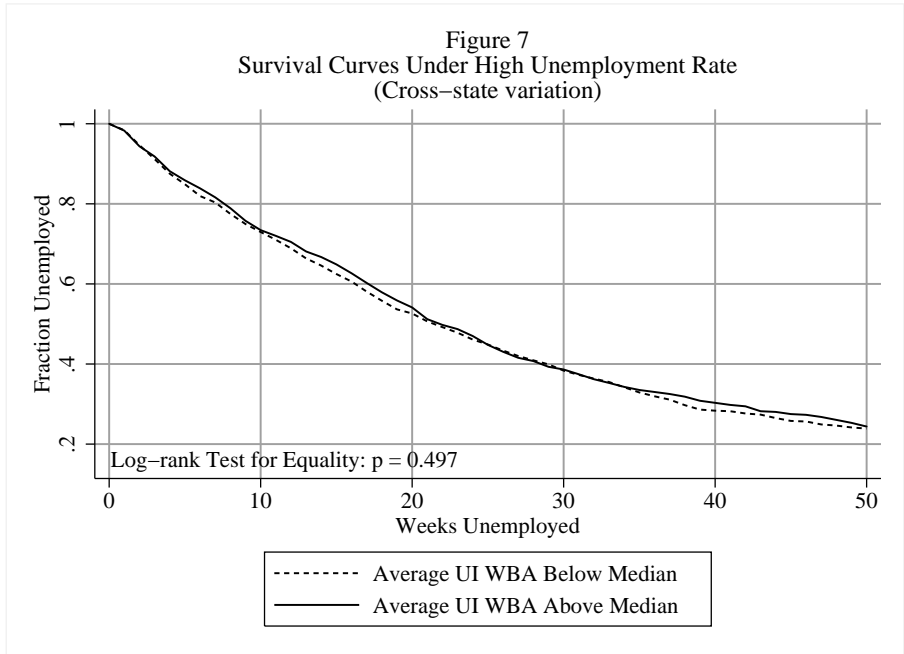
Notes: This figure is generated by calibrating the job search model in the main text with the following parameters. The wage distribution is degenerate with mean 0.25. The benefit level is set to 0.125. There is no discounting and interest rate is set to 0 (i.e., $r = \rho = 0$). The job offer arrival rate is $\lambda(e, \alpha) = \Lambda e^\alpha$, with $\Lambda = 0.0667$. The cost of search is ϕe^κ , with $\phi = 0.3$ and $\kappa = 1.1$. The job separation rate is $s = 0.0089$.



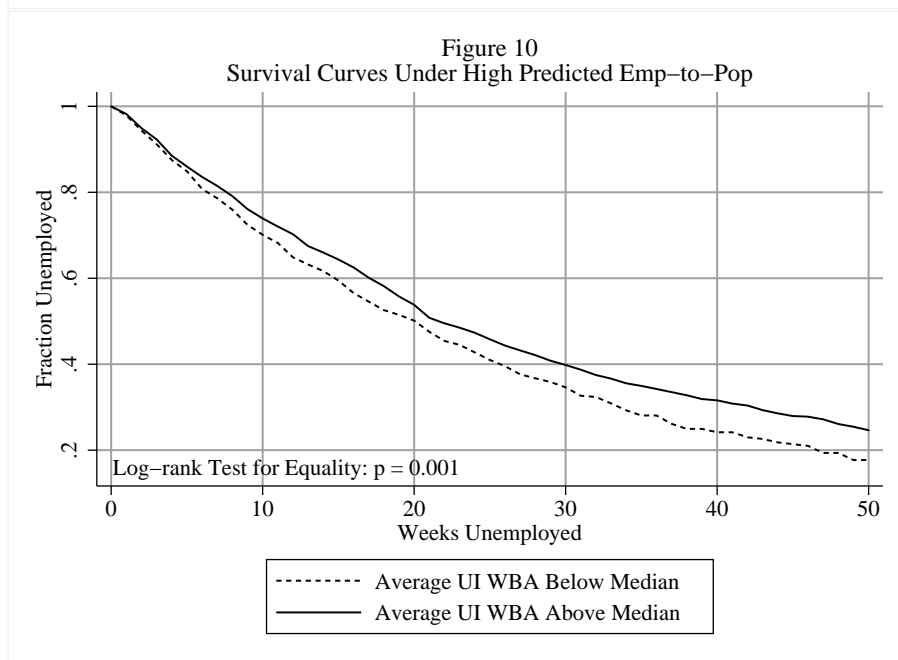
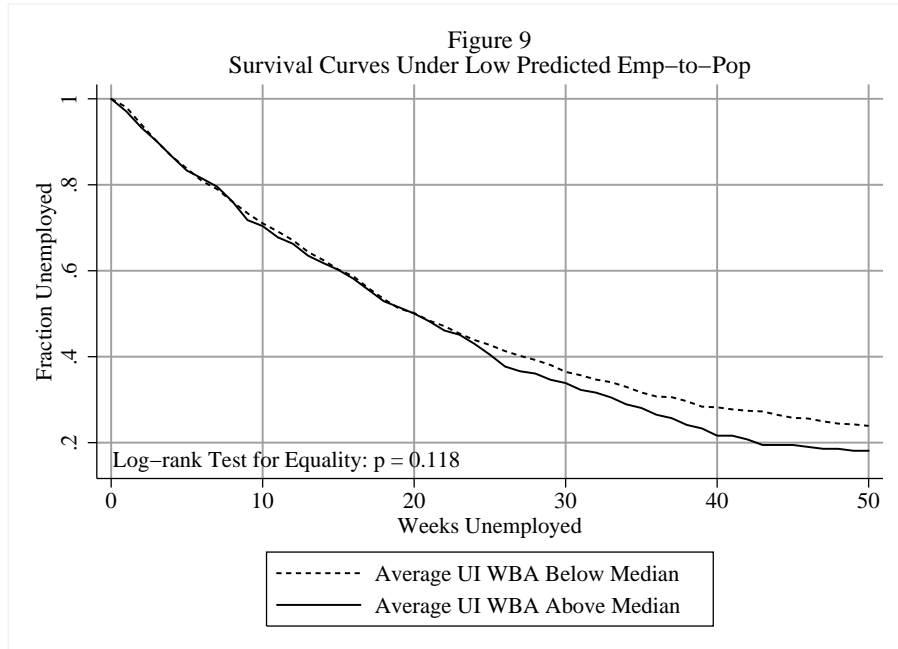
Notes: Data are individual-level unemployment spells from 1985-2000 SIPP. Each figure plots (Kaplan-Meier) survival curves for two groups of individuals based on whether or not Average UI Weekly Benefit Amount (WBA) in individual's state is above or below the median. The survival curves are adjusted following Chetty (2008), which parametrically adjusts for "seam effect" by fitting a Cox proportional hazard model with a seam dummy and then recovering the baseline hazard.



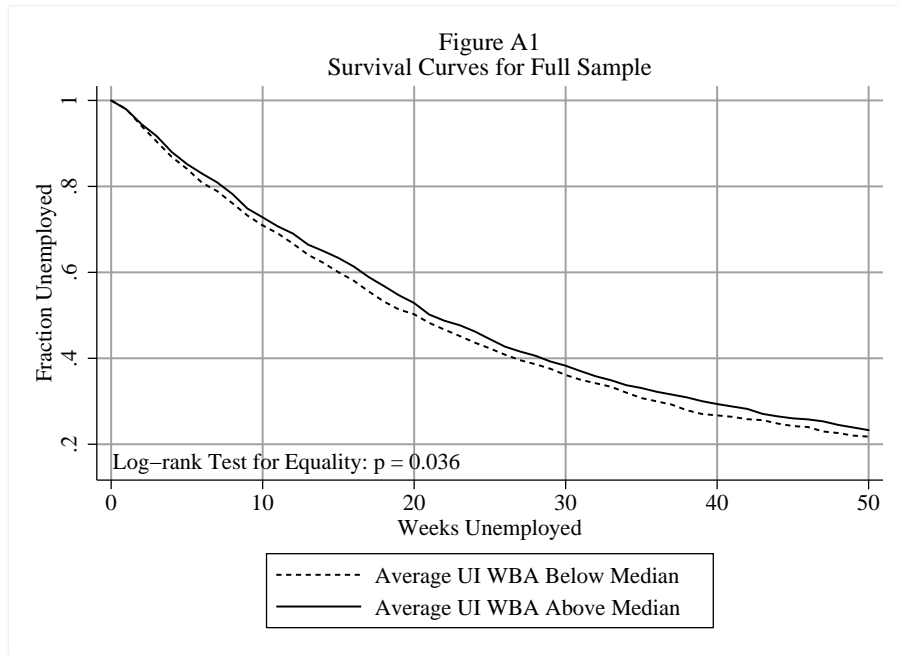
Notes: Data are individual-level unemployment spells from 1985-2000 SIPP. Each figure plots (Kaplan-Meier) survival curves for two groups of individuals based on whether or not Average UI Weekly Benefit Amount (WBA) in individual's state is above or below the median. The survival curves are adjusted following Chetty (2008), which parametrically adjusts for "seam effect" by fitting a Cox proportional hazard model with a seam dummy and then recovering the baseline hazard. The figures report results for sub-samples defined depending on whether the unemployment rate is above or below the median unemployment rate in the state during the sample period.



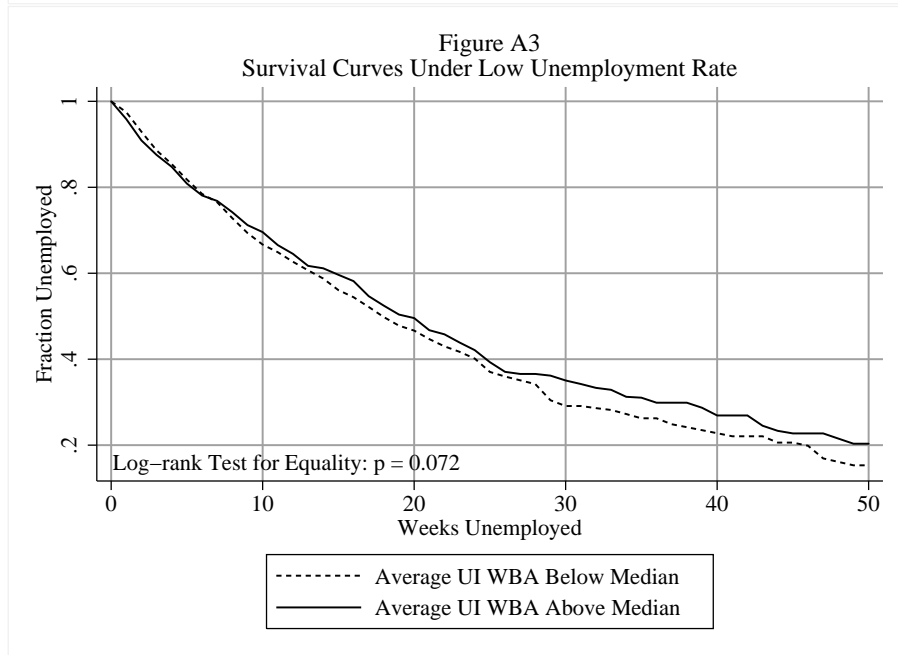
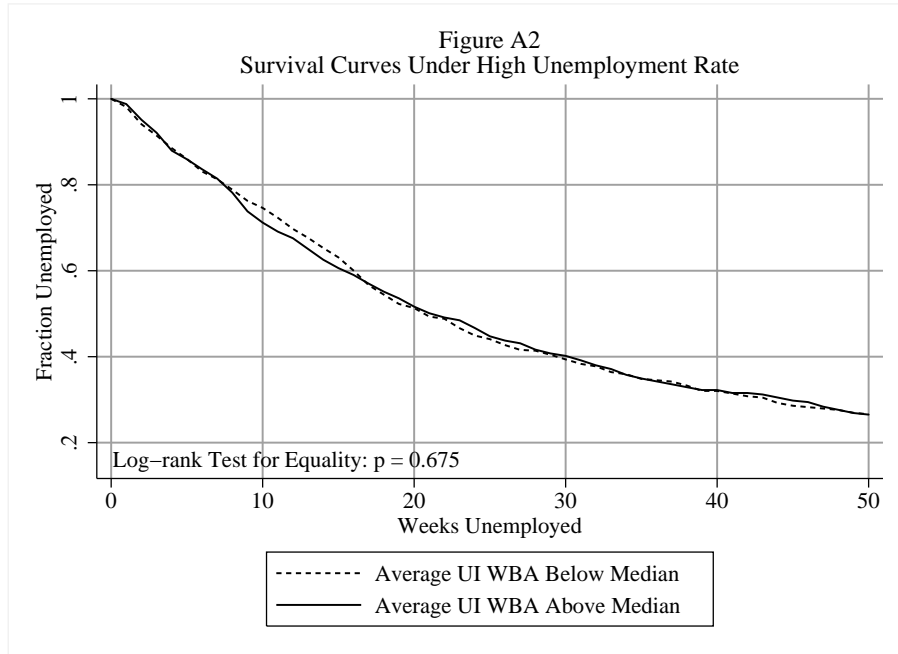
Notes: Data are individual-level unemployment spells from 1985-2000 SIPP. Each figure plots (Kaplan-Meier) survival curves for two groups of individuals based on whether or not Average UI Weekly Benefit Amount (WBA) in individual's state is above or below the median. The survival curves are adjusted following Chetty (2008), which parametrically adjusts for "seam effect" by fitting a Cox proportional hazard model with a seam dummy and then recovering the baseline hazard. The figures report results for sub-samples defined depending on whether the average unemployment rate in the state during the sample period is above or below the median across all states in the sample.



Notes: Data are individual-level unemployment spells from 1985-2000 SIPP. Each figure plots (Kaplan-Meier) survival curves for two groups of individuals based on whether or not Average UI Weekly Benefit Amount (WBA) in individual's state is above or below the median. The survival curves are adjusted following Chetty (2008), which parametrically adjusts for "seam effect" by fitting a Cox proportional hazard model with a seam dummy and then recovering the baseline hazard. The employment-to-population ratio is predicted following Bartik (1991); see text for details.



Notes: Data are individual-level unemployment spells from 1985-2000 SIPP. The figure plots (Kaplan-Meier) survival curves for two groups of individuals based on whether or not Average UI Weekly Benefit Amount (WBA) in individual's state is above or below the median. The survival curves are adjusted following Chetty (2008), which parametrically adjusts for "seam effect" by fitting a Cox proportional hazard model with a seam dummy and then recovering the baseline hazard.



Notes: Data are individual-level unemployment spells from 1985-2000 SIPP. In order to minimize liquidity effects, the sample is limited to individuals with net liquid wealth above the median. Each figure plots (Kaplan-Meier) survival curves for two groups of individuals based on whether or not Average UI Weekly Benefit Amount (WBA) in individual's state is above or below the median. The survival curves are adjusted following Chetty (2008), which parametrically adjusts for "seam effect" by fitting a Cox proportional hazard model with a seam dummy and then recovering the baseline hazard.