Consumption Commitments, Unemployment Durations, and Local Risk Aversion

Raj Chetty*

December 20, 2002

Abstract

Studies of risk preference have empirically established two regularities that are inconsistent with the canonical expected utility model: (1) risk aversion over small gambles greatly exceeds risk aversion over larger stakes and (2) insurance buyers play the lottery. This paper characterizes risk preferences both theoretically and empirically in a world with two consumption goods, one of which involves a commitment in that an adjustment cost must be paid when the good is sold. In this model, utility over wealth is more curved locally than globally: individuals are more risk averse with respect to moderate-scale income fluctuations than they are to large income fluctuations. Commitments also create a gambling motive. To test the empirical importance of commitments, I apply the labor-supply method of estimating risk aversion developed in Chetty (2002). Variations in unemployment insurance laws are used to estimate local curvature using both reduced-form hazard models and structural estimation of a dynamic job search model. I find that commitments significantly change preferences over wealth. The local coefficient of relative risk aversion is an order of magnitude larger than the global coefficient of relative risk aversion. Implications for optimal unemployment insurance are considered. When commitments are taken into account, the optimal rate of insurance rises from zero to 44% of pre-unemployment wages.

*E-mail: chetty@fas.harvard.edu. I am particularly indebted to Martin Feldstein, Gary Chamberlain, Larry Katz, Jerry Green, and Caroline Hoxby for their guidance and support. I thank Nava Ashraf, Julie Cullen, David Cutler, Amy Finkelstein, Ed Glaeser, Jon Gruber, Jens Hilscher, Emir Kamenica, David Laibson, Adam Looney, Casey Mulligan, Jonah Rockoff, Emmanuel Saez, Jesse Shapiro, Monica Singhal, Jeremy Tobacman, and seminar participants at Harvard University for helpful comments and discussions. Philippe Bouzaglou and David Lee provided excellent research assistance. Financial support from the National Science Foundation, Harvard University, NBER, and State Farm is gratefully acknowledged.
The United States spent approximately $200 billion on social insurance in 2000.\textsuperscript{1} The large size and continued growth of social safety nets in the US and other countries pose an interesting puzzle in public finance. Work by Baily (1978) and Gruber (1997) suggests that the optimal benefit level for unemployment insurance is lower than 20% of an individual’s pre-unemployment wage, and perhaps even 0.\textsuperscript{2} The reason is that empirical estimates of the distortionary costs of unemployment insurance tend to outweigh its consumption-smoothing benefits for reasonable levels of risk aversion. The difference between these theoretical predictions and actual benefit levels, which are around 50% of pre-unemployment wages, is striking.

This “social insurance puzzle” can be viewed as one manifestation of the general problem of explaining moderate-stake risk aversion using canonical expected utility theory.\textsuperscript{3} Rabin (2000) gives a powerful calibration theorem which shows that expected utility with strictly concave utility over wealth cannot explain observed risk preferences over small and large gambles simultaneously. To take one notable example, an expected utility maximizer who rejects a 50-50 bet of losing $1000 and winning $1010 will reject a 50-50 bet of losing $100,000

\textsuperscript{1}Definitions of “social insurance” vary; in this paper, I define social insurance as programs that insure income fluctuations. Unemployment Insurance costs amounted to $25 billion in 2000 (Source: Ways and Means Green Book 2000); Workers Compensation, $56 billion (National Academy of Social Insurance); and Survivors Benefits and Disability Insurance, $113 billion (Office of Management and Budget).

\textsuperscript{2}The zero replacement rate result assumes a coefficient of relative risk aversion less than 2, as estimated e.g., in Chetty (2002). Higher levels of risk aversion obviously increase the optimal replacement rate.

\textsuperscript{3}An alternative view is that the existence of social insurance reflects government inefficiency. Those who take this view may prefer to resolve the puzzle using a political economy model that explains why extensive social safety nets could emerge even if they do not improve social welfare. This approach is not considered here.
or winning any amount. In other words, the level of risk aversion implied by typical preferences over gambles involving moderate stakes – such as the demand for a substantial amount of unemployment, automobile, and non-catastrophic health insurance – translates into unrealistic levels of risk aversion over larger stakes.

An important assumption in Rabin’s calibration argument and in most existing models of risk preference is that agents consume one good, i.e. consumption is a composite commodity. This assumption is innocuous if agents can substitute freely among consumption goods at all times. In practice, however, substituting among goods within a period is costly. Many goods, such as housing and vehicles, involve “commitments” – a transaction cost must be paid to change consumption of these goods. Because of these adjustment costs, families tend to retain most of their commitments during small or short-lived income fluctuations. For instance, Gruber (1998) finds that less than 5% of the unemployed move out of their homes during a spell.

The central hypothesis of this paper is that incorporating consumption commitments into the expected utility framework can explain the social insurance puzzle as well as several other phenomena that are inconsistent with a one-good model. While the effects of adjustment costs have been studied in various contexts, their effects on preferences over wealth have not been fully explored. I develop a model in which agents consume two goods, housing and food.

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4 This claim requires that agents reject the $1000/$1010 bet at all wealth levels, but similar results are obtained when the range of wealth levels over which risk preferences are observed is restricted.
An adjustment cost must be paid to sell a house.\footnote{The adjustment cost need not be monetary; fixed psychological adjustment costs would have similar effects on preferences over wealth.} In this model, agents are relatively more risk averse with respect to moderate-scale income fluctuations than they are to large-scale income fluctuations. The intuition is easiest to see in a model with unrestricted borrowing and saving. In this world, agents rationally choose not to move out of their houses while wealth fluctuates within an \((s, S)\) band defined by their previous housing choice. Since the marginal utility of wealth diminishes more quickly over a smaller set of goods, the “local” curvature of utility over wealth within the \((s, S)\) band is greater than the “global” curvature of utility over wealth outside this band (see Figure 1). For example, if a family needs to reduce total expenditure by 10% and has pre-committed half of its income, they have to reduce spending on items such as food and clothing by 20%. The welfare costs of such a concentrated reduction in consumption can be large, driving up the benefits of insurance with the \((s, S)\) band. Consequently, the value of partial insurance for moderate-scale shocks can be \textit{larger} than the value of insurance against large shocks.

Commitments also rationalize gambling, by effectively generating non-concavities in utility over wealth. The motive for gambling arises because the marginal utility of wealth at \(S - \varepsilon\) is \textit{lower} than the marginal utility of wealth at \(S + \varepsilon\). This is because the marginal dollar is spent only on food at \(S - \varepsilon\), whereas it is allocated efficiently over both goods at \(S + \varepsilon\). Committed agents prefer bets that have big payoffs; fair gambles that have payoffs within the \((s, S)\) band are rejected. The model therefore offers a parsimonious explanation
for why insurance buyers also play the lottery. This basic lottery-insurance puzzle underlies many important economic phenomena. For example, by explaining the desire to gamble, the commitments model can also explain real wage rigidities: individuals may prefer a gamble in which they get fired with a small probability rather than take a reduction in wages with certainty.

I evaluate the empirical importance of commitments by testing the model against loss aversion, which is the leading existing explanation of moderate-stake risk aversion (see e.g. Rabin, 2000). The utility of loss averse agents is a function of changes in consumption from a reference point rather than levels of consumption. The utility benefit of increases in consumption is smaller than the utility loss of corresponding reductions. This creates a kink in utility at the reference point, making agents locally risk averse. The commitments model is not inconsistent with loss aversion. In fact, it is quite clear that certain types of behavior are not the result of commitments and must come from kinks in utility. For instance, the “first-order” risk aversion over very small stakes ($10-$100) documented e.g. by Kahneman and Tversky (1979) cannot arise from commitments because utility remains locally linear in this model.

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6 The lottery-insurance puzzle has posed problems for expected utility theory since its advent. Friedman and Savage (1948) propose nonconcave utility as an explanation for this behavior in defense of expected utility theory. The issue has been revisited more recently by Becker, Murphy, and Werning (1997) and Hartley and Farrell (2002).

7 More precisely, loss aversion must be coupled with “narrow bracketing” – treating individual decisions as isolated events – to explain local risk aversion. While loss aversion is typically a direct assumption about preferences, Brunnermeier (1997) shows that boundedly rational agents with neoclassical preferences can exhibit loss aversion.
However, I argue that the standard loss aversion model does not fit observed preferences over moderate wealth fluctuations unless it is augmented with commitments. The commitments model can be tested against the null hypothesis that individuals are loss averse (and have preferences over wealth that are unaffected by commitments) by exploiting a key difference between the two models: The standard loss aversion model does not predict that utility is highly curved away from the individual’s reference point, whereas the commitments model – which does not have an explicit notion of reference points – predicts high local curvature throughout the \((s, S)\) band.

Implementing this test requires a method of estimating the curvature of utility at various wealth levels. I apply the methodology of Chetty (2002), who shows that the curvature of utility can be estimated from data on labor supply choices. There is a direct connection between the coefficient of relative risk aversion and the elasticities of labor supply with respect to wages and unearned income, holding fixed the degree of complementarity between consumption and leisure. The degree of complementarity can in turn be inferred from data on consumption choices when employment is stochastic. Starting with the case of additive utility – where consumption and leisure are not complementary – Chetty shows that estimates of the coefficient of relative risk aversion are not very sensitive to plausible perturbations in the complementarity parameter. It follows that data on labor supply choices is sufficient to make inferences about risk aversion at specific wealth levels.\(^8\)

\(^8\)The reader is referred to Chetty (2002) for further details of the connection between risk aversion and labor supply.
In the present paper, I first show that the estimate of the coefficient of relative risk aversion in Chetty (2002), which is derived from labor supply responses to permanent income and wage changes, gives an upper bound for the global curvature of utility over wealth. I then show that the local curvature of utility can analogously be estimated using data on agents who tend to maintain their commitments while changing their labor supply in response to wage and unearned income fluctuations. One such group is the set of individuals who experience short spells of unemployment, during which they are unlikely to move out of their houses or sell their cars.9 I show that differences in unemployment insurance laws across states and time create the needed variation in effective wages (net of UI benefits) and unearned income. Importantly, the local curvature estimates from variations in UI laws arise from perturbations in wealth that lie either strictly below or above a loss-averse agent’s reference point, depending on the definition of reference point that is used. Since the curvature of utility is not estimated by “passing through” a kink at a reference point, evidence of high local curvature from this method supports the hypothesis that commitments are an important determinant of preferences over wealth.

To implement this strategy, I use a UI benefits simulation program that assigns each claimant a path of potential benefits based on his state of residence and date of unemployment. I combine this information with data on unemployment durations from the Survey of

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9While it may be surprising that I focus on individuals who become unemployed to estimate local curvature, it is difficult to find other “natural experiments” that cause changes in labor supply in response to small variations in unearned income and wages.
Income and Program Participation. I begin the empirical analysis by estimating reduced-form semiparametric hazard models for job-finding after unemployment. The estimates strongly suggest that elasticity of unemployment duration with respect to unearned income is large relative to the benefit elasticity, which is consistent with high local curvature. For an agent who is unemployed for five weeks, the income elasticity is 0.58 and the benefit elasticity is 0.89. Under strong assumptions that make unemployment durations the outcome of a static labor-leisure problem, these elasticities translate into a local coefficient of relative risk aversion of 9.55, assuming that the agent has no other source of income during unemployment. Introducing other sources of income into the model raises the estimate of local curvature further. The estimate of local curvature is an order of magnitude larger than the estimate of global risk aversion in Chetty (2002), which is approximately 1.

Unfortunately, the simple static approach has its limitations. For instance, it implies a negative compensated wage elasticity at many points during the spell. This suggests that a dynamic model of search behavior under uncertainty is necessary to fully capture the behavior of the unemployed. I structurally estimate a job search model in which unemployed individuals choose consumption and search effort in each week. Numerical dynamic programming methods are used to compute the optimal search effort path given a set of core parameters, including the local curvature of utility. A likelihood function for the observed unemployment duration data is constructed from the unemployment exit probabilities that come out of this model, and its structural parameters are estimated by maximizing the
The point estimate of the local coefficient of relative risk aversion in the search model is 7.37, and the null hypothesis that local and global curvature are equal can be rejected with a p value less than 0.01. This result is robust to a number of alternative specifications of the model. The large difference between local and global curvature suggests that commitments are an important determinant of preferences over wealth. A comparison of the predicted unemployment exit probabilities with the observed probabilities indicates that the local curvature estimate is identified from precisely the variation in UI laws that one would expect based on the reduced-form analysis.

The estimate of local curvature is used to construct an empirical utility over wealth with commitments. This utility escapes Rabin’s calibration critique while explaining a wide variety of other phenomena within a consequentialist expected utility framework. In a standard two-period model of unemployment insurance, the estimate of local curvature implies an optimal replacement rate for pre-unemployment wages of 44%, whereas the canonical one-good model predicts 0.10.

The remainder of the paper is organized as follows. Section 1 presents an overview of the commitments model and its main theoretical implications. Section 2 shows how labor supply data can be used to test this model. Section 3 outlines the unemployment insurance law variation, describes the data, and provides reduced-form estimates of the relevant

\footnote{Of course, the results do not imply that government provision of these benefits is optimal; other methods of insurance provision could improve welfare further.}
elasticities. These elasticities are mapped into an estimate of local curvature assuming a static model. The next section presents the dynamic search model used to estimate local curvature structurally. Section 5 summarizes details of the estimation procedure and reports the parameter estimates. The final section discusses some empirical implications of the commitments model and analyzes optimal unemployment insurance with consumption commitments.

1 Commitments and Risk Preferences

The leading example of a good involving a commitment is housing. Changing consumption of housing typically entails moving expenses and large transaction costs. Broker fees are around 5% of a home’s listing price, and early termination of a lease typically entails at least a one month penalty in rent. In addition, leaving a neighborhood and changing children’s schools could have direct welfare costs. More generally, any durable good involves some commitment, as market resale values are significantly lower than actual values because of asymmetric information and other market imperfections. The difference between the price of a new car and a one year old car suggests that the loss from reselling a durable is at least 10% of its value.11 Finally, a number of services require explicit contracts and monthly payments – examples include personal health insurance, health clubs, cellular phones, and

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11 This figure is based on author’s calculations from a sample of twenty 2001 model vehicles sold in the U.S. On average, a one-year old model with 10,000 miles in excellent condition sells for 23% less than the original sticker price. Assuming straight-line depreciation over ten years, the implied resale loss is 13%.
cable television.\textsuperscript{12} Informal evidence suggests that the cost of cancellation for these services is quite high, ranging from 50-100\% of the committed amount.\textsuperscript{13} Goods involving commitments comprise a significant portion of most households’ budgets.\textsuperscript{14} Table 1 shows that the typical household in the US allocates approximately half of its net-of-tax income to the costly-to-adjust goods described above.\textsuperscript{15}

Adjustment costs have been modeled formally since Scarf’s (1959) pioneering work on inventory decisions. In the consumption literature, Grossman and Laroque (1990) show that the low level of covariance between consumption and asset prices can be explained by transaction costs for a durable good. The subsequent literature on illiquid durables has tested this model and found that it provides a good description of many types of consumption decisions; individuals appear to change consumption of illiquid durables infrequently.\textsuperscript{16}

However, existing work has not fully explored the effects of commitments on preferences over

\textsuperscript{12}Note that late payments do not eliminate commitments. In the event of an income shock, the individual must still eventually pay the bill. The ability to make late payments should be viewed as an additional credit channel: it helps only by offering people longer smoothing horizons to dampen shocks.

\textsuperscript{13}For example, the cost of cancelling a one year contract for four leading cellular phone services in the Boston area ranges from $150-$200. Three major health clubs in the Boston area charge non-refundable initiation fees ranging from $300-$1500; another requires proof that one has moved more than 25 miles away from the club before a $700 one year contract can be terminated.

\textsuperscript{14}One may wonder why households choose to commit so much if the welfare costs of income fluctuations are large. Why not rent instead of purchase, in an effort to reduce commitments? This argument ignores market imperfections such as incomplete information, which make ownership advantageous. Glaeser and Shapiro (2002) argue that the agency problems associated with renting create strong incentives for homeownership. Agency costs also explain why it is often rational to purchase rather than rent cars, appliances, and furniture.

\textsuperscript{15}At least the first five categories in Table 1 could involve commitments. Certain expenditures, such as health care and education, may be committed or discretionary. As the exact fraction of committed expenditures is not crucial to the analysis that follows, this definition is left open to interpretation.

\textsuperscript{16}For instance, Attanasio (2000) estimates an \((s, S)\) model using data from the Consumer Expenditure Survey to show that car stocks are adjusted when non-durable consumption falls to \(1/3\) of its original level or doubles relative to the current value of the car (e.g. due to depreciation of the capital good).
wealth.\textsuperscript{17}

To do so, consider a \( T \) period economy with two consumption goods, housing \((h)\) and food \((f)\). Selling a house that provides \( h \) units of housing services per period entails a transaction cost \( kh \).\textsuperscript{18} Choose units so that the price of both \( f \) and \( h \) is fixed at 1. The agent’s income stream is as follows: Income is fixed at \( y_0 \) in period 0. In period 1, the agent earns stochastic income \( y_1 \sim G(y_1) \). Finally, in periods 2 through \( T \), the agent earns a fixed income of \( y_2 \). Let us abstract from capital market imperfections that prevent the agent from borrowing against future income. To reduce notation, assume that the agent does not discount future utility and that the interest rate is 0.

Suppose the agent has Cobb-Douglas preferences over the two consumption goods and an indirect utility function over wealth that is CRRA (in the absence of any commitments):

\[
u(f,h) = \frac{(f^\alpha h^{1-\alpha})^{1-\gamma_g}}{1 - \gamma_g}
\]

The reason that \( \gamma_g \) is used to denote the coefficient of relative risk aversion in this indirect utility will become clear shortly. Since there is no uncertainty after period 1, the agent’s consumption from period 1 onward is perfectly smooth. Therefore, his decision rule can be stated in terms of the present value of his wealth in period 1, \( W \). In particular, he follows

\textsuperscript{17}The model of this paper differs from that of Grossman and Laroque (1990) by introducing a second consumption good that does not involve an adjustment cost. The second good generates local risk aversion and non-concavity of utility over wealth, as shown below.

\textsuperscript{18}It does not matter whether \( k \) is a real monetary cost or a psychological cost incurred due to moving. Also, the results remain identical if the transaction cost is incurred when a good is purchased rather than sold (e.g. an initiation fee).
an \((s, S)\) rule for housing choice: there are two numbers \(s < S\) such that the agent stays in the home he previously purchased when \(W \in (s, S)\) and sells his house when \(W \notin (s, S)\).

If the agent does not move, his housing consumption remains fixed at \(h_1 = h_0\) and his food consumption is what remains of his wealth in each period: \(f_1 = \frac{W}{T} - h_0\). When the agent chooses to move out of his home, he pays the adjustment cost \(k\) and reoptimizes his consumption bundle, choosing

\[
f_1 = \frac{\alpha}{T}(W - kh_0), h_1 = \frac{1 - \alpha}{T}(W - kh_0)
\]

It follows that the agent’s utility over wealth in period 1 when he comes into the period with a house \(h_0\) is

\[
v(W; h_0) = \begin{cases} 
    c_1 h_0^{(1-\alpha)(1-\gamma^g)}(W - h_0)^{\alpha(1-\gamma^g)} & W \in (s, S) \\
    c_2 (W - kh_0)^{1-\gamma^g} & W \notin (s, S)
\end{cases}
\]

where \(c_1\) and \(c_2\) are constants.

Figure 1 illustrates the shape of indirect utility over wealth in period 1, after a commitment has been made.\(^{19}\) Since \(v(W)\) must be continuous in \(W\), the values of \(s\) and \(S\) can be determined by calculating the two values of \(W\) at which the value of staying in one’s house and moving are equal. We can now formally state the first major feature of risk preferences

\(^{19}\)We do not need to characterize the optimal choice of \(h_0\) because the results of this section apply for any value of \(h_0\).
with consumption commitments. Let $\gamma(w) = -\frac{vvwwww}{vw}$ denote the coefficient of relative risk aversion at wealth level $W$.

**Proposition 1 [Local Risk Aversion]:** Committed agents are more risk averse locally than globally:

$$\gamma(W) > \gamma(W') \quad \forall W \in (s, S), W' \notin (s, S)$$

The proof follows directly from computation of $\gamma(W)$ inside and outside the $(s, S)$ band.\(^20\)

The intuition for this result is straightforward: Small wealth shocks are accommodated over a small base of expenditure, with rapidly diminishing marginal utility. For example, if a family needs to reduce total expenditure by 10% because of an income shock and half of their income has been pre-committed, they have to reduce spending on adjustable items such as food and clothing by 20%. Larger shocks cause the individual to move out of his house, reducing the rate at which the marginal utility of money deteriorates. The same logic carries over to a model with multiple commitments: the agent abandons a larger fraction of his commitments as the wealth shock rises in magnitude, reducing the curvature of utility over wealth at wealth levels away from his reference point.\(^21\)

\(^{20}\) For more general specifications of $u(h, f)$, the analog of proposition 1 is that the difference in marginal utilities of wealth between some wealth outcomes $\{w_1, w_2\}$ within the $(s, S)$ band is larger than the difference in marginal utilities for some $\{w_1, w_2\}$ outside the $(s, S)$ band. The stronger statement that the coefficient of relative risk aversion is higher within the $(s, S)$ band than it is outside the $(s, S)$ band is not always true (consider quadratic utility).

\(^{21}\) Optimal policies in a model with more than one committed good are not simple; for instance, the initial guess that agent’s will have a set of $(s, S)$ bands that are progressively larger for goods with larger adjustment costs is not correct. This is because it can be optimal to abandon one large commitment instead of several smaller commitments. What remains true, however, is that utility will be most concave when the agent retains all his commitments and least concave when he does not retain any.
Proposition 1 sheds light on the “social insurance puzzle” that motivated this paper. To see this, suppose income in period 1 is the outcome of a binary lottery: \( y = s - \delta \) with probability \( \frac{1}{2} \) and \( y = S + \delta \) with probability \( \frac{1}{2} \). Now, suppose the agent is offered a fair partial insurance policy that pays \( d > 0 \) in the bad state and charges \( d \) in the good state. We can define the value of this partial insurance policy as the gain in expected utility from purchasing it:

\[
M(d; \delta) = \left\{ \frac{1}{2}v(s - \delta + d) + \frac{1}{2}v(S + \delta - d) \right\} - \left\{ \frac{1}{2}v(s - \delta) + \frac{1}{2}v(S + \delta) \right\}
\]

In a model without commitments, where \( v \) is smooth and strictly concave, \( \frac{\partial M}{\partial \delta} > 0 \) — that is, the value of the partial insurance policy is higher when the initial variance of income is greater. However, this result does not obtain when the agent has commitments. In particular, it can be shown that

\[
\exists \delta, d > 0 \text{ s.t. } M(d, \delta) < M(d, -\delta)
\]

In other words, the value of partial insurance may actually be lower when a mean-preserving spread of the initial wealth outcomes moves these outcomes outside the \((s, S)\) band in which the agent is committed to his house.\(^{22}\) This is because the value of partial insurance is

\(^{22}\text{The argument holds even though the \((s, S)\) band is endogenous to the level of insurance because the value function } v \text{ in the second period is smooth in } s \text{ and } S. \text{ An envelope condition arising from agent optimization guarantees that the changes in the optimal } s \text{ and } S \text{ will be second order with respect to perturbations in the partial insurance policy.}
augmented in situations where commitments are retained and income shocks are concentrated on a small fraction of consumption. For more catastrophic outcomes, a small increment in insurance may not be as valuable because the difference in marginal utility between the good and bad states is not as high.

The converse of this result is that expected utility maximizers with commitments can exhibit high degrees of “moderate stakes” risk aversion, in the terminology of Rabin (2000), while having “non-ridiculous” levels of risk aversion over larger stakes. This is in contrast to the implications of expected utility in the canonical one-good model, where Rabin gives a calibration argument showing that “anything but virtual risk neutrality over modest stakes implies manifestly unrealistic risk aversion over large stakes.” The intuition for Rabin’s result is that risk aversion over moderate stakes places a lower bound on the rate at which marginal utility of money deteriorates; under the assumption that utility is everywhere concave, risk preferences in the small therefore place constraints on feasible risk preferences in the large. These constraints turn out to be very strong: To take one striking example, an expected utility maximizer who rejects a 50-50 bet of losing $1000 and winning $1010 must reject a 50-50 bet of losing $100,000 or winning any amount.

This calibration argument does not apply to committed agents because there are non-concavities in utility over wealth at the threshold levels $s$ and $S$ where the agent abandons his commitment.\(^{23}\)

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\(^{23}\)While commitments could plausibly explain why people reject bets with $1000 stakes, they cannot explain why people rejects bets with $10 or $100 stakes. Explaining this type of “first-order”, documented e.g. by
Proposition 2 [Gambling Motive]: Utility over wealth is locally convex at \( y_1 = s \) and \( y_1 = S \):

\[ \exists \varepsilon > 0 \text{ s.t. } v_W(s - \varepsilon) < v_W(s + \varepsilon) \text{ and } v_W(S - \varepsilon) < v_W(S + \varepsilon) \]

The proof follows directly from evaluation of the two first derivatives. To see the intuition for why commitments generate a desire to gamble, consider an individual who is deciding whether to buy a candy bar that costs $1 or a fair lottery ticket for $1 that will pay $1 million if he wins. A one good model assumes that the agent will buy one million candy bars if he wins the lottery (or one million units of the composite commodity). This is not optimal because the marginal utility of candy is diminishing, and the agent would be better off getting one candy bar with certainty. However, in a world with commitments, the agent will buy more than just candy if he wins the lottery. While the $1 he has in hand cannot be spent to buy a better house or car, the $1 million dollars can.\(^\text{24}\)

Hence, committed individuals may buy insurance to reduce certain types of wealth fluc-

\(^\text{24}\)Friedman and Savage (1948) were the first to propose non-concave utility as an explanation of the lottery-insurance puzzle. Given economists’ predilection for diminishing marginal utility over goods, several models have been proposed as foundations for the Friedman-Savage utility. Kwang (1965) presents a model in which some goods are indivisible, creating an incentive to gamble. Flemming (1969) considers a model of upgrades in which agents consume an illiquid durable whose quality is continually improving over time; in this model, agents’ utilities over wealth have non-concavities at each of the points where it is optimal to upgrade. Both the Kwang and Flemming models do not predict high local curvature of utility specific to an agent’s initial wealth level, and can therefore be distinguished from the model of this paper using the empirical methodology discussed below.
tuations while playing the lottery for a chance at “a new lifestyle.” Agents are more likely to prefer lotteries that yield highly skewed outcomes (big payoffs) rather than those that yield small payoffs and do not give the opportunity to drop many commitments. This prediction fits with Golec and Tamarkin’s (1998) finding that racetrack bettors “crave skewness.”

Propositions 1 and 2 suggest that consumption commitments can potentially shed light on a number of important economic questions. There is a direct connection to models of decisions under uncertainty such as portfolio choice, insurance, entrepreneurial behavior, and executive compensation. Since the intertemporal elasticity of substitution is related to the curvature of utility over wealth when utility is time-separable, the model also has implications for consumption, savings, and labor supply over a lifetime. In addition, optimal tax and welfare policies are functions of the curvature of utility. Finally, the model is relevant for macroeconomic issues such as the existence of real wage rigidities. Committed individuals may prefer a gamble in which they get fired with a small probability rather than take a reduction in wages with certainty. Of course, the empirical importance of commitments for these positive and normative questions depends on the extent to which they raise the local curvature of utility over wealth. The remainder of the paper addresses this question.
2 Estimating Local and Global Curvature

While the commitments model offers an explanation for the difference between local and global risk aversion, other theories such as loss-aversion can also explain the same facts.\textsuperscript{25} The utility of loss averse agents is a function of changes in consumption from a reference point rather than levels of consumption. The utility benefit of increases in consumption is smaller than the utility loss of corresponding reductions. This creates a kink in utility at the reference point, making agents locally risk averse.

I will present evidence showing that the standard loss aversion model does not fit observed preferences over moderate wealth fluctuations (unless it is augmented with commitments). Given the flexibility of both models, it is difficult to determine whether commitments are the primary source of moderate-stake risk aversion by examining choices over gambles directly.\textsuperscript{26} However, there is one key testable difference between the two models: The standard loss aversion model does not predict that utility is highly curved \textit{away from} the individual’s reference point, whereas the commitments model (which does not have an explicit notion of reference points) predicts a high degree of curvature throughout the $(s, S)$ band.\textsuperscript{27}

\textsuperscript{25}Machina (1987) and Camerer (1992) review other non-expected utility theories that have been proposed as explanations for modest-scale risk aversion.

\textsuperscript{26}For instance, one might try to test whether more committed agents are willing to pay more to avoid gambles. However, the lack of exogenous variation in commitments makes it difficult to separate the effect of commitments on risk aversion from the underlying heterogeneity in preferences that led to the varying levels of commitments. In fact, Souleles (2002) finds that homeowners, who are presumably more committed, hold \textit{riskier} portfolios than renters. This should not be interpreted as evidence against commitments: rather, people who have purchased homes are precisely those who expect lower income uncertainty and therefore take greater risk in their investments.

\textsuperscript{27}In fact, Brunnermeier’s (1997) model of loss aversion as an explanation of local risk aversion predicts that utility is \textit{convex} below the reference. Other non-expected utility theories that explain local risk aversion
I test the commitments model against the null hypothesis that individuals are loss averse (and have preferences over wealth that are unaffected by commitments) by estimating the difference between local and global curvature.\textsuperscript{28} Implementing this test obviously requires a method of estimating the curvature of utility at various wealth levels. I apply the methodology of Chetty (2002), who shows that the curvature of utility can be estimated from data on labor supply choices. There is a direct connection between the coefficient of relative risk aversion and the elasticities of labor supply with respect to wages and unearned income, holding fixed the degree of complementarity between consumption and leisure. The degree of complementarity can in turn be inferred from data on consumption choices when employment is stochastic. Starting with the case of additive utility – where consumption and leisure are not complementary – Chetty shows that estimates of the coefficient of relative risk aversion are not sensitive to plausible perturbations in the complementarity parameter. It follows that data on labor supply choices can be used to make inferences about risk aversion at specific wealth levels.

In the present paper, I extend the labor supply methodology to estimate local and global curvature when agents have commitments.

(a) Global Curvature

Chetty (2002) uses estimates of labor supply responses to permanent changes in unearned

\textsuperscript{28} The results of this test should be interpreted carefully. Rejection of the null hypothesis implies that commitments are necessary to explain preferences over wealth. A richer model that allows for both commitments and loss-aversion could explain behavior better.
income and wages to infer the curvature of utility over wealth in a one-good model. In this section, I show that Chetty’s estimate of curvature gives an upper bound for global curvature in a model with commitments. To see this result, let us introduce a binary labor-leisure decision into the $T$ period model of housing and food consumption above. For expositional simplicity, assume that utility is additive in consumption and leisure, i.e. the cross-partial between consumption and leisure is 0.\footnote{See Chetty (2002) for an argument showing that relaxing this assumption changes the estimate of $\gamma$ by at most 0.25 for plausible values of the complementarity parameter. In the interest of space, I consider only the $u_{cl} = 0$ case in this paper.} I adopt the following notation: $l = 0$ when the agent does not work and $l = 1$ when he does; $\psi$ denotes disutility of working; $y$ is unearned income; and $w$ is the net-of-tax wage.

Let us focus on labor supply decisions in period 1, which are made after agents have made commitments in period 0. Suppose there is heterogeneity among agents in disutility of labor: $\psi \sim f_0(\psi)$. As a result of this heterogeneity, some agents who enter period 1 would have been working in the previous period and some would not. I will be interested in calculating the effects of (1) an increase in the wage $w$ and (2) a reduction in unearned income $y$ on labor supply in period 1. Since both of these changes only increase the incentive to work, agents who were already working prior to this change will not change their labor supply at all.\footnote{An interesting feature of the commitments model is that wage increases and decreases need not have effects of the same magnitude. The left and right derivatives of the labor supply function need not be equal because the non-convex adjustment cost makes the marginal agent who responds to a wage increase differ from the marginal agent who responds to a wage decrease. To compute curvature, we need to focus on the same set of marginal agents – this is why I consider the effects of an increase in $w$ and a decrease in $y$. However, in the empirical implementation, I do not distinguish between decreases and increases because $u_{cl} = 0$ case in this paper.} Hence, for our purposes, it is sufficient to focus on the set of agents who
were not working prior to these (unanticipated) perturbations.

To begin, assume that all individuals would move if they were to start working, i.e. $S < T(y + w) - kh_0$. Under this assumption, an agent will begin working in period 1 iff

$$\psi < \hat{\psi}(y, w) \equiv v(T(y + w) - kh_0) - v(Ty)$$

where $\hat{\psi}(y, w)$ is the disutility of labor for the agent who is indifferent between working and not working at wage $w$ and unearned income $y$. The fraction of agents who will begin working in period 1 at a given wage $w$ is

$$\theta(y, w) = \int_{0}^{\hat{\psi}(y, w)} f(\psi) d\psi$$

where $f(\psi)$ denotes the distribution of disutility of labor among agents who were not working in period 0. Now, consider the effect of a small increase in the wage $w$ on $\theta$:

$$\frac{\partial \theta^+}{\partial w} = f(\hat{\psi})v_W(T(y + w) - kh_0)$$

Similarly, the effect of small reduction in $y$ is:

$$\frac{\partial \theta^-}{\partial y} = f(\hat{\psi})\{v_W(T(y + w) - kh_0) - v_W(h_0, y - kh_0)\}$$

existing studies of labor supply have not documented systematic asymmetries in labor supply elasticities.
It follows that

\[
\frac{\partial \theta}{\partial y} \frac{\partial \theta}{\partial w} = \frac{v_W(T(y + w) - kh_0) - v_W(Ty)}{v_W(T(y + w) - kh_0)}
\]  (2)

This expression captures the connection between labor supply and risk aversion. It shows that the percent change in marginal utility of wealth from \(Ty\) to \(T(y + w) - kh_0\) is equal to the ratio of the income and wage effects on labor supply. The intuition for this relationship can be seen by considering two extreme cases. If agents are risk neutral, the marginal utility of wealth is constant, and the level of unearned income would not affect labor supply decisions; in this case, we would observe \(\frac{\partial \theta}{\partial y} = 0\). At the other extreme, if agents are very risk averse, utility is very curved, and a small reduction in unearned income creates a strong incentive to start working, resulting in a large income effect. More generally, the degree of risk aversion is directly related to the size of the income effect. Since the income effect confounds risk aversion with the density of disutility of labor \(f(\hat{\psi})\), we must normalize it by the wage effect to infer the shape of utility over wealth.

Note that the marginal utilities being compared in (2) are at points where the agent has fully optimized consumption over food and housing. Hence, the difference in marginal utilities that is observed through this experiment informs us about the global (unconstrained) curvature of utility over wealth. Translating (2) into a formula for the more familiar coefficient of relative risk aversion \(\gamma^d = -\frac{v_W}{v_w}W\) is a straightforward algebraic exercise. Let \(l\) denote the overall fraction of agents working in period 1 (including those who were already
working in period 0). It follows that

$$\frac{\partial \theta}{\partial y} \frac{\partial \theta}{\partial w} = \frac{\partial l}{\partial y} \frac{\partial l}{\partial w} = \frac{\varepsilon_{l,y}}{\varepsilon_{l,w}} \frac{w}{y}$$

where $\varepsilon_{l,y}$ denotes the elasticity of labor supply with respect to unearned income and $\varepsilon_{l,w}$ the elasticity of labor supply with respect to the wage. With this notation, assuming that the agent has CRRA utility over wealth when he does not have commitments, it is shown in the appendix that we obtain the following estimator for global risk aversion:\(^3\)

$$\bar{\gamma} = \frac{\log[1 - \varepsilon_{l,w} w]}{\log[1 + \frac{w - kh_0/T}{y}]}$$

(3)

This estimator applies when all of the agents sell their homes after they decide to start working. Let us now turn to the more general case in which there is also heterogeneity in the upper wealth cutoff for moving, $S$. This heterogeneity could arise due to variations in commitments and asset holdings (e.g. the type of house $h_0$ that the agent owns). In such an economy, the estimator $\bar{\gamma}$ is an average of the risk aversions of agents who change their labor supply.\(^3\)

Letting $\gamma^i(W)$ denote the (local) coefficient of relative risk aversion for the

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\(^3\)The functional form assumption is necessary because we are making an inference about second derivative properties of utility from a discrete change in wealth. CRRA utility over wealth is obtained when $u(h, f)$ is homothetic. If $u$ is not homothetic, $\gamma$ can be interpreted as the best-arc-fit (CRRA) utility function for the difference in marginal utility of wealth when the agent is working and not working. Loosely speaking, $\gamma$, is the “average” coefficient of relative risk aversion over this range, and $\bar{\gamma}$ gives an upper bound for this average.

\(^3\)See Chetty (2002) for a proof.
agents who do not move, we know that $\gamma^l > \gamma^g$ from Proposition 1. Hence, when some agents do not drop their commitments, there is a weight $\pi \in [0, 1]$ such that

$$\hat{\gamma}^g = (1 - \pi)\gamma^l + \pi \gamma^g \geq \gamma^g$$

If the gain in wealth from working is large, many agents will drop their commitments when they begin to work (i.e. $\pi$ will be large). In this case, $\hat{\gamma}^g$ will tightly bound the true $\gamma^g$.\footnote{Although individuals may not literally move out of their homes, they may buy a new car or new furniture using their higher incomes. In the two good model, $h$ should be interpreted as an aggregate of all these types of consumption.}

Chetty (2002) implements the estimator in (3) and a similar formula for intensive labor supply responses using a diverse set of wage and income elasticities from the existing literature on labor supply.\footnote{Since he is considering a one-good world, Chetty’s estimates effectively assume $k = 0$. However, a simple calibration shows that accounting for the effect of the adjustment cost does not affect the estimates of $\gamma^g$ much. Recalling that $h$ denotes the flow of services from a house, suppose the adjustment cost is 10% of the house’s value, $Th$. Then $kh = 0.1Th \Rightarrow k/T = 0.1$. Suppose the agent has committed 50% of his income prior to period 1, so $h = 0.5y$. Then the mean estimate of $\gamma^g$ rises from 1 to 1.02. The insensitivity comes from the fact that the adjustment cost is small relative to the change in lifetime wealth.} The measures of labor supply include hours worked, participation, and earned income. The source of identification for these elasticities includes welfare reforms, tax changes, cross-sectional variation, and large lottery payouts. Importantly, the labor supply changes induced by these variations tend to result in large, long-term changes in income. Agents are likely to move out of their $(s, S)$ bands when their adjustments in labor supply induce significant shifts in lifetime wealth. Consequently, the curvature estimate...
derived from these “static” labor supply elasticities reflects the global curvature of utility, $\gamma^g$.\footnote{One may be concerned that intensive labor supply responses will not create a sufficient change in income to cause agents to drop their commitments. In this case, using intensive labor supply responses could reduce $\pi$ and bias the curvature estimate upward, toward $\gamma^l$. In practice, this problem is mitigated by the fact that most labor supply changes, even on the intensive margin, tend to be lumpy – individuals work an extra day of the week when their wage rises, not an extra hour. Nonetheless, it is interesting to note that Chetty (2002) estimates $\gamma = 1$ when labor supply is measured by hours worked versus $\gamma = 0.3$ when using participation elasticities. This is consistent with the commitments model.}

Chetty finds an (unweighted) mean estimate from the static labor supply literature of $\tilde{\gamma}^g \approx 1$. In addition, he gives a calibration argument showing that if the uncompensated labor supply curve is upward sloping (as almost all studies of labor supply find), $\tilde{\gamma}^g < 1.25$. This allows us to identify a relatively tight bound for $\gamma^g$ despite the lack of consensus among labor economists about the magnitudes of static labor supply elasticities. The bottom line is that permanent increases in income do not appear to reduce labor supply much, implying that the taste for consumption deteriorates slowly with wealth. In other words, utility is not very curved globally.

**(b) Local Curvature in a Static Model of Unemployment**

To estimate local risk aversion, we need data on agents who maintain their commitments while changing their labor supply in response to wage and unearned income fluctuations. One such group is the set of individuals who experience short spells of unemployment and do not move out of their houses or sell their cars.\footnote{This group may not be representative of the general population. However, if anything, agents who experience unemployment should be less risk averse than the typical person, insofar as they selected into an occupation with greater risk of unemployment. This effect would only bias the estimate of local curvature downward from the population mean.} Variation in unemployment insurance
laws across states and time provide the needed variation in effective wages (net of UI benefits) and unearned income.

Job search during unemployment is intrinsically a dynamic process. However, we can derive a closed-form expression for local curvature under some strong assumptions that make the unemployed agent’s behavior the outcome of a static utility maximization problem. As in Moffitt and Nicholson (1982), let us model the unemployed agent’s problem as a labor-leisure choice within a single period of the $T$ period model discussed above. In particular, suppose the individual loses his job with probability $p$ at the beginning of period 1. Fix the individual’s labor supply at 1 in the state where he does not lose his job, and assume that there is no risk of unemployment in periods $2, ..., T$. While unemployed, the agent gets unemployment benefits of $b$ dollars from the government. Assume that (1) search is costless, (2) jobs are sufficiently plentiful that the individual can choose his labor supply deterministically, and (3) the agent is credit constrained, and cannot borrow from future periods’ income. Once we estimate the agent’s utility $u(h, f)$, the curvature of utility over wealth in a world without credit constraints can be calculated easily.

Let us begin with the case where the adjustment cost $k$ is sufficiently large that the unemployed agent will never choose to move out of his house in period 1. Intuitively, if $k$ is large, the agent will recognize that his income will rise again in subsequent periods and

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37 These assumptions are relaxed in the dynamic search model developed in section 6.
38 The reader may be concerned that we have only discussed the implications of commitments for utility over wealth in a world without credit constraints. The empirical strategy of this paper is to fully estimate the underlying utility over goods, $u(h, f)$, and then make statements about curvature of utility over wealth in the simple setting where there are no credit constraints.
will not find it optimal to pay a significant adjustment cost for a temporary gain in utility. In this case, the unemployed individual chooses \( l \) within period 1 by solving a standard labor-leisure choice problem:

\[
\max_l u(\overline{h}, y + b(1 - l) + wl - \overline{h}) - \psi(l)
\]

where \( \psi(l) \) is the disutility of supplying \( l \) units of labor in period 1. Following a derivation similar to that in section (a), it is shown in the appendix that the curvature of utility over food is

\[
\gamma_f = -\frac{u_{ff}}{u_f} = \frac{f}{w - b(1 - l + y/b)\varepsilon_{1-l,b} - (1 - l)\varepsilon_{1-l,y}}
\]

(4)

It is straightforward to map the estimate of \( \gamma_f \) into local curvature of utility over wealth in a world without credit constraints. It is shown in the appendix that the local curvature of utility over wealth is

\[
\hat{\gamma}'(W) \equiv -\frac{v_{WW}W}{v_W} = \gamma_f(1 + \overline{h})
\]

(5)

When \( \overline{h} = 0 \), agents have no commitments and the curvature of utility over wealth equals curvature of utility over the one consumption good. When \( \overline{h} > 0 \), the curvature of utility over wealth is higher than the curvature of utility over food. This is because a 1% increase in wealth causes food consumption to rise by more than 1%, increasing the elasticity of marginal utility with respect to wealth.
When some people move out of their homes while unemployed, $\hat{\gamma}^l$ gives a lower bound for local curvature, for the same reason that $\hat{\gamma}^g$ gives an upper bound for global curvature.

Two remarks about (5) are in order. First, note that a larger income elasticity (holding the benefit elasticity constant) implies a greater degree of local curvature. Hence, evidence that most of the response in unemployment durations to UI benefits comes through the income channel rather than the price channel would suggest that utility is highly curved locally. If an individual raises labor supply significantly solely to recoup a small loss of income, the lost income must have caused a sharp increase in his marginal utility of wealth.

Second, observe that loss aversion itself would not predict a high value for $\gamma^l$ when it is estimated from variations in UI benefits. If the agent is loss averse and uses pre-unemployment consumption as a reference, we should not necessarily expect a high value of $\gamma^l$ because the curvature of the value function is being estimated from variations in income that are strictly below his kink. In a more sophisticated model of loss aversion where the reference point moves along with the individual’s consumption, a natural assumption would be that the new reference is at the agent’s consumption in the unemployed state. However, since we are considering perturbations of income that would move the agent’s consumption above or below his starting point, which is at the kink, the standard loss-aversion model still does not predict a large value of $\gamma^l$. More generally, the only way in which loss aversion itself could generate a high estimate of $\gamma^l$ is if we “pass through” the kink at the reference when considering the effects of cross-state and time differences in UI benefit levels on behavior.
When we observe labor supply behavior at more than two UI benefit levels, there is no positioning of kinks that could generate a high estimate of curvature between three or more points on the utility function. Thus, a high estimate of $\gamma^l$ should be interpreted as evidence that commitments are an important determinant of preferences over wealth, irrespective of whether or not agents are also loss averse.

(c) Testing the Commitments Model

The derivations above suggest a straightforward test of the empirical significance of commitments:\(^{39}\)

$$\hat{\gamma}^l \geq 1.25 \geq \hat{\gamma}^g$$

In principle, one could also test whether more committed individuals have higher local curvatures than less committed individuals:

$$\frac{\partial \hat{\gamma}^l}{\partial (h/f)} \geq 0$$

This secondary test could be implemented using variations in pre-unemployment housing and vehicle expenses as the source of variation in degree of commitment.\(^{40}\)

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\(^{39}\)This test is imperfect because we are only able to identify bounds on local and global curvature. This test can reject the null hypothesis that commitments do not significantly augment local curvature, but it cannot be used to reject the alternative hypothesis. A purer test of the commitments model would be to look at the labor supply behavior of individuals conditional on their consumption decisions. Unfortunately, the lack of a panel dataset with high frequency information on both labor supply and consumption precludes such an approach.

\(^{40}\)A significant complication in this test is the lack of variation in $\pi$ that is credibly exogenous. As noted above, heterogeneity in preferences that is correlated with choice of commitments is likely to bias the results. The secondary test is not implemented in this paper for this reason and because data on pre-unemployment
One final note: In a model with more than one committed good, utility over wealth will have a spectrum of curvatures, ranging from local to global, as the degree of commitment falls. In this world, $\hat{\gamma}^g$ is an estimate of curvature near the right end of the spectrum, while $\hat{\gamma}^l$ is an estimate nearer to the left end.

That completes the description of the model and basic estimation strategy. The next section outlines the structure of unemployment insurance laws, describes the dataset, and gives some reduced-form empirical results.

3 Reduced-Form Empirical Analysis

3.1 Unemployment Insurance Laws

Unemployment Insurance is a federally mandated program run by individual states. After an involuntary job separation, workers are eligible for benefits from the state government if they have a sufficiently high amount of earned income during the past year. The program is large: In 1998, approximately seven million individuals in the United States received unemployment benefits and the program’s outlays exceeded $20 billion dollars.\(^4\)

In all states, the path of benefits for any claimant who is totally unemployed can be characterized by four basic parameters. First, there is a weekly benefit amount that is a function of the claimant’s past wage history and number of dependents. Second, there is a

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maximum potential duration for benefits, which also depends on the claimant’s past wage history. Third, many states have a waiting week that must be served before benefits are received. Finally, among the states with waiting weeks, some have a “retroactive date” after which the payment for the waiting week is returned. Figure 2 illustrates the variation in laws by graphing potential benefit paths for three claimants earning $400 per week in different states.

The rich variation in the path of benefits across states allows us to study how labor supply responds to variations in unearned income and the net-of-UI wage. The relationship between search behavior and weekly benefit amounts permits identification of an uncompensated price elasticity. One way of estimating the unearned income elasticity is to exploit the waiting period and retroactive date variations. After the first week of unemployment has passed, claimants in states without waiting weeks have effectively received a temporary lump-sum grant relative to claimants in states with waiting weeks. By comparing unemployment exit rates in the two groups of states, we can identify the effect of unearned income on search behavior.

While the waiting period and retroactive payment variation provides a transparent means of identifying unearned income effects, its use is limited by the fact that it creates only a one

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42 This waiting week typically applies only to the first spell if there are multiple spells within one year.
43 It may appear puzzling that a small change in UI benefits could have an income effect because the resulting change in lifetime income is trivial. The reason that temporary income fluctuations have an effect is that the unemployed tend to have very limited liquid assets and are likely to be credit constrained. For instance, Gruber (1997) shows that food consumption during unemployment is highly responsive to the level of UI benefits, confirming that small, transitory income variations do have substantial effects on behavior.
week difference in benefits, which corresponds to roughly $150 at the mean weekly benefit amount. On the other hand, variations in the weekly benefit amount can change potential income by up to $3000 over a spell. These large variations can also be exploited to identify an unearned income elasticity, but the logic is more subtle. This is because variations in weekly UI benefit amounts change unearned income and the net-of-UI wage simultaneously. Previous studies that have examined the effect of UI benefits on unemployment durations have not attempted to decompose this effect into income and price effects. Although it may appear that such a decomposition is impossible because higher UI benefits shift unearned income and the net-of-UI wage collinearly, it is made feasible by the finite duration of UI benefits.

To see how the effect of UI benefits can be decomposed into income and price effects, let us compare the search behavior of identical claimants earning $400 in Massachusetts and California at several points during an unemployment spell. In MA, the claimant receives approximately $250 per week for up to 26 weeks; in CA, he receives $150 per week. Let us first focus on individuals who have already exhausted their UI benefits. Massachusetts residents who have exhausted their benefits have accumulated $2600 more in benefits than their counterparts in California. Since there are no subsequent differences in UI payments, the higher UI benefits paid to those in MA can only affect search behavior after week 26 via the income effect. Now consider the week prior to benefit exhaustion. Here, higher UI benefits not only raise unearned income but also decrease the net-of-UI wage by $100 in
the next week. As a result, in week 25, MA residents have less incentive to search not only because they have more unearned income, but also because they have a lower wage. Finally, consider the first week of an unemployment spell. At this point, MA residents have a lower effective wage for the next 26 weeks but have not yet accumulated more unearned income than those in CA. Higher UI benefits affect behavior primarily through the price channel at the beginning of the spell.

These examples imply that we can make inferences about the underlying curvature of utility by comparing the effect of UI benefits on unemployment exit hazards over time using a dynamic model of search behavior. If an increase in UI benefits reduces the hazard rate much more early in a spell than later in a spell, the price effect must be of greater importance than the income effect, implying based on (4) that utility is not very curved. Conversely, if higher UI benefits continue to reduce hazard rates late into an unemployment spell, income effects must be important and utility must be highly curved locally.44

The strategy of using law variations as a “natural experiment” assumes that the variation in the path of benefits across states and time is exogenous. However, there may be unobservable heterogeneity in the labor market across states that is correlated with both UI laws and unemployment durations. This problem can be mitigated by controlling for observable characteristics such as the local unemployment rate, an individual’s wage, indust-

44A dynamic model of search is necessary because we need to take into account the fact that Michigan residents *anticipate* having higher unearned income while unemployed even at the beginning of the spell; using cumulative UI benefits as a measure of unearned income at a given point in the spell is therefore inappropriate.
try, occupation, education, job tenure, etc. In addition, the group of job losers who do not choose to take UI or are ineligible for UI can be used as a “control group” to test for unobservable variations in labor market conditions across states. The identifying assumption of the analysis that follows is that any remaining heterogeneity after conditioning on these controls is orthogonal to the variation in benefit paths.\textsuperscript{45}

3.2 Data

The data used in this study are from the 1985-1987 and 1990-1996 panels of the Survey of Income and Program Participation. The SIPP collects information from a sample of approximately 30,000 households every four months for a period of two to three years. The interviews I use span the period from the beginning of 1985 to the middle of 2000. At each interview, households are asked questions about their activities during the past four months. Most importantly for this study, each individual is asked to recall his labor force status for each week during the past four months; unemployed individuals are asked whether they received unemployment benefits in each month.\textsuperscript{46} Other data about the demographic and economic characteristics of each household member are also collected.

I make four exclusions on the original sample of job leavers to arrive at my final sample.

\textsuperscript{45} Variation in UI laws has previously been used to estimate the effects of unemployment benefit generosity on unemployment exit rates (Meyer, 1990), spousal labor supply (Cullen and Gruber, 2000), precautionary saving (Engen and Gruber, 1995), and consumption smoothing (Gruber, 1997).

\textsuperscript{46} The ability to identify UI takeup is an advantage of using the SIPP rather than the Current Population Survey. Another advantage is that the SIPP is a panel dataset, making it more suitable to measure unemployment durations. The CPS only gives a cross-section of ongoing spells.
First, following previous studies of UI, I restrict attention to prime-age males (over 18 and under 65). This exclusion allows to focus on the group of agents who were most likely to be active labor force participants and who were maintaining commitments with their income prior to unemployment. Second, I include only the set of individuals who report searching for a job at some point after losing their job, in order to eliminate individuals who have dropped out of the labor force from the analysis. Third, individuals who report that they were on temporary layoff at any point during their spells are dropped, since they might not have been actively searching for a job.\footnote{Katz and Meyer (1990) show that whether an individual considers himself to be on temporary layoff is endogenous to the duration of the spell; recall may be expected early in a spell but not after some time has elapsed since a layoff. Excluding temporary layoffs can therefore potentially bias the estimates. To check that this is not the case, I include temporary layoffs in some specifications of the model.} Finally, I exclude individuals who have less than three months work history within the survey because there is insufficient information to estimate pre-unemployment wages for this group.

The core sample used in the analysis is the remaining set of 4,457 individuals who report receiving UI benefits in the month they became unemployed. I focus on individuals who take up UI within one month because I do not model the takeup decision, assuming instead that all individuals receive UI benefits upon unemployment. The model therefore applies most directly to those who actually begin receiving benefits shortly after unemployment. As a robustness check, I also estimate models that include the additional 2,409 individuals who receive UI at some point in their spell but not in the month they become unemployed.

Table 2 gives summary statistics for the first-month UI claimants. The median UI
recipient is a high school graduate and has pre-UI gross annual earnings of $20,726 in 1990 dollars. Most claimants have limited financial assets before their unemployment spells begin: median liquid wealth net of unsecured debt is only $186.48. As in the general population, wealth inequality is extremely high in this group.

The key independent variables in the analysis are the four aspects of each state’s unemployment insurance law. The raw data for these variables were obtained from the Employment and Training Administration (various years), and supplemented with information directly from individual states. The computation of weekly benefit amounts is non-standard and deserves special mention. For much of the analysis, I use published state average benefits instead of each individual’s actual UI benefit amount, because computational limitations prevent the use of more detailed information in the estimation of the dynamic structural model. However, to check the reduced-form results, I also simulate each individual’s actual UI benefit level. Typically, unemployment benefits are a function of wages earned in the highest quarter during the “base period,” which is defined as the first four quarters of the five quarters preceding the date of unemployment. Unfortunately, reported wages fluctuate significantly in the SIPP and most claimants are not in the sample for a sufficiently long period to calculate their high quarter wages. Rather than attempting to assign each

48 The SIPP contains information about asset holdings only in certain “topical modules” which are administered once during the sample period. Consequently, pre-unemployment asset data is not available for the entire sample.

49 I am grateful to Julie Cullen and Jon Gruber for sharing their simulation programs, and to Suzanne Simonetta and Loryn Lancaster in the Department of Labor for providing detailed information about state UI laws from 1984-2000.
claimant a weekly benefit amount based on the wages he reports, I follow a two-stage procedure to predict weekly benefit amounts. I first predict the claimant’s pre-unemployment annual income using information on education, age, tenure, occupation, industry, and other demographics. The prediction equation for pre-UI annual earnings is estimated on the full sample of individuals who report a job loss at some point during the sample period. I then use the predicted wage as a proxy for the true wage, and assign the claimant unemployment benefits using the simulation program.

The mean weekly benefit amount is $168 and the mean replacement rate for pre-UI earnings is approximately 50%. There is considerable cross-state variation in unemployment benefits, from an average weekly benefit amount of $102 in Louisiana to $217 in Massachusetts in 1990. Fifteen percent of the UI claimants began receiving benefits immediately upon unemployment (i.e. did not have a waiting week), while another 16% received a retroactive payment for the waiting week after their spell extended past a certain length (typically 4 weeks). Thus, 31% of the individuals receive an extra week of benefits up-front.

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50 Since many individuals in the sample do not have a full year’s earning’s history before a job separation, I define the annual income of these individuals by assuming that they earned the average wage they report before they began participating in the SIPP. For example, individuals with one quarter of wage history are assumed to have an annual income of four times that quarter’s income.

51 This approach is needed to reduce attenuation bias in the benefit elasticity. See below for further details.

52 UI benefits are taxed; however, most individuals elect to pay taxes on UI benefits at the end of the year rather than have taxes withheld. I do not make an effort to account for taxes primarily because the noise in the pre-unemployment wage history is so large that it is impossible to identify UI benefits at the individual level precisely; most of the empirical analysis uses state average benefits instead. Moreover, insofar as income during the unemployment spell is the primary determinant of the credit-constrained unemployed (and not income obtained when taxes are filed), the untaxed benefit amount is what is relevant for our analysis.

53 Although some of this variation is due to differences in average wages across states, most of it is the result of differences in the UI laws themselves.
relative to claimants who live in states that have a waiting week and make no retroactive payments. Finally, the maximum potential duration is either 26 or 39 weeks for many claimants. This is because the federal government enacted an emergency program to extend unemployment benefits by thirteen weeks between November 1990 and February 1994 during a national recession. The recession also explains the mean state unemployment rate of 6.6%.

The mean spell lasts for 20.95 weeks before ending or being censored. Approximately one-fourth of all spells and 15% of spells that last for less than six months are right-censored, either because the survey ended before the individual found a job or because the individual stopped participating in the survey.\textsuperscript{54} The mean spell in this sample is considerably longer than mean unemployment duration of approximately 15 weeks reported by the US Department of Labor during the same period. The official figure is computed from the length of ongoing spells for the cross-section of unemployed individuals who report they are looking for work in the CPS. The official definition therefore excludes the spells of individuals who become discouraged and stop searching for work. Unfortunately, these individuals cannot be identified in the SIPP because of the lack of reliable information on search behavior. At a weekly frequency, reports of job search are frequently interspersed with reports that the individual is not looking for a job; moreover, individuals often find jobs after reporting that they were not looking for one. Therefore, the only feasible measure of the length of

\textsuperscript{54}Within the span of the SIPP interviews, 30\% of the claimants experience more than one job separation. I do not treat individuals with multiple spells differently from others. In the future, it would be interesting to investigate whether search behavior during second spells differs from behavior during first spells.
an unemployment spell is to count the weeks from job separation to either job finding or censoring.\textsuperscript{55} This remains a valid definition of the length of unemployment, but it should be distinguished from the more familiar measure, especially when the empirical results below are compared to those of other studies.\textsuperscript{56} Details of the procedure used to compute unemployment durations are given in an appendix.

Figure 3 gives a more complete picture of the pattern of unemployment spells by plotting a survival function for the UI recipients.\textsuperscript{57} In states with waiting weeks, there cannot be any individuals who have a one week unemployment spell and receive UI. Therefore, this figure shows the probability that a spell lasts for $t$ periods or longer given that it lasted for at least two weeks.\textsuperscript{58} The estimated survival rate at $t = 50$ is 14\%, indicating that many individuals become detached from the labor force for an extended period of time after experiencing a job separation that results in unemployment for more than one week.

The solid line in Figure 4 shows the empirical hazards corresponding to the survival curve in Figure 3.\textsuperscript{59} Though the hazard rate fluctuates throughout, there are sharp spikes at $t = 17$ and $t = 35$. These spikes reflect a reporting artefact known as the “seam effect,” which is common in longitudinal panels such as the SIPP. To see how the seam effect arises, recall

\textsuperscript{55}Cullen and Gruber (2000) use the same approach to calculate unemployment durations in the SIPP.

\textsuperscript{56}If one views “discouragement” as the corner solution of zero search effort, it would be appropriate to include these individuals in the estimation of the search model below.

\textsuperscript{57}The probability of survival until week $t$ is $s_t = \Pr(\text{dur} \geq t)$.

\textsuperscript{58}In the absence of censored data, the survival probability would simply be the fraction of spells that are ongoing at time $t$. Kaplan and Meier (1958) develop an extension of this approach to account for censored data.

\textsuperscript{59}The hazard rate at week $t$ is the probability that a spell ends in week $t$ conditional on it lasting for at least $t$ periods. That is, $h_t = \Pr(\text{dur} = t|\text{dur} \geq t)$.
that the SIPP data is collected by interviewing individuals every four months about their activities during the past four months, which is termed the “reference period.” Individuals tend to repeat answers about weekly job status. As a result, they under-report transitions in labor force status within reference periods and overreport transitions on the “seam” between reference periods. Hence, many spells of unemployment appear to last for exactly the length of one or two reference periods, which correspond to lengths of 17 and 35 weeks. The dashed hazard function plotted in Figure 4 shows the empirical hazards for spells that did not begin on a seam; the two spikes are no longer apparent.\textsuperscript{60} Given the magnitude of the seam effect, I explicitly model the recall problem in the structural estimation procedure.

3.3 Unearned Income and Benefit Elasticities

In this section, I estimate reduced-form hazard models to identify income and price effects. The reduced-form approach serves two purposes. First, the transparency of these estimates should comfort readers who are concerned about the detailed assumptions required to estimate the dynamic search model. Second, the reduced-form evidence shows that the inclusion of covariates does not affect the estimates of the key income and price elasticities. This is reassuring because computational limits prevent the inclusion of a large set of covariates in the structural approach.

(a) Reduced-Form Hazard Model Estimates

\textsuperscript{60}The remaining fluctuations in the hazard rate are consistent with the findings of Meyer (1990), who finds similar mean hazards using an administrative dataset of UI recipients.
The key law variations that drive the empirical analysis of this paper can be illustrated with two graphs. I first divide states into two categories: those that have average weekly benefit amounts above the mean and those below the mean. Figure 5 plots Kaplan-Meier survival curves for these two groups. As in Figure 2, the survival functions are plotted conditional on having a spell that lasts for more than one week. It is clear that individuals who receive more benefits have a lower job-finding hazard rate throughout the spell.

I then divide states into two different categories: those that have a waiting week and no retroactive payment and those that have either no waiting week or make a retroactive payment for the waiting week at a later date. Claimants in the latter set of states (INC = 1) get an extra week of benefits up front relative to claimants in the former group (INC = 0). Note, however, that if two individuals have the same maximum potential duration of benefits, the total weeks of benefits that both claimants can get remains the same; it is just that the individual in the INC = 1 state has one week of benefits shifted from the end to the beginning. One should therefore expect this shift of income to have a short-lived effect of the probability of survival. Figure 6 shows that this is indeed the case: Individuals in the INC = 1 states have lower job-finding hazards in the first month after unemployment, but begin to catch up to the other states in subsequent weeks.

I formalize the results shown in these graphs by estimating Cox survival models. An

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61 The states that make a retroactive payment for the waiting week are placed in the INC = 1 category because claimants in all of these states have an extra week of benefits after the retroactive date has passed. Of course, before the retroactive date, these claimants do not have a higher level of unearned income. This problem is resolved in the dynamic structural model estimated below.

62 Abrevaya and Hausman (1999) and Romeo (2001) argue that standard hazard models could yield biased
attractive feature of the Cox model is that it is semiparametric in that it does not impose any functional form on the baseline unemployment exit rates in each week. Consequently, the benefit and income effects are identified solely from the cross-state and time variation in UI laws. The Cox specification assumes a proportional-hazards form for the hazard rate $t$ weeks after unemployment:

$$h_t = \alpha_t e^{\beta X_t}$$

where $X_t$ denotes a set of covariates that can vary over time. Table 3 presents estimates of several specifications of this model. The coefficients reported are hazard ratios ($e^{\beta j}$), which can be interpreted as the ratio of the hazard when covariate $j$ equals $X_j + 1$ to the hazard when the covariate is $X_j$. All specifications only include unemployment spells that last for more than week for the reasons described above.

I begin by discussing the results in the simplest specification and then show the results of a series of robustness checks. In the first specification, as in most others, I use the average UI benefit level in a given state and year instead of each claimant’s individual UI benefit level. There are two reasons for this: first, since computational constraints limit the amount of heterogeneity that can be allowed in estimating the dynamic structural model, it is helpful to get a sense of whether using average benefits produces reasonable reduced-form results; second, the lack of precise data on pre-unemployment wages greatly hampers the ability to

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estimates when unemployment durations are mismeasured due to the seam effect. In the reduced-form empirics, I follow the standard approach because the estimates from the SIPP data appear to be consistent with those found in other studies that did not have mismeasured durations. I address the mismeasurement problem explicitly in the structural estimation procedure.
simulate each claimant’s UI benefit level correctly.

The estimates of specification (1), which are typical, indicate that a 10% increase in the weekly benefit amount reduces the hazard rate by 4.2% in the first week of an unemployment spell, matching the results of previous studies.\textsuperscript{63} This effect is statistically significant at the 1% level. The coefficient on the interaction between weeks unemployed and the weekly benefit amount is statistically insignificant; the point estimate implies that a 10% increase in the weekly benefit amount would reduce the hazard rate by 3.8% twenty weeks into an unemployment spell. The coefficient on the interaction between the benefit exhaustion dummy and benefit rate is also insignificant. These results suggest that higher UI benefits reduce search effort primarily via an income effect rather than a price effect. If the reduction in the net-of-UI wage were the primary source of changes in search behavior, higher benefits would reduce search effort much more at the beginning of a spell than in later weeks. Moreover, there would be a significant difference in the effect of higher benefits on job finding after benefits were exhausted, since UI benefits change search effort only via the income effect after exhaustion.

The next group of estimates reported in the table capture the time-varying effect of the INC variable defined above. The estimates indicate that the provision of an additional week of benefits at the beginning of a spell has large effects on short-run search behavior. An extra week of unearned income through UI benefits reduces the hazard rate by approximately

\textsuperscript{63}See e.g. Meyer (1990), who fits semi-parametric hazard models similar to those estimated here. Krueger and Meyer (2002) review the results of several studies on UI benefits and unemployment durations.
14% in the first two months after unemployment. However, as shown in Figure 6, the hazard rates of the INC = 1 group are 20% higher than those of the INC = 0 group in the third month. The difference between the INC coefficients in months 1 and 3 is statistically significant at the 5% level. The temporary nature of income effect is to be expected, since both groups ultimately are entitled to the same total benefits. It is just that the INC = 1 group temporarily has more cash because they receive an extra week of benefits near the beginning of their spells. After the fourth month, there is no discernible difference in the behavior of claimants in the INC = 0 and INC = 1 states.\textsuperscript{64}

Specification (2) adds a full set of controls, including industry and occupation dummies, year and month dummies, a 10 piece log-linear spline for the claimant’s pre-unemployment wage, and other demographic variables. The coefficients of the key UI law variables discussed above remain virtually unchanged. In specification (3), I check that simulating UI benefits for each individual rather than using the state average does not change the results. Note that in this specification, the UI benefit coefficients continue to be identified from variation in UI laws, and not the cross-sectional variation in benefit levels within a state that results from heterogeneity in wages. This is because we are controlling flexibly for pre-unemployment wages using the spline. To reduce measurement error from noise in the pre-unemployment wage history, individual weekly benefit amounts are generated using the two-stage procedure discussed above.\textsuperscript{65} The hazard ratio on UI benefits rises insignificantly and other coefficients

\textsuperscript{64}These hazard ratios are translated into an income elasticity and benefit elasticity in section (c) below.
\textsuperscript{65}When weekly benefit amounts generated from actual reported wages are used, the point estimate of the
remain unchanged.

In specification (4), I include state fixed effects, identifying the model from changes in state laws. The coefficients on UI benefit levels remain similar and statistically significant. Since only three states changed their waiting periods or retroactive provisions between 1985 and 2000, the standard errors on the INC coefficients rise in this specification. Hence, while the point estimates on the INC variables remain similar to those in specification (1), they are statistically indistinguishable.

Finally, specification (5) confirms that the income and price effects do not emerge for individuals who do not get UI.\textsuperscript{66} The point estimates are close to zero and the 95% confidence intervals do not include the point estimates of the income and benefits elasticities for the UI claimants. These results support the claim that the UI laws are causal.

Table 4 reports the results of additional robustness checks. In specification (6), I omit states that have waiting periods and later make a retroactive payment, identifying the INC coefficient exclusively from variations in whether a state has a waiting period or not. Conversely, in (7), I omit states that do not have a waiting period, thereby identifying the INC coefficient exclusively from variations in whether a state makes a retroactive payment or not.

\textsuperscript{66}Since takeup of UI is endogenous, a purer control group is the set of job losers who are ineligible for UI. This group consists of part-time workers who have very low levels of earned income before unemployment. I show results for individuals who do not takeup UI in order to have a control group that is similar to the UI claimants. However, results are similar for the group of job losers who are ineligible for UI.
The estimates of the INC coefficients are similar in both cases, supporting the claim that it is the extra week of benefits and not an omitted variable that is driving the estimates. In specification (8), I include temporary layoffs in the analysis, to check that the endogeneity of temporary layoff status with respect to spell duration is not biasing the results. In (9), I include the entire set of UI recipients, not just those who claimed UI in the month they became unemployed. While the estimate of the UI benefit effects remain the same, the estimates of the INC variables are now statistically insignificant. This is not surprising, since the one week waiting period has no impact on the benefit path of claimants who wait for more than a week to claim UI benefits anyway. In fact, it should be viewed as further evidence that the earlier specifications are identifying causal effects of the extra week of income on the behavior of claimants who take up UI immediately. Finally, in specification (10), I show that similar results are obtained if one focuses on the group of individuals with weekly wages above $400 who are affected by changes in the UI maximum benefit provision. This specification uses the maximum benefit levels in a state instead of average benefits to proxy for each individual’s UI benefit level.

To what extent is the endogeneity of the UI takeup decision biasing the estimates? There is no evidence of selection based on the waiting period or retroactive payment; regressing UI receipt on INC shows no relationship between the two variables. Moreover, even if there were a selection effect due to variations in the waiting period, we would expect it to mitigate rather than amplify the estimated income elasticity. If there is a fixed cost to UI takeup,
a waiting week should deter individuals who anticipate short unemployment spells, making
the average unemployment spell longer in the INC=0 states.

The endogeneity of takeup poses a more significant problem for the estimation of the
benefit elasticity. The elasticity of takeup with respect to the UI benefit rate is 0.2, which
exactly equals the estimate reported by Anderson and Meyer (1997). If the marginal in-
dividuals who decide to take up UI when benefits rise tend to have shorter unemployment
spells on average, the estimate of the benefit elasticity will be biased toward zero. However,
a rough calculation suggests that the magnitude of this bias is small. At the most, a 10%
increase in benefits may actually cause a 4.72% reduction in the hazard rate instead of the
4.2% that is estimated when the selection problem is neglected.67 This change of 0.52% lies
well within the 95% confidence interval for the estimate, indicating that selection effects are
likely to be dwarfed by the standard error of the estimates.

In summary, the reduced form evidence suggests quite strongly that UI benefits affect
search behavior more through an income effect rather than a price effect. Based on the
logic of section 2, this suggests that individuals’ utility functions are highly curved locally.
It should be noted that this is not the first study to find large labor supply responses

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67 Given the takeup elasticity of 0.2, note that a 10% increase in the UI benefit level will cause an additional 0.2% of the population to take up UI. The mean hazard rate of those who do not take up UI at current benefit rates is 1.3 times the mean hazard of those who do take UI. Note that the marginal individual who takes up UI is more likely to anticipate a long UI spell among the pool who are not taking up UI. Hence, we can derive an upper bound for the bias term by assuming that those who take UI as a result of the benefit increase have a hazard rate of 1.3 times the existing UI claimants. In particular, if the takeup rate is 50% before the change (as estimated e.g. in Andersen and Meyer, 1997) these claimants will artificially raise the estimated hazard ratio by \( \frac{0.002}{0.5} \times 1.3 = 0.52\% \).
to *transitory* income fluctuations. The literature on the “added worker effect” has also found that small reductions in income generate large labor supply responses among non-working spouses. Mincer (1962) found that married women’s labor supply responds twice as much to transitory fluctuations in husbands’ incomes as it does to permanent changes in husbands’ incomes. More recently, Cullen and Gruber (2000) exploit cross-state variation in unemployment insurance laws to identify an income elasticity for wives labor supply between -0.49 and -1.07. These income elasticities are considerably larger than the estimate of -0.1 obtained from the “static” labor supply literature, which is identified from longer term changes in unearned income.68

(c) Estimating Local Curvature in the Static Model

The hazard model estimates can be used to estimate $\gamma^l$ under the strong assumptions that make the individual’s unemployment duration the outcome of a static decision problem. The static analysis complements the more realistic dynamic search model primarily because of its transparency. Unlike the search model, it does not require any functional form assumptions, and it relies on reduced-form estimates where the source of identification is quite clear.

Recall that the key parameters in estimating curvature from labor supply behavior are the elasticity of unemployment duration with respect to unearned income and UI benefits. In the static model, I will exploit the variation in INC to identify the unearned income

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68 In fact, Cullen and Gruber (2000) remark that their high estimate of the income elasticity may be due to “fixed consumption commitments” during spells of unemployment. They do not, however, make inferences about the curvature of utility or welfare costs of unemployment from this elasticity.
effect, and, naturally, the variation in UI benefit levels to identify the benefit elasticity. A fully dynamic formulation of the search problem is required to infer curvature from the time-varying nature of UI benefit effect discussed above.

In addition to $\varepsilon_{1-l,y}$ and $\varepsilon_{1-l,b}$, the other inputs needed to compute local curvature using (5) are $\{b, y, w, l, \overline{\gamma}\}$. We also need to define the length of one “period,” which corresponds to the horizon over which the agent can smooth consumption freely. I will somewhat arbitrarily choose a smoothing horizon of three months, which does not seem unreasonable since the one week unearned income effect lasts for approximately three months. Consider an individual with a weekly wage of $400. Since replacement rates are typically 50% of pre-unemployment wages below the maximum thresholds, suppose his WBA is $200. I assume that the agent has no unearned income other than UI benefits ($y = 0$). Finally, based on the expenditures shown in Table 1, I assume that the agent has committed half of his pre-unemployment income ($\overline{\gamma} = 1$). Under these assumptions, (5) reduces to

$$\gamma' = \frac{2(\varepsilon_{1-l,y})}{1 - l (\varepsilon_{1-l,b} - \varepsilon_{1-l,y})}$$

To calculate the two elasticities, I choose an “initial” unemployment duration for the agent and consider the effect of perturbations in $b$ and $y$ on time unemployed. To begin, suppose an agent is unemployed for five weeks. The survival curve in Figure 6 shows that 82.7% of the sample remains unemployed after five weeks in states that have INC = 0. When the agent can choose his unemployment duration deterministically, we can calculate
how long he would have remained unemployed in an INC = 1 state by calculating the point at which 82.7% of the individuals remain unemployed in the INC = 1 group. The empirical survival curve plotted in Figure 6 implies that this point is 5.72 weeks. Since he has already received UI benefits for four weeks, the one week increase in UI benefits creates a 25% increase in unearned income; hence, $\varepsilon_{1-l,y} = \frac{0.72}{0.25} = 0.58$. A similar exercise using the survival curves in Figure 5 shows that $\varepsilon_{1-l,b} = 0.89$.

These calculations can be repeated for any given initial unemployment duration. The table below shows the elasticities and implied local curvatures for two cases early in an unemployment spell, when the one week up-front income grant is most likely to be viewed as an increase in unearned income by a credit-constrained individual.

<table>
<thead>
<tr>
<th>dur l</th>
<th>$\varepsilon_{1-l,y}$</th>
<th>$\varepsilon_{1-l,b}$</th>
<th>$\gamma^l$</th>
</tr>
</thead>
<tbody>
<tr>
<td>5</td>
<td>8/13</td>
<td>.58</td>
<td>.89</td>
</tr>
<tr>
<td>9</td>
<td>4/13</td>
<td>.73</td>
<td>.62</td>
</tr>
</tbody>
</table>

These two calculations illustrate the problems that arise when the static model of unemployment durations as labor-leisure choices is applied to analyze short duration spells. While the first calculation strongly supports the commitments hypothesis – suggesting that local curvature is an order of magnitude larger than global curvature – the second contradicts utility maximization. In the second case, the unearned income elasticity is larger than $\varepsilon_{1-l,b} = 0.89$.

\[\text{69 Since we only observe job finding at weekly intervals, I interpolate linearly between weeks to determine } S(t) \text{ at any time } t \text{ that is not an integer. Using the predicted survival curves based on the hazard model estimates rather than the empirical ones does not change the conclusions; the estimate of the income elasticity actually rises relative to the estimate of the benefit elasticity.}\]
the benefit elasticity, implying that the compensated wage elasticity is negative. But a rational agent cannot be more sensitive to a given change in unearned income than he is to an equivalent change in the UI benefit level. When the static analysis is applied at each of the weeks between 5 and 15, a negative compensated wage elasticity is obtained in seven of the eleven cases.

Why do we often find a larger income elasticity than benefit elasticity when applying the static model? The income elasticity is identified from a one week income grant at the beginning of a spell; but the benefit effect is estimated from increases in benefits that accrue week by week throughout a spell. If an agent can choose labor supply deterministically, it is irrelevant if extra benefits are obtained in week 3 or in week 10, provided that he has chosen to stay unemployed for 10 or more weeks. But in practice, job finding is stochastic, and there is some chance an agent will find and take a job before week 10. Consequently, higher benefits in week 10 are not worth as much as higher benefits in week 3. In other words, a 10% increase in UI benefits actually corresponds to less than a 10% increase in the benefit level relevant for a static model. This causes us to overestimate the income elasticity relative to the benefit elasticity. The magnitude of this bias is sufficiently large that we find the wrong sign on the compensated wage elasticity, indicating that the static model is a poor description of behavior at least for the first weeks after a job separation. Indeed, Moffitt and Nicholson (1982), who analyze unemployment durations using the labor-leisure model, themselves emphasize that the labor-leisure approach is only appropriate when analyzing
the behavior of the long-term unemployed.

Beyond this fundamental difficulty, there are at least two other reasons to use an explicitly dynamic formulation of the search problem to estimate \( \gamma' \). First, introducing dynamics allows us to exploit the time-varying UI benefit effect to identify income and price effects. Since this variation is much larger than the one week waiting period, estimates of \( \gamma' \) from this source are likely to be more reliable. Second, the dynamic model eliminates the need to make ad hoc assumptions about the agent’s smoothing horizon.

4 Dynamic Search Model

The structural estimation procedure implemented below involves three steps: (1) compute each agent’s optimal unemployment exit hazards using a search model, given the UI law he faces and a set of core preference parameters, \( \omega \); (2) use these hazards to compute the likelihood of observing a given set of unemployment durations for a given value of \( \omega \); and (3) estimate \( \omega \) by maximizing the likelihood function.\(^{(70)}\)

In canonical job search models, agents either choose a reservation wage or search effort in each period.\(^{(71)}\) I adopt the latter strategy primarily because it reduces the computational

\(^{(70)}\) Several studies have estimated structural parameters of job search models. Wolpin (1987) structurally estimates parameters of the wage offer distribution and the reservation wage. More recently, Paserman (2002) estimates hyperbolic preference parameters for unemployed agents. Existing structural models assume a linear utility function and therefore do not identify the curvature of utility. The introduction of a curved utility greatly complicates the estimation procedure because it makes the search problem analytically intractible.

\(^{(71)}\) Lippman and McCall (1979) and Atkinson and Micklewright (1991) reviews the theory and empirics of the job search literature.
intensity of the search problem in a manner that will be made clear shortly. This assumption does have some empirical basis: Wolpin (1987) and van den Berg (1990) find that reservation wages are very low and that job offers are accepted with a probability close to 1. This suggests that the primary disincentive effects of UI occur through a reduction of search intensity for job offers, rather than rejection of a larger fraction of offers after they are obtained.

I first outline the main features of the model, postponing discussion of the specific parametric assumptions that are made until the next section. I consider a discrete-time setting in which each period corresponds to one week, since unemployment duration data is available only at a weekly frequency. All dates that follow refer to time elapsed since the beginning of an unemployment spell and not to calendar time.

(a) Consumption. The agent has an infinite horizon and becomes unemployed exogenously with initial assets $a_0$ at time 0. Following a job separation, in each week $t$, he chooses food consumption, $f_t$, and housing consumption, $h_t$. Let us assume that agents do not abandon their commitments while searching for a new job. This simplifies the consumption decision to choosing $f_t$ in each period.\textsuperscript{72} Since $h_t$ is fixed at $\bar{h}$ throughout the search process, the housing argument of agent’s utility is suppressed and utility is denoted by $u(f_t)$ when the meaning is not ambiguous. Assume that $u$ is twice differentiable and strictly quasiconcave.

\textsuperscript{72}As shown in section 2b, if some agents do move out of their houses during unemployment spells, the estimate of local curvature will be biased downward. I am currently analyzing a richer model in which housing choice is also endogenous.
The agent has a time discount factor $\beta$ and savings earn a net interest rate $r = \frac{1}{\beta} - 1$. I model credit constraints by assuming that the agent can borrow up to a fixed amount $\mathcal{B}$. Letting $b_t$ denote unemployment benefits in week $t$ of the spell, the worker’s intertemporal budget constraint is

$$a_{t+1} = (1 + r)(a_t + b_t - \mathcal{B} - f_t)$$

The $\{b_t\}$ paths are heterogeneous across agents; it is the variation in UI benefit paths that allows us to estimate the agent’s utility function.

(b) Job Search. While unemployed, the agent chooses a level of search effort $e_t$ in each week to find a new job. Once the agent finds a job, it is permanent, and pays a wage of $w$ per week.73 The disutility of working is denoted by $\psi$.

Higher search effort translates directly into a higher probability of finding a job: the probability of finding a job in period $t$ is

$$p_t = \alpha_t e_t$$

The $\alpha_t$ parameters can be interpreted as the “baseline” probabilities of finding a job in week $t$ of an unemployment spell when one exerts 1 unit of search effort. These parameters are allowed to vary over time to capture the fluctuations in unemployment exit hazards evident in Figure 4. In addition, the $\alpha_t$ parameters may be heterogeneous across individuals:

\footnote{The model can be easily extended to allow for a downward trend in the agent’s potential wage as he remains unemployed for a longer period of time.}
for instance, they will depend on local economic conditions, as reflected e.g. by the state unemployment rate when an individual loses his job.

Search is costly: the agent incurs a disutility of $c(e)$ by expending $e$ units of search effort. Assume that $c$ is twice differentiable and convex.

Since our estimates of local curvature are identified by variations in short-run search behavior, a finite search horizon of $T = 55$ weeks is imposed on the model.\(^{74}\) This is a reasonable assumption since job finding hazards appear to be quite low after one year of unemployment in practice. If an individual reaches week $T + 1$ without finding a job, he receives a fixed welfare payment for the remainder of his life. In addition, I assume that the agent abandons his commitments in week $T + 1$, reoptimizing and choosing a new value of housing consumption after this point.\(^{75}\) Since the agent’s problem is stationary after week $T$, a closed-form solution for the value function at $T + 1$ can be obtained.

(c) Agent Optimization. The agent’s problem can be solved using dynamic programming. There are two state variables: asset holdings and the employment state.

When employed, the agent faces a stationary eat-the-pie problem without uncertainty. This problem has a simple solution: the agent consumes a constant fraction of his lifetime wealth in all subsequent periods. If the agent has assets $a_t$ and begins working in week $t$,

\(^{74}\)I choose 55 weeks since the third “seam” for many agents occurs at 52 or 53 weeks, raising the observed hazards at these points. See section (e) below for further details.

\(^{75}\)The decision to abandon commitments exactly at time $T$ is optimal if the adjustment cost for changing houses, $k$, is (1) sufficiently large that moving when one has a positive probability of finding a job is suboptimal and (2) sufficiently small relative to the welfare payment, $w$, that moving is rational after period $T$.
he consumes

\[ f_t^e = r a_t + w - \bar{h} \]

per week for the rest of his life. This implies that the agent’s value function when employed is the present value of his lifetime utility:

\[ V^e_t(a_t) = \frac{1}{1 - \beta} \{ u(ra_t + w - \bar{h}) - \psi \} \]

While the individual is unemployed, his value function in periods \( t \in \{1, \ldots, T\} \) is

\[ V^u_t(a_t) = \sup_{f_t, e_t} u(f_t) - c(e_t) + \beta \{ \alpha_t e_t V^e_{t+1}(a_{t+1}) + (1 - \alpha_t e_t) V^u_{t+1}(a_{t+1}) \} \tag{6} \]

If the agent remains unemployed after the search process ends, he faces another stationary eat-the-pie consumption problem in period \( T + 1 \). He consumes a constant fraction of his wealth for the remainder of his life. The agent sells his house in week \( T + 1 \), incurring a fixed transaction cost. Let \( w \) denote the agent’s weekly income net of (annuitized) transaction costs after period \( T \). Optimal housing consumption in each period after this point, \( h_w \), can be computed numerically. The agent’s value function in week \( T + 1 \) is

\[ V^u_{T+1}(a_{T+1}) = \frac{1}{1 - \beta} u(h_w(a_{T+1}), w - h_w(a_{T+1})) \tag{7} \]

When utility is non-linear, the Bellman equation in (6) does not have an analytical
solution. Given functional form assumptions on $u$ and $c$, the sequence $\{V_t^u(a_t)\}_{t=1}^T$ can be obtained by starting with the terminal equation (7) and using numerical backwards induction to solve the system in (6).

Solving the problem requires optimization over both effort and consumption in each week $t$ for every level of assets $a_t$ with which one could enter week $t$. This two-dimensional optimization problem can be reduced to a one-dimensional search for the optimal consumption path when unemployed, $\{f_t^u\}$, by deriving an analytic expression for $e_t(a_t)$ in terms of $f_t^u$ and the value functions at time $t+1$, $V_{t+1}(a_{t+1})$ and $V_{t+1}^u(a_{t+1})$. Since $c(e)$ is convex, $V_t^u(a_t)$ is concave in $e_t$ and the optimal effort level is given by the first order condition

$$c'(e_t) = \beta[V_{t+1}(a_{t+1}) - V_{t+1}^u(a_{t+1})] \alpha_t$$

By choosing a convenient functional form for $c(e)$, we can directly calculate the optimal effort level at time $t$ for a given level of $a_t$ and $f_t$, after we have computed the value functions at time $t+1$. This analytic simplification tremendously increases the speed of the algorithm, and is the benefit of formulating the model in terms of search effort rather than reservation wages.

The core parameters and functions required to solve the problem are

$$\omega = \{u, c, \{\alpha_t\}_{t=1}^{T-1}, \psi, \beta\}$$
Given values of \( \{w, \wbar, B, h, \{b_t\}_{t=1}^{T-1}\} \) and the initial asset level \( a_0 \), solving the dynamic programming problem yields an optimal policy \( \{f_t^u, f_t^e, e_t\}_{t=1}^{T-1} \) for an agent in terms of the core parameters \( \omega \). Indexing the individuals in the sample by \( i = 1, ..., I \), we can solve the model \( I \) times to obtain an optimal search effort path \( \{e_{it}\}_{t=1}^{T-1} \) for each agent. The \( i^{th} \) agent’s unemployment exit probabilities are

\[
p_t^i = \alpha_t^i c_t^i
\]

(d) Likelihood Function.\(^{76}\) I now compute the likelihood of observing a given dataset of unemployment durations for a core parameter vector \( \omega \) using the unemployment exit probabilities \( \{p^i(\omega)\}_{i=1}^I \). First, consider the subset of \( I_1 \) uncensored observations, each of which is observed from the date of unemployment to the date of job finding. Suppose individual \( i \) has an unemployment duration of \( d_i \) weeks. The probability of this event is the product of the probability that the individual does not find a job in weeks 1 to \( d_i \) and the probability that he finds a job in week \( d_i + 1 \):

\[
\Pr(d_i; \omega) = \prod_{t=1}^{d_i} (1 - p_t^i(\omega)) p_{d_i+1}^i(\omega)
\]

The likelihood function for the entire subset of \( I_1 \) individuals who have uncensored spells is

\[
L_1(\omega) = \prod_{i=1}^{I_1} \prod_{t=1}^{d_i} (1 - p_t^i(\omega)) p_{d_i+1}^i(\omega)
\]

\(^{76}\)Kiefer (1988) provides a review of maximum likelihood estimation of survival models.
For the censored spells, we know that the individual was unemployed for at least \( d_i \) weeks but do not know when he found a job. The probability of observing a censored spell of length \( d_i \) is therefore

\[
\Pr(d_i; \omega) = \prod_{t=1}^{d_i} (1 - p_{it}^i(\omega))
\]

and the likelihood function for the entire sample, in which the first \( I_1 \) individuals have uncensored spells and the remaining \( I_2 \) individuals have censored spells is

\[
L(\omega) = \prod_{i=1}^{I_1+I_2} \prod_{t=1}^{d_i} (1 - p_{it}^i(\omega)) \prod_{i=1}^{I_1} p_{d_i+1}^i(\omega)
\] (8)

When true unemployment durations are observed without measurement error, we can estimate the structural parameters of the search model by maximizing the likelihood in (8). However, the reduced-form evidence indicates that the “seam effect” described in section 3b creates a fairly substantial amount of measurement error in unemployment durations. The probability that an individual reports a transition on a seam is more than six times as large as the probability that he reports a transition within a reference period. One manifestation of this problem is the bunching of unemployment durations at 17 and 18 weeks. I now extend the model to allow for measurement error in unemployment durations.

(e) Mismeasurement in Unemployment Durations. When the true unemployment durations are latent, we need to maximize a likelihood function for reported durations rather than the likelihood for actual durations in (8). To develop a likelihood for reported durations, I
construct a mixture model to capture the process that generates the reporting errors. The parameters that govern this process are then estimated along with the core parameters in the search model.

Suppose that there are two types of individuals in the population: exact reporters, and seam reporters. Exact reporters always report job findings exactly on the week that they actually occur; for this group, there is no mismeasurement problem. Seam reporters never report job-findings within a four-month reference period; they effectively report all transitions on the seams between interviews. What remains is to specify which transitions are reported on which seam by seam reporters. Lacking prior information on the nature of the recall process, I make the assumption that seam reporters report job findings that occur in the first half of a reference period on the seam before that period and job findings in the second half of a reference period on the seam after that period. For example, suppose a seam-reporting agent has his interview seams at 17 weeks, 34 weeks, and 52 weeks after unemployment (recall that each reference period lasts for 17-18 weeks). If he finds a job between weeks 8 and 25 after unemployment, he reports the transition at week 17; if he finds a job between 26 and 43 weeks after unemployment, he reports the transition on the second seam; and if he finds a job between week 44 and week 55, the end of the search period in the model, he reports the transition on the third seam. Finally, if a spell ends in the first half of the

\footnote{This pattern would emerge if seam reporters report a constant employment status during the reference period based on whether they spent the majority of that reference period employed or unemployed.}

\footnote{Note that this is not the only pattern for the seam weeks because agents can report unemployment at any time within a reference period.}
first reference period (in this example, before the ninth week), I assume that all individuals correctly report the duration.

Given this model of recall errors, we can formulate a likelihood function for reported durations as follows. Let $\mu$ denote the fraction of seam reporters. The probability of observing a given duration $d$ is

$$Pr(d) = (1 - \mu) Pr(d|\text{exact reporter}) + \mu Pr(d|\text{seam reporter})$$

For exact reporters, the probability of observing $d$ coincides with the probability that the true duration of the spell is $d$. To see how the likelihood for the seam reporters is formulated, consider an agent who reports a duration $d$ that coincides with his first seam. Under our model of recall errors, we only know that the true duration lies in the second half of his first reference period (i.e. after $d - 8$) or the first half of his second reference period (before $d + 9$). The probability of observing $d$ is therefore equal to the probability that the true duration lies between $d - 8$ and $d + 9$. Hence, we can write a likelihood function for reported durations as follows:

$$\tilde{L}(\omega) = \prod_{i=1}^{I_1+I_2} \left[ (1 - \mu + \mu(d_i < \text{seam1}_i - 9)_{\text{onseam}_i}) \prod_{t=1}^{d_i} (1 - p_t(\omega)) \prod_{t=1}^{I_1} p_{d_i+1}^j(\omega) \right]$$

$$+ \mu_{\text{onseam}_i} \left( \prod_{t=1}^{\min(d_i+8,T)} (1 - p_t(\omega)) - \prod_{t=1}^{\max(d_i-9,1)} (1 - p_t(\omega)) \right)$$

(9)
where onseam\(_i\) is a dummy variable that indicates whether the agent’s duration \(d_i\) lies on one of his seams, seam1\(_i\) denotes the week on which agent \(i\)’s first seam occurs, and \((d_i < \text{seam1}_i - 9)\) is an indicator variable.

The measurement-error corrected estimator of \(\{\omega, \mu\}\) maximizes the likelihood function in (9). Asymptotic confidence intervals for the key parameters are obtained by inverting a maximum likelihood ratio test, i.e. constructing an interval such that the null hypothesis that the parameter lies outside this interval can be rejected with a given p value.\(^{79}\)

This completes the description of the search model and likelihood functions. The next section fills in the details of the estimation procedure and reports the results.

### 5 Structural Estimates of Local Risk Aversion

(a) Estimation Details

I choose the following functional form for \(u(h, f)\):

\[
u(h, f) = \frac{f^{1-\gamma_f}}{1 - \gamma_f} + \eta \frac{h^{1-\gamma_h}}{1 - \gamma_h}\]

The key parameter for our analysis is \(\gamma_f\). The nuisance parameters \(\gamma_h\) (curvature of utility over housing) and \(\eta\) (relative weight placed on housing in utility) cannot be identified from data on unemployment durations. However, these two parameters can be expressed in terms

\(^{79}\)The more common method of inverting the asymptotic information matrix could not be applied here because of problems in computing second derivatives numerically.
of $\gamma_f, \bar{h}$, and the global coefficient of relative risk aversion $\gamma^g = 1$ estimated in Chetty (2002). In other words, $\{\gamma_f, \eta, \gamma_h\}$ is homeomorphic to $\{\gamma_f, \bar{h}, \gamma^g\}$. Since both $\gamma^l(W)$ and $\gamma^g(W)$ vary with $W$ under this specification of $u(h, f)$, I will assume that $\gamma^g(W_0) = 1$ at the agent’s pre-unemployment wealth $W_0$ and report estimates of $\gamma^l(W_0)$ at the same point.

I also assume a power specification for the cost of effort:

$$c(e) = [u(\bar{h}, w - \bar{h}) - u(\bar{h}, w - 150)]e^{\theta}$$

This parametrization assumes that the elasticity of cost of effort with respect to effort is constant: $\varepsilon_{c,e} = \theta$. The multiplicative scalar $[u(\bar{h}, w) - u(\bar{h}, w - 150)]$ is included to aid in the interpretation of the estimates of the baseline job-finding probabilities $\{\alpha_t\}_{t=1}^{T-1}$. With this choice of units, the cost of one unit of effort is normalized as the welfare cost of losing $150$ of food per week.

Similarly, to aid in the interpretation of the disutility of labor parameter $\psi$, I define

$$\psi = u(\bar{h}, w - \bar{h}) - u(\bar{h}, w - \bar{h} - \psi_\$)$$

and report the estimate of $\psi_\$$, which can be interpreted as the disutility of labor expressed in dollars of food consumption lost from the initial level.

Since a set of dynamic programming problems must be solved at each step of the estimation algorithm, computational constraints limit the size of the parameter set that can
be estimated. However, it is important to retain a flexible specification of the baseline job-finding hazards to ensure that the $\gamma$ and $\theta$ parameters are being identified from the cross-sectional variation in UI laws rather than the time path of unemployment durations. I replace the nonparametric specification of the baseline job-finding probabilities $\{\alpha_t\}_{t=1}^{T-1}$ with a piecewise linear functional form. This parametrization can capture the variation in baseline hazards both flexibly and parsimoniously. The values of $\alpha_t$ at a set of break points $S \subset \{1, \ldots, T\}$ are included as parameters in the estimation set; the values of $\alpha_t$ in the other weeks are obtained by linearly interpolating between the break points. For instance, one can allow for separate pieces in each month of an unemployment spell by choosing $S = \{1, 5, 9, 13, \ldots, 55\}$. I check that the specification of $\{\alpha_t\}$ is sufficiently flexible by showing that the inclusion of additional pieces in the linear spline (expansion of $S$) does not affect the estimates of the key parameters.

To control for the effect of variations in local economic conditions on job-finding probabilities, I also allow the scale of the $\alpha_t$ parameters to depend on the state unemployment rate when individual $i$ loses his job:

$$\alpha^i_t = \alpha_t (1 + \alpha_u \text{unempdev}^i)$$

where unempdev$^i$ denotes the deviation in the state unemployment rate when agent $i$ becomes unemployed from the overall mean in the sample.

With this notation, the set of parameters to be estimated in the structural model is
reduced to:

\[ \tilde{\omega} = \{\gamma_f, \theta, \{a_t\}_{t \in S}, a_u\} \]

The remaining parameters \( \{\overline{h}, \overline{w}, \overline{B}, \beta, \psi_e\} \) are specified exogenously. The estimation algorithm can be run for different values of these parameters to check sensitivity.\(^80\)

Ideally, we would allow for heterogeneity in wages and benefit level \((w\) and \(\{b_t\}_{t=1}^{T-1}\)) across individuals by solving a separate dynamic programming problem for each agent who has a unique wage and benefit path. Since CPU time increases linearly in the number of dynamic programming problems solved, I instead consider a representative agent model in which each agent earns the mean wage in the sample gets the mean weekly benefit amount for his (state, year) pair.\(^81\)

(b) Parameter Estimates

Table 5 reports preliminary estimates of the search model. I begin by discussing a benchmark specification in detail, and then show the results of a few robustness checks.

The setup of specification (1), which includes the measurement-error correction, is as follows. Monthly break points are included in the linear spline for the \(\alpha\) parameters for the first five

\(^80\)While one would ideally estimate \(\beta\) and \(\psi_e\), it is not clear how these parameters can be identified given the available data. When \(\psi_e\) was included in the estimation set, the maximization algorithm did not converge. Fortunately, the maximum likelihood estimates of \(\gamma_f\) are not sensitive to the value of \(\psi_e\).

\(^81\)The dynamic programming algorithm was coded in MATLAB. The likelihood function was maximized using the FMINSEARCH algorithm; a concentrated likelihood over \(\gamma_f\) was formed to estimate a confidence interval for \(\gamma_f\). One run of the maximization algorithm took approximately fifty hours. I am grateful to the National Bureau of Economic Research for allocating computer resources to this project.
months and breaks every two months are included for the remainder of the search horizon:

\[ S = \{1, 5, 9, 13, 17, 25, 32, 40, 48, 55\} \]

I include more breaks early in the spell since weekly fluctuations in job search productivity are likely to diminish as a spell wears on. The weekly wage is fixed at \( w = 400 \). I assume that agents have committed \( h = 200 \) of their budgets to housing prior to unemployment and that they have pre-unemployment assets of \( a = 500 \). I fix \( \beta = 0.9^{1/50} \) to reflect a yearly discount rate of 10%. I set \( \psi = 100 \), so that the disutility of working equals the utility cost of consuming $100 of food per week instead of $200. The welfare payment after week 55 is set at \( \tilde{w} = 100 \).

Finally, since the agent must be able to afford his housing payment of $200 for 55 weeks, the level of \( \overline{B} \) reflects both the ability to make late payments on the commitments and actual borrowing channels such as credit cards. I set \( \overline{B} = 11,000 \), with the rationale that this covers the agent’s housing payments for a year and leaves approximately $5000 for food consumption over one year at the mean UI benefit level in the sample. This food budget is small in view of Gruber’s (1997) estimate that the average drop in food consumption is less than 10% at the mean UI benefit level. The drop in food consumption predicted by the model is much larger that 10% when \( \overline{B} = 11,000 \). Note that a higher level of \( \overline{B} \) raises the estimate of \( \gamma^l \) because it reduces the size of the consumption response associated with a change in UI benefits. The labor supply response to a change in unearned income depends
on the product of $\gamma^l$ with the resulting change in consumption. Therefore, generating a given labor supply response with a smaller consumption drop requires a larger $\gamma^l$.

Given these values for the exogenous parameters, the maximum likelihood estimate of the key local curvature parameter, $\gamma^l$, is 7.37 and the standard error is 2.07. Since the maximum likelihood estimate has a Normal asymptotic distribution, we can strongly reject the null hypothesis that commitments do not affect the curvature of utility ($\gamma^l = \gamma^g = 1$). Existing models of moderate-stake risk preferences such as loss aversion do not predict a difference between $\gamma^l$ and $\gamma^g$; hence, the estimate of $\gamma^l$ suggests that commitments are an important determinant of preferences over wealth.

The estimates of the other parameters in the model appear reasonable. The estimate of $\theta = \varepsilon_{c,e} = 1.54$ indicates that a 10% increase in effort results in a 15% increase in the disutility of search. Raising the probability of unemployment exit via extra search effort appears to be quite costly.

The estimates of the baseline job finding probabilities show that exerting search effort which costs the equivalent of losing $150 of food results in a 3-4% chance of finding a job in each week during the first six months after unemployment. Since losing $150 of food has significant welfare costs, these estimates suggest that the job search process is quite unpleasant. Figure 7 gives a more complete picture of the $\{\alpha_t\}$ estimates. The productivity of job search varies non-monotonically over time, but tends to fall sharply as the spell progresses, as one would expect. The negative duration dependence of search
productivity could arise because of deterioration of skills, exhaustion of job opportunities, or selection effects due to heterogeneity in skill level. Figure 7 also plots the average optimal search effort path. Search effort rises in initial weeks as agents deplete their assets and see the value of employment rising; however, in later weeks, as the productivity of search falls, the optimal level of search falls as well. Note that the kinks in the path of search effort come from the break points in the specification of the baseline job-finding hazards.

The estimate of $\alpha_u$ indicates that a 1 percentage point increase in the state unemployment rate lowers the job finding hazard by approximately 5%. It is reassuring that the structural estimate of the state unemployment rate coefficient almost exactly matches the reduced-form semiparametric hazard model estimate. Finally, the estimate of $\mu = 0.33$ indicates that approximately one third of the individuals in the sample are seam reporters who do not accurately report job transitions within a reference period.

I now turn to the results of some sensitivity checks. In specification (2), I estimate the model without including the measurement error correction by maximizing the likelihood for true durations in (8). The point estimate of $\gamma^t$ remains similar, but the standard error of the estimate rises. The bunching of unemployment durations at 17 weeks is picked up by the spike in the baseline hazard in week 17 ($\alpha_{17}$). Figure 8 uses the estimates of specification (2) to plot predicted survival curves for individuals in states that pay UI benefits below the mean and those that pay UI benefits above the mean. The empirical survival curves for
these two categories are also shown. Visual assessment of the fit indicates that the model captures both the time pattern of unemployment exit rates and the cross-sectional variation due to UI laws quite well. For comparison, Figure 9 replicates Figure 8 when $\gamma^l$ is held fixed at 2.5 while the likelihood function is maximized over all the other parameters. The predicted survival curves using these restricted maximum likelihood estimates continue to fit the overall time pattern of unemployment exit rates well; however, they fail to capture the extent to which higher UI benefits diminish search intensity (the gap between the two curves). These figures indicate that $\gamma^l$ is being identified from precisely the variation in UI laws that one would expect based on the intuition given in section 2. The fact that higher UI benefits persistently reduce unemployment exit rates throughout an unemployment spell, even after benefits have been exhausted, can only be explained by a high value of $\gamma^l$ in the dynamic search model.

Specification (3) replicates specification (1) but reduces the number of break points in the specification of the baseline hazards. The estimate of $\gamma^l$ remains virtually unchanged. This suggests that the specification of the baseline job-finding probabilities is sufficiently flexible that the key curvature parameter is being identified only from cross-sectional variations and not the overall time-pattern of unemployment exit rates. Finally, in specification (4), I check the sensitivity of $\gamma^l$ to the value of the borrowing constraint by setting $\beta = 10,000$. This

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82 I plot predicted survival curves for specification (2) rather than (1) because generating predictions from the measurement error model requires that we re-create the measurement error in the empirical survival curves. This process is complicated, since we need to take the distribution of seams in the population into account.
lowers the estimate of $\gamma^l$ to 6.15, which is well within the 95% confidence interval for the estimate in the benchmark specification.

Figure 10 summarizes the empirical content of the commitments model by plotting utility over wealth when $\gamma^l(W_0) = 7.37$ at the agent’s initial wealth level $W_0$. In this example, the agent has an annual income of $40,000, a prior commitment of 50% of income, and faces an adjustment cost of 7.5%. Assuming a 5% interest rate and an infinite horizon, this agent’s initial lifetime wealth is $W_0 = 800,000$. The committed region within the $(s, S)$ band includes roughly ±10% of lifetime wealth. Utility is considerably more curved locally than it is globally, as shown by the dashed curve, which extends the utility function within the committed region outside the $(s, S)$ band.

6 Implications of Commitments

This paper has characterized preferences over wealth in a world with two consumption goods, one of which involves a commitment in that an adjustment cost must be paid when the good is sold. In this model, utility over wealth is more curved locally than globally: individuals are relatively more risk averse with respect to moderate-scale income fluctuations than they are to large income fluctuations. Commitments also create a gambling motive.

The model was tested and calibrated using the labor-supply method of estimating risk aversion developed in Chetty (2002). Existing estimates of wage and income elasticities from the static labor supply literature imply an upper bound for global curvature of 1.25. In the
present study, I exploited differences in unemployment insurance laws across states and time to estimate the local curvature of utility.\textsuperscript{83} Both reduced-form estimates of semiparametric unemployment hazard models and structural estimates of a dynamic model of job search indicate that the local coefficient of relative risk aversion is an order of magnitude larger than the global coefficient of relative risk aversion. The empirical estimate of the difference between local and global curvature ultimately stems from two basic facts about labor supply: (1) Permanent increases in income do not reduce labor supply significantly, implying that the taste for consumption deteriorates slowly with wealth; but (2) temporary fluctuations in income do cause large labor supply responses, indicating that the marginal utility of wealth falls rapidly when substitution across consumption goods is hindered by prior commitments.

I conclude by discussing a few empirical puzzles that commitments could help explain:

1. [Risk Aversion and Insurance] Retaining a consequentialist, expected utility framework, the commitments model escapes Rabin’s (1999) critique by generating a utility that is not everywhere concave. The utility in Figure 7 rejects the 50-50 $1000/$1050 bet but accepts the $40000/$6.35 million one. In addition, it rationalizes the desire to insure moderate-scale income fluctuations (e.g. automobile and non-catastrophic health insurance).

2. [Gambling] The model explains why racetrack bettors prefer bets with long odds and

\textsuperscript{83}An interesting direction for future research is to formulate other “experiments” that can be used to evaluate the importance of commitments. For example, behavior during episodes of sharp inflation could provide another useful domain for inferring preferences over wealth.
large payoffs (Golec and Tamarkin, 1998). It is also consistent with the finding that higher incomes are less likely to play the lottery overall, but are relatively more likely to participate in lotteries that have very large payoffs (Clotfelter and Cook, 1987).

3. [Intertemporal Substitution] Commitments can reconcile limited intertemporal substitution, non-ridiculous degrees of risk aversion over large stakes, and time-separability of utility. When utility is time-separable, the elasticity of intertemporal substitution is the reciprocal of the curvature of utility over wealth. A large fraction of intertemporal substitution among nondurables and services is likely to occur while individuals are maintaining commitments. Since the relevant curvature parameter for intertemporal substitution lies between \( \gamma^g = 1 \) and \( \gamma^l \approx 7 \) in this case, an EIS significantly below 1 (as estimated e.g. by Hall (1988), Epstein-Zin (1991), and Vissing-Jorgensen (2002)) would be consistent with the risk preferences of committed agents.

4. [Labor Supply] Commitments can explain the “added worker effect” – the entry of spouses into the labor force when primary earners become unemployed – documented e.g. by Mincer (1962) and Cullen and Gruber (2000). When families have commitments, married women’s labor supply should be more responsive to transitory income fluctuations than permanent ones.

5. [Crime and Punishment] Becker (1968) showed that the optimal punishment strategy for risk averse agents combines large fines with low probabilities of punishment.
However, empirical evidence indicates that extensive resources are in fact devoted to enforcement (see e.g. Dickens et. al. (1989), who label this a “monitoring puzzle”). Criminals with commitments may be more sensitive to the probability of apprehension than to the severity of the punishment.

Finally, the answer to the applied question that motivated this paper – What is the optimal level of social insurance? – changes significantly when commitments are taken into account. To see this, let us compute the optimal level of unemployment insurance using the static model of unemployment developed in section 2.84 In this model, individuals make consumption and savings decisions in the first period, recognizing that they may be laid off with exogenous probability \( p \) at the end of the first period. In the second period, individuals choose their unemployment duration deterministically, subject to the level of UI benefits being provided.

Note that even though this model has two periods, it remains a static model of unemployment since agents do not make dynamic search decisions. The purpose of the first period is only to allow for distortions in saving as a result of changes in UI benefits. An analysis of the optimal path of UI benefits over time in a fully dynamic search model is outside the scope of this paper.

Let us adopt the following notation: \( s_1 \) denotes saving in period 1, \( l \) is labor supply in the unemployed state in period 2, \( w \) is the wage, \( \tau \) the UI tax rate, and \( b \) the UI benefit.

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84 This model is the same as Baily’s (1978) model of optimal UI, augmented with consumption commitments.
level. Assume that the adjustment cost $k$ for housing is large enough that it is not optimal to move when unemployed. If the individual is unemployed in the second period, he chooses labor supply by solving

$$\max_{l \in [0,1]} u(s_1 + (1 - l)b + lw(1 - \tau) - h, h) - \psi(l)$$

In the first period, the individual chooses housing $h$ and savings $s_1$ by solving

$$\max_{s_1, h} V_1 = u(w(1 - \tau) - s_1 - h, h) + (1 - p)u(s_1 + w(1 - \tau) - h, h) + pu(s_1 + (1 - l)b + lw(1 - \tau) - h, h)$$

The model is closed by a balanced-budget constraint:

$$\tau w(1 + p + (1 - p)l) = (1 - p)(1 - l)b$$

To find the optimal benefit rate $b^*$, we maximize the agent’s expected utility, taking into account his behavioral responses:

$$b^* = \arg \max_{b} V_1(s_1(b), h(b), l(b))$$

The optimal rate of benefits is chosen by trading off the benefits of smoothing consumption across states with the distortionary costs of UI. Letting $\frac{\Delta F}{f}(b)$ denote the percentage change in food (uncommitted) consumption between the employed and unemployed states,
the optimal $b^*$ is implicitly defined by

$$\frac{\Delta f(b^*)}{f(b^*)} \gamma^f \frac{f}{f + \bar{h}} = \epsilon_{1-l,b}$$ (10)

A sketch of the proof is given in the appendix. Gruber (1997) uses data from the PSID to estimate the drop in food consumption during unemployment as a function of the replacement rate $\frac{b}{w}$. The estimates from his baseline specification imply

$$\frac{\Delta f(b)}{f(b)} \approx .24 - .28\left(\frac{b}{w}\right)$$

This equation implies that consumption would drop by 24% in the absence of UI, and reflects the fact that consumption drops by approximately 10% at the mean UI replacement rate of 50%. Substituting this expression into (10) gives a closed-form solution for the optimal wage replacement rate $\frac{b^*}{w}$:

$$\frac{b^*}{w} = .86 - 3.57 \epsilon_{1-l,b} \frac{f + \bar{h}}{\gamma^f}$$ (11)

85 We do not need to know the elasticity of saving and housing choices with respect to UI benefits because these are reflected in the value of $\Delta f$. Formally, the agent’s utility maximizing behavior creates an envelope condition that allows us to ignore these second-order effects in calculating $b^*$.

86 As in other calculations of optimal UI, I take the consumption drop as exogenous and calculate the optimal level of UI conditional on this drop. However, one can imagine endogenizing the consumption drop itself. If agents are completely rational and forward looking, and the welfare costs of a reduction in consumption are high, one should not observe a lack of saving in the absence of UI. In other words, insofar as unemployment spells are relatively short relative to lifetimes, self-insurance should be a very good substitute for social insurance. Therefore, in a more general model where consumption is also chosen endogenously, social insurance would be optimal only if agents are myopic.
I set $\bar{h} = 0.5w$ and $f = 0.5w$. To make my results comparable to those of Gruber's, I set $\varepsilon_{1-l,b} = 0.432$. Using the estimate of $\gamma^l = 7.37$, (11) implies

$$\frac{b^*}{w} = 0.44$$

For analogous parameters in a one-good model ($\bar{h} = 0, \gamma^l = \gamma^g = 1$), Gruber finds that the optimal replacement rate is 0. When commitments are taken into account, a substantial amount of unemployment insurance becomes desirable despite the large elasticity of unemployment duration with respect to benefit rates.

Several caveats to this optimal unemployment insurance calculation deserve mention. First, I have excluded temporary layoffs from the analysis, and am ignoring distortions that UI may cause in the level of unemployment. If a more generous UI system raises the rate of unemployment (e.g. by subsidizing temporary layoffs), then the optimal level of benefits

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87 Since $f$ represents all non-committed expenditure in our model, one may be concerned that the drop in actual food consumption that Gruber estimates overstates the drop in non-committed consumption, biasing our calculation of $\frac{b^*}{w}$ upward. There are two lines of defense against this criticism: (1) Gruber (1998) uses data from the CEX to show that other types of uncommitted consumption also fall similarly to food consumption; (2) Even if certain types of consumption do not fall (e.g. clothing), it is not clear that this should lower the optimal replacement rate; rather, this may simply be evidence that clothing should also be thought of as “committed” during a short unemployment spell, raising $\bar{h}$, $\gamma^l$, and the optimal replacement rate.

88 Gruber derives this elasticity by multiplying Meyer’s elasticity for UI recipients of 0.9 by the UI takeup elasticity he estimates of 0.48.

89 From a positive standpoint, it is interesting to note that commitments appear to be the motivation for UI. In the Arizona State Benefit Adequacy Study, Burgess et. al. (1981) define a measure of “benefit adequacy.” They emphasize that “this measure of benefit adequacy is defined for each beneficiary within the context of his/her household circumstances. These circumstances are reflected by the total of necessary/obligated expenses for the entire household....”
will be lower.\textsuperscript{90} Second, this analysis ignores general-equilibrium effects caused by increased unemployment via higher UI benefits.\textsuperscript{91} If these effects are significant, the optimal level of benefits will again be lower. Third, the result is a statement about the optimal level of benefits and not the optimal method of provision of those benefits. For instance, individual UI savings accounts, as proposed by Feldstein and Altman (1998), may be a more efficient method of provision than government transfers.

Finally, as noted above, the analysis of optimal UI in a static model of unemployment is necessarily quite limited. Future research can use the calibrated search model of this paper to analyze optimal social insurance in a dynamic setting. For example, one can study the optimal path of UI benefits, recognizing that individuals will abandon more of their commitments as they remain unemployed for a longer period of time.\textsuperscript{92}

\textsuperscript{90} Feldstein (1978) and Topel (1983) present empirical evidence showing a strong relationship between the rate of temporary layoffs and UI benefits.  
\textsuperscript{91} Acemoglu and Shimer (1999) consider optimal unemployment insurance in a general equilibrium model.  
\textsuperscript{92} Werning (2002) shows that the optimal path of UI benefits in a one-good model is likely to be upward sloping, contrary to the existing structure of UI laws. This is because the value of insurance rises as agents deplete their savings. Commitments create a countervailing force in favor of a downward-sloping benefit path – as a spell continues, agents abandon more commitments, and the value of insurance falls.
References


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Appendix

(a) Global Curvature Upper Bound: Derivation of (3)

Given that
\[
\frac{\partial \theta}{\partial y} \frac{\partial \theta}{\partial w} = \frac{v_W(T(y + w) - kh_0) - v_W(Ty)}{v_W(T(y + w) - kh_0)}
\]

The agent is assumed to have CRRA utility over wealth when he is unconstrained (i.e. has no commitments):
\[v(W) = \frac{W^{1-\gamma g}}{1-\gamma g}\]

Under this assumption, (2) implies
\[-\frac{\partial \theta}{\partial y} \frac{\partial \theta}{\partial w} = \frac{(Ty)^{-\gamma} - (T(y + w) - kh_0)^{-\gamma}}{(T(y + w) - kh_0)^{-\gamma}}\]

Using the definitions of the elasticities in the text, taking logs and rearranging implies
\[\bar{\gamma}^g \equiv \frac{\log[1 - \frac{\epsilon_{\theta, w} w}{\epsilon_{\theta, w} y}]}{\log[1 + \frac{w - kh_0}{Ty}]}\]

(b) Local Curvature in a Static Model: Derivation of (4)

Recall that the agent solves
\[
\max_l u(\overline{h}, y + b(1 - l) + wl - \overline{h}) - \psi(l)
\]

The first order condition for this problem is
\[(w - b)u_f = \psi_l\]  \hspace{1cm} (12)

Implicit differentiation of (12) with respect to \(b\) and \(y\) yields
\[
\frac{\partial l}{\partial y} = -\frac{(w - b)u_{ff}}{(w - b)^2u_{cc} - \psi_{ll}}
\]
\[
\frac{\partial l}{\partial b} = \frac{u_f - (1 - l)(w - b)u_{ff}}{(w - b)^2u_{cc} - \psi_{ll}}
\]

This implies
\[
\frac{\partial l}{\partial y} = -\frac{(w - b)u_{ff}}{u_f} \frac{\partial l}{\partial b} - (1 - l)\frac{\partial l}{\partial y}
\]

Define the elasticities of unemployment duration with respect to unearned income and ben-
fits as follows:

\[
\varepsilon_{1-l,y} = -\frac{\partial l}{\partial y} \frac{y}{1-l} + b(1-l) \varepsilon_{1-l,y} = -\frac{\partial l}{\partial b} \frac{b}{1-l}
\]

With this notation, it follows that

\[
-\frac{u_{ff}}{u_f} f = \frac{f}{w - b(1 - l + y/b)\varepsilon_{1-l,b} - (1 - l)\varepsilon_{1-l,y}}
\]

(13)

To see how curvature of utility over wealth is related to curvature of utility over food in the world without credit constraints modeled in section 2, observe that agents will smooth their consumption perfectly, setting

\[
f + \bar{h} = \frac{W}{T}
\]

in each period. The indirect utility of having wealth \(W\) is

\[
v(W) = \sum_{t=1}^{T} u(\bar{h}, \frac{W}{T} - \bar{h})
\]

and it follows that

\[
g'(W) = -\frac{v_{WW} W}{v_W} = -\frac{u_{ff} W}{u_f T} = -(f + \bar{h}) \frac{u_{ff}}{u_f} = -\frac{u_{ff}}{u_f} f(1 + \frac{\bar{h}}{f})
\]

(c) Optimal Social Insurance in a Static Model: Derivation of (10)

The optimal benefit rate is computed by solving

\[
\max s_{1,h} V_1 = u(w(1-\tau) - s_1 - h, h) + (1-p)u(s_1 + w(1-\tau) - h, h) + pu(s_1 + (1-l)b + lw(1-\tau) - h, h)
\]

This problem is simplified by recognizing that \(\frac{\partial V_1}{\partial s} = 0, \frac{\partial V_1}{\partial h} = 0, and \frac{\partial V_1}{\partial \tau} = 0.\) These conditions follow from the Envelope theorem: since the individual has already maximized \(V\) given \(\{b, \tau\}, the first-order-conditions for maximization imply \(\frac{\partial V_1}{\partial s} = 0, etc.\) Hence, at the optimal \(b^*\)

\[
\frac{\partial V_1}{\partial b} = -[u_f(f_0, h) + (1-p)u_f(f_1^c, h) + pu_f(f_1^u, h))(\frac{\partial \tau}{\partial b} + p(1-l)u_f(f_1^u, h) = 0
\]

(14)
where $f_0$ denotes food consumption in period 0, $f^e_1$ denotes food consumption in period 1 when the agent is employed, and $f^u_1$ denotes the same when the agent is unemployed.

Starting with (14), if we follow Baily’s (1978) derivation (replacing $c$ by $f$), we obtain

$$\frac{\Delta f}{f}(b^*)\gamma_f = \varepsilon_{1-l,b}$$

Finally, using (5) gives

$$\frac{\Delta f}{f}(b^*)\gamma^l \frac{f}{f + h} = \varepsilon_{1-l,b}$$
Data Appendix: Measurement of Unemployment Duration

The SIPP reports the employment status of every individual over 15 years old for every week that they are in the sample. Weekly employment status (ES) can take the following values:

1. With a job this week
2. With a job, absent without pay, no time on layoff this week
3. With a job, absent without pay, spent time on layoff this week
4. Looking for a job this week
5. Without a job, not looking for a job, not on layoff

A job separation is defined as a change in ES from 1 or 2 to 3, 4, or 5. The duration of unemployment is computed by summing the number of consecutive weeks that ES $\geq 3$, starting at the date of job separation and stopping when the individual finds a job that lasts for at least one month (i.e. reports a string of four consecutive ES=1 or ES =2).

Individuals are defined as being on temporary layoff if they report ES = 3 at any point in the spell. They are included as “searching” if they report ES = 4 at any point during their spell.
NOTE – Figure plots utility over wealth in second period, after an agent has purchased a home in period 1. The transaction cost of selling the home makes it optimal not to move iff $W \in (s, S)$. The commitment affects preferences over wealth in two ways: (1) It makes agents locally risk averse by increasing the curvature of utility within the $(s, S)$ band; (2) It creates non-concavities in utility over wealth at $s$ and $S$, creating a gambling motive.
NOTE – Figure plots weekly unemployment insurance benefit payments in 1990 dollars for a claimant earning $400 per week over the entire year prior to unemployment. MI has no waiting period; TX has a retroactive payment for the waiting period after 4 weeks of unemployment; and CA has a waiting period and no retroactive provision. Following the notation used in the paper, MI and TX are INC = 1 states and CA is an INC = 0 state. For this claimant, all three states have a maximum duration for UI benefits of 26 weeks; since CA has a one week waiting period, benefits terminate in the 27th week after the job separation.
NOTE – Figure plots empirical survival probabilities for core sample of 4,475 unemployed men who took up UI. Survival probabilities are adjusted for censoring using the Kaplan-Meier correction. Data on unemployment durations is from the Survey of Income and Program Participation 1985-2000. The sample consists of all prime-aged males who (1) took up unemployment insurance benefits within one month of a job separation, (2) reported searching for a job at some point during the spell, (3) never reported that they were on temporary layoff, and (4) were unemployed for more than one week.
FIGURE 4

Empirical Hazards for UI Claimants

NOTE – Figure plots unemployment exit hazards for core sample of UI recipients. The hazard rate in week $t$ is defined as the probability that an individual who is unemployed for $t$ or more weeks finds a job in week $t$. In the full sample, the spikes at $t = 17$ and $t = 35$ reflect a reporting artefact in the SIPP known as the seam effect. The fact that this is a reporting error is evident by restricting attention to the 2,932 individuals who do not report losing their job on a seam between interviews; for this group, there are no spikes in the hazard rates at weeks 17 and 35. See Figure 3 for description of sample selection procedures.
NOTE – Figure plots survival probabilities for two categories of unemployment insurance claimants: those who live in states with average weekly benefit amounts (WBA) below the sample mean and those who live in states with WBAs above the mean. See Figure 3 for description of sample selection procedures.
NOTE – Figure plots survival probabilities for two categories of unemployment insurance claimants: those who live in states that have a one-week waiting period for UI benefits and make no retroactive payment for this week (INC = 0) and those who live in states that either have no waiting week or make a retroactive payment (INC = 1). Claimants in the INC = 1 states effectively receive an extra week of UI benefits at the beginning of the spell relative to claimants in INC = 0 states. See Figure 3 for description of sample selection procedures.
NOTE – Structural estimates of baseline job finding probabilities in the dynamic search model are plotted on right axis. These estimates correspond to specification (1) in Table 5. The baseline job finding probability, $\alpha_t$, in week $t$ is defined as the probability that an individual finds a job in that week when he exerts one unit of search effort. The cost of exerting one unit of search effort has been normalized to equal the welfare cost of losing $150$ of food in one week. On left axis, figure plots average level of search effort exerted by agents in the sample at the maximum likelihood estimates of the search model’s parameters.
NOTE – Figure shows empirical (solid) and predicted (dashed) survival curves for two categories of UI claimants: those who live in states that have weekly benefit amounts (WBA) below the mean and those who live in states that have WBAs above the mean. The predicted survival curves are based on specification (2) in Table 5, where the estimate of $\gamma^I = 7.07$. 
NOTE – Figure shows empirical (solid) and predicted (dashed) survival curves for two categories of UI claimants: those who live in states that have weekly benefit amounts (WBA) below the mean and those who live in states that have WBAs above the mean. The predicted survival curves are based on estimates of the search model without a measurement error correction in which the value of local curvature is constrained to equal $\gamma^l = 2.5$. All other parameters are unrestricted.
NOTE – The solid curve characterizes utility over wealth of a committed agent based on empirical estimates of the underlying utility function \( u(h,f) \) in the dynamic search model. The utility that is plotted has a local curvature of \( \gamma^l = 7.37 \), based on the empirical estimates of specification (1) in Table 5, and a global curvature of \( \gamma^g = 1 \). The figure assumes that the agent has committed half of his $40,000 income prior to entering this period to goods that have an adjustment cost of \( k = 7.5\% \). In addition, it assumes that the agent has an infinite horizon and does not face any credit constraints. The empirical estimates imply that such an agent will retain his commitment when his lifetime wealth is between \( s = $721,000 \) and \( S = $903,000 \), accommodating wealth fluctuations by adjusting consumption only over the uncommitted good in this region. The dashed curve extends utility in the committed region outside the \((s,S)\) band, showing what utility over wealth would look like if the agent never abandoned his commitments.
### TABLE 1
MEAN HOUSEHOLD EXPENDITURE SHARES

<table>
<thead>
<tr>
<th>Category</th>
<th>$5-10K</th>
<th>$30-40K</th>
<th>&gt;$70K</th>
<th>Overall Mean</th>
</tr>
</thead>
<tbody>
<tr>
<td>Shelter</td>
<td>0.20</td>
<td>0.20</td>
<td>0.21</td>
<td>0.20</td>
</tr>
<tr>
<td>Transport (excluding gas and maint)</td>
<td>0.12</td>
<td>0.16</td>
<td>0.16</td>
<td>0.16</td>
</tr>
<tr>
<td>Utilities, fuels, and public services</td>
<td>0.10</td>
<td>0.07</td>
<td>0.06</td>
<td>0.07</td>
</tr>
<tr>
<td>Health care</td>
<td>0.09</td>
<td>0.06</td>
<td>0.04</td>
<td>0.06</td>
</tr>
<tr>
<td>Education</td>
<td>0.03</td>
<td>0.01</td>
<td>0.02</td>
<td>0.02</td>
</tr>
<tr>
<td>Food</td>
<td>0.16</td>
<td>0.16</td>
<td>0.14</td>
<td>0.15</td>
</tr>
<tr>
<td>Apparel</td>
<td>0.06</td>
<td>0.05</td>
<td>0.06</td>
<td>0.15</td>
</tr>
<tr>
<td>Household supplies and furniture</td>
<td>0.06</td>
<td>0.06</td>
<td>0.10</td>
<td>0.06</td>
</tr>
<tr>
<td>Entertainment</td>
<td>0.04</td>
<td>0.05</td>
<td>0.06</td>
<td>0.05</td>
</tr>
<tr>
<td>Miscellaneous</td>
<td>0.03</td>
<td>0.02</td>
<td>0.02</td>
<td>0.02</td>
</tr>
<tr>
<td>Cash contributions</td>
<td>0.02</td>
<td>0.03</td>
<td>0.05</td>
<td>0.04</td>
</tr>
</tbody>
</table>

| Mean annual expenditure                    | $15,369| $32,609| $64,134| $35,930     |
| Mean take-home pay                         | $6,858 | $29,720| $90,748| $41,531     |

**NOTE** – The values are based on tabulations by the Bureau of Labor Statistics from the 2000 Consumer Expenditure Survey. Take-home pay is defined as gross income net of taxes and mandatory insurance/pension contributions.
<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Median</th>
<th>Deviation</th>
<th>Minimum</th>
<th>Maximum</th>
</tr>
</thead>
<tbody>
<tr>
<td>Age</td>
<td>37.00</td>
<td>35.00</td>
<td>11.08</td>
<td>18.00</td>
<td>64.00</td>
</tr>
<tr>
<td>1 = Married</td>
<td>0.61</td>
<td>1.00</td>
<td>0.49</td>
<td>0.00</td>
<td>1.00</td>
</tr>
<tr>
<td>Education (years)</td>
<td>12.07</td>
<td>12.00</td>
<td>3.06</td>
<td>0.00</td>
<td>18.00</td>
</tr>
<tr>
<td>Dependents</td>
<td>0.52</td>
<td>0.00</td>
<td>0.96</td>
<td>0.00</td>
<td>8.00</td>
</tr>
<tr>
<td>Pre-unemployment weekly wage</td>
<td>413.96</td>
<td>355.11</td>
<td>272.78</td>
<td>0.00</td>
<td>3,393.81</td>
</tr>
<tr>
<td>Annual pre-unemp family income</td>
<td>42,195.51</td>
<td>35,344.00</td>
<td>35,621.44</td>
<td>0.00</td>
<td>1,351,164.00</td>
</tr>
<tr>
<td>Pre-unemp home equity</td>
<td>32,053.37</td>
<td>9,440.32</td>
<td>49,494.13</td>
<td>-49,486.39</td>
<td>431,482.60</td>
</tr>
<tr>
<td>Pre-unemp liquid wealth</td>
<td>22,715.79</td>
<td>2,027.97</td>
<td>64,997.21</td>
<td>-8,973.50</td>
<td>1,153,178.00</td>
</tr>
<tr>
<td>Liquid wealth net of unsec. debt</td>
<td>18,441.26</td>
<td>186.32</td>
<td>66,588.72</td>
<td>-651,057.00</td>
<td>1,153,178.00</td>
</tr>
<tr>
<td>Weekly UI benefit amount</td>
<td>165.80</td>
<td>162.30</td>
<td>50.46</td>
<td>46.10</td>
<td>354.18</td>
</tr>
<tr>
<td>1 = Waiting period</td>
<td>0.86</td>
<td>1.00</td>
<td>0.35</td>
<td>0.00</td>
<td>1.00</td>
</tr>
<tr>
<td>1 = Retroactive payment</td>
<td>0.17</td>
<td>0.00</td>
<td>0.38</td>
<td>0.00</td>
<td>1.00</td>
</tr>
<tr>
<td>Maximum potential duration (weeks)</td>
<td>32.11</td>
<td>26.00</td>
<td>6.42</td>
<td>22.00</td>
<td>39.00</td>
</tr>
<tr>
<td>State unemployment rate</td>
<td>6.60</td>
<td>6.50</td>
<td>1.70</td>
<td>2.30</td>
<td>14.10</td>
</tr>
<tr>
<td>Weeks to job finding or censoring</td>
<td>20.89</td>
<td>16.00</td>
<td>21.03</td>
<td>2.00</td>
<td>171.00</td>
</tr>
<tr>
<td>1 = Censored</td>
<td>0.22</td>
<td>0.00</td>
<td>0.41</td>
<td>0.00</td>
<td>1.00</td>
</tr>
</tbody>
</table>

NOTE – The values are based on tabulations from SIPP panels spanning 1985-2000. The sample consists of the 4,457 prime-aged males who (1) took up unemployment insurance benefits within one month of a job separation, (2) reported searching for a job at some point during the spell, (3) never reported that they were on temporary layoff, and (4) were unemployed for more than one week. Wealth data are collected only in certain topical modules and are therefore available for only 2,530 observations. Liquid wealth is defined as total wealth minus home equity and vehicle equity. All monetary variables are in 1990 dollars.
## Table 3
### Semi-Parametric Hazard Model Estimates

<table>
<thead>
<tr>
<th>Variable</th>
<th>(1) Base Case</th>
<th>(2) With Simulated State Variable</th>
<th>(3) Controls</th>
<th>(4) Fixed Effects</th>
<th>(5) Non-UI</th>
</tr>
</thead>
<tbody>
<tr>
<td>Log State Avg. UI Weekly Benefit Amt</td>
<td>0.655 (0.151)</td>
<td>0.630 (0.118)</td>
<td>0.595 (0.178)</td>
<td>0.974 (0.097)</td>
<td></td>
</tr>
<tr>
<td>Wks Unemp * Log Avg. WBA</td>
<td>1.006 (0.008)</td>
<td>1.004 (0.008)</td>
<td>1.002 (0.009)</td>
<td>1.003</td>
<td></td>
</tr>
<tr>
<td>(Wks Unemp&gt;MaxDur) * Log Avg WBA</td>
<td>1.031 (0.020)</td>
<td>1.023 (0.020)</td>
<td>1.021 (0.021)</td>
<td>0.985</td>
<td></td>
</tr>
<tr>
<td><strong>Time-Varying Income Effects:</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>INC 1</td>
<td>0.886 (0.073)</td>
<td>0.882 (0.076)</td>
<td>0.876 (0.081)</td>
<td>0.895 (0.142)</td>
<td>1.046 (0.046)</td>
</tr>
<tr>
<td>INC 2</td>
<td>0.921 (0.103)</td>
<td>0.908 (0.084)</td>
<td>0.904 (0.085)</td>
<td>0.926 (0.026)</td>
<td>0.984 (0.042)</td>
</tr>
<tr>
<td>INC 3</td>
<td>1.144 (0.145)</td>
<td>1.110 (0.128)</td>
<td>1.110 (0.124)</td>
<td>1.130 (0.165)</td>
<td>1.058 (0.071)</td>
</tr>
<tr>
<td>INC 4</td>
<td>1.010 (0.115)</td>
<td>1.006 (0.099)</td>
<td>1.017 (0.100)</td>
<td>1.022 (0.022)</td>
<td>1.030</td>
</tr>
<tr>
<td>INC 5-12</td>
<td>1.030 (0.065)</td>
<td>1.027 (0.058)</td>
<td>1.068 (0.067)</td>
<td>1.060 (0.124)</td>
<td>1.095 (0.045)</td>
</tr>
<tr>
<td>State unemployment rate</td>
<td>0.960 (0.011)</td>
<td>0.951 (0.013)</td>
<td>0.956 (0.015)</td>
<td>0.995 (0.016)</td>
<td>0.976 (0.008)</td>
</tr>
<tr>
<td>1 = on seam</td>
<td>6.747 (0.260)</td>
<td>6.256 (0.213)</td>
<td>6.236 (0.212)</td>
<td>6.248 (0.214)</td>
<td>6.422 (0.201)</td>
</tr>
<tr>
<td>Age</td>
<td>0.983 (0.002)</td>
<td>0.986 (0.002)</td>
<td>0.982 (0.002)</td>
<td>0.982 (0.002)</td>
<td>0.987</td>
</tr>
<tr>
<td>1 = married</td>
<td>1.185 (0.048)</td>
<td>1.205 (0.047)</td>
<td>1.195 (0.049)</td>
<td>1.195 (0.049)</td>
<td>1.317 (0.034)</td>
</tr>
<tr>
<td>Years of schooling</td>
<td>1.013 (0.007)</td>
<td>1.022 (0.008)</td>
<td>1.011 (0.007)</td>
<td>1.011 (0.007)</td>
<td>1.030</td>
</tr>
<tr>
<td>1 = working spouse</td>
<td>1.063 (0.048)</td>
<td>1.060 (0.047)</td>
<td>1.057 (0.048)</td>
<td>1.057 (0.048)</td>
<td>0.930 (0.028)</td>
</tr>
<tr>
<td>Log Simulated UI WBA</td>
<td>0.696 (0.113)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Wks Unemp * Log Sim WBA</td>
<td>0.988 (0.006)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(Wks Unemp&gt;MaxDur) * Log Sim WBA</td>
<td>1.027 (0.020)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Sample Size</td>
<td>4457</td>
<td>4430</td>
<td>4430</td>
<td>4430</td>
<td>13108</td>
</tr>
</tbody>
</table>

**Note:** The sample for columns 1-4 consists of unemployed men from the SIPP who satisfy the restrictions described in Table 2. The sample for column 5 consists of unemployed men who did not report receipt of UI benefits at any point. Each column reports estimates of hazard ratios for a Cox model. Standard errors, reported in parentheses, are clustered by state. Notation for the time-varying income effect is as follows: INC 1 in states where individuals receive an additional week of benefits at beginning of spell. The INC 1 variable captures the effect of the INC dummy in month 1, INC 2 in month 2, etc. The on-seam variable indicates whether an individual is on a seam between SIPP interviews. Columns 1, 2, 4, and 5 use average weekly benefits in a (state, year) pair in place of the individual UI benefit rate. Column 3 uses simulated individual benefits based on the procedure described in the text. Columns 2-5 include industry, occupation, year, and month dummies, and a 10-piece log-linear wage spline. Column 4 also includes state dummies.
### Table 4
**Additional Hazard Model Estimates**

<table>
<thead>
<tr>
<th>Variable</th>
<th>(6) Retro Waiting Per</th>
<th>(7) Waiting Per</th>
<th>(8) Temp Layoffs</th>
<th>(9) All UI Recipients</th>
<th>(10) Maximum Benefits</th>
</tr>
</thead>
<tbody>
<tr>
<td>Log UI Benefit Level</td>
<td>0.659 (0.131)</td>
<td>0.608 (0.127)</td>
<td>0.650 (0.125)</td>
<td>0.743 (0.133)</td>
<td></td>
</tr>
<tr>
<td>Wks Unemp * Log UI Benefit Level</td>
<td>1.004 (0.009)</td>
<td>1.005 (0.010)</td>
<td>1.003 (0.008)</td>
<td>1.002 (0.007)</td>
<td></td>
</tr>
<tr>
<td>(Wks Unemp&gt;MaxDur) * Log UI Ben</td>
<td>1.013 (0.023)</td>
<td>1.012 (0.022)</td>
<td>1.021 (0.019)</td>
<td>1.016 (0.010)</td>
<td></td>
</tr>
<tr>
<td><strong>Time-Varying Income Effects</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>INC 1</td>
<td>0.893 (0.120)</td>
<td>0.862 (0.080)</td>
<td>0.877 (0.073)</td>
<td>0.902 (0.075)</td>
<td>0.700 (0.124)</td>
</tr>
<tr>
<td>INC 2</td>
<td>0.941 (0.123)</td>
<td>0.875 (0.084)</td>
<td>0.898 (0.081)</td>
<td>1.008 (0.072)</td>
<td>1.134 (0.148)</td>
</tr>
<tr>
<td>INC 3</td>
<td>1.168 (0.150)</td>
<td>1.057 (0.160)</td>
<td>1.123 (0.118)</td>
<td>1.128 (0.103)</td>
<td>1.387 (0.252)</td>
</tr>
<tr>
<td>INC 4</td>
<td>0.930 (0.147)</td>
<td>1.056 (0.110)</td>
<td>1.040 (0.099)</td>
<td>1.091 (0.097)</td>
<td>1.066 (0.119)</td>
</tr>
<tr>
<td>INC 5-12</td>
<td>0.993 (0.065)</td>
<td>1.038 (0.062)</td>
<td>1.022 (0.056)</td>
<td>1.083 (0.061)</td>
<td>1.119 (0.081)</td>
</tr>
<tr>
<td>State unemployment rate</td>
<td>0.952 (0.013)</td>
<td>0.948 (0.014)</td>
<td>0.956 (0.013)</td>
<td>0.961 (0.010)</td>
<td>0.956 (0.016)</td>
</tr>
<tr>
<td>1 = on seam &amp; 0.164 (0.234)</td>
<td>6.232 (0.237)</td>
<td>5.940 (0.197)</td>
<td>5.951 (0.203)</td>
<td>6.106 (0.286)</td>
<td></td>
</tr>
<tr>
<td>Age</td>
<td>0.982 (0.002)</td>
<td>0.983 (0.002)</td>
<td>0.983 (0.002)</td>
<td>0.983 (0.002)</td>
<td>0.983 (0.003)</td>
</tr>
<tr>
<td>1 = married</td>
<td>1.183 (0.054)</td>
<td>1.156 (0.052)</td>
<td>1.187 (0.048)</td>
<td>1.250 (0.040)</td>
<td>1.243 (0.077)</td>
</tr>
<tr>
<td>Years of schooling</td>
<td>1.011 (0.008)</td>
<td>1.009 (0.007)</td>
<td>1.013 (0.008)</td>
<td>1.015 (0.005)</td>
<td>1.004 (0.009)</td>
</tr>
<tr>
<td>1 = working spouse &amp; 1.077 (0.051)</td>
<td>1.112 (0.053)</td>
<td>1.063 (0.045)</td>
<td>1.026 (0.034)</td>
<td>1.031 (0.052)</td>
<td></td>
</tr>
<tr>
<td>Log State Max UI WBA</td>
<td>0.877 (0.171)</td>
<td>0.989 (0.011)</td>
<td>0.877 (0.171)</td>
<td>0.989 (0.011)</td>
<td></td>
</tr>
<tr>
<td>Wks Unemp * Log Max WBA</td>
<td>(0.051)</td>
<td>(0.053)</td>
<td>(0.045)</td>
<td>(0.034)</td>
<td>1.027 (0.035)</td>
</tr>
<tr>
<td>(Wks Unemp&gt;MaxDur) * Log Max WBA</td>
<td>Sample Size</td>
<td>3687</td>
<td>3837</td>
<td>4682</td>
<td>6966</td>
</tr>
</tbody>
</table>

**NOTE** – Each column reports estimates of hazard ratios for a Cox model with a different sample. Column 6 excludes states that make retroactive payments for their waiting weeks. Column 7 excludes states that do not have a waiting week. Column 8 adds individuals on temporary layoff to the core sample described in Table 2. Column 9 adds UI recipients who did not take up UI in the first month to the core sample. Column 10 includes only those in the core sample who report a weekly wage above $400. All columns include industry, occupation, year, and month dummies, and a 10-piece log-linear wage spline. Columns 6-9 use average weekly benefits in a (state, year) pair in place of the individual UI benefit rate. Column 10 uses the maximum weekly benefit amount in a given (state, year). Standard errors, reported in parentheses, are clustered by state. See Table 3 for descriptions of the INC and on-seam variables.
### TABLE 5
STRUCTURAL ESTIMATES OF SEARCH MODEL

<table>
<thead>
<tr>
<th>Parameter</th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Local curvature ($\gamma$)</td>
<td>7.3632</td>
<td>7.0668</td>
<td>7.3826</td>
<td>6.1538</td>
</tr>
<tr>
<td></td>
<td>(2.07)</td>
<td>(2.93)</td>
<td>(2.79)</td>
<td></td>
</tr>
<tr>
<td>Convexity of search effort ($\theta$)</td>
<td>1.5455</td>
<td>1.5216</td>
<td>1.5436</td>
<td>1.5672</td>
</tr>
<tr>
<td>Baseline Job Finding Probs:</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\alpha_1$</td>
<td>0.0393</td>
<td>0.0369</td>
<td>0.0378</td>
<td>0.0266</td>
</tr>
<tr>
<td>$\alpha_5$</td>
<td>0.0363</td>
<td>0.0341</td>
<td>0.0251</td>
<td></td>
</tr>
<tr>
<td>$\alpha_9$</td>
<td>0.0378</td>
<td>0.0334</td>
<td>0.0363</td>
<td>0.0261</td>
</tr>
<tr>
<td>$\alpha_{13}$</td>
<td>0.0349</td>
<td>0.0291</td>
<td>0.0356</td>
<td>0.0240</td>
</tr>
<tr>
<td>$\alpha_{17}$</td>
<td>0.0326</td>
<td>0.0369</td>
<td>0.0224</td>
<td></td>
</tr>
<tr>
<td>$\alpha_{21}$</td>
<td></td>
<td></td>
<td>0.0284</td>
<td></td>
</tr>
<tr>
<td>$\alpha_{25}$</td>
<td>0.0319</td>
<td>0.0270</td>
<td>0.0311</td>
<td>0.0219</td>
</tr>
<tr>
<td>$\alpha_{32}$</td>
<td>0.0282</td>
<td>0.0270</td>
<td></td>
<td>0.0198</td>
</tr>
<tr>
<td>$\alpha_{40}$</td>
<td>0.0208</td>
<td>0.0199</td>
<td>0.0222</td>
<td>0.0146</td>
</tr>
<tr>
<td>$\alpha_{48}$</td>
<td>0.0215</td>
<td>0.0192</td>
<td></td>
<td>0.0146</td>
</tr>
<tr>
<td>$\alpha_{55}$</td>
<td>0.0171</td>
<td>0.0178</td>
<td>0.0185</td>
<td>0.0125</td>
</tr>
<tr>
<td>State unemployment rate ($\alpha_u$)</td>
<td>-0.0460</td>
<td>-0.0445</td>
<td>-0.0459</td>
<td>-0.0775</td>
</tr>
<tr>
<td>Fraction of seam reporters ($\mu$)</td>
<td>0.3283</td>
<td>0.3258</td>
<td>0.3197</td>
<td></td>
</tr>
<tr>
<td>Log likelihood value</td>
<td>-12688.74</td>
<td>-13703.72</td>
<td>-12690.95</td>
<td>-12695.76</td>
</tr>
<tr>
<td>Sample size</td>
<td>4457</td>
<td>4457</td>
<td>4457</td>
<td>4457</td>
</tr>
</tbody>
</table>

**NOTE** – Each column reports estimates of the dynamic job search model. Standard errors, reported in parentheses if available, are derived by inverting a likelihood ratio test. Columns 1, 3, and 4 include a correction for mismeasurement of unemployment durations to account for the seam recall problem discussed in the text. In columns 1-3, exogenous parameters for the search model are as follows: weekly wage $w = 400$, commitment $\bar{h} = 200$, pre-unemployment assets $a = 500$, weekly discount rate $\beta = 0.9^{1/50}$, disutility of labor $\psi_S = 100$, welfare payment $\bar{w} = 100$, and borrowing/late-payment constraint $\bar{B} = 11,000$. In column 4, $\bar{B} = 10,000$. See notes to Table 2 for description of sample selection procedures.