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A Contribution to the Empirics of Reservation Wages†

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This paper provides evidence on the behavior of reservation wages over the spell of unemployment, using high-frequency longitudinal data on unemployed workers in New Jersey. In comparison to a calibrated job search model, the reservation wage starts out too high and declines too slowly, on average, suggesting that many workers persistently misjudge their prospects or anchor their reservation wage on their previous wage. The longitudinal nature of the data also allows for testing the relationship between job acceptance and the reservation wage, where the reservation wage is measured from a previous interview to avoid bias due to cognitive dissonance. (*JEL* J22, J31, J64)

The sequence of reservation wages over the spell of unemployment plays a central role in search theory. Starting with the seminal contribution of Mortensen (1977), several papers have studied the implications of different sources of nonstationarity for the path of the reservation wage over the spell of unemployment.[1] Yet, evidence on the behavior of reservation wages over the spell of unemployment remains scarce and relies mainly on cross-sectional data for those with varying lengths of unemployment at the time of the survey, which is susceptible to bias due to the evolving sample of unemployed workers over time. This paper attempts to fill this void by using repeated information on self-reported reservation wages over the course of an unemployment spell.[2] Specifically, we use data from our Survey of Unemployed Workers in New Jersey, which collected *weekly* information on

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[1] Mortensen analyzes the consequences of the limited duration of unemployment benefits for the optimal reservation wage path over the spell of unemployment. Other early contributions are Burdett and Vishwanath (1988), who consider learning about the job offer distribution, and Danforth (1979), who sets up a model with declining wealth over the spell of unemployment. Kasper (1967) provides early cross-sectional evidence.

[2] Devine and Kiefer (1991) review early evidence on reservation wages and find only two studies with repeated observations of the reservation wage of the same unemployed person, but for small and selected samples. More recently, Addison, Centeno, and Portugal (2009) provide longitudinal estimates of the effect of unemployment duration on the reservation wage, using *yearly* observations from the European Community Household Panel for the years 1994–1999, and find no significant relationship between unemployment duration and the reservation wage.

self-reported reservation wages for a period of up to 24 weeks for a sample of 6,025 unemployment insurance (UI) recipients in New Jersey, to examine the pattern of reservation wages over the spell of unemployment. In addition, we provide the first evidence on whether the self-reported reservation wage relative to the offered wage predicts job acceptance.

A small literature has previously examined the response of reservation wages to UI benefits by analyzing data on self-reported reservation wages, with mixed results. Feldstein and Poterba (1984) find a relatively large elasticity of reservation wages with respect to UI benefits, whereas DellaVigna and Paserman (2005) report a small effect of a dummy of UI receipt on the reservation wage. A burgeoning and complementary literature also indirectly infers the behavior of reservation wages by examining post-employment wages for those who become reemployed, often exploiting discontinuities in UI eligibility or duration across workers. These studies usually find small or insignificant effects of benefits on reemployment wages, and thus suggest that the reservation wage is relatively insensitive to UI benefit parameters (see Card, Chetty, and Weber 2007; and more recently, Schmieder, von Wachter, and Bender 2013; and Nekoei and Weber 2013). While these studies often use compelling identification strategies to estimate the effects of UI benefits on reemployment wages, they are potentially subject to selection bias in identifying reservation wage behavior over time because only a subset of unemployed workers are offered and accept employment, and those who do regain employment may have lower (and more steeply declining) reservation wages than those who remain unemployed. The longitudinal nature of our reservation wage data allows us to overcome this selection issue when analyzing the effect of unemployment duration on the reservation wage.[3]

One reason why it is important to empirically model reservation wages is that, in an influential article, Shimer and Werning (2007) showed that the elasticity of the reservation wage with respect to UI benefits reflects the optimality of UI benefits. They demonstrate that the after-tax reservation wage is a sufficient statistic for an unemployed worker's welfare, and show that if the pretax reservation wage is highly elastic with respect to UI benefits, it may be welfare improving to *raise* UI benefits. Based on the limited evidence available, they tentatively conclude that previous estimates of the UI benefit elasticity of the reservation wage imply large welfare gains of raising UI benefits above the level currently available in the United States. A better understanding of reservation wages can shed light on the optimality of UI benefits.

Our analysis in Section I starts by providing a stylized partial equilibrium model, where unemployed workers draw from a stationary distribution of wage offers and face a UI benefit of limited duration. This provides a benchmark against which to judge the observed adjustment in reservation wages over the spell of unemployment. Section II describes the survey design and Section III reports results from a cross-sectional analysis of reservation wages. We find no significant relationship

[3] Additionally, the reemployment wage reflects the reservation wage *and* the potential wage offer distribution and thus, even in the presence of precise estimates of the effect of UI on the accepted wage, the effect of UI on the reservation wage depends on the density of wage offers at the reservation wage and thus relies on assumptions about the dispersion of potential wage offers as well as the fraction of rejected job offers. These complications are not an issue for direct measures of the reservation wage.

between unemployment benefits and reservation wages, but financial circumstances such as severance payments and savings appear to be positively associated with the reservation wage. Consistent with search theory, we find that risk-averse workers set lower reservation wages. Section IV reports the main results from our longitudinal analysis. We find that reservation wages decline at a modest rate over the spell of unemployment, with point estimates ranging from 0.05 to 0.14 percent per week of unemployment, or 2.4 percent to 7.3 percent per year of unemployment. Moreover, our results indicate that the decline in reservation wages is concentrated among individuals age 51 and older, and those with savings of $10,000 or more at the start of the survey. We find no evidence of a decline in reservation wages for those who rely exclusively on UI benefits for income support. In comparison to the stylized model, reservation wages start out too high and decline too slowly, suggesting that many workers persistently misjudge their prospects or anchor their reservation wage to their previous wage.

Section V analyzes the job acceptance and rejection behavior of those who received job offers. Specifically, we model the likelihood of accepting a job offer as a function of the reservation wage, the offered wage, an indicator of whether the offered wage exceeds the reservation wage, and other variables. To avoid simultaneous reporting bias, we relate the reservation wage from a previous interview to the likelihood of accepting a job offer in a subsequent period, and thus minimize possible effects of cognitive dissonance. The results suggest that the self-reported reservation wage relative to the offered wage is efficacious for predicting future job acceptance, and is a stronger predictor than the pre-displacement wage. Indeed, there is a discrete rise in job acceptance if the offered wage equals or exceeds the reservation wage. We further find that workers are more likely to exit UI early if they report a relatively low reservation wage compared to their previous wage. Section VI provides some concluding observations on the implications of our findings.

I. Model

As a benchmark against which to judge our empirical estimates, we start by providing a simple model of unemployment and reservation wages. The model is based on Mortensen's (1977) canonical search model, but abstracts from endogenous search effort to focus on the reservation wage. The unemployed worker faces a constant arrival rate λ of job offers, and wages w are drawn from a wage offer distribution $F(w)$. The unemployed worker is assumed to have no savings, so consumption is equal to the unemployment benefit b when unemployed and equal to the wage w when employed. The value function $U(\cdot)$ of an unemployed worker is

$$(1) \quad U(t) = u\big(b(t)\big) + \beta \max_R \Big\{ U(t-1) + \lambda \int_R \big(W(x, m=0) - U(t-1)\big) dF(x) \Big\},$$

where t is the remaining duration of the unemployment benefit b, $u(\cdot)$ is the flow utility function, β is the discount factor, m is the number of months employed, and $W(w, m = 0)$ is the value of starting a job paying wage w for newly employed workers. Benefits are assumed to last for a maximum of T periods.

It is important to account for a minimum period when newly employed workers do not qualify for UI benefits, which has implications for the decline in reservation wages over the spell of unemployment. To see this, consider the extreme case where requalification for UI benefits is instantaneous: this will lead unemployed workers to reduce their reservation wage more strongly over the spell of unemployment as working on a new job will immediately qualify them for a full spell of unemployment benefits in the event of job loss on that new job. We introduce a qualifying period by assuming that unemployed workers do not qualify for UI benefits for the first $m < 6$ months of a job spell but then requalify for UI.

The value function for an employed worker on a job paying wage w who has yet to qualify for UI benefits (i.e., $m < 6 = \overline{m}$) is

$$(2)\quad W(w,m) = \mathrm{u}(w) + \beta\Big[(1-\delta)W(w,m+1) + \delta \mathrm{U}(0) + \lambda_e(1-\delta)\int_w \big(W(x,m+1) - W(w,m+1)\big)dF(x)\Big],$$

and for a worker who has qualified for benefits (i.e., $m \geq \overline{m}$) is

$$(3)\quad W(w,\overline{m}) = \mathrm{u}(w) + \beta\Big[(1-\delta)W(w,\overline{m}) + \delta \mathrm{U}(\mathrm{T}) + \lambda_e(1-\delta)\int_w \big(W(x,\overline{m}) - W(w,\overline{m})\big)dF(x)\Big],$$

where $\overline{m} = 6$ is the number of months it takes to requalify for unemployment benefits, δ is the exogenous separation probability, and λ_e the probability of receiving a job offer while employed. We assume that workers are not eligible for unemployment benefits if they quit their job. An unemployed worker's optimal choice is characterized by the following equation:

$$(4)\qquad W\big(R(t), m = 0\big) = U(t-1),$$

which implies that the reservation wage is set so the worker is indifferent between starting a new job at the reservation wage or remaining unemployed with *t-1* periods of benefits left. The model predicts that reservation wages will decline over the spell of unemployment as the unemployed worker approaches the point of UI exhaustion, and will remain constant thereafter, because the value associated with not working gradually declines as benefits approach the exhaustion point and remains constant thereafter.

We calibrate the model to derive quantitative predictions of how much one should expect reservation wages to decline over the spell of unemployment in this simple model. We view our model as a reduced form version of a model where unemployed workers are partially insured against the risk of exhausting UI (e.g., through savings, family and friends, or other government-provided insurance programs), and assume a drop in consumption at UI exhaustion of 31.3 percent, which is based on an average of estimates reported in Low, Meghir, and Pistaferri (2010) and Gruber (1997). Specifically, we computed the implied drop in food consumption associated with UI benefits falling from a replacement rate of 60 percent (the median in NJ) to

zero from Table 3 of Low, Meghir, and Pistaferri (2010) and from the instrumental variables estimates in Table 1 of Gruber (1997). Note that we adjusted Gruber's estimate to account for the fact that food consumption tends to respond less to income changes than other consumption items, by scaling his estimate according to the income elasticity of food consumption (0.61) reported by Blundell, Pashardes, and Weber (1993). We normalize the mean wage of the job offer distribution to 1 and calibrate the utility flow at the beginning of the unemployment spell to match the average unemployment duration in our sample (six months). We calibrate the model at the monthly frequency, with a discount factor β of 0.996 (about a 5 percent annual discount rate), a job offer arrival rate λ of 0.3 per month while unemployed, and a job offer arrival rate λ_e of 0.1 per month if employed. We assume a constant relative risk aversion (CRRA) utility function with a coefficient of relative risk aversion of 2. An employed worker may lose his job with a monthly probability of 0.02, and the wage offer distribution is assumed to be log normal with a standard deviation of 0.24. We take this value from Hall and Mueller (2013), who estimate the dispersion of wage offers for an individual job seeker using the same data.

Figure 1 shows the predicted decline in reservation wages over 23 months (99 weeks), which was the maximum duration of benefits during our survey period. The model implies a decline in the reservation wage relative to the past wage from 0.74 to 0.59 up to the point of UI exhaustion, which corresponds to a decline of around 0.23 percent per week of unemployment. The dashed line corresponds to the reservation wage if unemployment benefits last forever, while the dotted line corresponds to the reservation wage in an environment with no UI, and thus the consumption level is set to the same level as at UI exhaustion. Interestingly, with UI, the reservation wage declines below the dotted line. The reason for this result is that an unemployed worker can qualify for a new spell of unemployment benefits after a certain period and is thus willing to accept a lower wage offer.

An implication of the calibrated model is that the decline in reservation wages up to the point of UI exhaustion is approximately equal to the response of reservation wages to a permanent elimination of UI benefits, as the overall decline of the reservation wage corresponds approximately to the difference between the dashed and the dotted lines. The decline of reservation wages over the spell of unemployment is thus informative about the elasticity of the reservation wage with respect to UI benefits. Shimer and Werning (2007) showed that the elasticity of the reservation wage to UI benefits is an important statistic in the design of optimal UI policy. In particular, they showed that the after-tax reservation wage is a sufficient statistic of the unemployed worker's welfare. Thus, if the pretax reservation wage is highly elastic with respect to UI benefits, it may be welfare improving to *raise* UI benefits, whereas if the pretax reservation wage is inelastic to UI benefits, the optimal policy would be to reduce UI benefits.

Many factors that are not incorporated in this simple model probably would lead rational job searchers to *hasten* the pace of decline in their reservation wage over a spell of unemployment. One can classify these additional factors into three broad categories. First, some unemployed workers have positive savings at the start of the spell, and reduce their reservation wage as they reduce their savings (see Danforth 1979). Because a majority of unemployed workers have only a trivial amount of

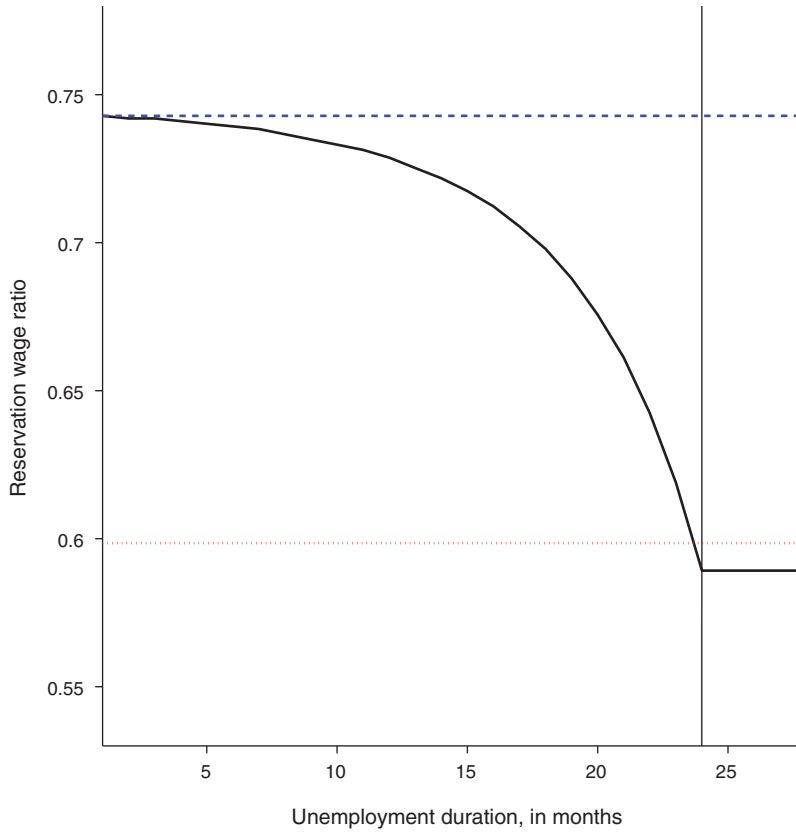


Figure 1. The Reservation Wage over the Spell of Unemployment in the Model

Notes: The solid black line in the figure shows the decline of the reservation wage over the spell of unemployment in the calibrated version of the model described in Section I. The dashed line shows the reservation wage path in an environment where unemployment benefits have unlimited duration, whereas the dotted line shows the reservation wage in an environment without unemployment insurance. Note that the reservation wage ratio is defined as the reservation wage divided by the previous wage.

savings, however, extending the model to include savings is an unnecessary complication for a large portion of the unemployed. A second factor involves learning about the distribution of potential wage offers or the probability of receiving a job offer (see Burdett and Vishwanath 1988). As unemployed workers update their beliefs regarding the potential wage offer distribution or the job offer arrival rate, they should also change their reservation wage. Such updating is likely to be negative due to human capital erosion, initial over confidence, and adverse signaling effects of long-term unemployment. Kroft, Lange, and Notowidigdo (2013) provide some evidence for the latter. As a consequence, the value of remaining unemployed, and thus the reservation wage, likely decline with unemployment duration. Third, the psychological costs associated with unemployment appear to increase over the spell of unemployment, increasing the cost of being unemployed.[4]

[4] See Clark et al. (2008) for evidence on the psychological effects of unemployment.

Lastly, allowing for endogenous search effort in the model would attenuate the effect of unemployment duration on reservation wages. We abstract from endogenous search effort, however, because it would also lead to the prediction that search effort increases over the spell of unemployment, whereas earlier research has found that search effort declines over the spell of unemployment (Krueger and Mueller 2011; Wanberg et al. 2012). Accommodating negative duration dependence in search effort would entail assuming some form of discouragement in the search process, or negative duration dependence in the wage offer distribution or the job offer arrival rate over time, but as discussed above, these alterations would hasten the pace of the decline in the reservation wage over the spell of unemployment.

Our benchmark search model with a non-stationary UI benefit neglects these additional elements, and thus likely understates the pace at which rational unemployed workers would be expected to adjust their reservation wage over the spell of unemployment. For this reason, we take the prediction that reservation wages should decline by about 0.23 percent per week over a 99 week spell of unemployment benefits as a lower bound for the rate predicted by conventional search models.

II. Data and Descriptive Statistics

We use data from our Survey of Unemployed Workers in New Jersey, in which we interviewed a sample of 6,025 unemployed workers each week for up to 24 weeks.[5] The sample frame for the study was drawn with a stratified random sampling procedure from the universe of unemployed workers in New Jersey as of September 28, 2009. The strata consisted of 8 intervals of duration of unemployment (0–2 weeks, 10–12 weeks, 20–22 weeks, 30–32 weeks, 40–42 weeks, 50–53 weeks, 60–69 weeks, and 70 and more weeks at the end of September 2009) and whether an e-mail address was on file. Individuals in the sample frame were invited to participate in the study for a period of 12 weeks, and the long-term unemployed were invited to participate for an additional 12 weeks at the end of the initial study period. Each week, participants were asked to complete a short online survey about their reservation wages and job offers, as well as their time use and job search activities. To encourage participation—and provide revealed preference evidence on discount factors—respondents were offered the choice between receiving a $20 visa gift card at the start of the study or a $40 visa gift card after 12 weeks (guaranteed regardless of participation in subsequent interviews).

The unemployment rate in NJ was stable over the survey period (October 2009 through March 2010), only fluctuating between 9.6 percent and 9.7 percent, and closely mirroring the national unemployment rate. Unemployment insurance in NJ is slightly more generous than in other states, replacing 60 percent of previous earnings up to a maximum weekly benefit of $584 in 2009. At the start of the survey period on October 13, 2009, the maximum duration of UI benefits was 79 weeks due to both federal and state-level extensions, up from the 26 weeks in normal times. On November 8, 2009, the federal Emergency Unemployment Compensation (EUC)

[5] See the Appendix in Krueger and Mueller (2011) for a more detailed description of the survey. The survey data can be downloaded at http://opr.princeton.edu/archive/njui/.

program was extended by another 20 weeks, increasing the maximum duration of UI benefit receipt to 99 weeks. Benefits temporarily lapsed for 640 workers in our sample who had exhausted benefits after 79 weeks but then qualified for an additional 20 weeks of benefits when Congress extended benefits to 99 weeks. These features of the UI program provide some exogenous variability in the generosity of UI benefits during our sample period.

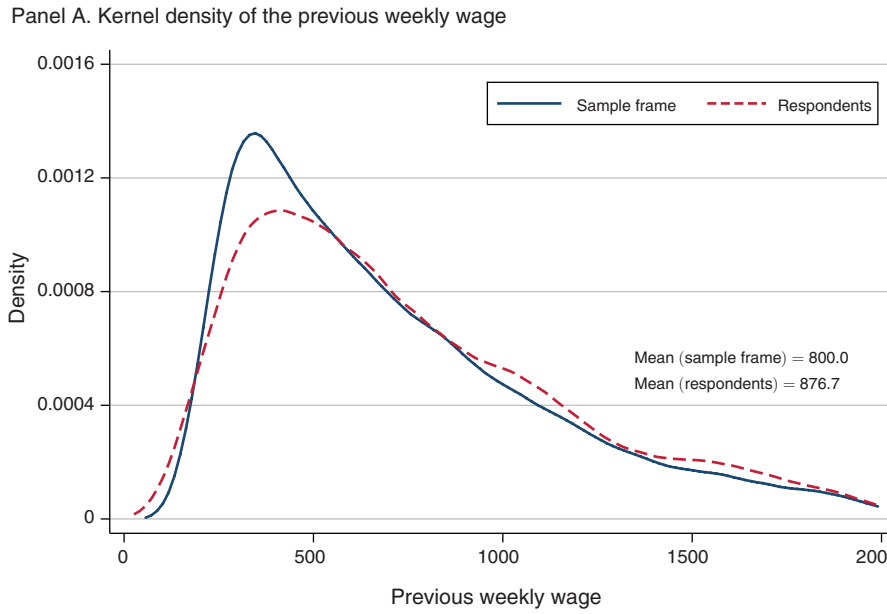
As discussed in Krueger and Mueller (2011), one concern with the survey is the low response rate. Only 10 percent of those who were contacted participated in the first interview, and respondents who participated in the first survey completed only around 40 percent of the subsequent weekly interviews. To address this issue, we constructed a set of survey weights, which adjust for the sampling probability of each strata as well as nonresponse based on demographic characteristics, such as age, gender, race, ethnicity, and educational attainment, for each calendar week during the survey period. Moreover, with updated UI records, Krueger and Mueller (2011) show that the Kaplan-Meier UI weekly exit rate of the respondents closely tracks that of the sample frame. In addition, we obtained access to administrative data on earnings prior to UI receipt for our sample, as well as updated administrative data on earnings in 2010 for those who were reemployed in New Jersey. Panel A of Figure 2 shows the kernel density of the prior wage for the respondents and the entire sample frame, using weights adjusting for different sampling probabilities and nonresponse. It is apparent from the figure that our respondent sample slightly overrepresents workers with higher wages on their previous job, but when we divide the average weekly wage in 2010 by the pre-unemployment wage for those who found jobs, the ratio is 0.90 for the respondents and 0.92 for the entire sample frame. Panel B of Figure 2 shows the kernel density of this ratio for the sample frame and the respondents, which look similar for both samples.[6] This suggests that, relative to their previous wage, respondents accepted similar wage offers as nonrespondents.

The reservation wage question was phrased, "Suppose someone offered you a job today. What is the lowest wage or salary you would accept (before deductions) for the type of work you are looking for?" This question is similar to the one used in the May 1976 Current Population Survey that was analyzed by Feldstein and Poterba (1984).[7] In addition, the survey asked participants whether they received any job offers during the last seven days, about the wage that they were offered, and whether they accepted the offer or not. Finally, we have access to administrative data on UI weekly benefit rates and wages on the prior job. As mentioned above, we also have administrative data on reemployment earnings in 2010 from payroll tax records, with the limitation that they are restricted to pay earned in New Jersey and may include earnings from part-time jobs during UI receipt.

The survey contains 39,201 interviews from a total of 6,025 unemployed workers. We restrict our sample to those aged 20–65 and exclude individuals once they accept a job offer. We follow Feldstein and Poterba (1984) and focus our analysis

[6] The estimated averages are obtained by using the same weights as in columns 3 and 4 of table 1 in Krueger and Mueller (2011). We applied the same thresholds for trimming on our survey measure of the reservation wage.

[7] The question in the Current Population Survey of May 1976 was, "What is the lowest wage or salary you would accept (before deductions) for this type of work?"



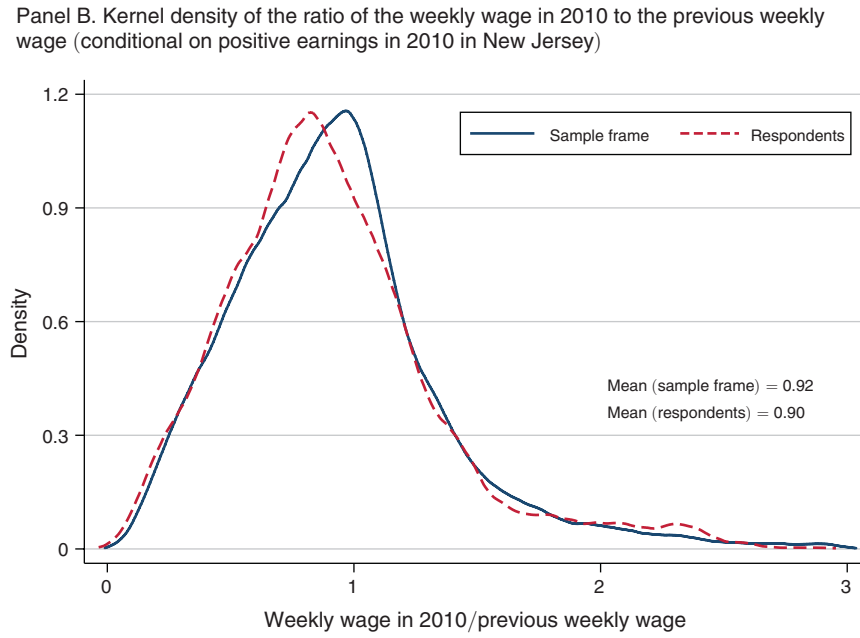


Figure 2. Comparison of Sample Frame and Respondent Sample

Notes: Both the weekly wage in 2010 and the previous wage are from administrative data. The previous weekly wage is computed from earnings in the base year, which are used to compute the unemployment benefits, whereas the weekly wage in 2010 is computed from New Jersey wage records (and thus earnings from other states are omitted). Weights are used to adjust for sampling probability and non-response.

throughout the paper on the reservation wage ratio, defined as the reservation wage divided by the previous wage, to control for individual-level heterogeneity.[8] Figure 3 shows the kernel density of the log of the reservation wage ratio for unemployed workers in their first interview in the survey. The cross-sectional distribution of the reservation wage ratios is close to log normal, the mean of the log of the ratio is −0.08, and the standard deviation is 0.37. Nearly 80 percent of newly unemployed workers reported a reservation wage ratio above 0.74, the optimal level from our calibrated model, suggesting either overly optimistic expectations or systematic misreporting of reservation wages. It is possible that the minimum wage imposes a lower bound on the reservation wage, but less than 1 percent of the sample reported an hourly reservation wage that was equal to or less than the NJ minimum wage at the time of the survey ($7.25).

To gain further insights into the determinants of the reservation wages, Appendix Table A1 provides some Mincerian wage regressions of the log hourly reservation wage, the log hourly previous wage, and the log hourly offered wage on observable characteristics. Columns 1 and 2 show that for respondents in their first interview, the coefficients on educational attainment and potential work experience are very similar for the log hourly reservation wage and the log hourly previous wage. These estimates are also similar to the values obtained when running the Mincerian wage regression in other datasets.

III. Reservation Wages, Unemployment Insurance, and Liquidity: A Cross-Sectional Analysis

A long-standing question is the extent to which reservation wages respond to aspects of the unemployment insurance program. Feldstein and Poterba (1984) analyzed a cross-sectional sample of 2,228 unemployed workers and found that the reservation wage ratio is positively associated with the benefit replacement ratio, with a 1 percentage point increase in the replacement ratio associated with an increase in the reservation wage ratio of between 0.13 and 0.42 percentage points. As pointed out by Shimer and Werning (2007), there are several shortcomings to their approach, as the coefficient on the replacement rate is biased toward one in the presence of measurement error in the previous wage or if there is an omitted third factor, such as local labor market conditions or other state-specific effects, that may be correlated with both the reservation wage ratio and the replacement rate. Moreover, even if pre-unemployment wages are measured without error, variation in the ratio could be driven to a large extent by other sources of randomness in earnings (e.g., random match quality).

Although the main novelty of our dataset is its longitudinal nature, it also permits another look at the relationship between the generosity of unemployment insurance

[8] We use the ratio of weekly wages because the administrative data do not contain information on hours. Hours worked on the last job were collected in the survey, but are likely to introduce measurement error in the measure of the previous wage. Following Feldstein and Poterba (1984), we trimmed observations with reservation wage ratios greater than three or smaller than one-third. In addition, we trimmed observations with weekly reservation wages greater than $8,000 or less than $100 and hourly reservation wages greater than $100 or less than $5. This procedure led us to exclude 2,692 observations out of a total of 36,514 observations.

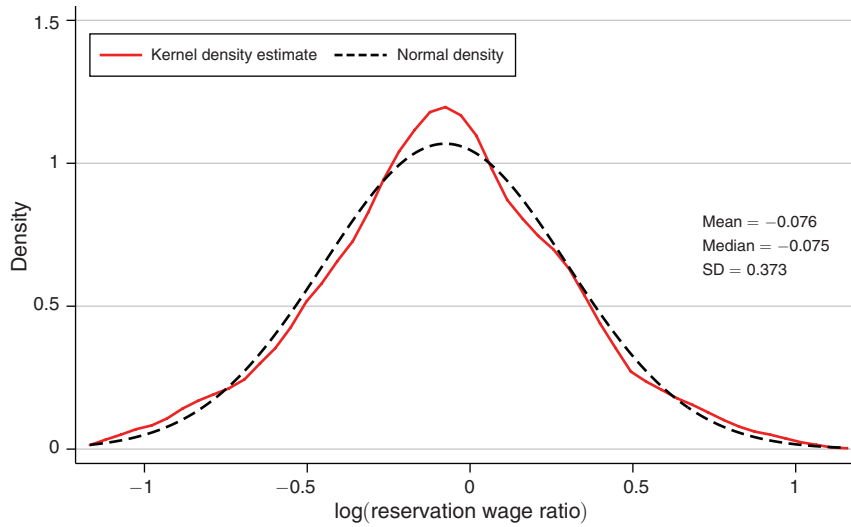


Figure 3. Kernel Density of the Ratio of the Reservation Wage to the Previous Wage

Note: kernel = epanechnikov, bandwith = 0.0600

benefits and reservation wages. The NJ dataset has several advantages over those used in the previous literature, as the sample size is larger, the survey is tailored to UI recipients, and we have access to administrative data on pre-unemployment wages and UI weekly benefits, which should reduce problems associated with measurement error. To be clear, UI benefits in NJ are a strict function of earnings in the base year (the first four quarters of the five quarters preceding the start of the unemployment spell), but as pointed out by Shimer and Werning (2007, 1159), it is "possible to exploit nonlinearities in benefit schedules [...] to obtain the desired variation." The maximum benefit constrains a third of our sample, and creates a nonlinear relationship.

The regression results in column 1 of Table 1 show that the effect of UI benefits on the log of the reservation wage ratio is negative but statistically insignificant, and the 95 percent confidence interval ranges from an elasticity of −0.16 to 0.07, below Feldstein and Poterba's estimates.[9] As explained above, the Feldstein-Poterba specification likely produces estimates that are biased upward. Indeed, if we estimate the same regression as in column 1 of Table 1, but include the log of the replacement rate instead of including the log of the weekly benefit and the log of the weekly previous as separate variables, we obtain an elasticity of 0.82 (with a standard error of 0.048). Note that this estimate is above Feldstein and Poterba's estimates, which most likely is due to differences in the sources of the variation in replacement rates in their and our data. Whereas Feldstein and Poterba use a sample that is representative of the United States and thus differences in replacement rates are driven to

[9] We include a dummy for observations where earnings were top-coded at $99,999 as our identification strategy relies on nonlinearities in the benefit schedule and thus may be sensitive to the methodology of imputing earnings above the top code. See Krueger and Mueller (2011) on how earnings above the top code were imputed.

Table 1—Regressions of the log of the Reservation Wage Ratio on UI Benefits, Savings, and Other Measures of Liquidity

	(1)	(2)	(3)	(4)
log(weekly benefit)	−0.047	0.015	−0.046	−0.031
	(0.059)	(0.067)	(0.060)	(0.060)
log(weekly previous wage)	−0.505	−0.559	−0.501	−0.522
	(0.047)***	(0.053)***	(0.048)***	(0.046)***
Received severance payment			−0.017	−0.020
			(0.018)	(0.018)
Severance payment amount (in \$)/10,000			0.009	0.006
			(0.004)**	(0.005)
Savings: \$10,000–\$24,999				0.014
				(0.023)
Savings: \$25,000–\$49,999				0.044
				(0.034)
Savings: \$50,000–\$99,999				−0.028
				(0.069)
Savings: \$100,000 or more				0.115
				(0.033)***
Access to \$5,000 in case of emergency				−0.005
				(0.015)
Access to at least one credit card				0.041
				(0.020)**
Spouse has job				0.018
				(0.026)
Incentive choice: \$40				0.010
				(0.018)
Degree of risk loving				0.011
(*0 = Unwilling to take risk; 10 = Very willing to take risks*)				(0.003)***
County unemployment rate	−0.003	−0.003	−0.001	−0.001
	(0.005)	(0.005)	(0.005)	(0.004)
Unemployment duration (weeks paid)	−0.002	−0.002	−0.002	−0.002
	(0.000)***	(0.000)***	(0.000)***	(0.000)***
Years of school	0.019	0.017	0.019	0.017
	(0.005)***	(0.005)***	(0.005)***	(0.005)***
Job tenure on previous job, in years	−0.003	−0.003	−0.004	−0.004
	(0.001)**	(0.001)*	(0.001)***	(0.001)***
Experience/10	0.110	0.133	0.109	0.114
	(0.024)***	(0.030)***	(0.024)***	(0.023)***
Experience^2/100	−0.023	−0.027	−0.023	−0.024
	(0.005)***	(0.007)***	(0.005)***	(0.005)***

(*continued*)

some extent by differences in statutory replacement rates and maximum benefits across states, replacement rates in our data from NJ are a nonlinear function of prior earnings.

As a further check on the specification, we estimated Probit models with a dummy equal to one if the person left UI within a month of the first interview and included the same explanatory variables as in columns 1 and 4 of Table 1. The coefficient on the log of the benefit rate was significant at the 5 percent level and the estimated elasticity ranged from −0.83 to −0.95. This is very close to the estimated elasticity of the duration of unemployment to the benefit level in Meyer (1990) and in the

Table 1—Regressions of the log of the Reservation Wage Ratio on UI Benefits, Savings, and Other Measures of Liquidity (*continued*)

	(1)	(2)	(3)	(4)
Female	−0.083	−0.079	−0.083	−0.086
	(0.016)***	(0.020)***	(0.017)***	(0.020)***
Married	−0.006	−0.019	−0.008	−0.031
	(0.020)	(0.021)	(0.020)	(0.026)
Number of children	−0.002	−0.007	−0.002	0.000
	(0.005)	(0.005)	(0.005)	(0.005)
Black	−0.021	−0.017	−0.021	−0.016
	(0.033)	(0.035)	(0.034)	(0.037)
Asian or other	0.017	−0.005	0.005	0.004
	(0.043)	(0.057)	(0.044)	(0.041)
Race not available	0.009	0.024	0.011	0.010
	(0.032)	(0.031)	(0.032)	(0.030)
Not Hispanic	0.026	0.022	0.031	0.033
	(0.026)	(0.029)	(0.026)	(0.025)
Ethnicity not available	−0.009	−0.021	−0.014	−0.009
	(0.027)	(0.025)	(0.026)	(0.025)
Dummies for industry and occupation (2-digit)	x	x	x	x
Dummy for imputed wage above top code	x	x	x	x
Dummies for unit of reported reservation wage	x	x	x	x
Observations	3,841	2,797	3,687	3,530
R^2	0.496	0.512	0.497	0.509

Notes: The weekly benefit includes the additional $25 ARRA payment. Survey weights are used. Universe: unemployed; no job offer yet accepted; age 20–65. The sample in column 2 only includes individuals with initial savings smaller than $10,000. Robust standard errors are in parentheses (clustered at the county level).

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

range of the estimates surveyed in Krueger and Meyer (2002), suggesting that our sample is not unusual, at least with respect to the observed relationship between UI exit and benefits.[10]

As explained above, the identification of the effect of UI benefits on reservation wages relies on the nonlinearity of the benefit schedule and thus, one may be worried that other factors that are correlated with the previous wage may lead to biases in the estimated coefficients. In particular, savings are likely to be positively correlated with both the previous wage and the reservation wage, which will impart a downward bias on the estimated coefficient on the benefit variable. When we exclude all individuals with savings of $10,000 or more in the regression reported in column 2, the coefficient on the log of the benefit rate becomes positive, but the point estimates remain small and insignificant. Moreover, our survey contains a rich set of controls for other measures of access to liquidity, and including them in the regressions reported in columns 3 and 4 does not meaningfully change the estimated

[10] Note that these estimates represent the microelasticity of UI and ignore potential general equilibrium effects; see Lalive, Landais, and Zweimüller (2013); Hagedorn et al. (2013); and Marinescu (2014) for efforts to estimate the demand-side as well as other general equilibrium effects of UI. One should also note that the microelasticity may be smaller in recessionary periods; see Kroft and Notowidigdo (2011) and Schmieder, von Wachter, and Bender (2012).

coefficient on the benefit replacement rate. These results are largely consistent with the literature on accepted wages, which usually finds small or insignificant effects (see the references in the introduction). While these studies rely on rigorous identification strategies using regression discontinuity designs for the subset of unemployed workers who found jobs, our cross-sectional analysis is a useful complement as it provides direct evidence on reservation wages for the full sample of unemployed workers.

The estimated effects of several of the covariates are of interest in their own right. The estimates in column 3 suggest that severance pay has a positive and significant relationship with the reservation wage. The regression in column 4 controls for additional variables, such as the amount of savings in the bank account, access to $5,000 in case of an emergency, access to at least one credit card and whether the worker's spouse has a job. In general, the sign on the coefficients of these variables tends to be in line with expectations, but the coefficients are not always significant. Notably, people with at least $100,000 in their checking or savings account have significantly higher reservation wages relative to previous pay, as do those with access to at least one credit card.[11]

The county-level unemployment rate appears to have little or no association with the reservation wage, suggesting that workers who are searching for jobs in more distressed areas do not take much account of local labor market conditions in setting their reservation wage.

Finally, we find no evidence that the reservation wage differs between individuals who chose the (delayed) $40 over the (immediate) $20 incentive pay, consistent with the findings of DellaVigna and Paserman (2005) who found that measures of impatience, such as smoking, bear no relationship to the reservation wage but instead tend to be associated with a lower intensity of job search. We do find, however, that the self-reported degree of risk aversion appears to have a strong and significant relationship with the reservation wage. Our regression results indicate that respondents who report themselves as the least willing to take risks compared with those who report themselves as the most willing to take risks (based on a linear zero to ten subjective scale) have an 11 percent lower reservation wage.[12] This is consistent with search theory, as risk averse workers would prefer to accept a job at a low wage than bear the risk inherent in remaining unemployed.

IV. Reservation Wages over the Spell of Unemployment

As workers exhaust their UI benefits and assets, the reservation wage is expected to decline. The longitudinal nature of our data permits a stronger test of this hypothesis, as it allows us to control for heterogeneity and sample selection biases potentially present in past analyses of cross-sectional data. We start by comparing our results to Feldstein and Poterba's (1984) cross-sectional findings. Table 2 reports the average ratio of the reservation wage to the pre-unemployment wage. The first two

[11] See Bloemen and Stancanelli (2001) for similar evidence on the effect of wealth on the self-reported reservation wage for a sample of 1,026 unemployed workers in the Netherlands.

[12] This question about risk aversion has been experimentally validated; see Dohmen et al. (2005).

Table 2—Reservation Wage Ratio by Duration of Unemployment

	All durations	Less than 5 weeks	5–9 weeks	10–14 weeks	15–19 weeks	20–24 weeks	25–49 weeks	50 + weeks
Feldstein and Poterba (1984): All job losers and leavers	1.07	1.11	1.09	1.04	1.06	1.04	1.02	0.99
Feldstein and Poterba (1984): Job losers	1.03	1.06	1.05	1.03	1.06	1.00	0.99	0.97
Krueger and Mueller: Cross-section (1st week)	0.99	1.04	1.02	1.01	1.00	1.06	0.95	0.94
Krueger and Mueller: Longitudinal estimate	0.99	1.00	1.00	1.00	0.99	0.99	0.98	0.97
Krueger and Mueller: Longitudinal estimate (full-time workers)	1.00	1.03	1.02	1.01	1.01	1.00	0.98	0.94

Notes: Survey weights are used. Universe: unemployed; no job offer yet accepted; age 20–65. Feldstein and Poterba's (1984) estimates are from a sample of 2,228 unemployed from the May 1976 Current Population Survey.

rows are directly taken from Feldstein and Poterba, who found that, on average, the reservation wage is slightly higher than the previous wage, and that the reservation wage is only slightly lower among workers with longer durations of unemployment. The second row, which shows their results for job losers, is probably more comparable to our sample of UI recipients.

The third row reports the average reservation wage ratio using just the entry week response to our survey to compare with the cross-sectional CPS data. The results are remarkably similar, although the economic environments were markedly different in the two time periods. (The national unemployment rate in May 1976 was 7.4 percent, down from 9.0 percent a year earlier, versus 9.6 percent in New Jersey in the fourth quarter of 2009, up from 6.5 percent a year earlier.) Across all durations, the reservation wage ratio is essentially equal to the previous wage, on average, in both samples.

In our cross-sectional data, the reservation wage ratio is 10 percentage points lower for workers with 50 or more weeks of unemployment than for those with less than 5 weeks of unemployment. The corresponding figure in Feldstein and Poterba is 9 percentage points.

The fourth row of Table 2 utilizes the longitudinal data. In particular, we regressed the log of the reservation wage ratio on unemployment duration and individual fixed effects, so we can examine how much the reservation wage falls as unemployment duration increases for a given set of job seekers. Specifically, we used the fixed effects estimates from column 3 of Table 3A and used the midpoint of each category to predict the log reservation wage ratio, and then exponentiated. The longitudinal estimates point to an even more gradual decline in the reservation wage with unemployment duration than that found by Feldstein and Poterba. The last row of Table 2 shows the same estimate, but for full-time workers only, with a slightly more pronounced decline than in the full sample.

The correlation in the reservation wage in adjacent weeks was 0.96, so at the individual level, the self-reported reservation wage was also relatively stable. (For reference, the correlation in earnings in a given period reported by individuals at different times is typically around 0.90; see Bound and Krueger 1991).

Table 3A—Regressions of log Weekly Reservation Wage Ratio on Duration of Unemployment, with and without Fixed Effects

Dependent variable: log(reservation wage ratio)	Week 1 (1)	Pooled cross section (2)	Fixed effects (3)	Fixed effects (4)
Unemployment duration, in weeks	−0.00193	−0.00125	−0.00056	−0.00046
	(0.00035)***	(0.00043)***	(0.00056)	(0.00080)
Lapse (before November 8)				0.01871
				(0.01423)
Exhausted UI				−0.01779
				(0.02794)
Weeks of UI left: 1–4				−0.00541
				(0.01333)
Weeks of UI left: 5–8				0.00772
				(0.01306)
Weeks of UI left: 9–12				0.00115
				(0.00884)
After extension of November 8				0.00056
				(0.00659)
Controls (see notes)	x	x		
Individual fixed effects			x	x
Reservation wage unit dummies	x	x	x	x
Mean of dependent variable	−0.08	−0.11	−0.10	−0.10
Observations	3,841	22,701	24,474	24,474
R^2	0.50	0.52	0.94	0.94

Notes: Survey weights are used. Universe: unemployed; no job offer yet accepted; age 20–65. The controls in columns 1 and 2 are the same as in column 1 of Table 1. Robust standard errors are in parentheses (clustered at individual level).

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

The top panel of Figure 4 shows the average reservation wage ratio by duration of unemployment for each of the sampled cohorts. The bottom panel shows the same graphs after removing individual means. Both of the reservation wage ratio graphs display little tendency for the reservation wage to decline over the spell of unemployment. A comparison of the cross-section to the longitudinal estimates suggests that, if anything, the cross-sectional estimates slightly overstate the decline in reservation wages over the duration of unemployment, contrary to the expectation that those with relatively low reservation wage ratios would return to work sooner than those with relatively high reservation wages, all else equal.

This conclusion is also borne out in the regression results presented in Table 3A, which regress the log reservation wage ratio on unemployment duration and various other variables. Columns 1 and 2 show a gradual decline in the reservation wage relative to the pre-unemployment wage over the spell of unemployment, and the fixed effects estimates in columns 3 and 4 indicate a statistically insignificant and trivially small change in the reservation wage as unemployment duration increases. Moreover, the reservation wage appears insensitive to periods of lapsed benefits before and after the point of exhaustion, and is unchanged after the November 8th extension of benefits from 79 to 99 weeks.

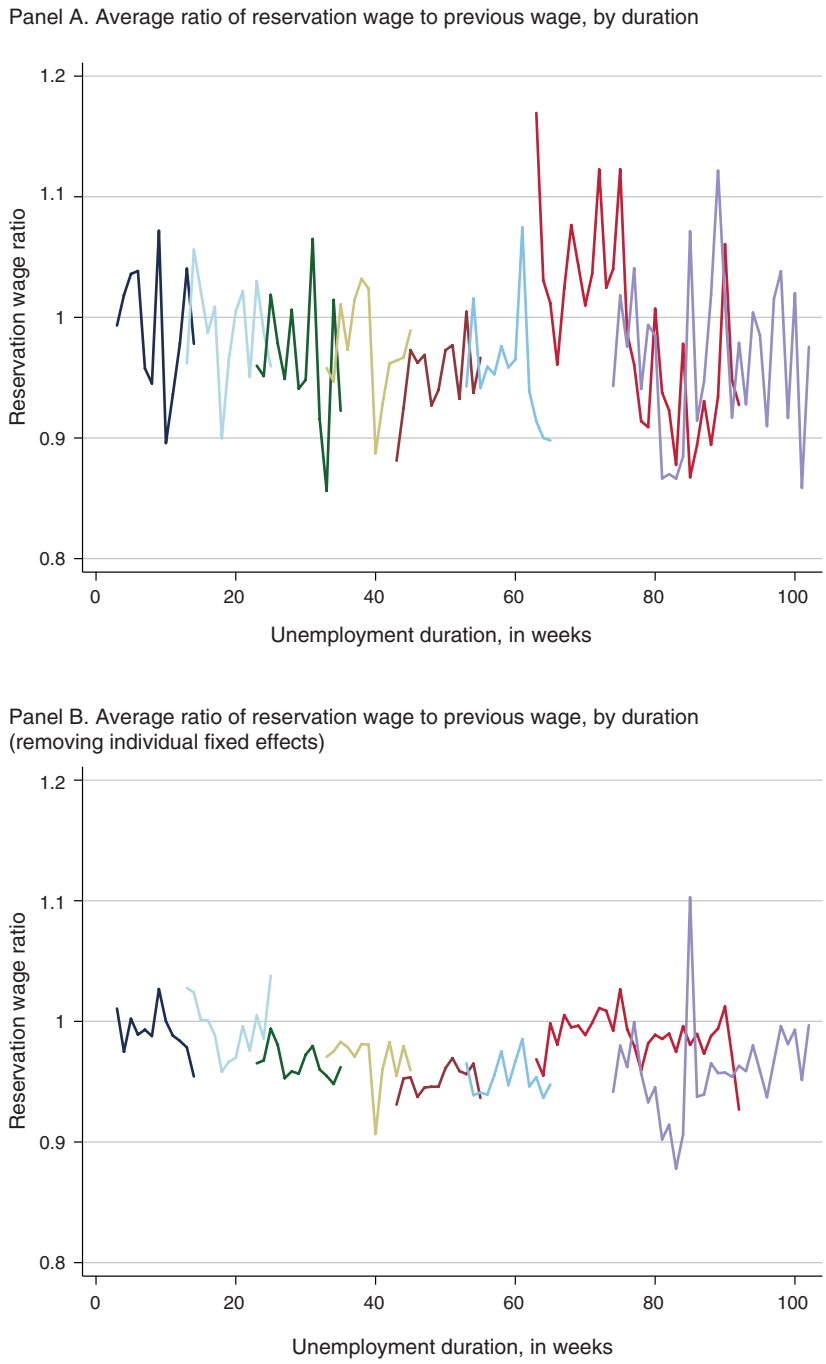


FIGURE 4

Notes: The eight lines show the average reservation wage ratios of the eight sampling strata. The sampling strata consisted of intervals of duration of unemployment prior to the start of the survey at the end of September 2009 (0–2 weeks, 10–12 weeks, 20–22 weeks, 30–32 weeks, 40–42 weeks, 50–53 weeks, 60–69 weeks, and 70 and more weeks). To remove individual fixed effects, we estimated a regression of the reservation wage ratio on dummies indicating duration of unemployment and individual fixed effects, and then plotted the estimated duration dummies normalized by the average fixed effect of each cohort.

Table 3B—Fixed-Effect Regressions of log Weekly Reservation Wage Ratio on Duration of Unemployment, for Different Subgroups

Dependent variable: log(reservation wage ratio)	Savings < $10,000 (1)	Savings > = $10,000 (2)	Age 20–50 (3)	Age 51–65 (4)	Savings > = $10,000 (age 20–50) (5)	Savings > = $10,000 (age 51–65) (6)
Unemployment duration, in weeks	0.00014	−0.00309	0.00045	−0.00260	−0.00182	−0.00423
	(0.00065)	(0.00114)***	(0.00072)	(0.00073)***	(0.00159)	(0.00159)***
Individual fixed effects	x	x	x	x	x	x
Reservation wage unit dummies	x	x	x	x	x	x
Mean of dependent variable	−0.06	−0.28	−0.05	−0.25	−0.23	−0.34
Observations	16,057	6,796	13,565	10,909	2,655	4,141
R^2	0.94	0.95	0.94	0.93	0.96	0.92

Notes: Survey weights are used. Universe: unemployed; no job offer yet accepted; age 20–65. Robust standard errors are in parentheses (clustered at individual level).

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

Table 3B provides additional estimates of the fixed effects specification for various subsamples. For those with $10,000 in savings or more at the start of the survey, or over age 50, we find a statistically significant, negative relationship between the reservation wage relative to previous pay and the duration of unemployment. For those who were both older than 50 *and* had more than $10,000 in savings at the start of the study, a spell of unemployment of 52 weeks is estimated to have reduced the reservation wage by 20 percent. The finding that the reservation wage is more sensitive to unemployment duration for those with some savings at the start is consistent with the idea that as workers draw on their assets during a spell of unemployment, they become less selective about which job they would accept. The apparent willingness of older workers to lower their reservation wage the longer they are unemployed is consistent with the view that job search is an investment: the cost of accepting a lower paying job is less for those who plan to spend less time in the labor market. Many older workers apparently *gradually* realize that they cannot find a job that pays as well as they expected and thus adjust their reservation wage downward, whereas younger workers are willing to maintain their reservation wage longer because it is more costly for them to accept a low paying job.

As a robustness check, we estimated the same regressions but for workers who indicated that they were looking for full-time work only. The results in Table 3C for this subsample are similar to those in Tables 3A and 3B, but show a statistically significant coefficient on unemployment duration, with the size of the coefficient nonetheless relatively small. The results in column 4 indicate that for full-time workers, the reservation wage declines by 7.3 percent over a period of a year. Finally, we also estimated a separate set of regressions where we excluded the 11 percent of unemployed workers who indicated that they expect to be recalled to their previous employer and the results in Tables 3A, 3B, and 3C were very similar.

It is not possible to separately identify the effect of calendar time and the effect of unemployment duration on the reservation wage as both are perfectly collinear once

Table 3C—Regressions of log Weekly Reservation Wage Ratio on Duration of Unemployment, with and without Fixed Effects (*full-time workers only*)

Dependent variable: log(reservation wage ratio)	Week 1 (1)	Pooled cross section (2)	Fixed effects (3)	Fixed effects (4)
Unemployment duration, in weeks	−0.00157	−0.00126	−0.00137	−0.00145
	(0.00036)***	(0.00047)***	(0.00060)**	(0.00070)**
Controls (see footnote)	x	x		x
Individual fixed effects			x	x
Reservation wage unit dummies	x	x	x	x
Mean of dependent variable	−0.07	−0.11	−0.10	−0.10
Observations	3,065	17,924	19,284	19,284
R^2	0.55	0.57	0.95	0.95

	Savings < \$10,000 (5)	Savings > = \$10,000 (6)	Savings > = \$10,000 (age 20–50) (7)	Savings > = \$10,000 (age 51–65) (8)
Unemployment duration, in weeks	−0.00074	−0.00321	−0.00193	−0.00404
	(0.00069)	(0.00148)**	(0.00224)	(0.00194)**
Individual fixed effects	x	x	x	x
Reservation wage unit dummies	x	x	x	x
Mean of dependent variable	−0.05	−0.29	−0.25	−0.34
Observations	12,700	5,329	2,129	3,200
R^2	0.95	0.95	0.96	0.92

Notes: Survey weights are used. Universe: unemployed; no job offer yet accepted; age 20–65. The controls in columns 1 and 2 are the same as in column 1 of Table 1. The controls in column 4 are the same as the ones shown in column 4 of Table 3A. The coefficient estimates of the controls in column 4 were of similar size as in Table 3A and all statistically insignificant. Robust standard errors are in parentheses (clustered at individual level).

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

we control for individual fixed effects. In particular, one may worry that seasonal factors or the state of the business cycle could confound duration effects. Several reasons, however, lead us to suspect that calendar time exerts little, if any, effect. First, in contrast to search intensity, the reservation wage is unlikely to be influenced by seasonal factors, such as holidays, as accepting a job at a low wage has a large opportunity cost in terms of foregone earnings over the entire duration of the job, whereas a lower search effort only *postpones* job finding. Our results are essentially unchanged if we drop observations in the weeks encompassing Thanksgiving and Christmas. Second, as mentioned, the unemployment rate in New Jersey was high but stable over the survey period: it remained in the narrow range of 9.6 percent to 9.7 percent over the entire survey period, and did not fall below 9.5 percent until December 2010. This suggests that unemployed individuals' job market prospects were stable at a low level over the period of the survey and thereafter. Moreover, when we added the county-level unemployment rate as an additional explanatory variable in the fixed effect regressions in Tables 3A and 3C, the size and (in)significance of the coefficient on duration was unaffected. The coefficient on the county unemployment rate was negative but insignificant. Finally, Elsby, Shin, and Solon

(2014) find that composition-adjusted real wages fell over the period from 2009 to 2012, which suggests that, if anything, one should expect reservation wages in New Jersey to have fallen by more than predicted by our model that assumed a stationary distribution of wage offers. Overall, these considerations suggest that temporal factors did not upwardly bias our estimate of the effect unemployment duration on reservation wages.

The relatively modest decline in reservation wages is difficult to reconcile with the calibrated search model, especially for those who lack personal savings.

Other Job Considerations.—After eliciting the reservation wage, the questionnaire asked, "How many minutes a day would you be willing to commute if you were offered a job at that salary?" Table 4 reports estimates of the same type of models as in Table 3, using commuting time as the dependent variable in place of the reservation wage. We find a statistically significant effect of unemployment duration on the willingness to accept a longer commute to work, but the relationship is modest. A 52 week increase in unemployment is associated with only a 4.6 minute increase (11 percent of the average) in the amount of time individuals said they would be willing to travel to work in our fixed effects models. Thus, as a practical matter, based on their responses, job seekers do not seem particularly willing to accept a job that requires a longer commute as their duration of unemployment increases.

The survey also asked the respondents an open-ended question on what type of job/occupation they were looking for.[13] The question was asked before the reservation wage question. We estimated the wage associated with each of the occupations based on the average log hourly wage for the occupation calculated from the outgoing rotation group data in the CPS for the years 2005–2009. We define the occupational reservation wage as the occupation paying the *lowest* wage among all those listed in the answer to the open ended question if more than one was listed. Table 5 shows the same type of regressions as in Table 3A but using the log of the occupational reservation wage as the dependent variable. The results in column 3 show a significant but quantitatively modest decline of the occupational reservation wage over the spell of unemployment. Over a 52 week period, the occupational reservation wage is predicted to decline by about 9 percent. This indicates that unemployed workers reduce their occupational aspirations over the spell of unemployment, but the size of the effect is only slightly higher than our estimate of the decline in the self-reported reservation wage.

Finally, an important question is whether the reservation wage falls around the time of UI benefit exhaustion. Column 4 in Tables 3 to 5 includes a dummy for having temporarily lapsed benefits in the period prior to the benefit extension of November 8, a dummy for having exhausted benefits (the full 99 weeks) and dummies for 1–4 weeks, 5–8 weeks, and 9–12 weeks prior to exhaustion. The estimated coefficients

[13] The exact wording was "Please describe in a few words the type of work you are looking for (for example: Electrical engineer, stock clerk, typist, farmer, …)." The coding of this question into three-digit occupational codes was performed by trained coders at the University of Wisconsin Survey Center, who have extensive experience with occupational coding based on work with the Wisconsin Longitudinal Survey. There were 35,166 records and 15,506 unique job titles associated with this item. The survey center produced occupational codes first according to the CEN2000 schemes and then converted them into the corresponding SOC codes.

Table 4—Regressions of Reservation Commuting Distance on Duration of Unemployment, with and without Fixed Effects

Dependent variable: Distance willing to commute, in min per day (for job paying the reservation wage)	Week 1 (1)	Pooled cross section (2)	Fixed effects (3)	Fixed effects (4)
Unemployment duration, in weeks	−0.022	0.008	0.089	0.111
	(0.029)	(0.025)	(0.039)**	(0.049)**
Lapse (before November 8)				1.166
				(1.642)
Exhausted UI				−1.216
				(1.851)
Weeks of UI left: 1–4				−0.796
				(1.246)
Weeks of UI left: 5–8				0.046
				(1.158)
Weeks of UI left: 9–12				−1.512
				(0.898)*
After extension of November 8				−0.012
				(0.436)
Controls (as in column 1 of Table 1)	x	x		
Individual fixed effects			x	x
Mean of dependent variable	41.5	40.7	40.8	40.8
Observations	4,068	24,304	26,995	26,995
R^2	0.21	0.27	0.90	0.90

Notes: Survey weights are used. Universe: unemployed; no job offer yet accepted; age 20–65. Robust standard errors are in parentheses (clustered at the individual level).

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

are small and insignificant, showing that there is no acceleration in the decline of reservation wages before the UI exhaustion point. Figure 5 illustrates this point as well by displaying that the reservation wage ratio in the weeks prior to and after exhaustion, after removing individual fixed effects. The graphs show that the ratio is stable in the weeks before exhaustion, in contrast to the predictions of our calibrated model in Section I, which shows a steeper drop in the reservation wage in the weeks prior to the exhaustion of UI benefits.

For each individual who reported a reservation wage on three or more occasions, we computed the unemployment duration-reservation wage gradient. Figure 6 presents a histogram of these individual-specific slopes. The slopes have a large mass at zero, and a mean of −0.06 percent per week, very close to the fixed effects estimate for the full sample in Table 3A. Only around one-third of the slopes are less than −0.23 percent per week, the benchmark from our calibration exercise. The standard deviation is sizable (3.8 percent), although it shrinks by about a quarter if we adjust for the fact that each slope is measured with sampling error. To make this adjustment, we subtracted off the average sampling variance of the individual slopes from the variance across the slopes. Overall, these results point to considerable heterogeneity, and a sizable proportion of individuals who are reluctant to reduce their reservation wage despite extended spells of unemployment.

Table 5—Regressions of the Occupational Reservation Wage on Duration of Unemployment, with and without Fixed Effects

Dependent variable: log(occupational reservation wage)	Week 1 (1)	Pooled cross section (2)	Fixed effects (3)	Fixed effects (4)
Unemployment duration, in weeks	−0.00117	−0.00015	−0.00180	−0.00232
	(0.00033)***	(0.00036)	(0.00094)*	(0.00152)
Lapse (before November 8)				−0.00700
				(0.01837)
Exhausted UI				0.03337
				(0.03232)
Weeks of UI left: 1–4				0.02225
				(0.03005)
Weeks of UI left: 5–8				0.00577
				(0.02034)
Weeks of UI left: 9–12				0.01021
				(0.01885)
After extension of November 8				0.00266
				(0.00921)
Controls (as in column 1 of Table 1)	x	x		
Individual fixed effects			x	x
Dummies for unit of reported reservation wage	x	x	x	x
Mean of dependent variable	2.97	2.97	2.97	2.97
Observations	3,715	21,856	24,272	24,272
R^2	0.47	0.43	0.84	0.84

Notes: Survey weights are used. Universe: unemployed; no job offer yet accepted; age 20–65. Robust standard errors are in parentheses (clustered at individual level).

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

A. *Sources of Support*

Section I discussed additional theories on why reservation wages decline over the spell of unemployment, beyond the decline implied by the limited duration of UI benefits. This subsection provides some additional evidence on these different sources and briefly reviews relevant evidence on search intensity.

Liquidity over the Spell of Unemployment.—Our findings raise the question of why reservation wages do not fall more steeply over the spell of unemployment, given that they seem to fall strongly for those with high initial savings. It is important to recognize, however, that relatively few workers in our sample entered unemployment with a significant amount of savings. Among those with less than three months of unemployment duration, 86 percent reported having less than $10,000 of savings, and 57 percent indicated that they had no savings at all. The survey also collected information on whether unemployed workers could raise $5,000 in the following week in the event of an emergency, and only 23 percent of those unemployed for less than three months responded affirmatively. This is consistent with Lusardi, Schneider, and Tufanon (2011) who find in a survey fielded in 2009 that 50 percent of Americans could either certainly or probably come up with $2,000 in 30 days, compared to 31 percent of unemployed individuals.

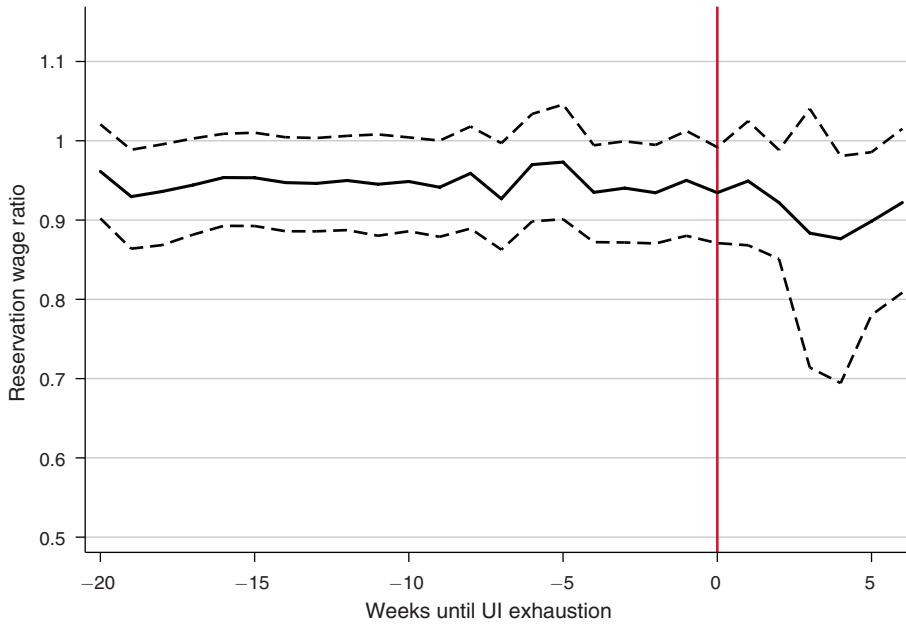


Figure 5. Average Ratio of Reservation Wage to Previous Wage, before and after UI Exhaustion (*removing individual fixed effects*)

Notes: To remove individual fixed effects, we estimated a regression of the reservation wage ratio on dummies indicating weeks until UI exhaustion and then plotted the estimated dummies normalized by the average fixed effect of those who have 20 weeks of UI left or less. The dashed lines show the 95 percent confidence interval, which was estimated by a bootstrap procedure with 5,000 draws. Note that the bootstrapped confidence intervals are not completely symmetric for weeks with small sample sizes. Universe: unemployed; no job offer yet accepted; age 20–65.

The question of access to $5,000 emergency funds was collected on a weekly basis, so we can examine how access to liquidity evolved over the spell of unemployment, controlling for individual effects. The results in column 1 of Table 6 show that the regression coefficient on duration of unemployment is significant at the 5 percent level, but the size of the coefficient is relatively small: an additional 25 weeks of unemployment is associated with a 5 percentage point increase in the likelihood of being liquidity constrained according to this measure. This decrease is mostly driven by the decline in savings and the selling of stocks as can be seen in the regressions in columns 2–6 of Table 6. Overall, these results suggest that financial circumstances only change modestly over the spell of unemployment as many unemployed workers are already financially constrained at the beginning of their unemployment spell. At the same time, these results heighten the question of why unemployed workers do not adjust their reservation wage more strongly, as most have no or only little access to savings to use to smooth consumption after exhausting UI benefits.

The Relationship between the Reservation Wage and Job Search Intensity.—The model presented in Section I does not include endogenous search effort, but it is a straightforward prediction that in such a model unemployed workers would increase their search effort as they approach the UI exhaustion point (see Mortensen 1977). In Krueger and Mueller (2011), however, we found with the same survey data that time

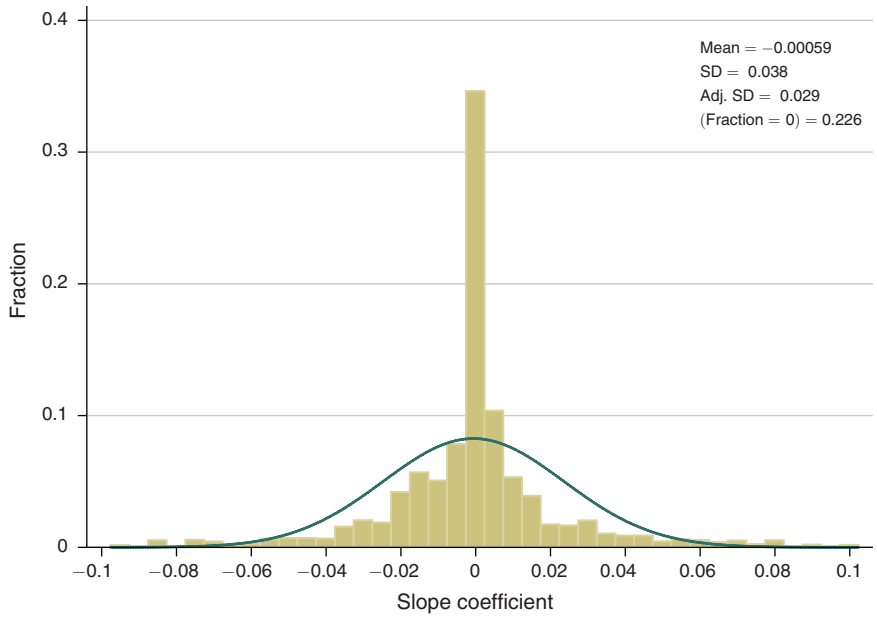


FIGURE 6. HISTOGRAM OF THE INDIVIDUAL SLOPE COEFFICIENTS OF THE LOG RESERVATION WAGE OVER THE SPELL OF UNEMPLOYMENT

Notes: We restrict the sample to those individuals with 3 observations or more of the reservation wage. The solid line shows the normal density with the mean and (unadjusted) standard deviation of the individual slope coefficients. To adjust for sampling variability, we also report an adjusted standard deviation, where we subtract from the variance the average of the squared standard errors of the individual slope coefficients, and then take the square root. The individual slopes were derived from a regression of the log reservation wage on the duration of unemployment for each individual. The sample size is 2,690.

TABLE 6—FIXED-EFFECT REGRESSIONS OF A DUMMY FOR ACCESS TO $5,000 ON DURATION OF UNEMPLOYMENT

Dependent variable:		Access to $5,000 through:				
	Access to $5,000	Savings	Selling stocks	Borrowing from friends or family	Borrowing (bank, credit card, home equity loan)	Other
	(1)	(2)	(3)	(4)	(5)	(6)
Unemployment duration, in weeks	−0.00188	−0.00107	−0.00031	−0.00029	0.00007	−0.00026
	(0.00074)**	(0.00051)**	(0.00015)**	(0.00062)	(0.00037)	(0.00028)
Controlling for individual f.e.	x	x	x	x	x	x
Mean of dependent variable	0.294	0.169	0.015	0.049	0.044	0.015
Observations	27,080	27,080	27,080	27,080	27,080	27,080
R^2	0.94	0.93	0.82	0.83	0.83	0.80

Notes: Survey weights are used. Universe: unemployed; no job offer yet accepted; age 20–65. Robust standard errors are in parentheses (clustered at individual level).

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

spent on job search activities declines over the spell of unemployment. The decline in search effort could be the result of discouragement in the search process or negative duration dependence in the wage offer distribution or in the likelihood of receiving a job offer. In contrast to Mortensen's model, these model alterations would also predict a positive correlation between search effort and the reservation wage, as both search effort and the reservation wage decline with duration of unemployment.

To further assess whether search effort has bearing on how reservation wages adjust to duration of unemployment, we use the survey data to test whether the reservation wage and search effort are correlated at the individual level. To be more specific, we estimated a fixed effect regression as in column 3 of Tables 3A and 3C, but included the amount of time spent on job search activities in the previous week as an additional explanatory variable (see Krueger and Mueller 2011 for details on how job search intensity is measured). We find a positive effect of search intensity that is statistically significant at the 5 percent level (Table 3A) and 1 percent level (Table 3C). The magnitude of the effect implies that 10 hours of additional search activity every week is associated with about a 0.5 percent increase in the reservation wage. This appears to be a relatively modest effect, but one should bear in mind that the correlation is the result of multiple forces, which go in opposite directions and potentially cancel each other out. Overall, these findings suggest that either job search discouragement, or negative duration dependence in the wage offer distribution or in the likelihood of receiving an offer, can explain part of the modest decline in reservation wages over the spell of unemployment. These empirical results thus exacerbate the tension with the predictions of our calibrated model, as the model abstracts from forces that lead to declining search effort over the spell of unemployment and assumes a constant job offer arrival rate throughout the spell of unemployment.

V. Validating Reservation Wages

For each week, the survey contains information on whether individuals received a job offer in the previous week, the number of job offers received, the wage offered, and whether the job was accepted or rejected. In the case of multiple job offers, the survey collected the wage of the best offer in the previous week. Among our sample of those aged 20–65, we have information on the wages of 1,499 job offers. Some 61.6 percent of job offers were accepted, 16.6 percent were rejected, and in the remaining 21.8 percent of cases the respondents were undecided. The survey did not collect information in subsequent interviews about the acceptance of earlier offers for which the respondent was undecided, but fortunately, we have access to administrative data and can test the extent to which the rate of UI exit in the weeks following the interview was similar among those who indicated that they accepted a job offer and those who indicated that they were undecided. We find that among those who accepted a full-time offer, 43 percent exited UI within one month and remained off the program, compared to 3 percent for those who rejected the offer and 19 percent for those who were undecided.[14] Thus, the undecideds are an intermediate group.

[14] We focus here on full-time offers because unemployed workers in NJ may work part-time and still receive UI benefits. There are two main reasons why an accepted full-time offer does not lead to early UI exit. First, it may

Past studies have not been able to assess the validity of self-reported reservation wage data because they only had access to cross-sectional data, with the exception of Holzer (1986), who provides some indirect evidence by showing that reservation wages in a given year are related to the subsequent duration of nonemployment and wages in the NLS-Y. Tables 7A and 7B present evidence of the likelihood that a worker accepts a job offer, depending on whether the offered wage was above or below the reservation wage. To avoid possible reporting bias due to cognitive dissonance, the reservation wage is taken from the most recent *prior* survey and is therefore not contemporaneous with the report on whether the job offer was accepted or rejected. The reservation wage and the offered wage, which could be reported in different time units (i.e., hourly wage versus weekly salary), were both converted into hourly earnings, which introduces some noise because the hours data entail some measurement error.

For the full sample, workers are 24 percentage points more likely to accept an offer that equals or exceeds the reservation wage than one that is below it, and 13 percentage points more likely to reject an offer that is below the reservation wage than one that equals or exceeds it. If we restrict the sample to offers of full-time jobs, the predictive power is somewhat stronger: workers are 30 percentage points more likely to accept offers that equal or exceed the reservation wage than one that is below it. The fact that 44 percent of the respondents accepted a job that paid less than the stated reservation wage suggests that the reservation wage is not a perfect measure, but it should be borne in mind that there is some noise in both the reservation wage and the offered wage because of measurement errors in the hours data. Moreover, nonwage job characteristics such as commuting distance or fringe benefits are also an important factor in the decision to accept an offer. If these characteristics are not perfectly reflected in the offered wage, we will observe acceptance of some offers that are below the reservation wage and rejection of some offers that exceed the reservation wage. Table 8 provides some further insight, by reporting the reason for rejecting a job offer. As would be expected, inadequate pay/benefits is an important factor in the decision to reject offers below the reservation wage, whereas unsuitable working conditions and transportation problems are more important in the decision to reject offers paying more than the reservation wage, though the sample size is too small to draw strong inferences on the relative importance of nonwage job characteristics.

Despite measurement error and nonwage amenities, Tables 7A and 7B suggest that the reservation wage has predictive power for the likelihood of accepting an offer. To further assess the predictive power of reservation wages, Figure 7 plots the probability of accepting a job (from a locally weighted regression) against the ratio of the job offer to the lagged value of the reservation wage. The figure shows that the likelihood of accepting a job offer increases with the ratio, but not monotonically so. There seems to be a jump when the offered wage equals the reservation wage and a small dip thereafter. The reason for the jump is that there is a large mass point of job offers that pay exactly the reservation wage (around 13 percent of job

lead to UI exit but after one month. We focus on one month because otherwise, the analysis is confounded by the presence of other offers. Second, we measure early UI exit by the most recent date of UI payment (as of June 12, 2010) and thus, potentially miss intermittent spells of employment.

Table 7A—Hourly Offered Wage below and above Hourly Reservation Wage

	Hourly offered wage < hourly reservation wage	Hourly offered wage >= hourly reservation wage
Accepted	50.5%	74.1%
Not accepted	23.6%	10.3%
Undecided	25.9%	15.7%
Observations	566	587

Note: Survey weights are used.

Table 7B—Hourly Offered Wage below and above Hourly Reservation Wage (*full time offers only*)

	Hourly offered wage < hourly reservation wage	Hourly offered wage >= hourly reservation wage
Accepted	44.4%	73.8%
Not accepted	24.2%	11.4%
Undecided	31.4%	14.8%
Observations	361	417

Note: Survey weights are used.

Table 8—Reasons for Not Accepting Offer, by Wage Offer below and above Reservation Wage

Reason for not accepting offer	Hourly offered wage < hourly reservation wage	Hourly offered wage >= hourly reservation wage	All
Inadequate pay/benefits	40%	1%	30%
Unsuitable working conditions	1%	28%	9%
Would not make use of my experience or skills	14%	4%	11%
Had insufficient experience or skills	1%	4%	2%
Insufficient hours/too many hours	7%	1%	6%
Changed plans	3%	1%	3%
Transportation problem	12%	19%	14%
Better offer	3%	10%	5%
Other	18%	31%	21%
Observations	121	42	163

offers; see Figure 8). When we exclude observations where the offered wage equals the reservation wage, the acceptance probability monotonically increases with the ratio (see the dashed line on Figure 7).

There are several potential explanations for the mass point of offers at the reservation wage, some of which we can rule out. First, unemployed workers may have had some information about a potential wage offer in the previous interview when they reported their reservation wage. Yet, we find that even for those offers where the most recent reservation wage was reported more than four weeks earlier, the percentage of wage offers equaling the reservation wage was 11 percent. Second, some unemployed workers expect to be recalled to their previous employer, and thus may set the reservation wage equal to the wage at their previous job. But 11 percent of those not expecting to be recalled reported a reservation wage exactly equal to the

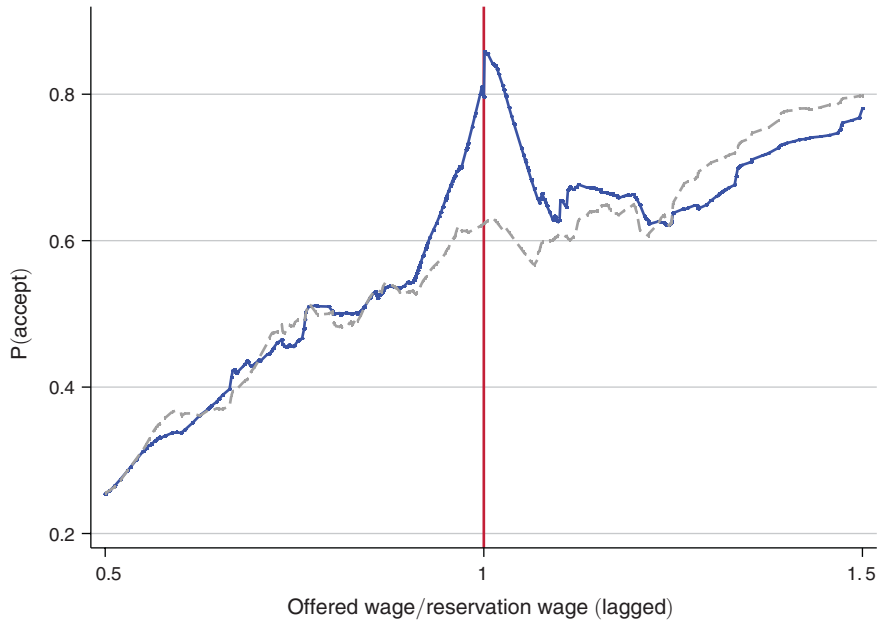


FIGURE 7. LOWESS OF ACCEPTING A JOB OFFER ON THE RATIO OF THE OFFERED WAGE TO THE RESERVATION WAGE

Notes: This graph shows the predicted value of a locally weighted regression (LOWESS) with bandwidth parameter 0.3 for a dummy of whether the job offer was accepted on the ratio of the hourly offered wage to the hourly reservation wage from the previous interview. The dashed line excludes observations where the offered wage equals the reservation wage. Sample: full-time offers, survey respondents aged 20–65.

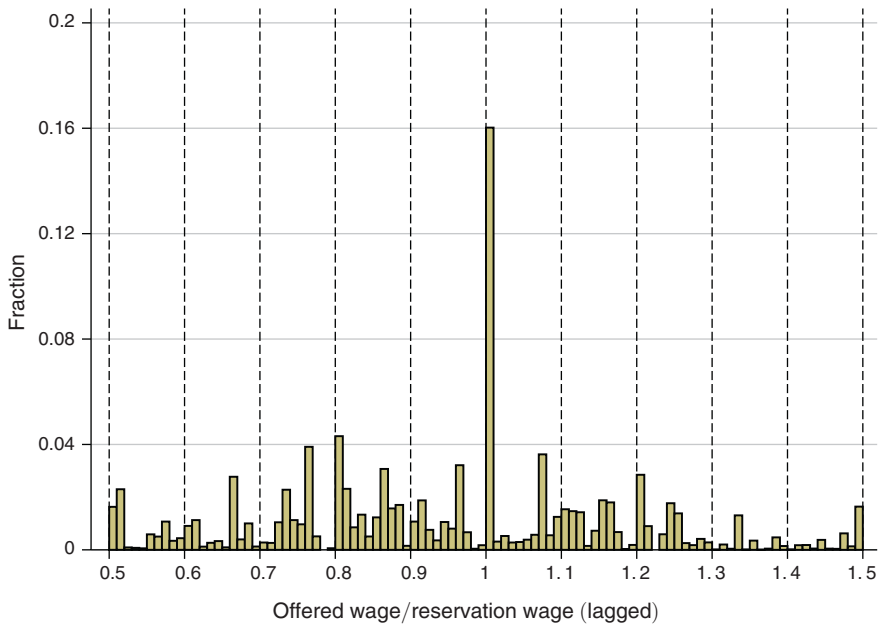


FIGURE 8. HISTOGRAM OF THE RATIO OF THE OFFERED WAGE TO THE RESERVATION WAGE

offered wage, so this cannot be the explanation either. Another possible explanation is that some unemployed workers search in labor markets with an extremely tight dispersion of potential wages. Lastly, it is possible that employers have some information regarding the unemployed worker's reservation wage, which would have important consequences for the unemployed worker's incentives to search. If a fraction of employers know and thus offer the reservation wage, unemployed workers have less of an incentive to search as these employers capture the entire surplus. In the limit, when all firms offer workers their reservation wage, unemployed workers have no incentive to search, the famous paradox raised by Diamond (1971).

Table 9 presents probit models where the dependent variable is 1 if an offer has been accepted and 0 if it has been rejected or a decision has not yet been made.[15] The first column indicates that workers are more likely to accept jobs that offer higher wages, less likely to accept a job if their reservation wage is higher, and more likely to accept part-time jobs. These regression results are in line with Figure 7 and show that the likelihood of accepting a job offer is driven by both variation in the offered wage and variation in the reservation wage. Moreover, the coefficient on the reservation wage is of similar magnitude to the coefficient on the offered wage, which suggests that the contribution of the reservation wage to the upward sloping curve in Figure 7 is substantial and almost as large as for the offered wage. Column 2 introduces a dummy variable that equals one if the hourly offered wage exceeds the hourly reservation wage, and zero otherwise. The results show that this threshold has a significant impact on the likelihood of accepting a job offer, although the interaction between part-time job status and being above or below the reservation wage indicates that the reservation wage cutoff does not predict an acceptance of part-time jobs. The value of the reservation wage itself does not predict whether a job is accepted when a binary indicator of whether the reservation wage exceeds the offered wage is included in column 2. The value of the offered wage, however, has predictive power conditional on whether the wage exceeds the reservation wage. One interpretation of the latter finding is that jobs that offer a higher wage can compensate for undesirable conditions that are not adequately reflected in the reservation wage (e.g., commuting distance), and thus, whether the reservation wage exceeds the offered wage is not a sufficient statistic for job acceptance or rejection. Nonetheless, the binary indicator of whether the reservation wage exceeds the offered wage is a strong predictor of job acceptance or rejection.

Recall that, on average, the reservation wage is close to the previous wage. If we use the previous wage in place of the reservation wage, however, the probit equation does not predict job acceptance as well as if we use the reported reservation wage (compare columns 2 and 3 of Table 9). Moreover, if we include a pair of indicators for whether the reservation wage and the previous wage exceed the offered wage in the same equation, the reservation wage indicator has the expected sign and is statistically significant at the 1 precent level, while the coefficient on the previous wage

[15] Note that we coded offers where job seekers were undecided as zeros because we estimated the fraction of rejections to be substantial for undecided cases as opposed to cases where the job seeker reported his decision. The probit estimates were qualitatively unchanged if we excluded observations in which the respondent was undecided as to whether to accept or reject the job offer; these results are available on request.

Table 9—Marginal Effects from Probit Model Regressions of Accepting a Job Offer

	(1)	(2)	(3)	(4)	(5)	(6)
Hourly offered wage >= Hourly reservation wage (lagged)		0.212		0.196	0.218	0.217
		(0.098)**		(0.061)***	(0.091)**	(0.089)**
Hourly offered wage >= Hourly reservation wage (lagged × Part-time job offer		−0.215			−0.222	−0.244
		(0.127)*			(0.123)*	(0.123)**
Hourly offered wage >= Hourly previous wage			0.152	0.106		
			(0.101)	(0.070)		
Hourly offered wage >= Hourly previous wage × Part-time job offer			−0.055			
			(0.126)			
Part-time job offer	0.127	0.210	0.152		0.212	0.194
	(0.064)**	(0.079)***	(0.079)*		(0.077)***	(0.076)**
log(lagged hourly reservation wage)	−0.278	−0.109			−0.109	−0.136
	(0.101)***	(0.122)			(0.126)	(0.146)
log(hourly offered wage)	0.367	0.216	0.185		0.227	0.194
	(0.084)***	(0.108)**	(0.091)**		(0.107)**	(0.112)*
log(hourly previous wage)			−0.029			−0.039
			(0.098)			(0.084)
Savings < $10,000					0.015	0.001
					(0.089)	(0.085)
Unemployment duration: 7–12 months					0.078	0.071
					(0.082)	(0.080)
Unemployment duration: 13–18 months					0.015	0.028
					(0.079)	(0.078)
Unemployment duration: 19 months or more					−0.063	−0.011
					(0.148)	(0.116)
Demographic controls						x
Mean of dependent variable	0.614	0.614	0.614	0.614	0.614	0.611
Observations	1,153	1,153	1,153	1,153	1,153	1,150
Pseudo R^2	0.054	0.071	0.051	0.052	0.076	0.120

Notes: Survey weights are used. Sample: respondents, age 20–65. The regression in column 6 includes additional controls for age, education, gender, marital status, race, and ethnicity (as in Table 11). Robust standard errors are in parentheses (clustered at individual level).

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

indicator is about half as large and insignificant with a t-ratio of 1.5 (see column 4). These results suggest that the reservation wage captures more information than the previous wage. Furthermore, column 5 indicates that savings and unemployment duration do not predict job acceptance conditional on the reservation wage. Because the reservation wage should be a sufficient statistic for workers' requirements, factors that affect their financial predicament or duration of job search should be irrelevant conditional on the reservation wage. The probit estimates are consistent with this prediction. Finally, the results in column 6 show that our results remain unaffected if we control for demographic characteristics, such as age, education, gender, marital status, race and ethnicity, as well as the log of the hourly previous wage.

In Table 10, we provide additional evidence on the validity of the reservation wage using administrative data on unemployment duration. More precisely, we take the first available report of the reservation wage ratio for each individual in our sample, and then estimate a probit model of the likelihood of exiting UI before UI exhaustion before mid-March 2010 (what we call "early" UI exit) using the initial report of the reservation wage ratio as a predictor variable.[16] The results in columns 1 and 2 of Table 10 show that a 1 percent increase in the reservation wage ratio is associated with a 0.5 to 0.6 percentage point lower probability of exiting UI early (or an elasticity at the average exit rate of about one third). This is consistent with the expectation that a higher reservation wage reduces the job finding rate and thus increases unemployment duration.

To probe whether these results are consistent with the regression results from the acceptance model in Table 9, it is useful to write the per-period job finding probability, P, as the product of the job offer rate λ, and the acceptance probability, A, as in the model in Section I. Note that this implies that the job offer probability is equal to the ratio of P to A. Assuming that the reservation wage (R) does not affect the offer probability, the response of the probability of finding a job to an increase in the reservation wage is thus equal to:

$$\frac{dP}{dln(R)} = \frac{P}{A}\frac{dA}{dln(R)},$$

which implies that the response of the acceptance probability should be the same as the response of the probability of finding a job, scaled by the ratio of P to A. In our data, the average rate of early UI exit is 0.173, whereas the average acceptance rate is 0.725 (0.616 plus 50 percent of the undecideds). Using the estimate of $\frac{dA}{dln(R)} = -0.278$ from column 1 of Table 9, the right-hand side of the equation above is -0.066, which is close to the estimates of -0.054 in column 1 of Table 10.[17]

One may also wonder at this point whether the results reported in Table 10 are consistent with the results reported in Tables 3A and 3C, where we found a slightly steeper decline of the reservation wage in the cross-sectional data than in the longitudinal data, where we controlled for individual fixed effects. Based on our results in Table 10, one would expect that individuals with low reservation wage ratios would find jobs more quickly and, therefore, estimates of the decline in the reservation wage in cross-sectional data should be biased upward, not downward. Note, however, that in a nonstationary environment, individuals laid off at different times over the business cycle may differ in their characteristics and reservation wage ratios. Indeed, in Krueger and Mueller (2011), we documented non-negligible "cohort effects" in terms of observable characteristics, such as pre-displacement

[16] We did not estimate a model with time until finding a job, as a large cohort of individuals exhausted UI at the end of March 2010 and thus had censored spells (i.e., all those who received an additional 20 weeks of UI benefits due to the extension of November 8, 2009, from 79 to 99 weeks of UI benefits).

[17] In regressions not reported here, we also included the individual slope of the reservation wage over the spell of unemployment from the start of the survey to November 30, 2009. The results indicate a negative effect of the individual slope of the reservation wage ratio on early UI exit, though the coefficient is significant only in the specification including controls in column 2 (with a t-ratio of 1.81).

Table 10—Probit Regressions (*marginal effects*): Leaving UI Early Related to the Reservation Wage Ratio

Dependent variable:	Leaving UI early (between first interview and March 14, 2010)	
	(1)	(2)
log(reservation wage ratio)	−0.054	−0.057
in first interview	(0.026)**	(0.027)**
Implied elasticity	[−0.312]	[−0.331]
Controls		x
Mean of dependent variable	0.173	0.172
Observations	4,356	4,348
Pseudo R^2	0.003	0.046

Notes: Survey weights are used. Sample: unemployed; no job offer yet accepted; age 20–65. The regression reported in column 2 includes additional controls for age, education, gender, marital status, race, ethnicity, and dummies for duration of unemployment at the start of the survey period (as in Table 11). The sample size is slightly lower for this sample because of missing information for some of the demographic controls. The dependent variable is a dummy variable, equal to 1 if the respondent left UI early, which is defined as exiting UI before March 14, 2010, without exhausting benefits and not starting a UI claim within the next 31 days of UI, and zero otherwise. We did not count as early UI exit respondents who left UI in weeks 51 or 52 of their spell, as in the majority of these cases UI benefits ended due to a test for continued eligibility for extended benefits (see Figure 2 in Krueger and Mueller 2011, for details). Robust standard errors are in parentheses.

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

industry and wages across the cohorts in our sampling frame. Moreover, Mueller (2012) documents that the composition of inflows in terms of demographics, industries and pre-displacement wages varies considerably over the US business cycle. It is thus possible that such cohort-heterogeneity may lead to a steeper decline in the reservation wage in the cross-section compared to the longitudinal estimates. Perhaps more importantly, we performed simulations based on our estimates in Table 10, and found that the upward bias in the cross-sectional estimate of Tables 3A and 3C should be very small, only around 0.04 percent in the reservation wage per week of unemployment, so changes in sample composition due to endogenous job finding are expected to exert a small effect on the decline of reservation wages in the cross-section. Intuitively, randomness in the job offer arrival rate and wage offer help to preserve a reasonably representative sample, despite the fact that workers with a lower reservation wage are more likely to accept a job and exit the sample over time.

Despite the predictive power of the reservation wage, close to half of those who were offered a job paying less than their reservation wage accepted it. As mentioned, these cases could represent measurement errors (e.g., because of errors in work hours), or weaknesses in the reservation wage data—or in the reservation wage concept itself. In Table 11, we estimate probit equations to examine factors that predict the likelihood of accepting a job that pays less than the reservation wage for the subset of searchers who were offered such jobs. The very long-term unemployed (more than 18 months) were more likely to accept jobs that paid below their stated reservation wage, although duration of unemployment did not matter for those who were unemployed 18 months or less. In addition, those who rated themselves as very

willing to take risks were less likely to accept jobs that paid below the reservation wage, but those who took the immediate $20 incentive payment over the delayed $40 payment were not more likely to accept such job offers. Lastly, less educated workers and African Americans were less likely to accept jobs offering less than their reservation wage. These results suggest that the mispredictions of the reported reservation wage are not random, although there is little evidence of systematic bias by unemployment duration for workers who are unemployed 18 months or less.

To further assess the magnitude of any bias in the reservation wage for the very long-term unemployed, we estimated a linear regression model of the difference between the accepted wage and the lagged reservation wage, including the same controls as above. Columns 3 and 4 of Table 11 report the results. Surprisingly, the log difference between the accepted and reservation wage is not significantly lower for the long-term unemployed, despite their higher likelihood of accepting a job below the reservation wage. The reason for this finding is that, for those unemployed more than 18 months, the likelihood of rejecting a wage offer above the reservation wage is also higher. Indeed, when we estimate the same regression as in columns 1 and 2, but for rejecting a job offer equal to or greater than the lagged reservation wage, we find that this likelihood is around 17 percent higher for those with more than 18 months of unemployment. We interpret this as evidence that, for the very long-term unemployed, the reservation wage is not necessarily a biased indicator, but rather that it is a noisier signal of the true underlying reservation wage.

VI. Conclusion

This paper presents new evidence on reservation wages over the spell of unemployment, using longitudinal data from weekly interviews of unemployed workers. Our findings suggest that reservation wages decline at a modest rate over the spell of unemployment, and that the decline is driven by older individuals and those with non-negligible savings. In particular, those with more than $10,000 in liquid assets at the start of the study were more likely to lower their reservation wage over the course of their spell of unemployment, as compared to those with less savings to start with, for whom the decline is not distinguishable from zero. A possible explanation for this finding is that people treat time-limited government social insurance benefits differently than their own personal savings, though we cannot rule out that differences in the responsiveness to UI exhaustion between those with and without liquid assets are responsible for the pattern that we observe. In any event, in contrast to the calibrated search model, which provides a benchmark for our estimates, we find that on average reservation wages start out too high and decline too slowly.

Indeed, our results suggest that unemployment insurance only has a limited impact on reservation wages. The calibrated model presented in this paper suggests that the decline of reservation wages over the 99 weeks of UI eligibility is roughly equal to the response of reservation wages to a permanent reduction of the UI replacement rate from 0.54 (the average in our sample) to a 0 replacement rate. Our estimates imply a decline of the reservation wage of between 2.4 and 7.3 percent over a period of a year, and thus suggest, at best, a modest effect of UI benefits on reservation wages. A small response of reservation wages to UI benefits is also consistent with

Table 11—Determinants of the Potential Bias in the Reported Reservation Wage

Dependent variable:	Probability of accepting job offer		log(accepted wage) - log(lagged reservation wage)	
Sample:	Offers below the reservation wage		Accepted offers	
Unemployment duration: 7–12 months	0.085	0.033	0.007	0.007
	(0.112)	(0.106)	(0.057)	(0.055)
Unemployment duration: 13–18 months	−0.002	−0.033	0.038	0.028
	(0.110)	(0.117)	(0.048)	(0.048)
Unemployment duration: 19 months or more	0.234	0.242	−0.044	−0.057
	(0.100)**	(0.106)**	(0.056)	(0.056)
Weeks since reservation wage reported	−0.051	−0.057	0.029	0.027
	(0.027)*	(0.027)**	(0.009)***	(0.008)***
Part-time job offer	0.230	0.184	−0.075	−0.078
	(0.087)***	(0.085)**	(0.046)	(0.044)*
Incentive choice: $40		−0.076		0.030
		(0.083)		(0.041)
Degree of risk loving (*0 = Unwilling to take risk; 10 = Very willing to take risks*)		−0.073		−0.001
		(0.019)***		(0.008)
Initial savings less than $10,000		0.029		0.123
		(0.109)		(0.057)**
Age/10	0.248	0.094	−0.086	−0.111
	(0.304)	(0.286)	(0.167)	(0.160)
$\text{Age}^2/100$	−0.034	−0.016	0.009	0.014
	(0.035)	(0.033)	(0.019)	(0.019)

(*continued*)

findings from our cross-sectional analysis, which finds no significant relationship between unemployment benefits and reservation wages, and the finding that reservation wages did not increase significantly in the weeks after the extension of UI benefits from 79 weeks to 99 weeks. These findings—read in the light of Shimer and Werning's optimal UI formula—challenge their tentative conclusion that it may be welfare improving to increase UI benefits beyond the currently available level, although it should be noted that benefits were provided for an unusually long period during our sample.[18]

It remains an open question *why* the reported reservation wage declines at a relatively modest rate over the spell of unemployment, particularly among those who have only minimal or no savings to cushion the drop in consumption at UI exhaustion. One possibility is that unemployed workers rely on other forms of insurance, such as family or friends, and thus experience a limited drop in consumption at the point of UI exhaustion. An alternative explanation, however, is that unemployed individuals face little dispersion in potential wage offers and thus may have little scope for reducing their reservation wage, even if the drop in consumption at UI exhaustion is large. (In our calibration exercise, however, increasing the standard deviation of log wage offers from 0.24 to 0.35 only trivially affects the predicted

[18] See Section 7 of the working paper version (Krueger and Mueller 2014) for a detailed discussion of how the decline in the reservation wage relates to Shimer and Werning's analysis of the optimality of UI benefits.

Table 11—Determinants of the Potential Bias in the Reported Reservation Wage (*continued*)

Dependent variable:	Probability of accepting job offer		log(accepted wage) - log(lagged reservation wage)	
Sample:	Offers below the reservation wage		Accepted offers	
Some college education	−0.017	−0.023	0.061	0.060
	(0.120)	(0.116)	(0.056)	(0.055)
College degree	0.157	0.191	−0.000	0.000
	(0.134)	(0.124)	(0.060)	(0.058)
Some graduate education	−0.014	0.096	−0.093	−0.067
	(0.171)	(0.150)	(0.077)	(0.074)
Graduate degree	0.283	0.314	0.146	0.145
	(0.140)**	(0.142)**	(0.061)**	(0.060)**
Female	0.117	0.071	0.002	0.002
	(0.131)	(0.121)	(0.071)	(0.069)
Married	−0.048	0.012	0.017	0.016
	(0.149)	(0.140)	(0.077)	(0.076)
Female × Married	−0.260	−0.288	−0.041	−0.021
	(0.176)	(0.167)*	(0.084)	(0.080)
Black	−0.146	−0.080	−0.032	−0.053
	(0.104)	(0.102)	(0.057)	(0.056)
Asian or other race	−0.239	-0.262	0.089	0.082
	(0.210)	(0.221)	(0.094)	(0.089)
Race not available	0.055	0.144	0.014	0.014
	(0.170)	(0.158)	(0.090)	(0.093)
Hispanic	−0.059	−0.090	−0.081	−0.110
	(0.141)	(0.131)	(0.077)	(0.074)
Ethnicity not available	0.349	0.298	0.091	0.087
	(0.139)**	(0.153)*	(0.117)	(0.119)
Mean of dependent variable	0.505	0.507	−0.021	−0.021
Observations	566	564	745	743
(Pseudo) R^2	0.121	0.176	0.125	0.145

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

decline in the reservation wage.) Finally, it is possible that many job seekers anchor their reservation wage on their previous wage or are persistently over confident about their prospects, which causes them to set their reservation wage too high and to reduce it too slowly compared to the optimal rate predicted by a search model. Distinguishing among these different explanations seems of primary importance for future research.

Our results are also in contrast with a recent paper by Brown, Flinn, and Schotter (2011), which found that in an experimental setting with a stationary environment, subjects' reservation wage declined strongly over time. The authors explain their novel finding with "non-stationary subjective costs" of time spent searching. While it is possible that the unemployed workers in our sample face increasing costs of remaining unemployed, our estimates of the decline of reservation wages over the spell of unemployment suggest at best a limited impact of these subjective costs on the reservation wage. Similarly, our estimates point toward a small degree of

updating on the potential wage offer distribution and/or the likelihood of receiving a job offer, as this should contribute toward a downward sloping path of the reservation wage over the spell of unemployment.

Lastly, our finding that the lagged reservation wage predicts job acceptance or rejection is an encouraging result for research based on self-reported reservation wages, as is the finding that the reservation wage conveys information beyond that in the previous wage. Including a question on unemployed workers' reservation wages in ongoing labor force surveys would thus seem to be a useful addition to existing labor market indicators.

Appendix

Appendix Table A1—Wage Regressions: Hourly Reservation Wage, Offered Wage, and Previous Wage

	Sample: first week of survey		Sample: interviews with job offers		
	log(hourly reservation wage)	log(hourly previous wage)	log(hourly reservation wage)	log(hourly offered wage)	log(hourly previous wage)
Years of school	0.093	0.098	0.076	0.072	0.080
	(0.004)***	(0.006)***	(0.013)***	(0.013)***	(0.026)***
Experience/10	0.258	0.296	0.182	0.166	0.272
	(0.028)***	(0.036)***	(0.069)***	(0.071)**	(0.093)***
Experience^2/100	−0.039	−0.038	−0.026	−0.027	−0.042
	(0.007)***	(0.008)***	(0.017)	(0.018)	(0.022)*
Female	−0.122	−0.111	−0.148	−0.160	−0.190
	(0.018)***	(0.024)***	(0.049)***	(0.048)***	(0.069)***
Married	0.091	0.117	0.131	0.081	0.130
	(0.021)***	(0.027)***	(0.054)**	(0.051)	(0.069)*
Black	−0.193	−0.173	−0.159	−0.292	−0.211
	(0.022)***	(0.033)***	(0.058)***	(0.054)***	(0.095)**
Asian or other	0.016	0.005	0.151	0.133	0.155
	(0.058)	(0.065)	(0.122)	(0.128)	(0.171)
Race not available	−0.095	−0.114	0.039	0.056	0.088
	(0.034)***	(0.045)**	(0.102)	(0.084)	(0.137)
Not Hispanic	0.055	0.004	0.181	0.321	0.273
	(0.031)*	(0.040)	(0.091)**	(0.078)***	(0.165)*
Ethnicity not available	0.066	0.087	−0.002	0.245	−0.054
	(0.038)*	(0.049)*	(0.103)	(0.085)***	(0.128)
Constant	1.272	1.169	1.472	1.406	1.281
	(0.068)***	(0.090)***	(0.202)***	(0.186)***	(0.340)***
Mean of dependent variable	2.825	2.866	2.841	2.731	2.827
Observations	4,133	4,133	1,216	1,216	1,216
R^2	0.386	0.315	0.321	0.310	0.272

Notes: Survey weights are used. Sample for columns 1–3: interviews with job offers; age 20–65. Sample for columns 4–5: unemployed; no job offer yet accepted; age 20–65.

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

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