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The Impact of Unemployment Insurance Benefit Levels on Recipiency

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This article studies the effect of unemployment insurance benefit levels on recipiency. Increasing benefit levels (as measured by the fraction of weekly earnings of the lost job that they replace) is found to significantly increase the probability of unemployment insurance recipiency among the eligible. There is some evidence, however, that the effect is smaller at high replacement rates. Cost increases resulting from take-up responses are found to be substantial for increases in the state's maximum benefit amount and for increases in the weekly benefit amount in low-replacement-rate states.

KEY WORDS: Dichotomous choice models; Replacement rates; Take-up rates.

The issue of cost figures prominently in legislative debates on changes in the unemployment insurance (UI) system. Assessing costs hinges on how individuals respond to changes in the UI system. Although Levine (1993) examined the effect of UI benefits on uninsured unemployment durations, the individual response that has received the most attention is the effect of changes in UI benefit levels on insured unemployment durations (see Devine and Kiefer [1991] for a survey of the literature). Not all unemployed workers eligible for UI benefits, however, actually apply to receive them. Yet take-up rate changes that arise from changes in the UI system will have an impact on the cost of such system changes. Nevertheless, the question of whether changes in UI benefit levels affect UI recipiency has received little attention in the literature (Blank and Card [1991] and Anderson and Meyer [1993] are exceptions).

This article investigates the impact of changes in UI benefit levels on UI recipiency using data from the 1984–1992 Current Population Survey's (CPS's) Displaced Worker Supplements (DWS). In the United States, the UI system is administered at the state level. The amount of benefits individuals are eligible to receive thus depends, in part, on their state of residence and the year in which they become unemployed. This state–year variation in benefit amounts is exploited to estimate the effect of UI benefit levels on the probability of take-up.

Increasing weekly benefit amounts, when measured as a fraction of weekly earnings of the lost job that they "replace," are found to significantly increase the probability that a displaced worker will receive unemployment benefits. There is some evidence, however, that suggests that the effect is smaller at high replacement rates. This result differs from that of Blank and Card (1991), who, using a microdata sample from the Panel Survey of Income Dynamics, found no significant impact of replacement rates on take-up rates. This finding, however, is consistent with what Blank and Card (1991) found using state-level data and with what Anderson and Meyer (1993) found using administrative data collected from six states in the late seventies and early eighties.

This article also examines the cost implications of these take-up responses to UI benefit changes. The cost impacts of take-up responses to an increase in weekly benefit amounts are found to be larger in low-replacement-rate states than in high-replacement-rate states. All states set a maximum weekly benefit amount. When adjusting UI benefits, states typically change the maximum weekly benefit amount rather than weekly benefit amounts directly. For such changes, only individuals eligible to receive the maximum, who tend to have lower replacement rates, are affected. Cost increases resulting from the take-up response to an increase in the state's maximum benefit amount are found to be substantial when compared to nonbehavioral or direct cost increases. When averaged across all states, the cost increase resulting from take-up responses to a 10% increase in the state's maximum benefit amount range from 26% to 30% of the direct cost increase for white-collar workers and from 19% to 20% of the direct cost increase for blue-collar workers.

The remainder of the article is organized as follows. Section 1 describes the data. Section 2 reports the empirical results of the article. Estimates are obtained separately for blue-collar and white-collar workers. Section 3 presents estimates of the cost impact of take-up responses to benefit-level changes. Section 4 contains some concluding remarks.

1. DATA

The data used in this study are derived from the 1984, 1986, 1988, 1990, and 1992 CPS DWS. To limit the amount of recall bias, individuals are included in the sample only if they were displaced in the two years preceding the survey. In addition, only individuals between the ages of 20 and 61 who were displaced from nonagricultural, full-time jobs due to plant closure, slack work, abolishment of a position, or the end of seasonal work are included in the sample.

This study focuses on the determinants of UI recipiency among eligible individuals. Although the DWS asks respondents whether or not they received UI benefits after losing their jobs, it neglects to ask respondents either about their eligibility status or their benefit levels. In addition, the DWS does not survey all individuals potentially eligible for UI benefits. For example, it does not survey those who are laid off and subsequently recalled. None of the job terminations included in the sample, however, disqualify an individual for UI benefits.

In most states, eligibility for UI benefits (as well as the weekly benefit amount) depends on the employment and/or earnings history of an individual in the year prior to unemployment (the "base period"). For example, eligibility may depends on the highest quarterly earnings in the base period. In the DWS only years of tenure and (usual) weekly earnings in the lost job are reported. To minimize the number of ineligibles in the sample, individuals are included only if they report having at least one year of tenure in their lost job. In addition, state earnings requirements for eligibility are converted into a minimum weekly earnings amount. Individuals are then included in the sample only when their reported weekly earnings in the lost job exceed this minimum amount for their state of residence.

The amount of UI benefits an individual is eligible to receive is imputed from the U.S. Department of Labor (various issues) using reported weekly earnings in their lost job and state of residence (Portugal and Addison [1990] employed a similar imputation method). No adjustments for dependents are made. In many states, weekly benefits depend on high-quarter earnings. Measurement error will then occur to the extent that reported usual weekly earnings in the lost job differs from (the weekly equivalent of) high-quarter earnings.

Individuals reporting zero weeks of joblessness following a job displacement are also included in the sample. This is done to account for the possibility that an increase in UI benefits will reduce the amount of on-the-job search and hence increase the probability of incurring a positive spell of joblessness. This, in turn, may lead to an increase in the probability of take-up.

The final sample size is 8,504. The take-up rate for the full sample is approximately .65. Clearly the effort to exclude ineligible individuals also resulted in some eligible individuals being excluded because, although lower than the sample, the take-up rate of those reporting less than one year of tenure is .40, whereas the take-up rate of those reporting weekly earnings less than the state's minimum weekly earnings requirement is .25.

There is substantial variation in take-up rates over time. Table 1 presents sample UI take-up rates by year of displacement. UI recipiency rates fell from 76.7% in 1982 to 58.2% in 1987. Using March CPS data through 1987, Blank and Card (1991) also documented such a drop, although their take-up levels differed. Some of this difference in take-up levels can be attributed to the inclusion of those reporting zero weeks of joblessness in this sample. If these individuals are excluded, then take-up rates between the two studies are similar. For the remainder of the 1980s, the UI recipiency rate remained stable, hovering around 59%. In the early nineties, however, the take-up rate began to rise and was nearly 70% in 1991.

Table 1. Take-up Rates Among UI Eligible: 1982–1991

	Take-up rates (× 100)			
Year	Overall	Blue-collar	White-collar	% blue-collar
1982	76.7	80.4	68.9	67.5
1983	68.4	73.6	59.5	63.5
1984	61.0	64.1	56.4	59.2
1985	60.6	65.4	52.3	62.0
1986	61.9	68.7	55.1	50.1
1987	58.2	61.9	53.7	54.2
1988	59.6	58.0	61.1	50.1
1989	58.9	60.1	57.4	57.1
1990	62.0	64.6	59.4	49.6
1991	69.6	70.8	68.9	52.6

NOTE: This table reports the fraction of individuals eligible for unemployment insurance benefits who actually receive benefits. The data are from the 1984–1992 CPS DWS, described in the text. Fractions are calculated using CPS household weights.

Table 1 also breaks down yearly UI recipiency rates by the blue-collar/white-collar status of the lost job. In the early years of the sample, the take-up rates of white-collar workers were substantially lower than the take-up rates of blue-collar workers. For example, between the years 1982 and 1986, the take-up rate averaged 59% for white-collar workers and 71% for blue-collar workers. Between 1987 and 1991, however, take-up rates rose to 61%, on average, for white-collar workers and fell to 65%, on average, for blue-collar workers. Thus the gap narrowed by 8% between the two periods.

The last column of Table 1 shows that the mix of eligible displaced workers has also changed substantially over the 1982–1991 period. Between 1982 and 1986, 61% of displaced workers were blue collar. This percentage fell to 53% between 1987 and 1991 (also see Farber 1993).

Weekly benefit amounts are measured as the fraction of weekly earnings of the lost job that they replace. The average replacement rate for the sample is .426. There is considerable variation in replacement rates across states. Figure 1 presents a scatterplot of state average take-up rates by state average replacement rates. The correlation between the replacement and take-up rate, .28, is significantly different from 0 at the 5% level. Thus states with higher than average replacement rates also tend to have higher than average take-up rates.

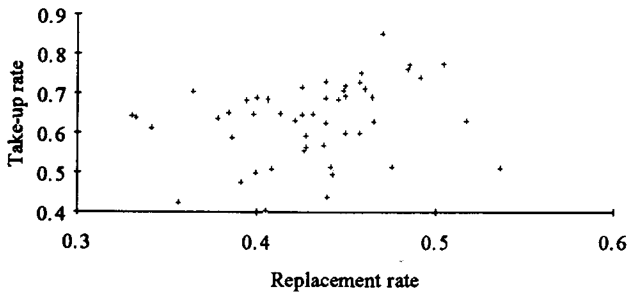

Figure 1. State Average Take-up Rates by State Average Replacement Rates.

Table 2. Summary Statistics

Variable	White-collar (n = 3,627) Mean	White-collar Standard deviation	Blue-collar (n = 4,877) Mean	Blue-collar Standard deviation
Replacement rate*	.411	.122	.438	.106
State unemployment rate (%)	7.073	2.210	7.511	2.500
Waiting week	.860		.845	
Displaced due to plant closing	.353		.415	
Displaced due to slack work	.395		.476	
Displaced due to position abolished	.242		.082	
Displaced because seasonal job ended	.010		.036	
Expected to be displaced	.512		.532	
Tenure in lost job (years)	5.174	5.647	5.664	6.279
Log of previous weekly earnings (1985 dollars)	5.909	.539	5.741	.503
Reside in S.M.S.A.	.788		.653	
Female	.491		.236	
Nonwhite	.098		.147	
Children present	.425		.489	
Young child present (0–5 years)	.180		.222	
Age (years)	37.092	10.177	36.130	10.600
Married	.609		.632	
Head of household	.608		.681	
Less than 12 years schooling	.113		.354	
More than 12 years schooling	.574		.190	

(Continued)

Table 2. (Continued)

Variable	White-collar (n = 3,627) Mean	White-collar Standard deviation	Blue-collar (n = 4,877) Mean	Blue-collar Standard deviation
Region of residence				
New England	.094		.090	
Middle Atlantic	.140		.122	
East North Central	.137		.159	
West North Central	.078		.083	
South Atlantic	.145		.155	
East South Central	.038		.071	
West South Central	.110		.104	
Mountain	.110		.095	
Pacific	.148		.121	
Year of displacement				
1982	.094		.138	
1983	.100		.129	
1984	.082		.093	
1985	.098		.118	
1986	.102		.082	
1987	.097		.081	
1988	.078		.057	
1989	.084		.083	
1990	.114		.086	
1991	.151		.133	
Industry of lost job				
Manufacturing	.249		.463	
Mining	.025		.044	
Construction	.038		.189	
Transportation	.066		.075	
Trade	.259		.114	
Finance, insurance, real estate	.095		.009	
Services	.248		.100	
Public administration	.020		.006	

NOTE: The data are from the 1984–1992 CPS DWS described in the text.
*The replacement rate is defined as weekly unemployment insurance benefits divided by weekly earnings in the lost job.

Section 2 presents estimates of the effect of the replacement rate on the probability of take-up using the individual level data from the DWS. Summary statistics for the explanatory variables used in the analysis, broken down by blue-collar/white-collar status of the lost job, are presented in Table 2.

2. EMPIRICAL RESULTS

2.1 Logistic Estimates

Before presenting the main empirical results of the article, the take-up decision is discussed briefly (for a more detailed model of take-up decisions, see Anderson and Meyer [1993]). An individual will file a claim for UI benefits when it leads to an increase in his/her expected discounted utility. The costs of take-up include both the (time) costs associated with filing a claim for UI benefits and any psychological (stigma) costs connected with receiving UI. The benefit of receiving UI depends on the level of weekly benefits, the duration of benefits, and the probability distribution of unemployment durations. In this article, a reduced-form approach is taken that focuses primarily on the effect of benefit amounts on take-up. Other regressors are included in the model to control for the other factors that affect take-up.

In preliminary estimates (which for the sake of brevity are not reported here), using the 1986–1992 DWS (the 1984 DWS is excluded due to differences in the manner in which joblessness durations were determined), the explanatory variables used in this study are found to have significantly different effects on the expected log joblessness durations of white-collar and blue-collar workers. Because the expected duration of joblessness affects the decision to file a claim, white-collar and blue-collar workers will be analyzed separately.

Tables 3 and 4 report the estimates from a logistic probability model for white-collar and blue-collar workers, respectively. Columns (1)–(3) report estimates that control for regional fixed effects, and columns (4)–(6) report estimates that control for state fixed effects. The estimates in columns (1) and (4) assume that the replacement rate has a linear effect on the log-odds of take-up. Although economic theory predicts that an increase in benefits should have a positive effect on the log-odds of take-up, it need not be linear. Thus, in columns (2) and (5), a quadratic term for the replacement rate is also included in the model. Finally, columns (3) and (6) include four linear spline variables, $\max(0, .1(1+j)-r)$,

Table 3. Logistic Estimates: White-Collar Sample

Row	Variable	(1)	(2)	(3)	(4)	(5)	(6)
1.	Replacement rate	2.171	7.199	−1.027	2.232	6.712	−.716
		(.594)	(1.905)	(1.628)	(.735)	(2.021)	(2.111)
2.	Replacement rate squared	—	−6.426	—	—	−6.016	—
			(2.320)			(2.529)	
3.	Max(0,.2—replacement rate)	—	—	−6.656	—	—	−6.556
				(4.899)			(5.024)
4.	Max(0,.3—replacement rate)	—	—	−1.831	—	—	−1.621
				(3.399)			(3.456)
5.	Max(0,.4—replacement rate)	—	—	2.228	—	—	1.853
				(2.679)			(2.787)
6.	Max(0,.5—replacement rate)	—	—	−4.481	—	—	−3.822
				(2.411)			(2.829)
7.	State unemployment rate	.105	.106	.103	.085	.083	.082
		(.025)	(.025)	(.025)	(.034)	(.034)	(.034)
8.	Waiting week	.111	.108	.106	.356	.359	.343
		(.113)	(.113)	(.114)	(.582)	(.582)	(.582)
9.	Displaced due to slack work	.651	.646	.647	.645	.644	.643
		(.089)	(.089)	(.089)	(.091)	(.090)	(.091)
10.	Displaced due to position abolished	.440	.439	.434	.443	.442	.439
		(.092)	(.092)	(.093)	(.094)	(.094)	(.094)
11.	Displaced because seasonal job ended	.300	.310	.308	.255	.257	.255
		(.368)	(.373)	(.371)	(.378)	(.381)	(.380)
12.	Expected to be displaced	−.236	−.242	−.242	−.232	−.239	−.240
		(.073)	(.073)	(.073)	(.074)	(.074)	(.074)
13.	Tenure in lost job	.007	.006	.007	.005	.004	.005
		(.007)	(.007)	(.007)	(.008)	(.008)	(.008)
14.	Log of previous weekly earnings	.481	.548	.595	.505	.527	.583
		(.141)	(.144)	(.146)	(.165)	(.166)	(.169)
15.	Reside in S.M.S.A.	−.201	−.216	−.209	−.150	−.164	−.160
		(.096)	(.097)	(.097)	(.107)	(.107)	(.107)
16.	Female	.103	.105	.111	.130	.131	.135
		(.091)	(.091)	(.091)	(.093)	(.093)	(.093)
17.	Nonwhite	.206	.212	.207	.252	.258	.251
		(.125)	(.125)	(.125)	(.128)	(.128)	(.129)
18.	Children present	.184	.189	.192	.187	.187	.191
		(.098)	(.098)	(.098)	(.100)	(.100)	(.100)
19.	Young child present	.010	.012	.008	−.006	−.001	−.005
		(.117)	(.118)	(.118)	(.120)	(.120)	(.120)
20.	Age	.046	.044	.044	.039	.039	.039
		(.029)	(.029)	(.029)	(.030)	(.030)	(.030)
21.	Age squared/10	−.003	−.003	−.003	−.002	−.002	−.002
		(.004)	(.004)	(.004)	(.004)	(.004)	(.004)
22.	Married	−.189	−.184	−.190	−.183	−.183	−.188
		(.089)	(.089)	(.090)	(.091)	(.091)	(.091)
23.	Head of household	−.333	−.338	−.339	−.319	−.327	−.327
		(.093)	(.093)	(.093)	(.094)	(.094)	(.094)
24.	Less than 12 years schooling	−.056	−.060	−.052	−.027	−.033	−.027
		(.139)	(.140)	(.140)	(.142)	(.143)	(.143)
25.	More than 12 years schooling	−.137	−.143	−.147	−.135	−.143	−.146
		(.086)	(.086)	(.086)	(.087)	(.087)	(.088)
26.	Regional controls	yes	yes	yes	no	no	no
27.	State controls	no	no	no	yes	yes	yes
28.	Year controls	yes	yes	yes	yes	yes	yes
29.	Industry controls	yes	yes	yes	yes	yes	yes
30.	Intercept	.438	.438	.438	.442	.443	.443
	Likelihood value	−2,286	−2,281	−2,280	−2,260	−2,257	−2,255

NOTE: Standard errors are in parentheses. This table reports logistic estimates of the probability of unemployment benefit receipt. The data are from the 1984–1992 CPS DWS described in the text.

$j = 1, \ldots, 4$, in addition to the linear term in which r denotes the replacement rate.

Although the effect of the replacement rate on the probability of UI take-up is significant at the 5% level for both blue-collar and white-collar workers in models that include only a linear term, the results in columns (2), (3), (5), and (6) of Tables 3 and 4 show that the nonlinear terms are (jointly) significant at the 5% level; the exception is column (6) of Table 4, in which the linear spline terms are jointly significant at the 10% level.

Table 4. Logistic Estimates: Blue-Collar Sample

Row	Variable	(1)	(2)	(3)	(4)	(5)	(6)
1.	Replacement rate	2.157	8.533	−1.497	2.298	7.593	−2.674
		(.527)	(2.073)	(1.453)	(.628)	(2.246)	(2.002)
2.	Replacement rate squared	—	−7.979	—	—	−6.855	—
			(2.503)			(2.803)	
3.	Max(0,.2—replacement rate)	—	—	−2.311	—	—	.159
				(6.279)			(6.354)
4.	Max(0,.3—replacement rate)	—	—	−.119	—	—	−1.372
				(3.624)			(3.704)
5.	Max(0,.4—replacement rate)	—	—	−1.269	—	—	.568
				(2.563)			(2.727)
6.	Max(0,.5—replacement rate)	—	—	−3.571	—	—	5.727
				(2.158)			(2.747)
7.	State unemployment rate	.102	.103	.103	.113	.113	.111
		(.021)	(.021)	(.021)	(.030)	(.030)	(.031)
8.	Waiting week	.022	.007	−.005	−.397	−.402	−.396
		(.105)	(.105)	(.105)	(.460)	(.461)	(.461)
9.	Displaced due to slack work	.587	.586	.585	.581	.581	.580
		(.075)	(.075)	(.075)	(.076)	(.076)	(.076)
10.	Displaced due to position abolished	−.030	−.032	−.029	−.034	−.039	−.034
		(.124)	(.124)	(.125)	(.126)	(.126)	(.127)
11.	Displaced because seasonal job ended	.460	.453	.454	.429	.431	.434
		(.175)	(.174)	(.174)	(.178)	(.177)	(.178)
12.	Expected to be displaced	−.001	.000	.000	−.016	−.017	−.016
		(.067)	(.067)	(.067)	(.068)	(.068)	(.068)
13.	Tenure in lost job	.016	.015	.015	.014	.013	.014
		(.007)	(.007)	(.007)	(.007)	(.007)	(.007)
14.	Log of previous weekly earnings	.831	.875	.891	.884	.889	.911
		(.115)	(.116)	(.117)	(.129)	(.130)	(.130)
15.	Reside in S.M.S.A.	−.143	−.158	−.151	−.114	−.126	−.121
		(.076)	(.076)	(.076)	(.084)	(.084)	(.084)
16.	Female	.244	.260	.259	.260	.273	.272
		(.097)	(.097)	(.097)	(.099)	(.099)	(.099)
17.	Nonwhite	.196	.197	.194	.209	.208	.203
		(.100)	(.100)	(.100)	(.105)	(.105)	(.105)
18.	Children present	−.141	−.141	−.138	−.135	−.137	−.133
		(.088)	(.088)	(.088)	(.090)	(.090)	(.090)
19.	Young child present	.189	.182	.179	.181	.178	.175
		(.100)	(.100)	(.100)	(.102)	(.102)	(.102)
20.	Age	.039	.039	.040	.039	.039	.040
		(.025)	(.025)	(.025)	(.025)	(.025)	(.025)
21.	Age squared/10	−.002	−.002	−.002	−.002	−.002	−.002
		(.003)	(.003)	(.002)	(.003)	(.003)	(.003)
22.	Married	.195	.188	.190	.187	.181	.182
		(.082)	(.082)	(.082)	(.083)	(.083)	(.084)
23.	Head of household	−.191	−.191	−.190	−.193	−.194	−.195
		(.086)	(.086)	(.086)	(.088)	(.088)	(.088)
24.	Less than 12 years schooling	.038	.040	.040	.051	.051	.049
		(.084)	(.084)	(.084)	(.086)	(.086)	(.086)
25.	More than 12 years schooling	−.117	−.108	−.112	−.120	−.111	−.115
		(.090)	(.090)	(.090)	(.092)	(.092)	(.092)
26.	Regional controls	yes	yes	yes	no	no	no
27.	State controls	no	no	no	yes	yes	yes
28.	Year controls	yes	yes	yes	yes	yes	yes
29.	Industry controls	yes	yes	yes	yes	yes	yes
30.	Intercept	.876	.878	.878	.886	.888	.888
	Likelihood value	−2,765	−2,760	−2,760	−2,737	−2,734	−2,733

NOTE: Standard errors are in parentheses. This table reports logistic estimates of the probability of unemployment benefit receipt. The data are from the 1984–1992 CPS DWS described in the text.

Figures 2 and 3 plot the probability of take-up versus the replacement rate for the white-collar and blue-collar estimates, respectively, when the other explanatory variables are evaluated at their sample means. Although both the quadratic and linear spline models suggest that at high replacement rates an increase in the replacement rate leads to a decline in the take-up rate, the effect is insignificantly different from 0 for replacement rates above .5 in the linear spline model for both the white-collar and blue-collar estimates.

Although Blank and Card (1991) focused primarily on explaining the recent decline in take-up rates, they reported results concerning the effect of replacement rates on take-

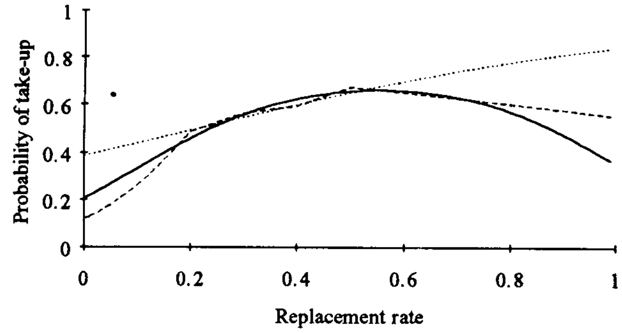

Figure 2. Estimated Probability of Take-up by Replacement Rates: White-Collar Sample. The dotted, solid, and dashed lines are based on logistic models in which replacement rate has a linear, quadratic, and linear spline effect on the log-odds of take-up, respectively. All other explanatory variables are evaluated at their sample means.

up rates. Using a microdata sample from the Panel Survey of Income Dynamics, Blank and Card (1991) found no significant effect of the replacement rate on the probability of take-up. This insignificance, however, may stem from their assignment of state average replacement rates to individuals. Thus, after controlling for region effects, little variation in the replacement rate remained.

Using a sample derived from the Continuous Wage and Benefit History project, Anderson and Meyer (1993) estimated an elasticity of the probability of take-up with respect to weekly benefit amount of .46 for their mass lay-off sample. The comparable elasticity estimates for this article, which are based on the estimates in columns (1) of Tables 3 and 4, are .35 for white-collar workers and .26 for blue-collar workers. Much of the difference between the elasticity estimates of this article and Anderson and Meyer's (1993) article can be attributed to the lower take-up rate in their sample (.41 versus .60 for white-collar workers and .68 for blue-collar workers). Using this lower take-up probability raises the estimated elasticities to .53 for white-collar workers and .56 for blue-collar workers. Elasticity estimates based on the estimates in columns (2) and (3) of Tables 3 and 4, however, are substantially lower than the elasticities reported by Anderson

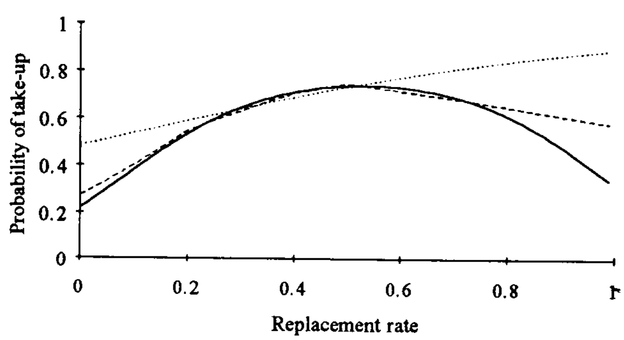

Figure 3. Estimated Probability of Take-up by Replacement Rates: Blue-Collar Sample. The dotted solid, and dashed lines are based on logistic models in which the replacement rate has a linear, quadratic, and linear spline effect on the log-odds of take-up, respectively. All other explanatory variables are evaluated at their sample means.

and Meyer (1993) (.15 and .13, respectively, for white-collar workers and .11 and .11, respectively, for blue-collar workers) even after accounting for the difference in take-up rates (.22 and .19, respectively, for white-collar workers and .18 and .18, respectively, for blue-collar workers).

Many of the other estimates reported in Tables 3 and 4 appear to reflect differences in expected joblessness durations. For example, white-collar workers who were displaced either because of slack work or because their positions were abolished are significantly more likely to file a claim than white-collar workers displaced because of plant closings. No significant difference in take-up rates, however, exists between those displaced because seasonal jobs ended and those displaced because of plant closings. Estimates obtained from the 1986–1992 DWS using a log-normal duration model that adjusts for right censoring show that white-collar workers who lost their jobs either because of slack work or because their positions were abolished have significantly longer joblessness durations than those who lost their jobs because of plant closings. Similar to the findings for take-up rates, there is no significant difference between the joblessness durations of white-collar workers who lost their jobs because seasonal jobs ended and those who lost their jobs because of plant closings.

Blue-collar workers displaced either because of slack work or because of seasonal jobs ending are significantly more likely to file a claim than those who lost their jobs because of plant closings. The take-up rates of blue-collar workers who lost their jobs because of abolished positions and those who lost their jobs because of plant closings, however, do not significantly differ. Again, this may reflect the fact that blue-collar workers who lost their jobs either because of slack work or because seasonal jobs ended have significantly longer joblessness durations than those who lost their jobs because of plant closings. The joblessness durations of those who lost their jobs because of abolished positions, on the other hand, do not significantly differ from those who lost their jobs because of plant closings.

As is found with the log-joblessness-duration estimates, the effect of the explanatory variables on the probability of take-up significantly differs between blue-collar and white-collar workers. Likelihood ratio tests of equal coefficients (other than the intercept) reject the null at the 1% level both in models that control for regional effects ($\chi^2(45) = 103.5$, $\chi^2(46) = 99.7$, and $\chi^2(49) = 99.8$, for estimates which include linear, quadratic, and linear spline terms for the replacement rate, respectively) and in models that control for state effects ($\chi^2(87) = 144.3$, $\chi^2(88) = 139.0$, and $\chi^2(91) = 139.7$, for estimates that include linear, quadratic, and linear spline terms for the replacement rate, respectively).

2.2 Robustness Checks

This subsection reports the results of several robustness checks that are performed to assess the sensitivity of the results reported previously to model specification. Estimates are obtained with several other dichotomous choice models including the Weibull, probit, and linear probability models.

In addition, alternative linear spline models in which the knot points are fixed at the first, second, and third quartiles of the replacement rate are estimated. In both cases, the results regarding the effect of the replacement rate on the probability of take-up are unchanged.

For logistic probability models that include the weekly benefit amount instead of the replacement rate, the estimated elasticity of the probability of take-up with respect to the weekly benefits is .32, .28, and .27 for white-collar workers and .22, .20, and .21 for blue-collar workers when weekly benefits are assumed to have a linear, quadratic, and linear spline effect on the log-odds of take-up, respectively, and the knot points of the linear spline are fixed at \$50, \$100, \$150, and \$200. Although the elasticity estimates derived from models including only a linear term for the weekly benefit amount are similar to those reported previously for the replacement rate, the elasticity estimates for models with quadratic and linear spline terms increase by over .1. Once differences in take-up rates are accounted for, the elasticity estimates produced by these models are similar to those reported by Anderson and Meyer (1993) irrespective of how the effect of weekly benefits is modeled.

Nearly 40% of those with less than one year of tenure in their lost job report receiving UI benefits. So, estimates are also obtained using a sample that includes these individuals. Although the estimated elasticity of the probability of take-up with respect to the replacement rate increases by .14 for the white-collar sample, when the replacement rate is assumed to have a linear effect on the log-odds of take-up, elasticity estimates derived from the other model specifications are similar.

Because the weekly benefit amount is a function of past earnings, it may be difficult to disentangle the effects of UI benefit levels from past history. To check whether the findings reported previously for replacement rates are due to the assumption that weekly earnings in the lost job have a log-linear effect on the log-odds of take-up, logistic estimates are also obtained using models in which weekly earnings in the lost job have a cubic effect on the log-odds of take-up. Using the models from columns (1)–(3) of Tables 3 and 4 but replacing the log-linear term for weekly earnings in the lost job with a cubic, the elasticity estimates of the probability of take-up with respect to the replacement rate are .23, .15, and .13 for white-collar workers and .17, .12, and .12 for blue-collar workers. Only the elasticity estimates of models that assume that the replacement rate has a linear effect on the log-odds of filing a claim differ appreciably. The linear spline terms for the replacement rate, however, are no longer jointly significant, and the quadratic terms for the replacement rate are now significant only at the 10% level.

An alternative way to disentangle the effect of previous earnings from the effect of the replacement rate on the probability of take-up, as well as to control for the potential endogeneity of the replacement rate, is to instrument for the replacement rate in the probability-of-take-up equation using an instrument that is not a function of past earnings. Both two-stage least squares (TSLS) and generalized two-stage least squares (GTSLS) estimates of the conditional probability of UI recipiency are obtained using the state's maximum benefit amount as an instrument for the replacement rate. The latter method is a two-step method similar to that proposed by Goldberger (1964) for linear probability models, which adjusts for the heteroscedasticity that results from a dichotomous dependent variable. Table 5 presents the TSLS and GTSLS estimates of the effect of the replacement rate on the probability of take-up. Ordinary least squares (OLS) and Goldberger's generalized least squares (GLS) estimates are also presented for comparison purposes. The coefficient for the replacement rate remains significant and actually increases in size in all but one case when the state's maximum benefit amount is used as an instrument. For the OLS and TSLS estimates, the Hausman (1978) test for endogeneity fails to reject the null of no endogeneity at convectional levels both for the white-collar ($\chi^2(46) = .285$) and blue-collar ($\chi^2(46) = 2.52$) samples.

Table 5. Two-Stage Least Squares Estimates of the Effect of the Replacement Rate on Take-up

Row	Sample	OLS (1)	TSLS (2)	GLS (3)	GTSLS (4)
1.	White-collar	.469	.681	.452	.686
		(.130)	(.257)	(.122)	(.247)
2.	Blue-collar	.433	.763	.490	.408
		(.100)	(.224)	(.084)	(.197)

NOTE: Standard errors are in parentheses. The estimates in rows 1 and 2 include the same independent variables as the estimates reported in column (1) of Tables 3 and 4, respectively. Column (1) presents linear probability model estimates for the probability of UI receipt. Column (2) presents SLS estimates for the probability of UI receipt when the state's maximum benefit amount is used as an instrument for the replacement rate. Column (3) presents GLS estimates for the probability of UI receipt. Column (4) presents GTSLS estimates for the probability of UI receipt when the state's maximum benefit amount is used as an instrument for the replacement rate. The data are from the 1984–1992 CPS DWS described in the text.

3. THE COST IMPACT OF UI TAKE–UP RATE CHANGES

This section examines the cost implications of take-up responses to UI benefit changes. All states specify both maximum and minimum benefit levels. Denote these levels by M and m, respectively. Assuming constant weekly earnings in the base period, eligible individuals are then paid a fraction, f, of their weekly earnings in the lost job subject to these maximum and minimum amounts. Let b denote the weekly benefit amount. An individual's weekly benefit amount is then determined by $b = \min(M, \max(m, fw))$, where w denotes weekly earnings in the lost job. Usually, states change M, m, and/or f rather than b directly. This section presents point estimates of the costs arising from take-up responses to 10% increases in M and b.

Recall that the replacement rate is defined by $r = b/w$ and let $\mathbf{x}$ be a k-dimensional vector of the other variables which affect take-up. The probability of UI receipt given r and $\mathbf{x}$ is denoted by $p(r, \mathbf{x})$. Let $\Delta c(\mathbf{x})$ represent the expected change in costs due to a change in UI benefits for an individual with the vector of characteristics $\mathbf{x}$. Then,

$$\Delta c(\mathbf{x}) = p(r_2, \mathbf{x})E(d \mid r_2, \mathbf{x})b_2 - p(r_1, \mathbf{x})E(d \mid r_1, \mathbf{x})b_1, \quad (1)$$

where d denotes the duration of insured unemployment and the subscripts 1 and 2 refer to the UI parameter values before and after the change, respectively. Equation (1) can be rewritten as

$$\begin{aligned}\Delta c(\mathbf{x}) = p(r_1, \mathbf{x})E(d \mid r_1, \mathbf{x})(b_2 - b_1) \\ + E(d \mid r_1, \mathbf{x})b_2(p(r_2, \mathbf{x}) - p(r_1, \mathbf{x})) \\ + p(r_2, \mathbf{x})b_2(E(d \mid r_2, \mathbf{x}) - E(d \mid r_1, \mathbf{x})). \quad (2)\end{aligned}$$

The first term in (2) denotes the nonbehavioral or direct-cost impact of a change in UI benefits, whereas the second and third terms represent the cost changes due to the take-up rate and expected insured duration response, respectively. An estimate of the average change in cost per eligible individual is then

$$\begin{aligned}\Delta\bar{c} = n^{-1}\sum_{i=1}^{n}\widehat{p}(r_{1i}, \mathbf{x}_i)\widehat{E}(d \mid r_{1i}, \mathbf{x}_i)(b_{2i} - b_{1i}) \\ + \widehat{E}(d \mid r_{1i}, \mathbf{x}_i)b_{2i}(\widehat{p}(r_{2i}, \mathbf{x}_i) - \widehat{p}(r_{1i}, \mathbf{x}_i)) \\ + \widehat{p}(r_{2i}, \mathbf{x}_i)b_{2i}(\widehat{E}(d \mid r_{2i}, \mathbf{x}_i) - \widehat{E}(d \mid r_{1i}, \mathbf{x}_i)), \quad (3)\end{aligned}$$

where hats denote estimated values. Thus the average cost change due to take-up responses, when measured as a fraction of the average direct cost change, equals

$$\frac{\left(\sum_{i=1}^{n}\widehat{E}(d \mid r_{1i}, \mathbf{x}_i)b_{2i}(\widehat{p}(r_{2i}, \mathbf{x}_i) - \widehat{p}(r_{1i}, \mathbf{x}_i))\right)}{\left(\sum_{i=1}^{n}\widehat{p}(r_{1i}, \mathbf{x}_i)\widehat{E}(d \mid r_{1i}, \mathbf{x}_i)(b_{2i} - b_{1i})\right)}. \quad (4)$$

Tables 6 and 7 present estimates of these take-up cost changes for a 10% increase in weekly benefit amounts and for a 10% increase in the state's maximum benefit amount. For the latter change, only individuals at the state maximum are affected. In this data, approximately 45% of blue-collar workers and 56% of white-collar workers were eligible for the maximum benefit amount in their state.

For the estimates in Tables 6 and 7, $\widehat{p}$ is based on the logistic model estimates reported in columns (1)–(3) of Tables 3 and 4, and $\widehat{E}$ is based on estimates from a log-normal duration model that adjusts for right censoring. Because d refers to insured durations rather than total joblessness durations, when the expected joblessness duration of an individual exceeds 26 weeks $\widehat{E}(d \mid x)$ is fixed at 26. Take-up cost estimates based on (4), which are measured as a fraction of direct costs, are not very sensitive to this assumption. Tables 6 and 7 also present, in parentheses, the average change in the probability of take-up,

$$\Delta\bar{p} = n^{-1}\sum_{i=1}^{n}[\widehat{p}(r_{2i}, \mathbf{x}_i) - \widehat{p}(r_{1i}, \mathbf{x}_i)],$$

for a 10% increase in the state's weekly benefit amount and for a 10% increase in the state's maximum benefit amount.

Rows 1 and 2 of Tables 6 and 7 present estimates of the take-up cost increase from a 10% increase in the weekly benefit amount and from a 10% increase in the state's maximum benefit amount, respectively, when the sums in (4) are taken over individuals from all states. Rows 3 and 4 of Tables 6 and 7 present estimates of take-up cost increases that are based only on individuals living in Indiana, whereas rows 5 and 6 present estimates of take-up cost increases that are based only on individuals living in New Jersey. Indiana was chosen because of its relatively low average replacement rate (.332). New Jersey, on the other hand, was chosen because of its relatively high average replacement rate (.492).

Estimates of the take-up cost increase resulting from a 10% increase in weekly benefits depends to a large extent on how the effect of the replacement rate on the probability of take-up rate is modeled. When only a linear term for the replacement rate is included in the model, a 10% increase in weekly benefits is estimated to increase the probability of take-up by .019 for white-collar workers and .018 for blue-collar workers. This increase translates into a take-up cost increase of \$68.77 (\$81.14) per eligible blue-collar (white-

Table 6. Estimate Cost Impact of the Take-up Rate Change for 10% Increases in Benefits and State Maximum Benefit Amounts: White-Collar Sample

		Model		
Row	Sample	Linear (1)	Quadratic (2)	Spline (3)
1.	$b_2 = 1.1\ b_1$: All states	.301	.181	.138
		(.019)	(.010)	(.007)
2.	$M_2 = 1.1\ M_1$: All states	.255	.273	.302
		(.008)	(.009)	(.009)
3.	$b_2 = 1.1\ b_1$: Indiana	.286	.287	.251
		(.018)	(.014)	(.014)
4.	$M_2 = 1.1\ M_1$: Indiana	.272	.240	.310
		(.014)	(.015)	(.016)
5.	$b_2 = 1.1\ b_1$: New Jersey	.340	.088	.093
		(.022)	(.004)	(.005)
6.	$M_2 = 1.1\ M_1$: New Jersey	.252	.247	.252
		(.009)	(.009)	(.010)

NOTE: This table reports the estimates of the cost impact for 10% increases in benefit levels or 10% increases in the state's maximum benefit amount measured as a fraction of nonbehavioral or direct cost increases. The estimates in columns (1)–(3) are based on the logistic probability model estimates reported in columns (1)–(3) of Table 3, respectively. Average take-up rate changes reported in parentheses.

Table 7. Estimated Cost Impact of the Take-up Rate Change for 10% Increases in Benefits and State Maximum Benefit Amounts: Blue-Collar Sample

		Model		
Row	Sample	Linear (1)	Quadratic (2)	Spline (3)
1.	$b_2 = 1.1\ b_1$: all states	.239	.088	.017
		(.018)	(.005)	(−.001)
2.	$M_2 = 1.1\ M_1$: all states	.186	.185	.199
		(.005)	(.005)	(.005)
3.	$b_2 = 1.1\ b_1$: Indiana	.186	.195	.174
		(.012)	(.012)	(.012)
4.	$M_2 = 1.1\ M_1$: Indiana	.163	.226	.215
		(.011)	(.015)	(.015)
5.	$b_2 = 1.1\ b_1$: New Jersey	.300	−.113	−.106
		(.023)	(−.009)	(−.010)
6.	$M_2 = 1.1\ M_1$: New Jersey	.246	.087	.067
		(.008)	(.003)	(.002)

NOTE: This table reports the estimates of the cost impact for 10% increases in benefit levels or 10% increases in the state's maximum benefit amount measured as a fraction of nonbehavioral or direct cost increases. The estimates in columns (1)–(3) are based on the logistic probability model estimates reported in columns (1)–(3) of Table 3, respectively. Average take-up rate changes reported in parentheses.

collar) worker which amount to 23.9% (30.1%) of the direct cost increase.

Take-up cost estimates are considerably lower, however, when the replacement rate is assumed to have a nonlinear effect of the log-odds of take-up. For example, when the replacement rate is modeled using a quadratic, estimates of the take-up rate increase following a 10% increase in weekly benefits drop to .01 for white-collar workers and .005 for blue-collar workers. This increase translates into a take-up cost increase of $25.44 ($48.78) per eligible blue-collar (white-collar) worker, which amounts to 8.9 (18.1%) of the direct cost increase.

There is more consistency across model specifications when estimating the take-up cost increase from a 10% increase in the state's maximum benefit amount. This occurs because, on average, individuals at the state maximum have lower replacement rates (.359 for blue-collar workers and .338 for white-collar workers) and at low replacement rates all models predict that increasing the replacement rate increases the probability of filing a UI claim. When averaged over individuals in all states, estimates of the take-up cost increase range from 18% to 20% of the direct cost increase for blue-collar workers and from 26% to 30% of the direct cost increase for white-collar workers across the different model specifications.

When based on logistic models in which weekly earnings in the lost job have a cubic effect on the log-odds of take-up, estimates of the take-up cost increase for blue-collar (white-collar) workers range from 7% to 16% (11% to 20%) of the direct cost increase for a 10% increase in weekly benefits, and from 11% to 15% (17% to 27%) of the direct cost increase for a 10% increase in the state's maximum benefit level. Again, take-up cost increases are larger for white-collar workers and, although not unambiguously, for increases in the state's maximum benefit amount.

For a low-replacement-rate state such as Indiana, estimates of the take-up cost increase are fairly stable across the different model specifications, both for a 10% increase in weekly benefits and for a 10% increase in the state's maximum benefit level. Again, this occurs because at low replacement rates the different model specifications yield similar estimates of the take-up response to such benefit increases. Estimates of take-up cost increases range from 25% to 29% of direct cost increases for white-collar workers and from 17% to 20% of direct cost increases for blue-collar workers for a 10% increase in weekly benefits. For a 10% increase in the state's maximum benefit level, estimates of take-up cost increases range from 24% to 31% of direct cost increases for white-collar workers and from 16% to 23% of direct cost increases for blue-collar workers.

For a high-replacement-rate state like New Jersey, the estimate of the take-up cost increase due to a 10% increase in weekly benefits depends to a large extent on the model specification, with larger estimates being produced by those models in which the replacement rate is restricted to have a linear effect on the log-odds of take-up. The take-up cost estimates for a 10% increase in the state's maximum benefit level also depend on model specification for blue-collar workers. For white-collar workers, however, all models estimate that the cost increase due to take-up responses amounts to one-quarter of the direct cost increase. Finally, note that estimates of the cost increase due to take-up responses for both Indiana and New Jersey do not depend on whether weekly earnings in the lost job is assumed to have a log-linear or cubic effect on the log-odds of take-up.

To summarize, the results of this section suggest that changes in take-up behavior can have a substantial cost impact when a state increases its maximum benefit level or when a low-replacement-rate state, such as Indiana, increases weekly benefits directly. For high-replacement-rate states such as New Jersey, however, the cost impact of the take-up response to a direct increase in weekly benefits is small.

4. CONCLUSIONS

This article analyzed the determinants of UI recipiency among those eligible for unemployment insurance benefits. Using a sample from the CPS DWS, it was found that raising the replacement rate increases the likelihood of UI receipt. There is some evidence, however, of a smaller effect at high replacement rates. Thus the cost impact of take-up responses to a uniform increase in weekly benefits is larger in low-replacement-rate states such as Indiana than in high-replacement-rate states such as New Jersey. Because the replacement rate of individuals eligible for the state's maximum benefit level tends to be low, the cost increase resulting from take-up responses to an increase in the state's maximum benefit level is a substantial fraction of the direct cost increase.

The cost of a change in benefit rules to an unemployment insurance system depends on several factors, including how such a change affects the lay-off policies of firms, the take-up rates among those who are laid off, and the expected duration of benefit receipt among those who receive UI benefits. This article focused on just one of these factors, take-up rates, which has been ignored for the most part. Although the effect of benefit changes on unemployment durations has been analyzed extensively by researchers of unemployment insurance, it may be useful in future empirical research to jointly model UI recipiency and unemployment durations, because unobserved determinants of take-up and unemployment durations may be correlated. Such an endeavor may lead to better predictions about the cost-related impact of benefit changes.

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