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Bonus Impacts on Receipt of Unemployment Insurance

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This chapter presents estimates of how a reemployment bonus affected three different outcomes of policy interest: weeks of unemployment insurance (UI) receipt during the benefit year, dollars of UI received during the benefit year, and the percentage of persons who exhaust their benefits. An assessment of how the reemployment bonus affected reemployment wages is made in Chapter 5.

To provide a deeper understanding of the reemployment bonus effects, this chapter also presents impact estimates based on slightly more involved computations. As shown in Chapter 2 in the discussion of the results of randomization (pp. 56–57), the characteristics of the treatment and control groups were quite similar in the Washington and Pennsylvania experiments; however, numerous statistically significant differences existed between the control and treatment groups in Illinois. Since the control and treatment groups were not completely homogenous, "regression adjustment" was done to account for differences between the treatments and controls in observable characteristics and improve the precision of treatment impact estimates.

Taken together, the Illinois, Pennsylvania, and Washington experiments examined a wide variety of bonus levels and qualification period lengths. In the third section of this chapter, we examine the marginal response to variations in these parameters. The analysis is done using a "continuous variables model" with the bonus amount and qualification period entered linearly.[1] Results are presented for the experiments separately and then for a pooled sample using evidence from all three.

To examine how the bonus offers changed behavior regarding the timing of unemployment, we present a hazard analysis of leaving insured unemployment. We then report treatment impacts in the three experiments for various subgroups: treatment impact estimates are presented by gender, age, race, previous industry, area unemployment rate, and weekly benefit entitlement. Some evidence from the Washington experiment about the effectiveness of the reemployment bonus for dislocated workers is also presented. The final section offers some conclusions.

Evidence from the three experiments suggests that a cash reemployment bonus offer can be expected to modestly shorten the average spell of insured unemployment, but the likely impacts seem to be too small for the bonus offer to be an optimum strategy for reducing UI costs. A full cost-benefit analysis investigating this question is provided in Chapter 7. No particular population subgroups had distinctly stronger reactions to the bonus than other groups; however, the bonus was more effective at times and places where the unemployment rate was lower.

TREATMENT IMPACTS ON INSURED UNEMPLOYMENT

The three experiments differed in the number and type of treatments examined and in the samples given a bonus offer. Figure 4.1 summarizes the characteristics of the treatment designs for the three experiments (which are more fully described in Chapter 2).

While there was only a single claimant treatment in the Illinois experiment—$500 for reemployment within 11 weeks on a job held 4 months—the availability of Federal Supplemental Compensation (FSC) was ended about halfway through the enrollment period. This resulted in a "natural experiment" in which 53 percent of the claimants offered a bonus in the Illinois experiment had a maximum entitled benefit duration of 38 weeks, while the others had a maximum duration of 26 weeks. According to Davidson and Woodbury (1991), the different entitled durations lead to significantly different response to the bonus offer. We present estimates separately for the different duration entitlement groups.

Figure 4.1 Design of the Reemployment Bonus Offers

A. Illinois Job Search Incentive Experiment

Bonus amount	Qualification period
\$500	11 weeks

B. Pennsylvania Reemployment Bonus Experiment

	Qualification period	
Bonus amount	6 weeks	12 weeks
3 × WBA	Low bonus/short qual.	Low bonus/long qual.
6 × WBA	High bonus/short qual.	High bonus/long qual.

C. Washington Reemployment Bonus Experiment

	Qualification period	
Bonus amount	(0.2 × potential UI duration) + 1 week	(0.4 × potential UI duration) + 1 week
2 × WBA	Low bonus/short qual.	Low bonus/long qual.
4 × WBA	Med. bonus/short qual.	Med. bonus/long qual.
6 × WBA	High bonus/short qual.	High bonus/long qual.

As described in Chapter 2, the four treatments in Pennsylvania involved either a low or high bonus amount and either a short or a long qualification period. The low bonus (set at three times the weekly benefit amount [WBA]) had a mean of \$502, while the high bonus (set at 6 × WBA) had a mean of \$1,000. The short qualification period was set at 6 weeks and the long qualification period was set at 12 weeks. With WBAs ranging from \$35 to \$260, bonus offers in Pennsylvania ranged from \$105 to \$1,560. For all treatments in Pennsylvania, the reemployment period was set at 16 weeks.

In Washington, there were three formulas for the bonus amount (low, medium, and high) and two formulas for the qualification period (short and long). The bonus amounts were figured as 2, 4, or 6 times the WBA, and the qualification periods were computed as either 20 percent or 40 percent of the entitled duration of benefits plus one week.[2] The mean low bonus amount offered was \$307, the mean medium amount was \$615, and the mean high amount was \$925. The mean short qualification period was 7 weeks, and the mean long qualification period was 12 weeks. The bonus offers in Washington ranged from $110 to $1,254. The reemployment period for all treatments was set at four months.

For the Pennsylvania and Washington experiments we present estimates of the impact of each separate offer. For each of those experiments, we give an estimate of the mean effect of the short-qualification-period offers, the long-qualification-period offers, and the different bonus level offers, as well as an estimate for each experiment of the overall mean impact of all treatments.

Unadjusted Treatment Impacts

For a classically designed experiment involving random assignment and large sample sizes, treatment impact estimates may be computed as the simple difference between treatment and control group means of outcome measures. Simple experimental impact estimates are not constrained by econometric models and may be very convincing and easy to understand if randomization was effective in creating homogenous control and treatment group samples. This simplicity is one of the fundamental appeals of experiments for program evaluation.

Table 4.1 presents unadjusted estimates of the response to the bonus offer. The overall mean estimates indicate that the bonus offer reduced weeks of UI benefit receipt by 1.15 weeks in Illinois, 0.62 weeks in Pennsylvania, and 0.34 weeks in Washington (Woodbury and Spiegelman 1987; Corson et al. 1992; O'Leary, Spiegelman, and Kline 1995). For Illinois and Pennsylvania, the null hypothesis of no impact can be rejected at the 95 percent confidence level. The estimates also show that the bonus offer reduced UI benefits received by $194 in Illinois, $81 in Pennsylvania, and $22 in Washington. The bonus offer also reduced the probability of UI benefit exhaustion by 3.3 percent in Illinois, 0.8 percent in Pennsylvania, and 1.0 percent in Washington.[3]

This pattern of results—bonus impacts in Illinois being markedly larger than those in Pennsylvania, and impacts in Pennsylvania being somewhat larger than those in Washington—holds up throughout the comparison of various measures of treatment impact across the three experiments. The relative impacts from the three experiments may be examined in comparison to the parameters of the bonus offers. The

Table 4.1 Unadjusted Differences between Experimental and Control Group Means in the Illinois, Pennsylvania, and Washington Reemployment Bonus Experiments[a]

	Insured weeks	UI compensation ($)	Exhaustion rate	Sample size[b]
ILLINOIS				
Control	20.1 (0.19)	2,786 (33)	0.472 (0.008)	3,952
Treatment	–1.15** (0.27)	–194** (46)	–0.033** (0.011)	4,186
Control—FSC-elig.	21.6 (0.30)	3,094 (51)	0.490 (0.011)	2,106
Treatment—FSC-elig.	–1.78** (0.41)	–316** (70)	–0.054** (0.015)	2,337
Control—FSC-inelig.	18.3 (0.23)	2,520 (41)	0.447 (0.012)	1,600
Treatment—FSC-inelig.	–0.71** (0.34)	–90 (59)	–0.015 (0.018)	1,589
PENNSYLVANIA				
Control	14.9 (0.18)	2,388 (36)	0.277 (0.008)	3,392
Low bonus/short qual.	–0.42 (0.34)	–26 (68)	0.012 (0.014)	1,395
Low bonus/long qual.	–0.41 (0.28)	–44 (56)	–0.001 (0.012)	2,456
High bonus/short qual.	–0.51* (0.31)	–66 (61)	–0.001 (0.013)	1,910
High bonus/long qual.	–0.95** (0.27)	–146** (54)	–0.026** (0.011)	3,073
Mean bonus/short qual.	–0.47 (0.26)	–49 (52)	0.004 (0.011)	3,305
Mean bonus/long qual.	–0.71** (0.23)	–100** (46)	–0.015 (0.010)	5,529
Low bonus/mean qual.	–0.41 (0.25)	–38 (50)	0.003 (0.011)	3,851
High bonus/mean qual.	–0.78** (0.24)	–115** (47)	–0.016* (0.010)	4,983
Mean bonus/mean qual.	–0.62** (0.22)	–81* (43)	–0.008 (0.009)	8,834

(continued)

Table 4.1 (continued)

	Insured weeks	UI compensation ($)	Exhaustion rate	Sample size[b]
WASHINGTON				
Control	14.3	2,066	0.239	3,082
	(0.19)	(34)	(0.008)	
Low bonus/short qual.	–0.05	30	0.011	2,246
	(0.30)	(52)	(0.012)	
Med. bonus/short qual.	–0.19	5	–0.004	2,348
	(0.30)	(51)	(0.012)	
High bonus/short qual.	–0.62*	–69	–0.012	1,583
	(0.33)	(58)	(0.013)	
Low bonus /long qual.	–0.50*	–58	–0.030**	2,387
	(0.29)	(51)	(0.011)	
Med. bonus/long qual.	–0.14	12	–0.014	2,353
	(0.30)	(51)	(0.012)	
High bonus/long qual.	–0.73**	–86	–0.023*	1,535
	(0.34)	(58)	(0.013)	
Low bonus/mean qual.	–0.28	–16	–0.010	4,633
	(0.25)	(43)	(0.010)	
Med. bonus/mean qual.	–0.16	9	–0.009	4,701
	(0.25)	(43)	(0.010)	
High bonus/mean qual.	–0.67**	–77	–0.018	3,118
	(0.27)	(47)	(0.011)	
Mean bonus/short qual.	–0.25	–5	–0.000	6,177
	(0.24)	(41)	(0.009)	
Mean bonus/long qual.	–0.42*	–39	–0.022**	6,275
	(0.24)	(41)	(0.009)	
Mean bonus/mean qual.	–0.34	–22	–0.011	12,452
	(0.22)	(38)	(0.008)	

[a] Standard errors in parentheses. * = Statistically significant at the 90% confidence level; ** = statistically significant at the 95% confidence level.

[b] The Illinois sample sizes for FSC-eligible and FSC-ineligible do not sum to the total Illinois sample size. This discrepancy is due to the FSC eligibility conditions, which differed depending on the date of the benefit claim, and our desire to use the largest possible sample in the full sample analysis. Further details are given in Chapter 3.

average bonus offer as a multiple of the average WBA was 3.6 in Illinois, 4.7 in Pennsylvania, and 3.8 in Washington. The average qualification period was 11.0 weeks in Illinois, 9.0 weeks in Pennsylvania, and 8.4 weeks in Washington. Pennsylvania had both an average bonus which was a higher multiple of the average WBA and a longer average qualification period than did Washington. Pennsylvania also had an average bonus which was a higher multiple of the average WBA than Illinois, but the average qualification period in Pennsylvania was two weeks shorter than the uniform 11-week qualification period in Illinois.

A variety of bonus levels and lengths of qualification periods were tested in Pennsylvania and Washington. The results found in each of these separate experiments tends to corroborate the conclusion which might be drawn from the above comparison of overall mean responses across experiments. For Pennsylvania, combining treatments with similar WBA multiples yields two groups that differ in the level of the bonus offer but have the same mean qualification period; a similar exercise yields three bonus level groups for Washington. For both Pennsylvania and Washington, the only combinations estimated with statistical significance are the high bonus treatments combined across qualification periods. For Pennsylvania, the high bonus offers had a mean impact of –0.71 weeks across qualification periods, while in Washington the mean impact of the high bonus treatments was –0.67. In both experiments, these impact estimates are greater than the point estimates for the lower bonus multiples. When the responses to the short and long qualification periods in Pennsylvania and Washington are compared, the estimated response to the long qualification period is statistically significantly greater than for the short qualification period. The largest and most statistically significant impact estimates in both Pennsylvania and Washington were for the high-bonus-multiple/long-qualification-period treatments, with the effect on UI compensation being –0.95 weeks in Pennsylvania and –0.73 weeks in Washington.

An interesting sub-analysis involves the natural experiment with varying entitled duration of benefits in Illinois. The average bonus response of –1.15 weeks of benefits was made up of a response of –1.78 weeks for those eligible for FSC (38 weeks of benefit entitlement) and –0.71 weeks for those not eligible (26 weeks of benefit entitlement). The average response of –0.71 for the FSC-ineligible sample in Illinois is close to the response observed in Pennsylvania and Washington, where the entitled duration of benefits was also similar.

When the Illinois FSC-ineligible results are compared with the findings in Pennsylvania and Washington, the patterns of bonus offer impacts on the UI exhaustion rate within and across the three experiments are also similar. The mean FSC-ineligible impact was –1.5 percent in Illinois, –0.8 percent in Pennsylvania, and in –1.0 in Washington. The overall impact on exhaustions in Illinois was –3.3 percent, with the impact for the FSC-eligible group in Illinois an astounding –5.4 percent. For the high bonus/long qualification treatments, the impact on exhaustion of benefits was –2.6 percent in Pennsylvania and –2.3 percent in Washington. In Illinois and for the high bonus/long qualification treatments in Pennsylvania and Washington, the reductions in exhaustion rate were statistically significant, indicating that some participants dramatically changed behavior by moving from benefit exhaustion to relatively short-term unemployment. The same general pattern of bonus offer impacts is observed on dollars of UI in the benefit year; however, the results in Pennsylvania and Washington were relatively weak compared with estimated impacts on the other outcomes.[4]

Adjusted Treatment Impacts

If treatment-control differences in outcome variables are due to factors other than the treatment, a simple comparison of means may not be adequate to identify treatment effects. As discussed in Chapter 2, for the Washington experiment there were no more differences between treatment and control groups in observable characteristics than would be expected to result from a random assignment process. Unfortunately, the variables on which there were the most pronounced differences—WBA and base period earnings (BPE)—may have affected the measurement of outcomes of interest.

An alternative procedure for computing the simple difference between treatment and control means on an outcome variable of interest that yields the same result involves estimating

$$Y = a + \mathbf{T}\boldsymbol{B} + u, \tag{4.1}$$

by ordinary least squares regression. In this equation the intercept, a, is the mean value of the outcome variable, Y, for the control group. **T** is a matrix of dummy variables representing the treatments, and u is a normally distributed mean zero error term. The parameter vector $\boldsymbol{B}$ contains estimates of the simple differences between treatment and control means on the outcome variable.

The model used to estimate treatment impacts while holding other factors constant is a straightforward generalization of Equation 4.1. The specification for computing adjusted treatment impacts involves adding terms for concomitant variables and is referred to as a covariance model. In the present case it takes the form

$$(4.2) \quad Y = a + \mathbf{T}\boldsymbol{B} + ZC + u,$$

where the introduction of concomitant variables, Z, into the model reduces the experimental error caused by differences in the observable characteristics between the control and treatment groups.[5] When estimating this model, each concomitant variable is included as a deviation from its own mean, thereby allowing interpretation of the intercept as the mean value of the outcome variable Y for the control group given the mean observable characteristics of the whole sample.[6] By this approach, the vector $\boldsymbol{B}$ yields estimates of treatment impacts adjusted for differences across observations in the characteristics Z.

Investigation into the causes of the lack of homogeneity in the observable characteristics of the experimental and control groups in the Washington sample indicated that treatment impacts estimated without adjusting for the heterogeneity of the groups are likely to be biased. Variations in the WBA accounted for most of the important heterogeneity across the Washington groups. In particular, it was found that omission of WBA when estimating treatment effects on dollars of UI compensation in Washington is likely to yield biased impact estimates. Nonetheless, to avoid any possibility of omitted variable bias in estimating adjusted treatment impacts, a full set of concomitant or adjustment variables is included as covariates in estimation of treatment effects for Illinois, Pennsylvania, and Washington.[7]

The introduction of adjustment variables as covariates reduced the standard errors on parameter estimates, with a greater effect on dollars of UI compensation than the other outcomes (Table 4.2). The mean

Table 4.2 Adjusted Differences between Experimental and Control Group Means in the Illinois, Pennsylvania, and Washington Reemployment Bonus Experiments[a]

	Benefit weeks	UI compensation ($)	Exhaustion rate	Sample size[b]
ILLINOIS				
Control	20.0 (0.19)	2,763 (28)	0.468 (0.008)	3,952
Treatment	–1.04** (0.26)	–150** (40)	–0.026** (0.011)	4,186
Control—FSC-elig.	21.45 (0.30)	2,996 (45)	0.485 (0.011)	2,106
Treatment—FSC-elig.	–1.46** (0.41)	–228** (62)	–0.042** (0.015)	2,337
Control—FSC-inelig.	18.18 (0.24)	2,458 (37)	0.446 (0.012)	1,600
Treatment—FSC-inelig.	–0.65* (0.33)	–57 (51)	–0.009 (0.017)	1,589
PENNSYLVANIA				
Control	14.9 (0.18)	2,400 (31)	0.274 (0.008)	3,358
Low bonus/short qual.	–0.62* (0.34)	–99* (58)	0.001 (0.014)	1,388
Low bonus/long qual.	–0.35 (0.28)	–67 (48)	0.000 (0.012)	2,432
High bonus/short qual.	–0.44 (0.30)	–99* (52)	0.001 (0.013)	1,890
High bonus/long qual.	–0.82** (0.26)	–133** (45)	–0.014 (0.011)	3,038
Mean bonus/short qual.	–0.53** (0.26)	–99** (45)	0.000 (0.011)	3,278
Mean bonus/long qual.	–0.62** (0.23)	–104** (39)	–0.008 (0.010)	5,470
Low bonus/mean qual.	–0.43* (0.25)	–75** (43)	0.002 (0.010)	3,820
High bonus/mean qual.	–0.69** (0.24)	–123** (40)	–0.009 (0.010)	4,928
Mean bonus/mean qual.	–0.58** (0.21)	–102** (37)	–0.004 (0.009)	8,748

(continued)

	Benefit weeks	UI compensation ($)	Exhaustion rate	Sample size[b]
WASHINGTON				
Control	14.35	2,099	0.241	3,082
	(0.19)	(30)	(0.007)	
Low bonus/short qual.	–0.05	22	0.008	2,246
	(0.29)	(46)	(0.011)	
Med. bonus/short qual.	–0.23	–28	–0.005	2,348
	(0.29)	(45)	(0.011)	
High bonus/short qual.	–0.72**	–117**	–0.017	1,583
	(0.32)	(51)	(0.013)	
Low bonus/long qual.	–0.58**	–112**	–0.032**	2,387
	(0.29)	(45)	(0.011)	
Med. bonus/long qual.	–0.28	–44	–0.019*	2,353
	(0.29)	(45)	(0.011)	
High bonus/long qual.	–0.75**	–135**	–0.020	1,535
	(0.33)	(52)	(0.013)	
Low bonu/mean qual.	–0.33	–47	–0.013	4,633
	(0.24)	(39)	(0.010)	
Med. bonus/mean qual.	–0.25	–36	–0.012	4,701
	(0.24)	(38)	(0.009)	
High bonus/mean qual.	–0.74**	–126**	–0.018*	3,118
	(0.27)	(42)	(0.010)	
Mean bonus/short qual.	–0.29	–33	–0.003	6,177
	(0.23)	(37)	(0.009)	
Mean bonus/long qual.	–0.51**	–92**	–0.024**	6,275
	(0.23)	(36)	(0.009)	
Mean bonus/mean qual.	–0.40*	–63*	–0.014*	12,452
	(0.21)	(33)	(0.008)	

[a] Standard errors in parentheses. * = Statistically significant at the 90% confidence level; ** = statistically significant at the 95% confidence level.

[b] The Illinois sample sizes for FSC-eligible and FSC-ineligible do not sum to the total Illinois sample size. This discrepancy is due to the FSC eligibility conditions, which differed depending on the date of the benefit claim, and our desire to use the largest possible sample size in the full sample analysis. Further details are given in Chapter 3.

regression-adjusted response across all treatments to a bonus offer was a reduction of \$150 in Illinois, \$102 in Pennsylvania, and \$63 in Washington. The Illinois and Pennsylvania estimates are significant at the 95 percent confidence level, while the Washington estimate is significant at the 90 percent confidence level. The adjusted impact estimates are somewhat larger for Pennsylvania and Washington and somewhat smaller in Illinois than the unadjusted estimates, but in no case is the adjusted estimate significantly different from the unadjusted estimate.

Regression adjustment of impact estimates has the greatest effect on the high bonus/long qualification treatments in Pennsylvania (–\$133) and Washington (–\$135). Both estimates are statistically significant and larger than the unadjusted estimates, but they are not significantly different from the unadjusted estimates.[8]

Regression-adjusted estimates of the bonus impact on weeks of UI and benefit exhaustion were only modestly changed from the unadjusted estimates. The overall regression-adjusted mean estimates indicate that the bonus offer reduced UI benefit receipt by 1.04 weeks in Illinois, 0.58 weeks in Pennsylvania, and 0.40 weeks in Washington. The mean regression-adjusted estimate of the reduction in the probability of UI benefit exhaustion was 2.6 percent in Illinois, 0.4 percent in Pennsylvania, and 1.4 percent in Washington.

For Washington, where lack of homogeneity was most severe, regression adjustment allowed detection of statistically significant impacts on compensation and weeks for three of the individual treatments at the 95 percent level of confidence, whereas not a single treatment impact on dollars of compensation was estimated with significance before regression adjustment.

A CONTINUOUS VARIABLES MODEL OF TREATMENT RESPONSE

In both Pennsylvania and Washington the bonus amount was defined in terms of the WBA. Since the WBA varies widely across claimants, the bonus amount may be used as a continuous variable when estimating bonus impacts. In Pennsylvania the qualification period was set at either 6 or 12 weeks, while in Washington it was defined as either 20 or 40 percent of the entitled duration of benefits plus 1 week. Therefore, the qualification period may also be used as a continuous variable. The variation in the bonus amount and qualification period in Pennsylvania and Washington is sufficient to allow estimation of the impact of incremental changes in the values of these parameters of a bonus offer. This section presents estimates from continuous variables models of bonus impacts for the separate Pennsylvania and Washington experiments and for these experiments combined with data from the Illinois experiment.

The continuous variables models represent a bonus offer as a given dollar amount paid for reemployment within a certain number of weeks. (These models are called continuous variables models because the treatments are represented by continuous variables rather than discrete indicator variables.) These models allow estimation of the effect of incremental changes in the bonus amount and the qualification period. The estimating equations have the following general linear specification:

(4.3) $Y = a + b_1B + b_2Q + ZC + u,$

where, B is the bonus amount in dollars, Q is the qualification period length in weeks, and Z is a set of concomitant or adjustment variables centered around their mean.[9] The parameters to be estimated are a, b_1, b_2, and C. The random error term, u, is assumed to be normally distributed with mean zero. In Equation 4.3, B and Q take on the value of zero for members of the control group. Including the WBA as a control variable improves the within-treatment homogeneity of B and results in the parameter b_1 being estimated using variation in B across the treatments. The same effect is achieved in estimating the parameter on Q by including the entitled duration of benefits. Parameter estimates for Equation 4.3 are presented in Table 4.3 for five samples. The dependent variable for each regression model is weeks compensated in the benefit year.

In the models estimated separately on the Pennsylvania and the Washington samples, none of the parameter estimates on B nor Q are statistically significant. For the sample formed by pooling data from the two experiments, however, the standard errors on the estimated parameters on B and Q are smaller and the parameter on the bonus amount, B, is estimated with statistical significance. The pattern and magnitude of results is consistent across the Pennsylvania and Wash-

Table 4.3 Continuous Variables Models of Bonus Impacts (UI Weeks)[a]

Sample[b]	$100 Increase in bonus amt.	One-week increase in qual. period	Sample size
Pennsylvania	–0.035 (0.030)	–0.028 (0.026)	2,106
Washington	–0.044 (0.031)	–0.023 (0.024)	15,534
Pennsylvania and Washington (pooled)	–0.039* (0.021)	–0.026 (0.018)	27,640
Pennsylvania, Washington, and Illinois (pooled)	–0.031 (0.021)	–0.047** (0.016)	35,778
Pennsylvania, Washington, and Illinois FSC-inelig. (pooled)	–0.027 (0.021)	–0.034** (0.017)	30,829

[a] Standard errors are in parentheses. * = Statistically significant at the 90% confidence level in a two-tailed test; ** = statistically significant at the 95% confidence level in a two-tailed test.

[b] The Pennsylvania estimates were computed in a regression model which also included cohort indicators, office indicators, and demographic and economic variables. The Washington estimates were computed in a regression model which also included the WBA and the entitled duration of benefits. The pooled sample estimates were estimated in regressions which included cohort indicators, office indicators, indicators for Washington, Pennsylvania, and Illinois FSC where possible, and demographic and economic variables.

ington samples; a higher bonus amount and a longer qualification period both reduce the weeks of UI drawn in the benefit year by similar magnitudes.

Evidence from the pooled Pennsylvania and Washington sample suggests that a $100 increase in a bonus offer will reduce weeks of UI compensation in the benefit year by an average of 0.039 weeks, and a 1-week increase in the qualification period will reduce weeks of UI compensation in the benefit year by 0.026. These parameters, estimated on the combined data, have sampling errors about 50 percent smaller than parameters estimated on the separate samples. Using results of the pooled estimation, response to the bonus offer in Pennsylvania and Washington may be summarized by saying that with a 12-week qualification period, it would take a bonus of nearly $1,800 to induce an average 1-week decline in insured unemployment among those offered a bonus. To appreciate the meaning of this, recall that about 12.9 percent of those offered a bonus in Pennsylvania and Washington actually were paid a bonus (Table 3.1 in Chapter 3).[10] This means that, on average, it cost about $232 in bonus payments to save a week of UI benefit payments, which averaged about $159.[11]

For Illinois, because there is no variation in either the bonus amount or the qualification period, the separate influence of these parameters may not be estimated using a sample with data only from Illinois. Nonetheless, these effects may be inferred by combining the Illinois data with that from the other experiments. Adding the Illinois data to the combined Pennsylvania and Washington sample yields a total sample of nearly 36,000 UI beneficiaries. As seen in Table 4.3, combining the Illinois data with that from Pennsylvania and Washington results in an estimated increased effect of the qualification period and a reduced effect of the dollar bonus amount. In a formal statistical sense, however, the parameter estimates on the dollar bonus amount and the qualification period are not affected by adding the Illinois data to the Washington and Pennsylvania data.[12]

It was previously noted that the Illinois beneficiaries who were FSC-ineligible, and therefore had an initial benefit entitlement of 26 weeks, responded to the bonus offer in much the same way as beneficiaries in Pennsylvania and Washington. When the FSC-eligible group is excluded from the three-state sample, the results of estimating the continuous variables model indicate that the qualification period has a diminished impact on benefit duration. This impact is more in line with the combined Pennsylvania and Washington impact; however, even with the FSC-eligible group excluded, the impact of the dollar bonus amount for the three-state sample remains somewhat smaller than for the Pennsylvania and Washington sample. Even though there appears to be some underlying differences between Illinois and the other two experiments, the results of estimating the continuous variables model on the combined sample excluding the Illinois FSC-eligible group probably provides the best summary of bonus response across the three experiments.

In Table 4.4 we present estimates of the reduction in weeks of UI receipt for four hypothetical bonus offers based on parameter estimates for the sample formed by pooling data from Pennsylvania and Washington together with the FSC-ineligible sample from Illinois. The fol-

Table 4.4 The Predicted Impacts[a] of Four Hypothetical Bonus Offers on UI Receipt, Based on a Continuous Variables Model Estimated on the Pooled Pennsylvania, Washington, and Illinois Data[b]

Hypothetical bonus offer	Amount of the bonus offer ($)	Bonus qual. period (weeks)	Impact on weeks of UI benefits
1	500	6	–0.34*** (0.09)
2	1,000	6	–0.47*** (0.16)
3	500	12	–0.54*** (0.15)
4	1,000	12	–0.68*** (0.17)

[a] Predicted impacts are based on the regression-adjusted impact estimates presented in Table 4.3 for the pooled Pennsylvania, Washington, and Illinois FSC-ineligible sample, which included 30,829 observations.
[b] Standard errors of predicted impacts in parentheses. *** = Statistically significant at the 99% confidence level in a two-tailed test.

lowing four hypothetical bonus offers are examined: 1) $500 for reemployment within 6 weeks, 2) $1000 for reemployment within 6 weeks, 3) $500 for reemployment within 12 weeks, and 4) $1000 for reemployment within 12 weeks. For wage and benefit levels prevailing during the 1980s, these four hypothetical bonus offers span the range of policy-relevant bonus options reasonably approximated by use of a linear model.

As shown in Table 4.4, weeks of UI benefit receipt are predicted to decline for each of the separate hypothetical bonus offers. Furthermore, the impact on UI receipt of a hypothetical bonus offer increases as the bonus amount or the duration of the bonus increases. All of the impact estimates computed using this simulation methodology are significant at the 99 percent confidence level. The least generous hypothetical bonus offer—$500 with a six-week qualification period—would reduce UI receipt by one-third of a week. Since the continuous variables model involves a linear relationship between the bonus parameters and UI receipt, doubling the bonus amount and duration simply doubles the impact of the bonus offer on UI receipt. As shown in Table 4.4, the offer with a $1,000 bonus amount and 12-week qualification period has twice the impact of the offer with a $500 bonus amount and 6-week qualification period. Hypothetical bonus offers 2 and 3, which combine a high dollar amount with a short qualification period or a low dollar amount with a long qualification period, would be likely to reduce UI receipt by a greater amount than the least generous offer but less than the most generous offer.

The offers reviewed in Table 4.4 are summarized graphically along with all other possible dollar bonus amounts for the 6- and 12-week qualification periods in Figure 4.2. The figure makes plain the linear nature of the impact response surface estimated. It is also clear from the diagram that, for any given dollar bonus amount, there will be a greater response the longer the qualification period and, for any given qualification period, there will be a greater response the higher the dollar bonus amount. The illustration does, however, mask the seemingly different way that Illinois beneficiaries respond to the parameters of the bonus in comparison with benefit recipients in Pennsylvania and Washington.[13]

THE TIMING OF TREATMENT IMPACTS

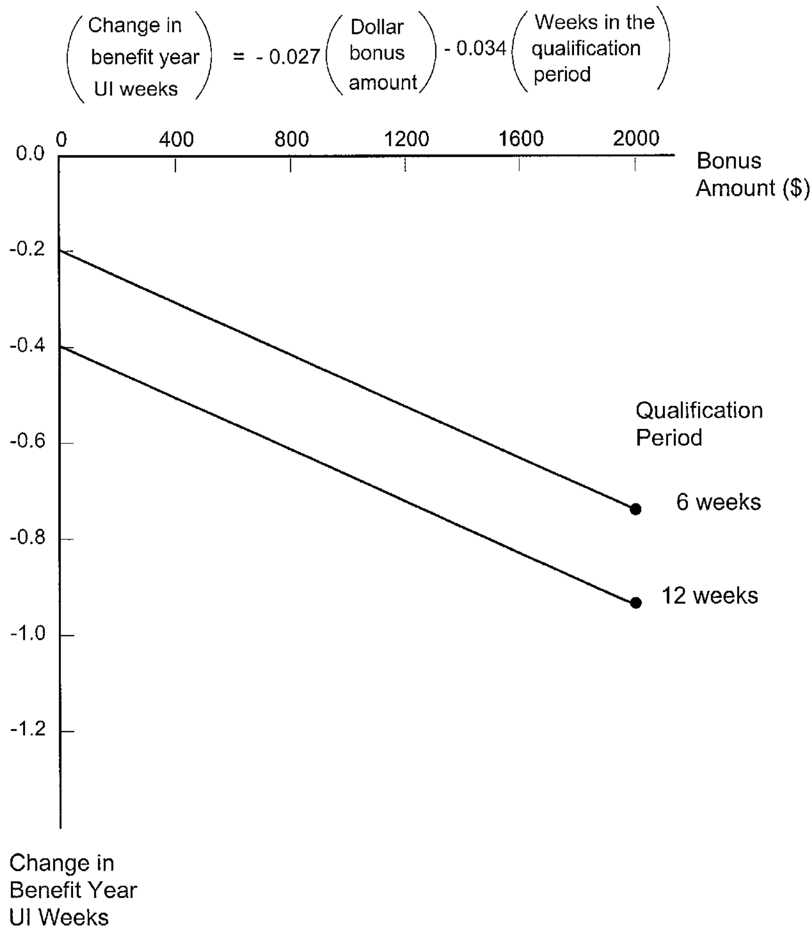
We now consider the impact of bonus offers on the time pattern of ending receipt of UI benefits. To understand this pattern, we examine the cumulative UI exit rate.[14] Since UI continued claim forms must be filed every two weeks by claimants wishing to continue benefit receipt, we analyze patterns over time measured in two-week intervals.

Estimates of cumulative UI exit rates through the end of any time period, t, are computed by dividing the difference between the initial risk pool of UI claimants and the remaining number of UI claimants at the start of time period $t+1$ by the initial risk pool of UI claimants. The algebraic formula for the computation is simply:

$$(4.4) \quad \mathrm{c}(t) = (R_0 - R_{t+1})/R_0$$

In this formula R_0 is the initial number of UI benefit recipients in the separate treatment and control groups, and R_{t+1} is the number of UI

Figure 4.2 Linearized Effect of the Bonus Offer on Benefit Year UI Weeks



recipients in these groups at the start of time period $t + 1$. Throughout our discussion of the timing of treatment response, we focus on the first spell of UI benefit receipt that followed the initial claim for UI benefits. In none of the three experiments did as many as 5 percent of initial claimants have multiple spells of UI benefit receipt during their benefit year.

Cumulative exit rates for treatment and control groups for each of the three experiments are reported in Table 4.5 observed at two-week intervals for unemployment spells between 1 and 39 weeks in duration. This number of periods is considered because the longest entitled duration of benefits among the experiments was 38 weeks for FSC-eligible claimants in Illinois.

To clearly understand the ideas involved consider an example from Table 4.5. The cumulative UI exit rate through the fifth week in the benefit year for control subjects in Pennsylvania was 0.403. This means that out of 8,834 initial UI recipients, 40.3 percent of the control group had stopped drawing UI benefits by the start of the sixth week of their claim. The comparable percentage among those offered a bonus in Pennsylvania was 42.4. The difference in these cumulative UI exit rates was found to be statistically significant.

There is a common pattern in the cumulative UI exit rates across the three experiments. Cumulative exit rates are higher among the treatment groups, at least through the first several weeks of UI receipt. However, the general level of UI exit rates is considerably lower in Illinois than in either Pennsylvania or Washington and tends to be slightly lower in Pennsylvania than in Washington.[15] Four factors might account for these differences: 1) differences in the labor market structures of the three states (that is, differences in the mix of industries and occupations), 2) differences in demand for labor among the three states, 3) differences among the three states in the administration of UI, and 4) differences among the three states in the characteristics of workers who are eligible for UI.

An effective bonus offer would lead claimants in treatment groups to have higher UI exit rates than controls during the period over which the bonus could be earned. Comparison of exit rates for the control and treatment groups in Illinois clearly shows that exit rates were higher during the bonus qualification period for UI claimants who were offered the bonus.

Table 4.5 Cumulative Hazard Rates for Leaving Unemployment Insurance

Illinois	Control	Treatment	FSC-eligible control	FSC-eligible treatment
Initial risk set	3,952	4,186	2,106	2,337
Spell length (weeks)				
0–1	0.127	0.150**	0.138	0.154
2–3	0.197	0.230**	0.215	0.249**
4–5	0.252	0.287**	0.271	0.307**
6–7	0.298	0.343**	0.322	0.370**
8–9	0.339	0.389**	0.364	0.417**
10–11	0.372	0.427**	0.396	0.458**
12–13	0.406	0.460**	0.422	0.484**
14–15	0.432	0.488**	0.443	0.509**
16–17	0.457	0.513**	0.466	0.531**
18–19	0.483	0.537**	0.487	0.548**
20–21	0.515	0.559**	0.508	0.569**
22–23	0.541	0.584**	0.528	0.590**
24–25	0.578	0.616**	0.556	0.615**
26–27	0.816	0.833**	0.661	0.709**
28–29	0.825	0.843**	0.674	0.724**
30–31	0.835	0.853**	0.690	0.740**
32–33	0.841	0.859**	0.702	0.749**
34–35	0.851	0.869**	0.720	0.767**
36–37	0.863	0.877*	0.744	0.780**
38–39	0.998	0.997	0.996	0.995

Pennsylvania	Control	Mean treatment	High bonus/ long qual.
Initial risk set	3,392	8,834	3,073
Spell length (weeks)			
0–1	0.258	0.271	0.273
2–3	0.343	0.355	0.357
4–5	0.403	0.424**	0.428**
6–7	0.456	0.486**	0.487**
8–9	0.515	0.538**	0.542**
10–11	0.564	0.587**	0.591**
12–13	0.608	0.635**	0.649**
14–15	0.647	0.671**	0.681**
16–17	0.681	0.701**	0.713**
18–19	0.708	0.726*	0.740**
20–21	0.734	0.750*	0.763**
22–23	0.759	0.773*	0.786**
24–25	0.787	0.797	0.810**
26–27	0.986	0.989*	0.991*
28–29	0.993	0.993	0.993
30–31	0.994	0.994	0.994
32–33	0.995	0.995	0.995
34–35	0.995	0.996	0.996
36–37	0.996	0.996	0.997
38–39	0.997	0.996	0.997

(continued)

Table 4.5 (continued)

Washington	Control	Mean treatment	High bonus/ long qual.
Initial risk set	2,702	11,052	1,358
Spell length (weeks)			
0–1	0.206	0.196	0.221
2–3	0.333	0.337	0.364*
4–5	0.431	0.437	0.464**
6–7	0.506	0.508	0.526
8–9	0.564	0.563	0.588
10–11	0.611	0.613	0.653**
12–13	0.644	0.650	0.684**
14–15	0.679	0.681	0.720**
16–17	0.715	0.720	0.758**
18–19	0.748	0.756	0.781**
20–21	0.784	0.792	0.816**
22–23	0.813	0.821	0.838*
24–25	0.838	0.850	0.862**
26–27	0.862	0.873	0.888**
28–29	0.885	0.899**	0.909**
30–31	0.982	0.982	0.980
32–33	0.989	0.989	0.987
34–35	0.993	0.991	0.992
36–37	0.994	0.993	0.993
38–39	0.995	0.994	0.994

[a] * = Cumulative hazard rate significantly different from control group in a two-tailed test at the 90% confidence level; ** = cumulative hazard rate significantly different from control group in a two-tailed test at the 95% confidence level.

For the Pennsylvania experiment, the cumulative UI exit rate for the treatment groups taken as a whole was higher than for the control group through most of the bonus qualification period. Compared with Illinois, cumulative impacts appeared to be smaller, but were nevertheless big enough to produce a 0.62-week reduction in insured unemployment (as reported in Table 4.1). The cumulative impacts for the high bonus/long qualification treatment in Pennsylvania were modestly larger over the first 14 weeks following the start of UI benefits (Table 4.5, Pennsylvania, high bonus/long qualification).

In the Washington experiment, there is scant statistical evidence that bonus-offered workers gave up their UI benefits at a faster pace than did the control group. When all treatments are taken together, UI exit rates were higher for bonus-offered workers than for controls only by the week 28–29 period. This is consistent with the small (and statistically insignificant) 0.3-week reduction in insured unemployment that can be seen in Table 4.1 for the mean bonus offer. Isolating the most effective treatment in Washington, the high-bonus/long-qualification treatment (Table 4.5) there is some evidence that UI exit rates of the bonus-offered workers were higher than the exit rates of the control group. These results are consistent with the somewhat larger reduction in insured unemployment for this group (0.73 week) that can be seen in Table 4.1.

The basic conclusion to be drawn from the analysis of cumulative exit rates is that the Illinois and Pennsylvania bonus offers taken as a whole both acted to increase UI exit rates during their respective bonus qualification periods. There is also some evidence that the high bonus offers in the Washington experiment increased UI exit rates. These time patterns of cumulative exit rates for treatment and control subjects are presented graphically in Figure 4.3. Treatment impacts on cumulative exit rates are presented graphically in Figure 4.4, where the estimated impact is equal to the difference between the cumulative exit rate for the treatment group and the cumulative exit rate for the control group.

Cumulative UI exit rates are also useful for examining the extent to which the workers who took advantage of bonus offers were individuals who otherwise would have exhausted their benefits. For potential benefit exhausters, each bonus offer accepted would generate a large reduction in observed unemployment. In Illinois, for example, the reduction would be from 27 weeks of insured unemployment to at most 11 weeks. On the other hand, for workers who would have otherwise had relatively short durations of UI benefit receipt, each bonus offer accepted might generate only a relatively small reduction in observed unemployment.

After 11 weeks, 37.2 percent of the claimants in the Illinois control group had stopped receiving UI benefits, whereas 42.7 percent of the Illinois treatment group had stopped receiving benefits (Table 4.5). In other words, at the end of the bonus qualification period, an additional 5.5 percentage points (42.7 – 37.2) of the bonus-offered workers had ended their UI receipt as compared with the controls. Moreover, this 5.5 percentage point difference diminishes only slightly after the end of the bonus qualification period. Even immediately before exhaustion of regular UI benefits (weeks 24–25), the difference is nearly 4 percentage points. In other words, the control group fails to catch up with the treatment group in terms of exit from UI. This failure of the control and treatment cumulative exit rates to converge suggests that a large proportion of the workers who responded to the Illinois bonus offer would have had long spells of insured unemployment in the absence of the bonus offer. These results are consistent with estimates in Table 4.1, which show that the Illinois bonus reduced the UI benefit exhaustion rate by over 3 percentage points overall (and by over 5 percentage points among FSC-eligibles). Figures 4.3 and 4.4 illustrate that cumulative UI exit rates for treatment subjects remain above controls at 26 weeks and at 38 weeks for the FSC-eligible group. We conclude that the Illinois bonus reduced the duration of UI spells among workers who tended to have relatively long expected durations of UI receipt.

The evidence from Pennsylvania exhibits a different pattern. The difference between the control and combined treatment cumulative exit rates peaks at between 2 and 3 percentage points during weeks 6–7 and 12–13 (Table 4.5 and Figures 4.3 and 4.4). These peaks in cumulative UI exit coincided with the end of the short (6-week) and the long (12-week) bonus qualification periods. For the high bonus/long qualification treatment alone, the difference is larger, peaking at about 4 percentage points in weeks 12–13. Unlike for Illinois, however, the control and treatment cumulative exit rates do converge near the end of the qualification period. In terms of leaving UI, the control group caught up with the combined treatment groups by weeks 24–25 (the

Figure 4.3 Cumulative UI Exit Rates by Treatment and Control Groups

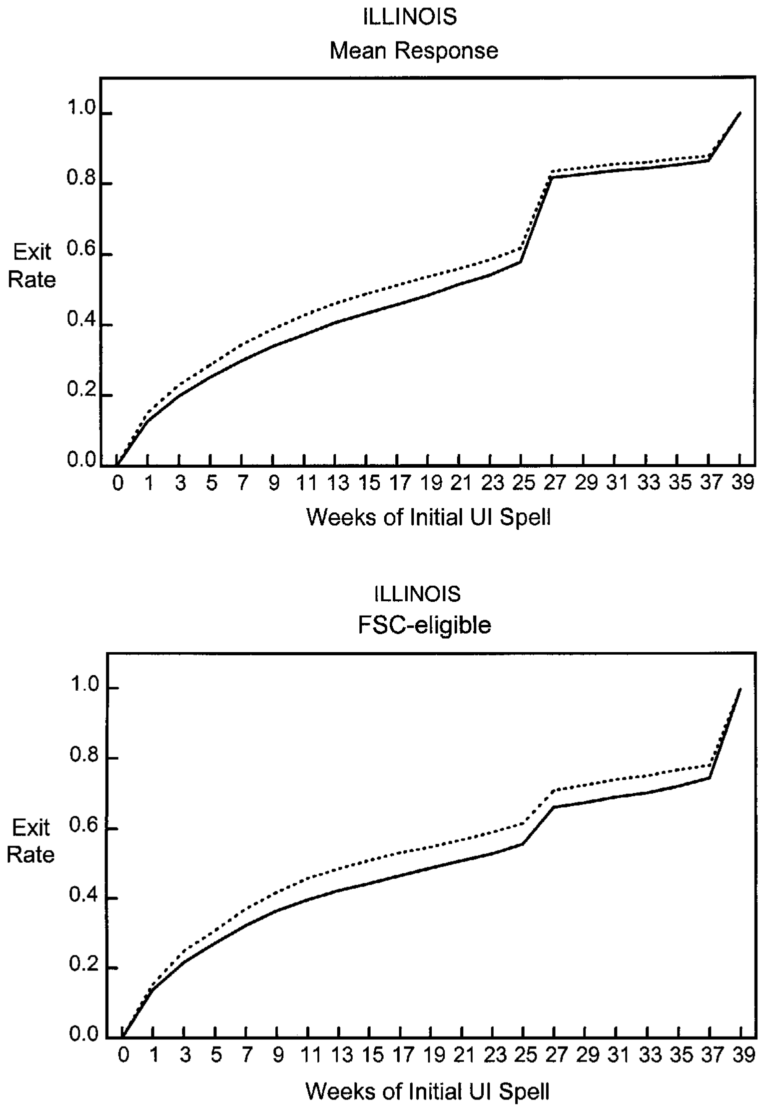
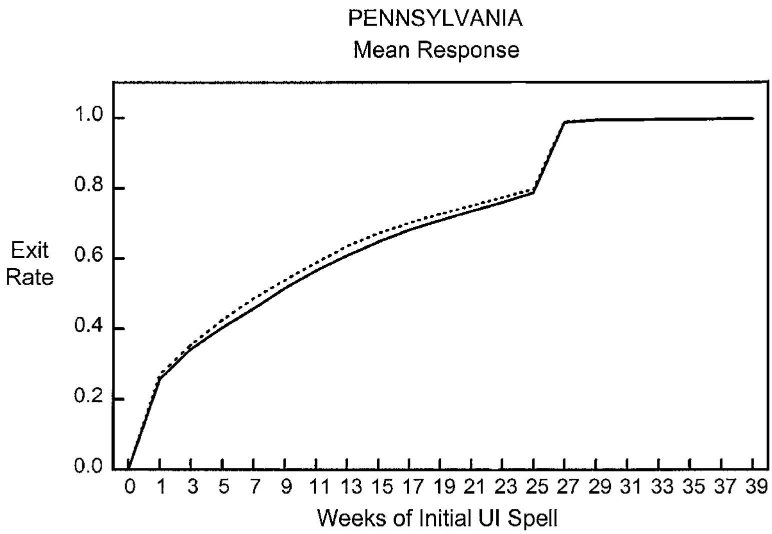
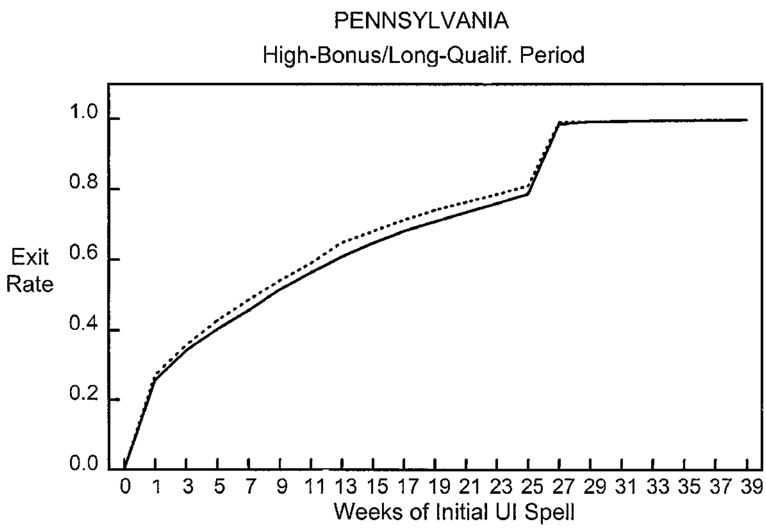


Figure 4.3 (cont.)





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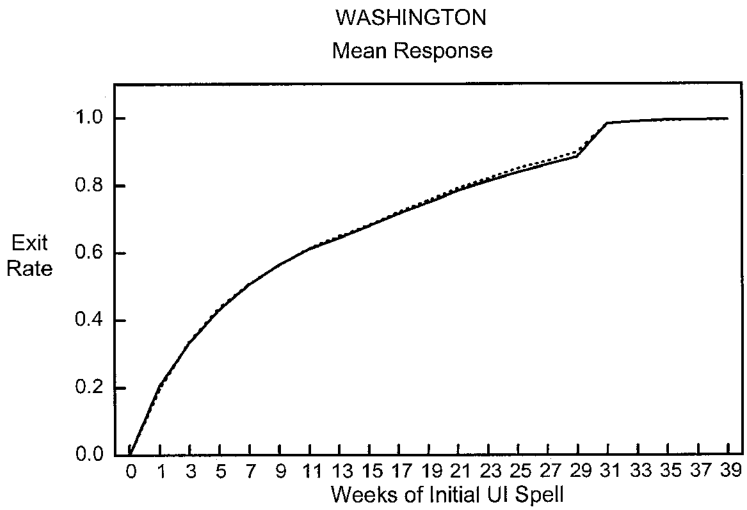


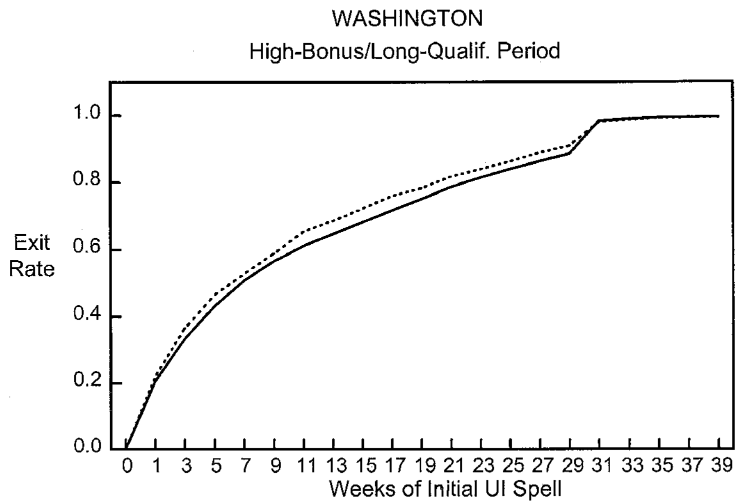


Figure 4.4 Simple Treatment Impacts on Cumulative UI Exit Rates

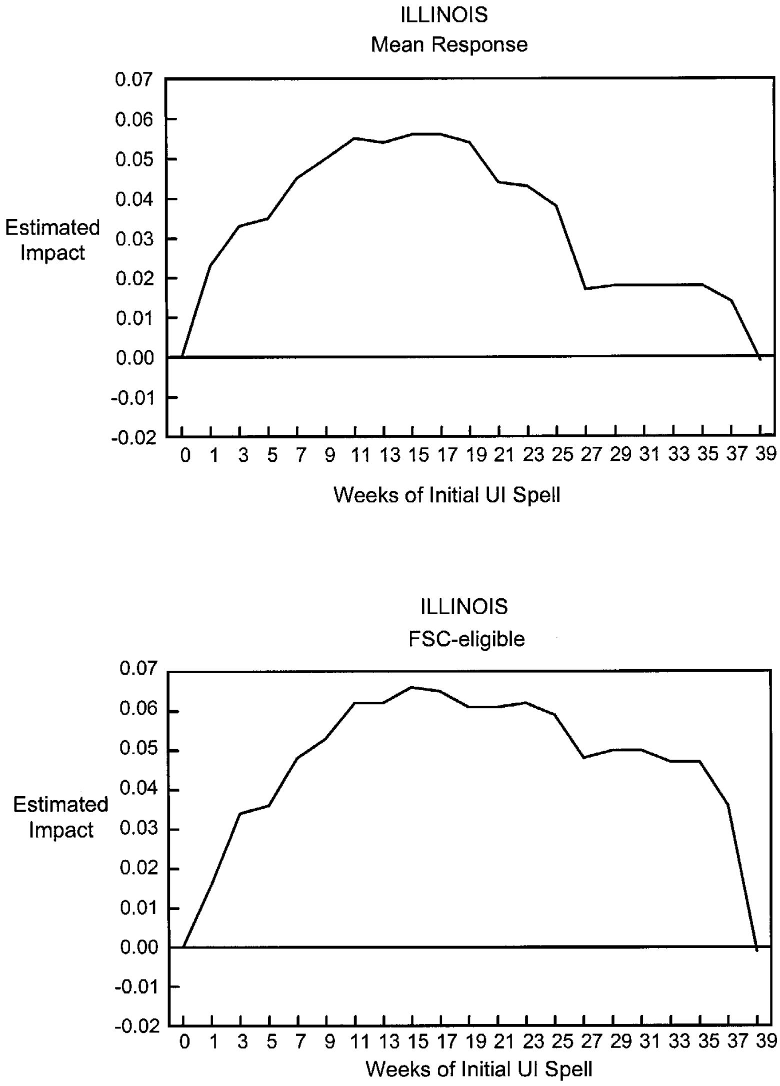
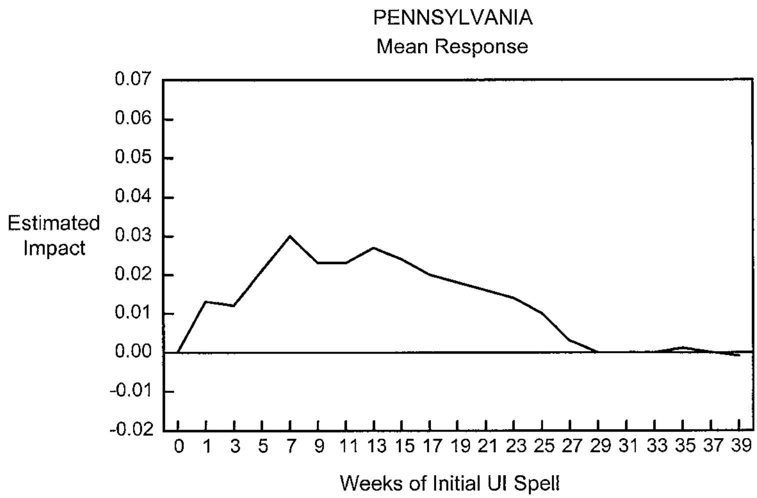


Figure 4.4 (cont.)



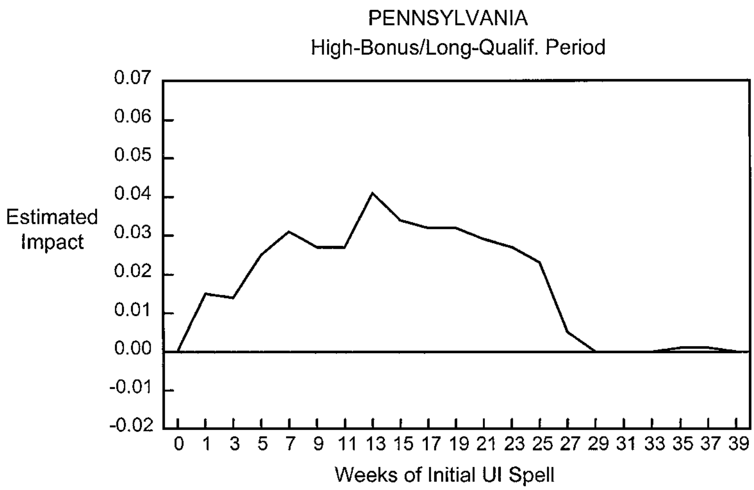
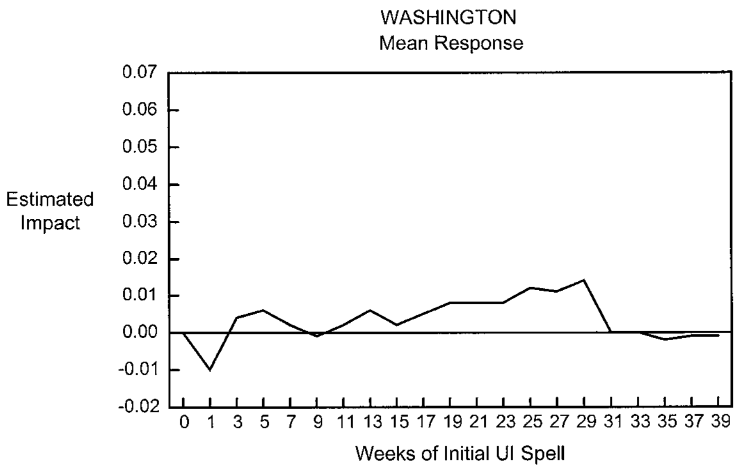
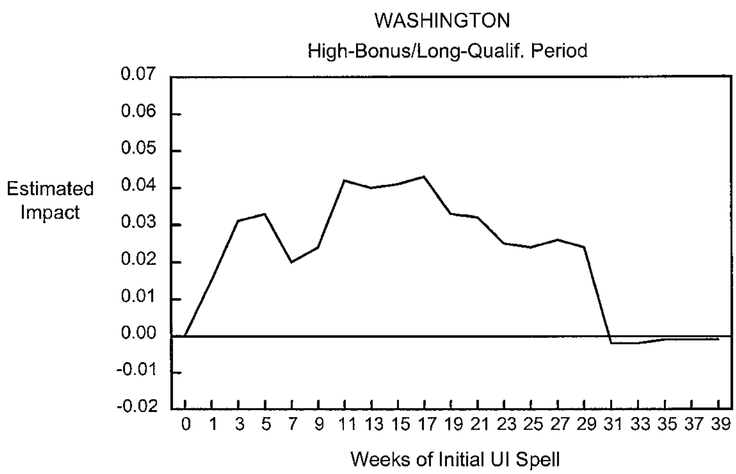


Figure 4.4 (cont.)





time period before exhaustion), but not until weeks 28–29 (just after benefit exhaustion) for the high-bonus/long-qualification treatment in Pennsylvania.

The decaying impact of the Pennsylvania treatments on the cumulative UI exit rate shows that much of the bonus-induced reduction in UI receipt occurred because the bonus effect was concentrated among claimants who would have had relatively short spells of UI receipt. The exception is the high bonus/long qualification treatment in Pennsylvania, which reduced the probability of a worker exhausting UI benefits by 2.6 percent as reported in Table 4.1. The point can also be seen in Table 4.5, which shows that in weeks 24–25 there is still a 2.3 percentage point difference between the cumulative exit rates of the control group and the high bonus/long qualification treatment group. Overall, though, the Pennsylvania treatments appear to have affected the behavior mainly of workers with relatively low expected durations of UI receipt.

The cumulative UI exit rate results for Washington show a weaker, but similar, pattern to that seen in Pennsylvania. For all treatments combined, the bonus initially creates an insignificant difference between the cumulative UI exit rates of the control and the experimental groups which diminishes before weeks 28–29 (the period before exhaustion). The exception is the high bonus/long qualification treatment, for which the cumulative UI exit rate remains 2.4 percentage points greater than controls in weeks 28–29 (Table 4.5). This is consistent with the significant 2.3 percentage point reduction in UI exhaustion reported in Table 4.1 for those with the high bonus/long qualification offer. The evidence indicates that the average Washington bonus offer reduced the UI receipt of workers who would have experienced relatively short spells of UI receipt in any event. The exception presented for Washington is the high bonus/long qualification offer, which reduced benefit exhaustion.

IMPACTS OF THE BONUS OFFER ON POPULATION SUBGROUPS

There are two main reasons for examining treatment impacts by population subgroup. One is to provide information to policymakers who may consider targeting a reemployment bonus program to certain groups, such as dislocated workers or older UI claimants. Another is to identify possible biases in the effects—a program that benefits only one gender or certain ethnic groups may not be considered good policy even if it is cost-effective.

This section reports on treatment impacts for 12 subgroups defined by binary variables for six characteristics: gender, age, race, industry, area unemployment, and the WBA level. The dummy variables specified for the analysis were a variable indicating whether a claimant was female; an age variable indicating whether the claimant was 35 years old or over; a race variable indicating whether a claimant was black; an industry variable indicating whether the claimant's previous job was in manufacturing; an unemployment variable indicating whether the local unemployment rate was below 5 percent; and a variable indicating whether the claimant qualified for the maximum WBA.

It is standard practice to examine program impacts by gender, age, and race. We investigated differential treatment impacts by industry, area unemployment, and WBA for various reasons. Because manufacturing is a shrinking sector, workers leaving manufacturing are more likely than others to be displaced workers. Accordingly, examining impacts for manufacturing workers separately from others may shed light on the value of the reemployment bonus as a way of aiding displaced workers. Because the reemployment bonus could be implemented either as an ongoing or as a countercyclical program, it was important to examine whether the bonus had an impact only when reemployment prospects were good. Finally, existing evidence suggests that the disincentive effects of UI are greatest for workers whose prior earnings were relatively low; these claimants who typically have WBAs below the state maximum receive benefits which tend to replace a larger share of lost earnings. Accordingly, it seemed important to examine whether workers below the maximum WBA had greater treatment impacts than other workers. If so, then the reemployment bonus might be an effective way of mitigating the disincentive effects of the UI system for workers who are most adversely affected.

We estimate all subgroup treatment impacts in a single regression model. This means that the treatment response for each subgroup is estimated controlling for the influence of all other subgroup characteristics. For example, the model estimates the treatment impacts associated with being black controlling for the fact that blacks are less likely to be at the maximum WBA, less likely to claim benefits in areas with unemployment rates below five percent, and so on. If we did not proceed in this way, any differential impact associated with being black might be a result not of being black but rather of other characteristics that blacks possess.

The equation estimated for each outcome of interest is a straightforward generalization of Equation 4.2:

$$(4.5) \quad Y = a + \mathbf{T}\boldsymbol{B} + \mathbf{Z}\boldsymbol{C} + \mathbf{G}\boldsymbol{D} + \mathbf{GT}\boldsymbol{E}' + \mathbf{GZ}\boldsymbol{F}' + \mathbf{ZT}\boldsymbol{H}' + u$$

where Y is the outcome measure (either weeks of insured unemployment, dollars of UI compensation, or the UI benefit exhaustion rate); **T** is the matrix of treatment dummies; **Z** is a matrix of concomitant variables in deviation form; **G** is the matrix of dummy variables which code for membership in a subgroup; a is the intercept; ***B***, ***C***, ***D***, ***E***, ***F***, and ***H*** are conformable parameter vectors, and u is a normally distributed random error term with mean of zero. Note that Equation 4.5 is written to include concomitant variables, **Z**, that may have been correlated with assignment to treatment and control groups. In the models estimated for all three experiments, the concomitant variables, **Z**, were entered linearly and interacted with all subgroup dummies.[16]

Equation 4.5 specifies a complete one-way interaction model—that is, all first-order products of subgroup dummy variables, treatment indicators, and concomitant variables with each other are included. This model allows simultaneous estimation of all subgroup treatment impacts, but imposes linear restrictions on those estimates.[17]

Table 4.6 displays the estimated bonus impacts on weeks of insured unemployment, dollars of UI benefits received, and the UI exhaustion rate for each of the 12 subgroups. Inspection of these results reveals one main finding: there is virtually no difference between any pair of subgroups shown that is both statistically signifi-

Table 4.6 Impacts of the Treatments on UI Receipt by Subgroup[a]

	Illinois				Pennsylvania				Washington			
Subgroup	Benefit weeks	$ of UI compensation	Exhaust. rate	Sample size	Benefit weeks	$ of UI compensation	Exhaust. rate	Sample size	Benefit weeks	$ of UI compensation	Exhaust. rate	Sample size
Males	–0.94**	–135**	–0.029*	4,519	–0.40	–90*	0.001	7,237	–0.69**	–98**	–0.012	9,471
	(0.36)	(55)	(0.015)		(0.29)	(49)	(0.012)		(0.28)	(47)	(0.011)	
Females	–1.37**	–207**	–0.029*	3,619	–0.91**	–123**	–0.015	4,845	0.06	4	–0.018	6,063
	(0.40)	(61)	(0.017)		(0.36)	(61)	(0.015)		(0.35)	(51)	(0.014)	
Aged <35 yr.	–1.18**	–177**	–0.031**	5,045	–0.87**	–136**	–0.010	6,682	–0.29	–30	–0.012	8,169
	(0.34)	(51)	(0.014)		(0.30)	(51)	(0.012)		(0.30)	(47)	(0.012)	
Age 35 yr. and over	–1.04**	–151**	–0.026	3,093	–0.29	–63	0.000	5,544	–0.52*	–82*	–0.017	7,365
	(0.44)	(66)	(0.018)		(0.32)	(55)	(0.014)		(0.31)	(49)	(0.012)	
Black	–0.89*	–197**	–0.017	2,122	–1.37**	–202*	–0.015	1,412	–0.48	45	–0.011	695
	(0.53)	(80)	(0.022)		(0.63)	(108)	(0.026)		(1.01)	(160)	(0.040)	
Non-black	–1.21**	–156**	–0.033**	6,016	–0.50**	–90**	–0.004	10,670	–0.39*	–59*	–0.015*	14,839
	(0.31)	(47)	(0.013)		(0.23)	(39)	(0.009)		(0.21)	(34)	(0.008)	
Manufacturing	–0.74	–134*	–0.015	2,084	–1.81**##	–257**#	–0.027	3,111	–0.71	–104	0.000	3,505
	(0.53)	(81)	(0.022)		(0.43)	(74)	(0.018)		(0.45)	(71)	(0.016)	
Nonmanufacturing	–1.26**	–178**	–0.034**	6,054	–0.19	–50	0.002	8,971	–0.31	–40	–0.019**	12,029
	(0.31)	(47)	(0.013)		(0.25)	(43)	(0.010)		(0.24)	(38)	(0.009)	
Low unemployment	–1.33**	–209**	–0.040	1,459	–1.05**#	–167**	–0.015	5,332	–1.04**	–203**##	–0.048**##	5,328
	(0.63)	(96)	(0.026)		(0.33)	(56)	(0.014)		(0.37)	(58)	(0.015)	
Not low unemployment	–1.08**	–158**	–0.027**	6,679	–0.25	–52	0.002	6,750	–0.06	23	0.003	10,206
	(0.29)	(44)	(0.012)		(0.29)	(50)	(0.012)		(0.26)	(41)	(0.010)	

Maximum WBA	–1.34**	–236**	–0.032	2,780	–1.03	–195	–0.010	2,377	0.99*	155*	0.001	5,252
	(0.65)	(99)	(0.027)		(0.70)	(120)	(0.029)		(0.53)	(83)	(0.021)	
Not maximum WBA	–1.02**	–132**	–0.027*	5,358	–0.50*	–81*	–0.004	9,849	1.11**	–161**##	–0.022*	10,282
	(0.40)	(62)	(0.017)		(0.27)	(46)	(0.011)		(0.33)	(51)	(0.013)	

[a] Standard errors in parentheses. * = Statistically significant at the 90% confidence level for a two-tailed test; ** = statistically significant at the 95% confidence level for a two-tailed test; # = statistically different from the complementary subgroup at the 90% confidence level for a two-tailed test; ## = statistically different from the complementary subgroup at the 95 percent confidence level for a two-tailed test.

cant at conventional confidence levels and consistent across the three experiments. The implication of this finding is quite striking. The reemployment bonus has a remarkably even impact on various subgroups of workers, whether delineated by gender, age, race, sector of employment, local unemployment rate, or maximum WBA eligibility.

The evidence suggests some differences that may be worth noting, however. For example, the evidence from Illinois and Pennsylvania (but not from Washington) suggests that women may be slightly more responsive than men to the reemployment bonus. This is consistent with research on female labor supply, which has found higher labor supply elasticities for women than for men, meaning that women's attachment to the labor force is more flexible than men's.

Results from Illinois and Pennsylvania (but again not from Washington) also suggest that younger workers may be somewhat more responsive than older workers to the reemployment bonus. An explanation for younger workers' relatively larger response is similar to that given for women—younger workers' labor force attachment is more flexible than older workers'.

There is no evidence to suggest that the reemployment bonus had a different impact on African Americans compared to other racial or ethnic subgroups. Impact estimation for each experiment was also done using a further disaggregation by race and ethnicity. The results suggested that the response of blacks differed from the overall mean response to a bonus offer by more than any other racial or ethnic subgroup. Yet, there were no significant differences between any of the more finely defined groups. In short, the impact of the reemployment bonuses did not vary by race or ethnicity.

In Pennsylvania and Washington, claimants whose previous jobs were in manufacturing had a stronger response to the bonus offer than workers whose previous job was not in manufacturing, with the difference in impact across groups being statistically significant in Pennsylvania. It may be the case that claimants who had lost a job in manufacturing tended to wait longer for recall, so that the bonus has more possibility to alter search behavior. However, this result is not obtained in Illinois, where the impact was even slightly greater for workers whose previous jobs were outside of manufacturing. Overall, there is no clear evidence for differential impacts across these subgroups.

The results from both the Pennsylvania and Washington experiments (although not from Illinois) suggest that UI claimants in low unemployment areas tended to respond more strongly to the bonus than did claimants in high unemployment areas. We have defined *low unemployment* as a 5 percent rate of total unemployment or lower, since this is at or below the rate previously considered to be associated with nonaccelerating inflation.[18]

That the bonus response appears to be greater in areas of low unemployment suggests that the bonus offer is more effective when increased search effort is more likely to generate additional job offers. This is an issue of labor demand and appeals to the notion that, in a labor market in which job prospects are neither good nor improving, the bonus impact can be expected to be low. That the bonus was effective in Illinois, where the unemployment rate was relatively high on average, may be a result of the improving labor market in northern Illinois during the experiment. We return to this issue in Chapter 6, which addresses the problems of interpreting the implications of experimental results for policy implementation.

It was expected that because the earnings replacement rate is relatively higher for those below the WBA maximum, the initial intensity of job search would be correspondingly low, and the reemployment bonus would have a greater impact. There is no evidence from the three experiments that claimants below the maximum WBA responded more strongly to the bonus than did claimants at the maximum WBA. In fact, in the Pennsylvania experiment (where the greatest difference in impact between the two groups occurred), those at the maximum WBA reduced weeks of insured UI by twice as much as those below the maximum.

SUMMARY

Field experiments were conducted in Illinois, Pennsylvania, and Washington in the 1980s to test the theory that the average duration of insured unemployment could be shortened if cash reemployment bonuses were offered to unemployment insurance beneficiaries. Claimants who were randomly assigned to treatment were told they would be given a lump sum cash payment if they started a new job by a certain date and stayed working full time for at least four months. In each of the three experiments, a randomly selected control group was compared with the various treatment groups.

While the first experiment, conducted in Illinois, yielded very encouraging results, the subsequent field tests provided less support for the bonus idea. The relatively weak response to the bonus offer in Pennsylvania and Washington led to a reexamination of the powerful Illinois results. It was discovered that within the designed experiment, a second experiment had unintentionally taken place. In 1984 as Illinois was recovering from a major recession, the availability of Federal Supplemental Compensation was terminated. This resulted in about half of the claimants studied having 38 weeks of UI benefit eligibility, with the remainder being eligible for only 26 weeks of regular UI benefits. It turns out that the mean bonus response of –1.15 weeks in Illinois was made up of a response of –1.78 weeks for those FSC-eligible and –0.54 weeks for those FSC-ineligible. The average response of –0.54 for the FSC-ineligible sample in Illinois is close to the response observed in Pennsylvania and Washington where the entitled duration of benefits was also similar.

Among the individual treatments, the impact on weeks of UI benefits ranged from –0.05 for the low bonus/short qualification offer in Washington to –1.78 for the bonus offer to FSC-eligible claimants in Illinois. Impacts for Pennsylvania tended to fall between those for Illinois and Washington. Overall a cash bonus can be expected to modestly shorten spells of insured unemployment—the mean effect of the offers made in the three states yielded about a one-half week reduction in weeks of UI benefits.

Evidence from a sample created by pooling data from Illinois, Pennsylvania, and Washington indicated that either lengthening the duration of the search period or increasing the bonus amount would lead to a greater reduction in weeks of UI benefits collected. Simulations based on data pooled from the three experiments suggest that offering a reemployment bonus with a $1,000 dollar amount, a 12-week qualification period, and a four-month reemployment period would result in an average reduction of 0.66 weeks of UI benefits.

The degree of response to the bonus offer was also examined for important subgroups within the sample. Results from Pennsylvania and Washington suggest that UI claimants in low unemployment areas and claimants whose prior employment was in manufacturing tended to respond more strongly to the bonus. However, close inspection of subgroup results reveals one main finding: there is no difference between any pair of subgroups shown that is both statistically significant at conventional confidence levels and consistent across the three experiments. The implication of this finding is quite striking—the reemployment bonus has a remarkably even impact on various subgroups of workers, whether delineated by gender, age, race, industrial sector of employment, level of local unemployment, or level of the weekly benefit amount.

Notes

1. Decker and O'Leary (1992) used the continuous variables model to analyze pooled Pennsylvania and Washington data. A summary article on the pooled analysis is Decker and O'Leary (1995).
2. One week was added in the computation of the qualification period in Washington to adjust for the presence of a waiting week. The *waiting week* is the first week after filing for benefits that a claimant would otherwise be eligible for compensation. It does not reduce total entitlement but does postpone payment, thereby discouraging casual entry into the system. The Washington experiment added a week to the qualification period because the treatments were intended to have qualification periods which were 20 and 40 percent of the compensable period.
3. The New Jersey UI reemployment experiment tested a declining bonus. A similar treatment was tried as part of the Pennsylvania experiment, but it is not reported in Tables 4.1 and 4.2 because it is not comparable to other treatments in Illinois, Pennsylvania, and Washington. The impact of the Pennsylvania declining bonus offer, which started at $6 \times$ WBA and declined by 10 percent of the original amount per week, had insignificant point estimate of a 0.2 weeks reduction in UI benefits.
4. See Chapter 6, p. 199, for an explanation of why the FSC-eligible and FSC-ineligible responses differ.
5. Netter and Wasserman (1974, Chapter 22) give a concise discussion of covariance analysis.
6. The effect of centering around the mean is that the intercept takes the value of the outcome measure for a hypothetical person in the sample who was not exposed to the experimental treatment and whose exogenous characteristics are at the mean value for each of the characteristics across the total sample (control and experimental groups combined). Therefore, the control group means reported in Table 4.2 are slightly different from the unadjusted values given in Table 4.1. Control group means may also differ because sample sizes differ due to missing data on

concomitant variables for some observations. The treatment effect is the impact of the treatment on the outcome measure for that hypothetical individual.

7. The appendix table (Table A4.1) presents complete regression results for each experiment from a model of weeks of UI benefits with a single treatment indicator and all concomitant variables. The full final project reports are Corson et al. (1992) for Pennsylvania, Spiegelman, O'Leary, and Kline (1992) for Washington, and Spiegelman and Woodbury (1987) for Illinois.
8. Because of the upper limit on dollars and weeks of UI compensation, a type of censoring is at work. Tobit models which explicitly accounted for censoring were therefore estimated. Tobit models yielded virtually the same results as the ordinary least squares models and are therefore not reported. Since the outcome in the exhaustion equation is a zero/one binary variable, a probit model for exhaustion was estimated. Again the results were extremely close to those computed by ordinary least squares.
9. The concomitant variables included when computing the estimates reported in Table 4.3 were gender, age, race, previous industry, the WBA, entitled maximum duration of benefits, base period earnings, work-search exempt status, enrollment site, and (for Pennsylvania) the quarter of enrollment into the experiment.
10. In Pennsylvania, 10.6 percent of 8,864 bonus offers were paid; in Washington, 14.6 percent of 12,452 offers were paid.
11. The mean WBA for the treatment-assigned beneficiaries was $166 in Pennsylvania and $154 in Washington.
12. While the design of the Illinois experiment prevents formal testing of differences in the parameters across all three samples, the "two-sigma rule" suggests the response is unchanged by adding the Illinois data since the parameter estimates are within twice the standard error of those estimated on the combined Pennsylvania and Washington data alone. Formal analysis of variance tests were performed by separately removing Pennsylvania and Washington from the full sample of data combined from the three experiments, and the parameter estimates were not found to be statistically different in either case. These tests also provide some justification for the exercise of estimating a bonus response surface on pooled data.
13. Estimation of a quadratic response surface for the pooled sample suggested that, for low dollar bonus amounts or very long qualification periods, the bonus could lead to increased UI benefit durations. The quadratic model included the dollar bonus amount, weeks in the qualification period, the bonus amount squared, the qualification period squared, and the bonus amount multiplied by the qualification period along with the concomitant variables included in the linear continuous variables model. Results from the quadratic model suggest that Figure 4.2 is an appropriate representation of response when the bonus amount is in the range $700 to $1,700 and the qualification period is between 5 and 12 weeks. The quadratic model suggests the minimum response occurs for a bonus amount of $700 with a 9.7-week qualification period.
14. More formally, the cumulative exit rate is computed by what Kiefer (1988) referred to as an integrated hazard function. In the present application, the hazard

function yields conditional UI exit rates which are not examined here since they are subject to selection bias resulting from response to the bonus offer by treatment group members. The cumulative hazard is not subject to such bias since the base for computation, the initial risk pool, does not change between periods.

15. The UI exit rates for the Illinois controls are roughly 50 to 70 percent of the UI exit rates in Pennsylvania and roughly 40 to 60 percent of the UI exit rates in Washington.
16. The concomitant variables, Z, included when performing subgroup impact estimation were WBA, entitled duration of benefits, base period earnings, work-search exemption status, an industry missing indicator, and five indicator variables for the quarter of enrollment in Pennsylvania.
17. Treatment impacts for a particular subgroup were computed as the sum of the parameter estimated on the product of the subgroup dummy variable and the treatment indicator added to the sum of estimates of parameters on subgroup dummies interacted with treatment indicators multiplied by their respective population shares (i.e., the proportion of the population having that characteristic). In each computation, parameter estimates for the complement to the subgroup of interest were omitted. For example, expressing population shares as the expected value of the subgroup dummy variables, $E(G)$, the subgroup impacts would be computed as ß + $E(G)\hat{E}$, where ß and $\hat{E}$ are the least squares estimates of $\boldsymbol{B}$ and $\boldsymbol{E}$ from Equation 4.5.
18. Rosen and Quandt (1988, p. 54) estimated the nonaccelerating inflation rate of unemployment (NAIRU) as 5.6 percent over the period 1932 to 1983, using a disequilibrium model of the aggregate labor market and setting the inflation rate at 5 percent.

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Appendix

Definitions of variables in Appendix Table A4.1. Site indicators differ acrossthe experiments (see site definitions below).

Intercept	constant term
Treatment	dummy variable 1 if treatment assigned, else zero
WBA	weekly benefit amount
Female	dummy variable 1 if female, else zero
AGELT35	dummy variable 1 if less than 35 years, else zero
AGEGT54	dummy variable 1 if greater than 54 years, else zero
Black	dummy variable 1 if Black, else zero
Hispanic	dummy variable 1 if Hispanic, else zero
Other non-White	dummy variable 1 if not White, Black or Hispanic, else zero
Entitled duration	Illinois, dummy 1 if 26 weeks; Pennsylvania, dummy one if 15 weeks; Washington number of weeks of UI entitlement
BPE	base period earnings
Manufacturing	dummy variable 1 if prior job in manufacturing, else zero
SIC missing	dummy variable 1 if SIC code missing, else zero
Search exempt	dummy variable 1 if search exempt, else zero

Site 1 - Illinois, Springfield South; Pennsylvania, Coatesville; Washington, Auburn
Site 2 - Illinois, Springfield North; Pennsylvania, Philadelphia (North); Washington, Renton
Site 3 - Illinois, Danville; Pennsylvania, Philadelphia (Uptown); Washington, Lynnwood
Site 4 - Illinois, Kankakee; Pennsylvania, Reading; Washington, North Seattle
Site 5 - Illinois, Rockford East; Pennsylvania, Lewistown; Washington, Rainier
Site 6 - Illinois, Rockford West; Pennsylvania, Butler; Washington, Everett
Site 7 - Illinois, Peoria; Pennsylvania, Connellsville; Washington, Bellevue
Site 8 - Illinois, Moline; Pennsylvania, McKeesport; Washington, Bellingham
Site 9 - Illinois, Galesburg; Pennsylvania, Erie; Washington, Bremerton
Site 10 - Illinois, West Town; Pennsylvania, Pittston; Washington, Mt. Vernon
Site 11 - Illinois, Mt. Prospect; Pennsylvania, Scranton; Washington, Olympia
Site12 - Illinois, Waukegon; Pennsylvania, Lancaster; Washington, Lewis County
Site 13 - Illinois, Villa Park; Washington, Aberdeen
Site 14 - Illinois, Aurora; Washington, Cowlitz County
Site 15 - Illinois, Woodlawn; Washington, Spokane

Site 16 - Illinois, South Chicago; Washington, Moses Lake
Site 17 - Illinois, Chicago Heights; Washington, Wenatchee
Site 18 - Illinois, Bedford Park; Washington, Yakima
Site 19 - Illinois, Chicago Hotel and Restaurant workers; Washington, Sunnyside
Site 20 - Illinois, Chicago Professional and Sales workers; Washington, Tri-cities
Site 21 - Illinois, Ravenswood; Washington, Walla Walla
Site 22 - Illinois, Evergreen Park
Site missing - Illinois

NQ1 - enrolled in Pennsylvania in third quarter 1988
NQ2 - enrolled in Pennsylvania in fourth quarter 1988
NQ3 - enrolled in Pennsylvania in first quarter 1989
NQ4 - enrolled in Pennsylvania in second quarter 1989
NQ5 - enrolled in Pennsylvania in third quarter 1989

Table A4.1 Regression-Adjusted Treatment Impact Equations

	Illinois		Pennsylvania		Washington	
Variable	Parameter estimate	Std. error	Parameter estimate	Std. error	Parameter estimate	Std. error
Intercept	20.00**	0.19	14.91**	0.18	14.35**	0.19
Treatment	–1.04**	0.26	–0.58**	0.21	–0.40*	0.21
WBA	0.02**	0.00	0.02**	0.00	0.03**	0.00
Female	0.38	0.27	1.04**	0.21	1.41**	0.19
AGELT35	–2.00**	0.28	–1.52**	0.21	–1.91**	0.18
AGEGT54	—	—	3.19**	0.34	1.27**	0.31
Black	3.96**	0.40	–1.90**	0.39	1.77**	0.45
Hispanic	1.86**	0.57	–1.57**	0.55	–0.09	0.40
Other non-White	0.82	0.93	–1.81	1.25	–0.34	0.40
Entitled duration	–2.83**	0.27	–1.82**	0.91	0.15**	0.02
BPE	0.08**	0.02	–0.07**	0.01	–0.03**	0.01
Manufacturing	0.68**	0.32	–0.48**	0.23	–1.61**	0.21
SIC missing	1.19**	0.42	0.16	0.34	0.87	1.98
Search exempt	—	—	–0.60*	0.31	–3.77**	0.22
Site 1	–2.46**	0.86	–1.15**	0.52	–1.99**	0.95
Site 2	–2.15**	0.88	1.36**	0.49	–2.44**	0.97
Site 3	2.35**	0.87	—	—	–3.80**	0.99
Site 4	0.04	0.95	–1.75**	0.51	–1.36	0.95
Site 5	–0.25	1.01	0.38	0.59	–0.90	0.92
Site 6	–0.74	0.94	1.00*	0.54	–2.06**	0.96
Site 7	0.55	0.78	1.85**	0.55	–2.79**	0.96
Site 8	0.14	0.82	1.12**	0.52	–1.45	1.01
Site 9	1.15	1.45	0.60	0.52	–0.30	1.04
Site 10	0.44	0.80	1.83**	0.58	–2.03**	1.00
Site 11	–5.78**	0.68	–0.10	0.52	–0.60	1.01
Site 12	–1.08	0.95	–2.38**	0.52	–3.01**	1.08
Site 13	–2.85**	0.79	—	—	–1.10	1.05
Site 14	–2.74**	0.84	—	—	–3.45**	1.06
Site 15	–0.69	0.70	—	—	–0.69	0.94
Site 16	–0.16	0.67	—	—	–2.92**	1.05
Site 17	0.37	0.78	—	—	–1.95**	0.99

Table A4.1 (continued)

Variable	Illinois		Pennsylvania		Washington	
	Parameter estimate	Std. error	Parameter estimate	Std. error	Parameter estimate	Std. error
Site 18	0.40	0.71	—	—	–1.01	0.94
Site 19	2.30	3.41	—	—	–0.15	1.02
Site 20	0.02	0.70	—	—	–0.62	0.99
Site 21	–0.51	0.77	—	—	—	—
Site missing	0.87	1.48	—	—	—	—
NQ1	—	—	–2.13**	0.97	—	—
NQ2	—	—	–1.44**	0.41	—	—
NQ3	—	—	–1.98**	0.40	—	—
NQ4	—	—	–1.80**	0.41	—	—
NQ5	—	—	–1.49**	0.40	—	—
Sample size	8,138		12,106		15,534	
R^2	0.08		0.05		0.06	

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