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Do Alternative Base Periods Increase Unemployment Insurance Receipt Among Low-Educated Unemployed Workers?

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ABSTRACT

Unemployment Insurance (UI) is the major social insurance program that protects against lost earnings resulting from involuntary unemployment. Existing literature finds that low-earning unemployed workers experience difficulty accessing UI benefits. The most prominent policy reform designed to increase rates of monetary eligibility, and thus UI receipt, among these unemployed workers is the Alternative Base Period (ABP). In 2009 the American Recovery and Reinvestment Act sought to increase use of the ABP, making ABP adoption a necessary pre-condition for states to receive their share of the $7 billion targeted at UI programs. By June 2012, 40 states and the District of Columbia had adopted the ABP despite the absence of an evaluation of ABP efficacy using nationally representative data. This paper analyzes Current Population Survey data from 1987-2007 to assess the efficacy of the ABP in increasing UI receipt among low-educated unemployed workers. We use a natural-experiment design and logistic regression models to capture the combined behavioral and mechanical effects of the policy change. We find no association between state-level ABP adoption and individual UI receipt for all unemployed workers. However, among part-time unemployed workers with less than a high-school degree, adoption of the ABP is associated with a 3.4 percentage point increase in the probability of UI receipt.

INTRODUCTION

Unemployment Insurance (UI) is the major social insurance program in the U.S. protecting against lost earnings incurred during involuntary unemployment. The program has dual aims: to smooth the consumption of temporarily unemployed workers and to stabilize the macro-economy during recessions. The program is designed to serve involuntarily unemployed workers with sufficient labor force attachment. Existing literature, however, finds that low-earning unemployed workers experience difficulty accessing benefits, despite the fact that regressive UI taxes are paid on their earnings (Anderson & Meyer, 2006). Low-earning workers may have difficulty accessing UI for several reasons: 1) failure to apply for benefits; 2) non-monetary ineligibility; and 3) monetary ineligibility. Recent efforts to expand access to UI among low-earning workers have largely focused on expanding monetary eligibility. The implicit assumption behind policy changes targeting monetary ineligibility is that there are some unemployed workers for whom monetary ineligibility is the sole barrier to benefit receipt.

The most prominent policy reform designed to increase rates of monetary eligibility, and thus UI receipt, among low-earning unemployed workers is the Alternative Base Period (ABP). When an ABP is used, state UI offices shift the window during which they examine earnings for eligibility, looking at the four most recent completed quarters. In contrast, under the standard base period, states examine a four quarter period that could have ended as much as six months prior to the job separation. By raising monetary eligibility rates, ABP proponents hope to increase rates of UI receipt among this population.

In the current study, we assess the efficacy of the ABP in increasing rates of UI receipt among low-educated unemployed workers using a natural-experiment design and logistic regression models that capture the combined behavioral and mechanical effects of the policy change. We use nationally representative data from the Current Population Survey for the years between 1988 and 2008 to gain information about UI receipt between 1987 and 2007 (1988 was the first year that a question regarding UI receipt in the year prior appeared in the CPS-ASEC). We exploit the temporal variation in ABP adoption in a number of states to test whether its adoption increases the probability that an unemployed worker will receive UI. We find no association between state-level ABP use and individual UI receipt for the broad population of unemployed workers. However, among part-time unemployed workers with less than a high-school degree, use of the ABP is associated with a 3.4 percentage point increase in the probability that an unemployed worker will access UI.

BACKGROUND

UI Eligibility Among Low-Earning Unemployed Workers

Previous work has shown that low-earning (both in terms of wage rates and average work hours) unemployed workers are less likely to receive UI than their higher-earning counterparts (Government Accountability Office, 2006). This is partly because they have lower rates of eligibility than higher-earning workers, and may also be partially a result of low-earning eligible workers being less likely to apply for benefits (Shaefer, 2010). Despite lower rates of UI receipt, the wages of low-earning workers are subject to UI taxes, even regressively so (Anderson and Meyer, 2006). In recent years, there has been some interest in reforming UI eligibility rules to make it easier for these workers to access benefits.

There are two types of eligibility criteria for UI: non-monetary and monetary. Most non-monetary requirements relate to the circumstances surrounding a worker's job separation, including the reason for job loss and search for future employment. These rules are meant to ensure that the worker separated from employment through no fault of his or her own and that the worker is an active member of the labor force. Typically, to be eligible, workers must have left employment due to layoff, plant closing, or some other involuntary reason, without cause, and be looking for work.

There is evidence that low-earning workers more often voluntarily leave jobs, in some cases because their personal circumstances, such as inadequate transportation and dependent care responsibilities, constrain them from working (General Accounting Office, 2000). Low-earning workers are also disproportionately clustered in industries that avoid formal lay-offs, making non-monetary eligibility difficult to achieve (General Accounting Office, 2000). Many employers in food service and retail sectors, for example, follow a practice termed "work loading," keeping employees on the payroll but reducing scheduled hours to zero so that formal lay-off is avoided (Lambert, 2008). A number of existing studies suggest that non-monetary requirements may be the key eligibility barrier to UI access for unemployed workers with low earnings or short or sporadic work histories (Holzer, 2000; O'Leary & Kline, 2008; Rangarajan et al, 2002; Shaefer and Wu, 2011).

The ABP policy change, however, focuses on *monetary* eligibility. Monetary eligibility generally requires a state-specific minimum of earnings (or, in two states, minimum of work hours) from any qualifying employer, with the goal of ensuring applicants have an adequate record of labor force attachment. These requirements vary, but generally fall between $1,000 and $3,500 earned over four quarters. Table 1 displays the monetary eligibility requirements for the 50 states plus the District of Columbia in 2007, the most recent year analyzed in our study. Many states also have a high quarter requirement, with a minimum requirement within a single quarter, and some states require two quarters of positive earnings.

{{Place Table 1 About Here}}

Until recently, the base period used by most states to determine eligibility included earnings in the first four of the previous five completed quarters. Historically, this fifth "lag quarter" was necessary for states to process earnings data. The exclusion of as much as six months of an unemployed worker's most recent earnings may pose a challenge for unemployed workers with low earnings or short or sporadic work histories. Unemployed workers with low wages need to work more hours to meet earnings requirements than higher wage unemployed workers. Thus, a low-earning worker who returned to the workforce eight months prior to a lay-off could find himself or herself ineligible for UI, even though he or she had substantial earnings during that period. Further, on average, low-earning workers have shorter job tenures than their more highly educated counterparts, making the exclusion of recent wages a more consequential issue for this group.

The Alternative Base Period

The ABP shifts the “window” in which earnings requirements are examined. Rather than excluding the most recent completed quarter, the ABP includes this quarter and drops the first quarter of the standard base period, during which workers with short work tenure may not have been earning wages. Figure 1 illustrates the two base periods that could be used in determining monetary eligibility for an unemployed worker who lost a job and filed for benefits in, for example, February of 2012.

{{Place Figure 1 About Here}}

Under the standard base period, this particular worker would have the five most recent months of earnings excluded. Under the ABP, only two months of recent earnings would be excluded. The premise of the ABP is that if unemployed workers can count more recent earnings, they will be more likely to be monetarily eligible. In most states, UI applicants are given two “chances” to monetarily qualify under the ABP, first using the standard base period and then the ABP. By giving unemployed workers two chances to qualify monetarily, and by allowing the unemployed to count more recent earnings, policy makers surmise that low-earning unemployed workers who were on the margins of monetary eligibility under the old system will now be more likely to qualify for, and thus receive, UI. Beyond the mechanical effect of increasing monetary eligibility rates among applicants, an ABP may also increase rates of UI application among unemployed workers who would not have otherwise applied (O'Leary, 2011). As individuals learn about the new rules, they may think they would be more likely to meet eligibility criteria, and thus be more likely to apply.

While the logic of the ABP is based on the premise that workers may have difficulty meeting monetary eligibility criteria, examining Table 1, we see that minimum base period earnings requirements are low. For example, in Michigan, the 2007 minimum base period earnings requirement was $2,997. A full-time worker working at the state minimum wage[1] would earn eligibility in just eleven weeks. In Kentucky, a state with a lower minimum earnings requirement (and a lower minimum wage), a full-time minimum wage earner could achieve monetary eligibility in eight weeks. In two states (Washington and Oregon), applicants must have completed a minimum number of hours of employment, and these actually translate to relatively high thresholds compared to other states. Still, for most workers with regular attachment to the labor force, monetary eligibility may not be a significant barrier to UI access.

Despite these concerns, adoption of the ABP has been widespread: as of 2007, eighteen states and the District of Columbia had adopted an ABP. In 2009 the American Recovery and Reinvestment Act sought to increase use of the ABP, making ABP adoption a necessary pre-condition for states to get any of their share of the $7 billion targeted at UI programs. By June 2012, 40 states plus the District of Columbia had adopted an ABP.

Previous Research and the Current Study

Despite the widespread adoption of the ABP, we are unaware of any study that uses nationally representative data to assess its efficacy in increasing UI receipt. A few studies have

[1] This calculation uses the Michigan minimum wage from the first half of 2007- $6.95.

used state-specific administrative data or simulations based on survey data to estimate the impact of an ABP on monetary eligibility rates and UI receipt. Two studies look only at monetary eligibility. Rangarajan and Razafindrakoto (2004) report on simulations based on survey data from Mathematica's National Evaluation of the Welfare-to-Work Grants Program. They estimate that an ABP would increase monetary eligibility among women who left welfare for work by 4 to 9 percent. Stettner, Boushey and Wenger use the Survey of Income and Program Participation and estimate that universal adoption of the ABP in the late 1990s and early 2000s would have increased monetary eligibility of separated workers by 6 percent, with low-wage workers disproportionately affected.

Vroman (2008) used UI administrative data from Ohio from 1967-2007. He found that in 2006-2007, 6 percent of UI claimants in Ohio accessed the UI program through the ABP, and that applicants were widely aware of this policy change. Ohio has relatively high monetary eligibility thresholds, which may increase the impact of the ABP. Vroman's sample was also limited to UI applicants, and so cannot speak to the ABPs impact on previous non-applicants. Finally, Vroman does not look at earnings records longitudinally, so does not take into account UI applicants who would have re-applied when their standard base period shifted.

O'Leary (2011) conducted simulations using administrative data from Kentucky. He estimated that ABP adoption would increase the proportion of monetarily eligible applicants by 2.82 percent and would increase the proportion of UI beneficiaries by 2.21 percent as a result of the ABP, compared to Vroman's 6 percent. The gap between Vroman's and O'Leary's results may stem from the fact that Ohio's monetary eligibility requirements are relatively high while Kentucky's are relatively low (see table 1). Further, O'Leary looks at earnings records longitudinally, taking into account UI applicants that would have reapplied when their standard base period shifted.

On the whole, existing studies suggest that an ABP should increase the likelihood that unemployed workers are eligible for UI, if only slightly. This effect should be most concentrated among workers with low earnings or short or sporadic work histories. None of the studies reviewed above, however, use nationally representative data to evaluate the effects of existing ABPs on UI receipt of low-earning unemployed workers. These studies are either designed only to assess monetary eligibility; are confined to administrative data from one or a few states; and/or are based on simulation only. These studies therefore miss the behavioral effect of drawing previous non-applicants into the programs.

The ABP is the most prominent policy meant to increase UI receipt among low-earning workers. In order to evaluate the policy's effectiveness in increasing UI receipt nationally, it is necessary to use survey data because administrative data do not include key indicators, including educational attainment, work hours, and other key demographic characteristics. Further, there is no nationally representative source of UI administrative data, limiting the generalizability of analyses using such data.

The current study uses data from the Current Population Survey Annual Social and Economic Supplement, stratifying a sample of unemployed workers by education level, full-time/part-time status and other factors. We take advantage of the temporal variation in implementation of an ABP by eighteen states plus the District of Columbia between 1987 and 2007. These states differ widely by region, population size and demographics, industrial base, and other factors. By using a parsimonious natural experiment design, we test whether adoption of an ABP increases the probability that low-educated unemployed workers will access UI.

DATA AND METHODS

Data

The Current Population Survey (CPS), a monthly survey of approximately 60,000 households, is a major source of labor market statistics for the US. The CPS offers a nationally representative multistage stratified sample of the non-institutionalized population. Detailed labor market and demographic data are collected on all respondents aged 16 years and older. The Annual Social and Economic Supplement provides annualized data for the preceding year on numerous labor market and public program participation outcomes. Data were extracted from the Integrated Public Use Microdata Series. In this series, CPS data from the Annual Supplement between 1962 and 2007 were integrated and variables were "harmonized" (coded identically) to be consistent over time (King et al., 2011).

The CPS offers a larger sample and more uniform data across our complete study period than other large nationally representative surveys. Underreporting of public benefits in household surveys is a concern (Meyer, Mok and Sullivan, 2009), and will prove to be a limitation of our study. However, Meyer et al. (2009) find that the CPS-ASEC reporting rates for UI benefit dollars are relatively high compared to other programs. All but three years in our study sample have a UI dollars reporting rate of 75 percent or above. Further, use of a natural experimental design with inclusion of year controls should mitigate concern that under-reporting could drive the results of a study using our method, assuming that rates of underreporting do not co-vary with ABP adoption.

We restrict our sample to adults ages 18 to 64 who report during a calendar year that they both 1) worked for pay and 2) experienced a spell of unemployment of at least two weeks. By limiting our sample to individuals who both worked for pay and experienced a spell of unemployment, we hope to restrict our sample to workers with reasonable labor force attachment.[2] Our sample includes unemployed workers from all 50 states plus the District of Columbia who experienced unemployment between 1987 and 2007. Our resultant sample consists of 164,131 respondents.

Hypotheses

The CPS data do not allow us to determine an unemployed worker's non-monetary eligibility status, nor whether he or she applied for UI. Our objective is to test whether the implementation of ABP is associated with increased UI receipt for unemployed low-educated workers, accounting for the combined the mechanical effect of increasing monetary eligibility among applicants, and for the behavioral effect of drawing otherwise non-applicants into the program. Because the CPS data allow us to measure education levels more precisely than earning levels, we stratify by education level when testing the following hypothesis:

> H_1: ABP use at the state level is associated with increased UI receipt among unemployed workers with less than a high school diploma.

Because low-educated part-time workers have lower quarterly earnings than their full-time counterparts, we also test the following hypothesis:

[2] Ideally, we would drop labor force entrants who transition from out of the labor force to unemployed and then become employed. The harmonized CPS-ASEC data do not allow us to do this, and we cannot connect our data to the basic CPS monthly survey. This is a limitation of our analysis.

H_2: ABP use at the state level is associated with increased UI receipt among unemployed workers with less than a high school diploma who worked part-time prior to job separation.

Model

To determine whether the implementation of the Alternative Base Period is associated with an increase in the probability that a low-educated unemployed worker will receive UI, we take advantage of the natural experiment created by the gradual state-by-state implementation of the Alternative Base Period. States that adopted an ABP between 1987 and 2007 are reported in Table 2.

{{Place Table 2 About Here}}

These states vary on characteristics such as region, dominant industry, union density, and political orientation of state legislature. In addition, the variation in year of implementation captures variation in the economic cycle over time. This variation arguably creates a natural experiment: it is as if the state and year of implementation had been randomly selected. Further, because we use state and year controls, our approach is robust against spurious factors that could influence implementation decisions such state-specific levels of UI receipt and UI recipiency rates at a given point in the business cycle, unless these factors co-vary with ABP adoption.

In order to determine the effect of ABP use on the probability that an unemployed worker will receive UI, we use logistic regression models. Parameter estimates have been converted to the average marginal effect and therefore can be interpreted similarly to output from linear probability models. The main specification is:

$$PR\ (UI|Unemp)_{i,j,t} = \Phi(\beta_1 ABP_{j,t} + \lambda X_{i,j,t} + \theta_{j,t})$$

The dependent variable is a dichotomous measure where 1 = UI receipt and 0 = no UI receipt in the year of an unemployment spell for individual *i* in year *t* in state *j*. ABP use at the state level is the independent variable of interest, a dichotomous measure of whether the unemployed worker's state of residence *j* in year *t* used an Alternative Base Period where 1 = ABP use and 0 = no use of ABP. X is a vector of individual demographic characteristics, which include a set of age dummies, a categorical measure of educational attainment, sex, race and ethnicity, marital status, and full-time status. θ includes state and year controls as well as the state-year unemployment rate. We also include in this vector state-year controls for minimum eligibility thresholds, which may co-vary with ABP adoption, and state minimum wages, which may impact the likelihood of monetary eligibility for a low-earning worker. We further ran a sensitivity test with a control for state-year UI program part-time work search requirements (which allow UI beneficiaries to search for part-time work only under certain conditions). This variable had no effect on our results in regards to ABPs. This indicator was not included in the main specification because of concerns that it had been imprecisely measured.[3]

[3] There is no official source of information on adoption of part-time work search requirements by state-year, prior to information provided by ETA in the early 2000s. Thus, we contacted state UI research offices for states that had this requirement before that point to inquire about the year the policy was adopted. Most state UI research offices did not keep a record of the year of policy implementation. Thus the data we received from them were generally based on legal research (despite the fact that many states changed policy prior to statute change) and the memory of the agencies' most senior employees, which are of course subject to recall bias. Taking into account these limitations, there appears to be no relationship between changes in work search requirements and ABP implementation. Given the imprecision of measurement and the fact that the indicator had no impact on our ABP estimate, we did not include it in our main model.

By including state and year dummies, the ABP indicator comes to represent the effect of adoption of an ABP within states, over time, reducing the risk that the variable is spuriously capturing associations between the ABP and other state-level characteristics that may impact UI receipt. Φ represents the logistic distribution. We use person-level probability weights and cluster our standard errors by state to account for the CPS' stratified sample design.

We first run the model on the full population (n=164,131), and then stratify our sample to determine the effect of ABP on subpopulations of interest. We first stratify by education-level: less than high school (n=33,867); high school only (n=63,127); some college (n=45,454); bachelor's degree and higher (n=21, 683). Within the less than high school population, we stratify further to full-time (n=24,373) and part-time unemployed workers (n=9,494).

We employ stratified models rather than models with interaction effects because the stratified models allow us to more clearly interpret the sub-group effects and more accurately model confounding variables for sub-populations who have vastly different experiences in the labor market, without requiring numerous interactions. In this case, we are particularly interested in the impact of the ABP on low-educated unemployed workers, and would not expect the policy to greatly impact higher-educated unemployed workers, who have extremely high rates of monetary eligibility.

RESULTS

Effect of Alternative Base Period by Education Level

Table 3 shows our model run on the full population of unemployed workers, stratified to sub-groups by education level and full-time/part-time status (prior to separation).

{{Place Table 3 About Here}}

The first column shows that, for the general population, we find no association between use of the ABP and the probability that an unemployed worker will receive Unemployment Insurance in any significant way—in fact the point estimate is essentially zero. Our small standard error (.007) suggests that we are estimating a precise zero. The following four columns show that when we stratify our sample only by education level and not by hours worked prior to job separation, we do not see a statistically significant effect of ABP use. In other words, we find no support for the hypothesis that the Alternative Base Period is associated with increased probability of UI receipt, even for unemployed workers with less than a high school degree.

Turning to columns 6 and 7, we show the association between the ABP and UI receipt for unemployed workers who were previously employed part-time and unemployed workers who were previously employed full-time. While there is no significant association between an ABP and the probability that a worker previously employed full-time will receive UI, there is a significant effect for workers previously employed part-time at the .05 significance level. According to our results, ABP use is associated with an increased probability that an unemployed worker with less than a high school degree, previously employed part-time, will receive UI; we see an increased probability of 3.4 percentage points. This is the only significant effect that we find associated with the implementation of the ABP in our main models[4].

[4] In many states, workers who seek re-employment at the part-time level are ineligible to receive UI; interestingly, despite this non-monetary barrier to UI receipt, it appears that the ABP is most helpful to this population.

When measuring the effect of ABP, including individual-level demographic characteristics and state-level controls in our models neither substantively affects our parameter estimates nor our standard errors (as discussed in more detail below). Despite this, we include these variables in our main specifications as a consistency check with other studies. We find that—across subpopulations—older workers are more likely to receive UI. Across all subgroups, full-time workers are more likely to receive UI and higher unemployment rates are positively associated with UI receipt. Unemployed workers with a high school degree or some college are significantly more likely to receive UI than their counterparts with less than a high school diploma, but a college degree is not associated with a higher probability of UI receipt, after controlling for other factors in the model. In most models, blacks are less likely to receive UI than whites; Hispanic unemployed workers are less likely to receive UI than white non-Hispanics; and female unemployed workers are less likely to receive UI than male unemployed workers. Further, across most subgroups married unemployed workers are more likely to receive UI than unmarried unemployed workers.

Demographic characteristics are less strongly associated with probability of UI receipt among workers with less that a high-school diploma that worked part-time prior to job separation than among other subgroups, though use of the ABP is more predictive of UI receipt among this group. We see, too, that the state unemployment rate is less strongly associated with increased probability of UI receipt among workers with less than a high school diploma (with a parameter estimate of 0.012, compared to 0.016 or 0.017 for all other education levels). This weaker countercyclical effect appears to be driven primarily by those unemployed workers with less than a high school diploma who worked part-time prior to job separation.

Sensitivity Analyses

Because we see significant differences in probability of UI receipt by demographic characteristics—e.g. women and racial/ethnic minorities being less likely to receive UI in the broad population, we perform sensitivity analyses to examine the effect of ABP on UI receipt for these subgroups in particular[5]. Because we are primarily interested in whether the ABP is associated with the probability that less-educated unemployed workers will receive UI, we examine the effect of the ABP on subpopulations of unemployed workers with less than a high school diploma. Table 4 reports the point estimates and standard errors for the ABP coefficient when we stratify within the group of unemployed workers with less than a high school education to run our model on subgroups by race/ethnicity[6] and sex.

When examining the full population of unemployed workers with less than a high school education, we do not see a significant effect of ABP on UI receipt for any racial/ethnic subgroup, nor for males or females. When we restrict to unemployed workers who worked part-time prior to employment separation, we find a positive association with significance at the .05 level for subpopulation of black unemployed workers and a positive association with significance at the .05 level for male unemployed workers. We find no evidence that the ABP is associated with increased UI receipt among Hispanic unemployed workers with less than a high school diploma, even among those working part-time before job separation.

These sensitivity analyses suggest that the association of ABP with UI receipt among Black and male unemployed workers without a high school diploma who worked part-time hours drives the significance reported in the initial model reported in column 6 of Table 2. It is

[5] Full output available upon request

[6] Asians and Native Americans were excluded from this analysis due to small sample size; unemployed workers who indicated more than one racial group or reported an "other" racial category were excluded from this analysis because of within-group heterogeneity among these populations.

interesting to note that while the ABP has the strongest effect for a racial group with a UI recipiency rate that is lower than average (black unemployed workers) it has the strongest effect for the gender group that already has a higher rate of UI receipt than average (male unemployed workers).

Table 5 reports on a final sensitivity analysis in which we compare the ABP point estimates from our fully specified model to a reduced model that includes only the ABP indicator with state and year controls. If what we are capturing is truly a natural experiment, we would expect that the point estimates would be essential un-moved by the exclusion of the other covariates. In fact, that is what we find. In all cases, the point estimates in the reduced model are substantively similar to the one with the full set of covariates. This offers additional evidence that our specification is capturing a natural experiment.

INTERPRETATION OF RESULTS

Our results suggest that implementation of the ABP is associated with an increase in the probability that low-educated, part-time unemployed workers will receive UI benefits of 3.4 percentage points, significant at the .05 level. Taking data from the most recent year in our sample, 2007, we estimate that universal adoption of an ABP would be associated with UI receipt for approximately 20,000 additional unemployed part-time workers without a high-school diploma than would be true in the absence of an ABP. However, we estimate the total number of workers who experienced unemployment over the course of 2007 to be 11 million. Our estimates suggest that—compared to universal use of the standard base period—universal adoption of the ABP would extend new UI coverage to a group of unemployed workers whose size is two tenths of a percent of the total unemployed population.

Though our results suggest that the ABP would extend UI coverage to an appreciable number of unemployed workers, we only find an effect for part-time unemployed workers without a high school degree. Our results demonstrate that in the general population of unemployed workers, the ABP does not help enough workers to make a difference in the probability that a randomly selected unemployed worker will receive UI; only one subgroup of unemployed workers examined, less-educated workers who were previously employed part-time, sees any significant change in levels of UI receipt associated with the policy change. Because part-time workers work fewer hours, and because less-educated workers have lower hourly wages, this group is likely to have difficulty achieving monetary eligibility[7]. Thus, it makes sense that this group would benefit most from use of the ABP.

Our point estimates and significance levels may appear to be slightly lower than results from previous studies reviewed in the background section of this paper. However, the studies we reviewed that used state-level administrative data examined the change in levels of UI receipt among UI applicants only, and in the case of Vroman, the study was conducted in a state with high monetary eligibility requirements. Our study examines the full population of unemployed workers, including non-applicants; thus our estimates should be expected to be substantially lower. O'Leary found that in Kentucky, an ABP would increase the UI receipt of *applicants* by 2.21 percent. Applying our statistically insignificant point estimate for the association between an ABP and UI receipt among all unemployed workers, we find that the estimate is consistent with a 1 percent increase in UI receipt among all unemployed workers.

Thus, our results seem reasonable when compared to the O'Leary estimate. Our study, however, goes beyond previous evaluations to offer an estimate of the effect of ABP

[7] We also ran our analysis for part-time workers who had more than a high-school diploma and found no significant effect.

implementation on actual UI receipt at the population level. Our findings suggest that adoption of an ABP to increase levels of monetary eligibility alone may not be an effective strategy for raising UI recipiency rates among low-educated or low-earning workers, broadly. This implication is consistent with previous work that has suggested that non-monetary eligibility requirements and rates of application may be important barriers to UI access for the broad group of low-earning unemployed workers (Gould-Werth and Shaefer, forthcoming; Holzer, 2000; O'Leary & Kline, 2008; Rangarajan et al, 2002; Shaefer, 2010; Shaefer and Wu, 2011). Our findings suggest further policy change would be necessary to substantially impact UI coverage for the broad group of low-educated unemployed workers.

CONCLUDING COMMENTS

We find that the ABP policy change increases access to Unemployment Insurance for only a small fraction of the workforce: low educated part-time workers. However, today many employers are scheduling low-educated workers for variable hours and using business models that incorporate a shorter average job tenure than has been common historically (Government Accountability Office 2007, Kalleberg 2009). Our study suggests that for this group of low-educated workers who have difficulty scheduling a sufficient number of hours and who experience short job tenure, the ABP is a helpful intervention: because of their shorter work tenure and difficulty amassing enough work hours to qualify for UI, the elimination of the lag quarter and ability to "try twice" to qualify is most helpful to this group of workers.

However, low rates of coverage continue to exist among the broad swath of low-educated workers. Our results thus indicate that the implementation of ABP should be coupled with other interventions to make sure that the UI program is fulfilling its intended purpose for all workers.

Moving forward, further research should investigate other barriers to UI receipt for low-earning unemployed workers. Future studies, both qualitative and quantitative, could examine barriers to application and the nature of non-monetary eligibility among this group. The results of these studies would provide information about the potential need for other interventions to increase UI access for low-educated unemployed workers, such as employer-filed claims or an individual Unemployment Insurance Savings Account system. Such interventions, if found to be necessary, could be effectively coupled with the ABP. The U.S. economy has changed dramatically since the Unemployment Insurance system was established in 1935. In the context of the modern economy, further policy change is necessary if we hope to extend the program to all unemployed workers who lose their job through no fault of their own.

TABLES AND FIGURE

Table 1: Monetary Eligibility Earnings Thresholds in 2007 by State (in dollars)

Alabama	2290	Montana	1982
Alaska	1000	Nebraska	2592
Arizona	2250	Nevada	600
Arkansas	1917	New Hampshire	2800

California	1125	New Jersey	2860
Colorado	2500	New Mexico	>1548
Connecticut	780	New York	2400
Delaware	920	North Carolina	4113.24
District of Columbia	1950	North Dakota	2795
Florida	3400	Ohio	4000
Georgia	1680	Oklahoma	1500
Hawaii	130	Oregon*	1000
Idaho	1658	Pennsylvania	1320
Illinois	1600	Rhode Island	2840
Indiana	2750	South Carolina	900
Iowa	1730	South Dakota	1288
Kansas	2880	Tennessee	>1560
Kentucky	2994	Texas	2072
Louisiana	1200	Utah	2800
Maine	3612	Vermont	2677
Maryland	900	Virginia	>2700
Massachusetts	3000	Washington*	
Michigan	2997	West Virginia	2200
Minnesota	1250	Wisconsin	1590
Mississippi	1200	Wyoming	2600
Missouri	2100		

Source: Department of Labor Comparison of State Unemployment Insurance Laws

*Washington is the only state with a base period hours requirement rather than a base period earnings requirement. In 2007 they required 680 base period hours. Oregon has an hours requirement in addition to an earnings requirement.

Table 2: ABP Implementation Dates by State

State	Year
Vermont	1986
Washington	1988
Ohio	1989
Maine	1993

Rhode Island	1993
Massachusetts	1994
New Jersey	1996
North Carolina	1998
New York	1999
Wisconsin	2000
Michigan	2001
New Hampshire	2001
Connecticut	2003
Georgia	2003
Washington, D.C.	2003
Hawaii	2004
New Mexico	2004
Virginia	2004
Oklahoma	2005

Dates are taken from the state comparison of UI laws, ARRA letters, and other published sources. When effective dates occurred after June 30th, we rounded to the next calendar year.

Table 3: Receipt of Unemployment Insurance by Education Level and Part-Time/ Full Time Status
Unemployed Workers, aged 18-64
Logit models, average marginal effects, standard errors in parentheses

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	All Unemployed workers	Bachelor's Degree and Higher	Some College	High School Graduates	Less Than High School	Less than High School, Part-time	Less than High School, Full-time

Alternative Base Period	0.003	-0.006	-0.004	0.010	0.011	0.034**	-0.001
	(0.007)	(0.013)	(0.01)	(0.01)	(0.014)	(0.014)	(0.016)
Ages 26-35	0.181***	0.248***	0.203***	0.161***	0.160***	0.111***	0.178***
	(0.004)	(0.012)	(0.005)	(0.006)	(0.007)	(0.01)	(0.008)
Ages 36-45	0.242***	0.338***	0.252***	0.217***	0.227***	0.139***	0.261***
	(0.004)	(0.014)	(0.008)	(0.007)	(0.01)	(0.012)	(0.011)
Ages 46-55	0.275***	0.345***	0.271***	0.257***	0.274***	0.147***	0.321***
	(0.006)	(0.016)	(0.008)	(0.008)	(0.009)	(0.01)	(0.012)
Ages 55-64	0.301***	0.329***	0.284***	0.291***	0.315***	0.165***	0.377***
	(0.006)	(0.02)	(0.009)	(0.011)	(0.01)	(0.014)	(0.012)
H.S. Graduate	0.058***						
	(0.008)						
Some College	0.047***						
	(0.008)						
College +	-0.005						
	(0.008)						
Female	-0.030***	-0.046***	-0.035***	-0.036***	-0.008	0.009	-0.015
	(0.007)	(0.008)	(0.009)	(0.006)	(0.009)	(0.007)	(0.011)
Black	-0.053***	0.003	-0.030**	-0.079***	-0.053***	-0.023*	-0.065***
	(0.010)	(0.014)	(0.013)	(0.012)	(0.01)	(0.012)	(0.014)
Hispanic	-0.052***	-0.056***	-0.040***	-0.066***	-0.048***	-0.015	-0.059***
	(0.013)	(0.012)	(0.011)	(0.014)	(0.012)	(0.011)	(0.015)
Married	0.060***	0.008	0.047***	0.079***	0.072***	0.030***	0.086***
	(0.003)	(0.006)	(0.005)	(0.004)	(0.008)	(0.006)	(0.011)
Full-Time	0.222***	0.227***	0.216***	0.226***	0.206***		
	(0.005)	(0.012)	(0.008)	(0.008)	(0.006)		
Unemployment Rate	0.016***	0.017***	0.016***	0.016***	0.012***	0.004*	0.014***
	(0.002)	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.004)
Minimum Base Period Earnings	0.000*	0.000	0.000***	0.000	0.000	0.000***	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
State Minimum Wage	0.000	0.008	-0.002	-0.006	0.001	-0.002	0.003
	(0.003)	(0.01)	(0.006)	(0.005)	(0.008)	(0.01)	(0.010)
Pseudo R^2	0.135	0.098	0.134	0.136	0.176	0.147	0.13
N	164,127	21,683	45,454	63,123	33,867	9,494	24,373

Source: Authors' analysis of Current Population Survey Data (King et al, 2010)
State and year controls included in model, but omitted from table
* p<0.1, ** p<0.05, *** p<0.01

Table 4. Association of ABP with UI Receipt among Unemployed Workers (ages 18-64) with Less Than a High School Degree by Race/Ethnicity and Gender
(Logit Models, Average Marginal Effects)
Standard Errors in Parentheses, N in Third Row

	All	Part-Time	Full-Time
White	0.011	0.024	0.005
	(0.017)	(0.018)	(0.019)
	27,349	7,433	19,913
Black	0.007	0.060**	-0.035
	(0.024)	(0.024)	(0.034)
	4,416	1,377	2,937
Hispanic	0.025	0.04	-
	(0.021)	(0.032)	-
	12,049	2,494	9,439
Male	0.015	0.037**	0.008
	(0.014)	(0.018)	(0.016)
	21,797	4,968	16,829
Female	0.011	0.033	-0.013
	(0.02)	(0.016)	(0.029)
	12,070	4,526	7,544

* p<0.1, ** p<0.05, *** p<0.01
Source: Authors' analysis of Current Population Survey Data (King et al 2010)
Note: for Hispanic full-time workers with less than a high school diploma, our model does not converge, perhaps because of collinearity between ABP status and state in this geographically concentrated sample

Table 5: Comparison of Alternative Base Period Effects
Full model and reduced model omitting all controls except state and year fixed effects
Logit models, average marginal effects, standard errors in parentheses

	Full Model	**Reduced Model**
All Unemployed Workers	0.003	0.008
	(0.007)	(0.009)
BA+	-0.006	-0.003
	(0.013)	(0.014)

Some College	-0.004	0.004
	(0.010)	(0.013)
HS Dip	0.010	0.021
	(0.010)	(0.013)
Less than HS	0.011	0.001
	(0.014)	(0.014)
Less than HS part-time	0.034**	0 .034**
	(0.014)	(0.015)
Less than HS full-time	-0.001	-0.011
	(0.016)	(0.014)
White less than HS-all	0.011	-0.004
	(0.017)	(0.017)
part-time	0.024	0.019
	(0.018)	(0.019)
full-time	0.005	-0.012
	(0.019)	(0.018)
Black less than HS-all	0.007	0.006
	(0.024)	(0.023)
part-time	0.060**	0.065**
	(0.024)	(0.025)
full-time	-0.035	0.000
	(0.034)	(0.016)
Hispanic less than HS-all	0.025	-0.001
	(0.021)	(0.029)
part-time	0.04	0.043
	(0.032)	(0.027)
Male less than HS-all	0.015	0.004
	(0.014)	(0.016)
part-time	0.037**	0.044**
	(0.018)	0.021
full-time	0.008	-0.005
	(0.016)	(0.018)
Female less than HS-all	0.011	-0.003
	(0.02)	(0.019)
part-time	0.033	0.018
	(0.016)	(0.022)
full-time	-0.013	-0.014
	(0.029)	(0.026)

* p<0.1, ** p<0.05, *** p<0.01
Source: Authors’ analysis of Current Population Survey Data (King et al 2010)

Figure 1: Comparison of Base Periods

First Quarter	Second Quarter	Third Quarter	Fourth Quarter	Lag Quarter	Filing Quarter
October-December 2010	January-March 2011	April-June 2011	July-September 2011	October-December 2011	January-March 2012

Standard Base Period

First Quarter	Second Quarter	Third Quarter	Fourth Quarter	Lag Quarter	Filing Quarter
October-December 2010	January-March 2011	April-June 2011	July-September 2011	October-December 2011	January-March 2012

Alternative Base Period

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