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 precondition for states to receive their share of the $7 billion targeted at UI programs. By January 2013, 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 study analyzes Current Population Survey data from 1987 to 2011 to assess the efficacy of the ABP in increasing UI receipt among low-educated unemployed workers. We used a natural-experiment design to capture the combined behavioral and mechanical effects of the policy change. We found 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 was associated with a 2.8 percentage point increase in the probability of UI receipt. © 2013 by the Association for Public Policy Analysis and Management.

INTRODUCTION

Unemployment Insurance (UI) is the major social insurance program in the United States that protects workers 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, shows that low-earning unemployed workers experience difficulty accessing UI benefits (Government Accountability Office, 2007; Shaefer, 2010). Low-earning workers may have difficulty accessing UI for several reasons: (a) failure to apply for benefits, (b) nonmonetary ineligibility, and (c) 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 that target

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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 the ABP is used, state UI offices shift the window of time during which they examine earnings for eligibility, and look at the four most recently completed quarters. In contrast, under the standard base period, state UI offices 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.

This study assesses the efficacy of the ABP in increasing rates of UI receipt among low-educated unemployed workers[1] using a natural experiment design that captures the combined behavioral and mechanical effects of the policy change. We used nationally representative data from the Current Population Survey Annual Social and Economic Supplement (CPS ASEC) administered annually from 1988 through 2012, which reports on UI receipt between 1987 and 2011 (1988 was the first year that a question regarding UI receipt appeared in the CPS ASEC; King et al., 2010). We exploited the temporal variation in ABP adoption across states to test whether its adoption increased the probability that an unemployed worker would receive UI. We found no association between state-level ABP use and individual UI receipt for the overall population of unemployed workers. However, among part-time, unemployed workers with less than a high school degree, use of the ABP was associated with a 2.8 percentage point increase in the probability of UI receipt.

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 low-earning unemployed workers have lower rates of eligibility than higher earning workers, and may also be partly a result of low-earning eligible workers being less likely to apply for benefits (Gould-Werth & Shaefer, 2012). Despite lower rates of UI receipt, the wages of low-earning workers are subject to UI taxes, even regressively so (Anderson & Meyer, 2006). In the past two decades, 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: nonmonetary and monetary. Most nonmonetary 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 was separated from employment through no fault of his or her own and that the worker is an active labor force participant. Typically, to be eligible for UI, workers must (a) be looking for work and (b) have left employment either due to an involuntary reason, such as layoff or plant closing, or for cause but not for misconduct.

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

[1] We use education level as the best available indicator of workers' long-term earnings potential available in the CPS ASEC.

Accounting Office, 2000). Low-earning workers are also disproportionately clustered in industries that avoid formal layoffs, making nonmonetary eligibility more difficult to achieve (General Accounting Office, 2000). Many employers in the food service and retail sectors, for example, follow a practice termed "work loading," which keeps employees on the payroll but reduces scheduled hours, sometimes to zero, so that formal layoff is avoided (Lambert, 2008). A number of existing studies suggest that nonmonetary 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, Razafindrakoto, & Corson, 2002; Shaefer, 2010; Shaefer & 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 that 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 2011, the most recent year analyzed in our study. Many states also have

Table 1. Monetary eligibility earnings thresholds in 2011 by state (in dollars).

State	Earnings threshold	State	Earnings threshold
Alabama	>2,340	Montana	2,305
Alaska	2,500	Nebraska	3,770
Arizona	2,250	Nevada	402
Arkansas	2,870	New Hampshire	2,800
California	1,125	New Jersey	2,900
Colorado	2,500	New Mexico	1,750
Connecticut	600	New York	2,400
Delaware	720	North Carolina	3,985
District of Columbia	1,950	North Dakota	2,795
Florida	3,400	Ohio	4,000
Georgia	1,760	Oklahoma	1,500
Hawaii	130	Oregon[a]	1,000
Idaho	2,340	Pennsylvania	1,320
Illinois	1,600	Rhode Island	2,960
Indiana	4,200	South Carolina	4,455
Iowa	1,454	South Dakota	1,288
Kansas	3,330	Tennessee	>1,560
Kentucky	1,001	Texas	2,220
Louisiana	1,200	Utah	3,100
Maine	4,228	Vermont	3,084
Maryland	864	Virginia	2,700
Massachusetts	3,500	Washington[b]	
Michigan	4,307	West Virginia	2,200
Minnesota	1,250	Wisconsin	1,890
Mississippi	1,200	Wyoming	3,257
Missouri	2,250		

Notes: Oregon and Washington calculate base period requirements differently from the other states listed.

[a]Oregon has a base period hours requirement in addition to an earnings requirement.

[b]Washington is the only state with a base period hours requirement instead of a base period earnings requirement. In 2011, Washington required 680 base period hours.

Source: "Comparison of State Unemployment Insurance Laws," by the U.S. Department of Labor, 2011. Retrieved from: http://workforcesecurity.doleta.gov/unemploy/statelaws.asp a high-quarter earnings requirement, with a minimum earnings requirement within a single quarter, and some states require two quarters of positive earnings.

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 do. Thus, a low-earning worker who returned to the workforce eight months prior to a layoff could find himself or herself ineligible for UI, even though he or she had substantial earnings during that period. Because low-earning workers tend to have shorter job tenures than their more highly educated counterparts, the exclusion of recent wages is a more consequential issue for this group (General Accounting Office, 2000). Finally, in states that have high-quarter wage requirements, excluding the most recent completed quarter of earnings can mean excluding the quarter with the highest earnings (especially for workers with short or sporadic work histories), making the high-quarter wage requirement more difficult to satisfy.

The ABP

The ABP shifts the window during which earnings requirements are examined. Rather than excluding the most recently 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 to determine monetary eligibility for an unemployed worker who lost a job and filed for benefits in, for example, February 2013.

Under the standard base period illustrated in Figure 1, the hypothetical worker who filed in February 2013 would have his or her 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 addition, if the hypothetical worker satisfied the high-quarter wage requirement only between October and December 2012, he or she would be monetarily ineligible under the standard base

First Quarter	Second Quarter	Third Quarter	Fourth Quarter	Lag Quarter	Filing Quarter
October to December 2011	January to March 2012	April to June 2012	July to September 2012	October to December 2012	January to March 2013

Standard Base Period

	First Quarter	Second Quarter	Third Quarter	Fourth Quarter	Filing Quarter
October to December 2011	January to March 2012	April to June 2012	July to September 2012	October to December 2012	January to March 2013

Alternative Base Period

Figure 1. Comparison of Base Periods.

period (which excludes this quarter), but eligible under the ABP (which includes this quarter).

In most states, under the APB, UI applicants are given two chances to monetarily qualify: first using the standard base period and then using the ABP. By giving applicants two chances, and by allowing them to count more recent earnings, policymakers surmise that those who were on the margins of monetary eligibility under the old system will be more likely to qualify for, and thus receive, UI. Beyond the mechanical effect of increasing monetary eligibility rates among applicants, the ABP may also increase rates of UI application among those who would not have otherwise applied (O'Leary, 2011). As individuals learn about the new rules, they may think they are 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, Table 1 shows that minimum base period earnings requirements are low in most states. In Montana, the state with the median base period earnings requirement in 2011 of $2,305, a full-time worker working at the federal minimum wage would earn eligibility in roughly nine weeks. Even in North Carolina, which had the highest base period earnings requirement, a minimum-wage worker with full-time hours could achieve monetary eligibility after as little as four months of work. In two states (Washington and Oregon), UI 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 this, adoption of the ABP has been widespread: Between 1986 and 2008, 19 states and the District of Columbia adopted the ABP. In 2009, the American Recovery and Reinvestment Act (ARRA) sought to increase use of the ABP, making ABP adoption a necessary precondition for states to get any of their share of the $7 billion targeted at UI programs. By June 2011—the most recent year in our study period—36 states plus the District of Columbia had adopted the ABP. To date, 40 states and the District of Columbia use the 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 used state-specific administrative data or simulations based on survey data to estimate the impact of the ABP on monetary eligibility rates and UI receipt. Two studies looked only at monetary eligibility (Rangarajan & Razafindrakoto, 2004; Stettner, Boushey, & Wenger, 2005). Rangarajan and Razafindrakoto (2004) reported on simulations based on survey data from Mathematica's National Evaluation of the Welfare-to-Work Grants Program. They estimated that the ABP would increase monetary eligibility among women who left welfare for work by 4 to 9 percentage points. Stettner, Boushey, and Wenger (2005) used the Survey of Income and Program Participation (SIPP) and estimated that universal adoption of the ABP in the late 1990s and early 2000s would have increased monetary eligibility of separated workers by 6 percentage points, with low-wage workers disproportionately affected.

Vroman (2008) used UI administrative data from Ohio from 1967 to 2007. He found that in 2006 and 2007, more than 6 percent of UI claimants accessed the UI program through the ABP, and that over time applicants became familiar with the new program rules associated with this policy change. Ohio has relatively high monetary eligibility thresholds that may increase the impact of its ABP. Vroman's sample was also limited to UI applicants, and so cannot speak to the ABP's impact on previous nonapplicants. Finally, Vroman did not look at earnings records longitudinally, so he did not take into account applicants who accessed UI through the ABP but who—even in the absence of the ABP—would have reapplied and received UI 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. 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 looked at earnings records longitudinally, taking into account applicants who would access UI through the ABP, but who would have reapplied when their standard base period shifted and would have received UI even in the absence of the ABP.

On the whole, existing studies suggest that the 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, used nationally representative data to evaluate the effects of existing ABPs on UI receipt of low-earning unemployed workers. These studies were designed to assess monetary eligibility only; were confined to administrative data from one or a few states; or were based on simulation only. Further, these studies miss the behavioral effect of drawing previous nonapplicants into the program.

To evaluate the ABP's effectiveness in increasing UI receipt, it is necessary to use survey data, because administrative data generally do not include indicators for educational attainment, work hours, and other key demographic characteristics. Further, there is no nationally representative source of UI administrative data.

The current study used data from the CPS ASEC, stratifying a sample of unemployed workers by education level, full-time or part-time status, and other factors. We took advantage of the temporal variation in ABP implementation by 19 states plus the District of Columbia between 1986 and 2008, plus another 17 states that adopted the ABP between 2009 and 2011 in response to the UI modernization incentives included in the ARRA.[2] These states differ widely by region, population size and demographics, industrial base, and other factors. By using a parsimonious natural experiment design, we tested whether adoption of the ABP increases the probability that low-educated unemployed workers will access UI.

DATA AND METHODS

Data

CPS is a monthly survey of approximately 60,000 households that offers a nationally representative, multistage, stratified sample of the noninstitutionalized population. Detailed labor market and demographic data are collected on all respondents aged 16 years and older. The CPS ASEC 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 (IPUMS). In this series, data from the annual supplement, now called ASEC, between 1962 and

[2] California and Nebraska approved ABPs in 2011, but they were not implemented until after June 30, 2011. Because June 30 was our cutoff for inclusion in a calendar year, these two states were excluded from our sample.

2007 were integrated and variables were "harmonized" (coded identically) to be consistent over time (King et al., 2010).

The CPS ASEC offers a larger sample and more uniform data across the complete study period than other nationally representative surveys. Underreporting of public benefits in household surveys is a concern (Meyer, Mok, & Sullivan, 2009), and is a limitation of this study. It is possible that low-earning UI recipients might be the most likely to underreport benefits, because the amount and duration of benefits are likely to be the shortest among these recipients. Unfortunately, little is currently known about the characteristics of households most likely to underreport UI benefits. However, Meyer, Mok, and Sullivan (2009) found that the CPS ASEC reporting rate for UI benefit dollars (see the Appendix[3]) is the highest among peer surveys, averaging 79.2 percent across the period 1987 through 2007 (the latest year in their analysis), as opposed to 77 percent (with fewer years available) for the Panel Study of Income Dynamics (PSID); 74.5 percent for the SIPP; and 51.2 percent for the Consumer Expenditure Survey (CEX). Over the course of the study period, the CPS ASEC reporting rate first rose from 76.7 percent in 1987 to 94.5 percent in 1995, stayed high during the late 1990s, and then slowly fell during the 2000s. Thus, there is no consistent trend across the study period in changing rates of reporting. All told, the CPS ASEC was the most reliable source of data available for this study. Finally, use of a natural-experimental design with inclusion of year controls should mitigate concern that underreporting could drive the results of a study using this method, assuming that rates of underreporting do not covary with ABP adoption.

We restricted our sample to adults ages 18 to 64 who reported during a calendar year that they both (a) worked for pay and (b) 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 restricted our sample to workers with reasonable labor force attachment.[4] Our sample of 206,991 respondents included unemployed workers from all 50 states plus the District of Columbia who experienced unemployment between 1987 and 2011.

Hypotheses

The CPS ASEC data did not allow us to determine unemployed workers' nonmonetary eligibility status, nor whether they applied for UI. Our objective was to test whether ABP implementation was associated with increased UI receipt for low-educated unemployed workers, which would account for both the mechanical effect of increasing monetary eligibility among applicants and the behavioral effect of drawing those who would otherwise not apply into the program. Because the CPS ASEC data on education levels are better indicators of long-term earnings than reported earnings during a year with unemployment, we stratified by education when testing the following hypothesis:

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

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

[3] All appendices are available at the end of this article as it appears in JPAM online. Go to the publisher's Web site and use the search engine to locate the article at http://www3.interscience.wiley.com/cgi-bin/jhome/34787.

[4] Ideally, we would have dropped labor force entrants who transitioned from out of the labor force to unemployed and then became employed. The harmonized CPS ASEC data did not allow us to do this, and we could not connect our data to the basic CPS monthly survey. This is a limitation of our analysis.

Table 2. ABP implementation years by state, 1986 to 2011.

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
District of Columbia	2003
Hawaii	2004
New Mexico	2004
Virginia	2004
Oklahoma	2005
Illinois	2008
Montana	2009
Nevada	2009
West Virginia	2009
Alaska	2010
Arkansas	2010
Colorado	2010
Delaware	2010
Idaho	2010
Iowa	2010
Kansas	2010
Minnesota	2010
Oregon	2010
South Dakota	2010
Tennessee	2010
Maryland	2011
South Carolina	2011
Utah	2011

Notes: Dates are taken from the state comparison of UI laws, ARRA letters, and other published sources. When effective dates occurred after June 30, we rounded to the next calendar year. California and Nebraska approved ABPs in 2011, but the ABP was not implemented until after June 30, so these two states were excluded from our sample.

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

Models

To determine whether ABP implementation is associated with increased probability that a low-educated unemployed worker will receive UI, we took advantage of the natural experiment created by the gradual state-by-state implementation of the ABP. States that implemented the ABP between 1987 and 2011 are listed in Table 2.

We collected information on ABP implementation by state-year from a number of sources, including UI modernization letters, previous research on ABPs, and, in some cases, primary analysis of state laws. If ABP implementation in a given year occurred after June 30, we rounded its initial implementation to the next calendar year.

The ABP states vary on characteristics such as region, dominant industry, union density, and political orientation of state legislature. Roughly half the states listed in Table 2 adopted the ABP on their own prior to the announcement of ARRA's UI modernization provisions, which included incentives for use of the ABP. The other half of these states adopted the ABP between 2009 and 2011, likely electing to use the ABP in response to the ARRA incentives. Variation in the year of implementation over the 25-year study time period captures variation across economic cycles and changes in states' political and economic circumstances. This variation arguably creates a natural experiment: It is as if the state and year of implementation had been randomly selected. Further, because we used state and year controls, our approach is robust against spurious factors that could have influenced implementation decisions, such as state-specific levels of UI receipt and UI recipiency rates at a given point in the business cycle, unless these factors covary with ABP adoption.

To determine the effect of ABP use on the probability that an unemployed worker would receive UI, we used linear probability models. We also estimated our models using logistic regression, reported as average marginal effects, and the results were substantively similar (results available from the authors upon request). Our primary specification is

$$UI_{i,j,t} = \beta ABP_{j,t} + \lambda \mathbf{X}_{i,j,t} + \gamma_j + \theta_t + \varepsilon_{i,j,t}. \quad (1)$$

The dependent variable, *UI*, 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, *ABP*, is the key independent variable,[5] a dichotomous measure of whether the unemployed worker's state of residence *j* in year *t* used the ABP where 1 = ABP use and 0 = no use of ABP.[6] **X** is a vector of individual demographic characteristics and characteristics of the state in which the individual resides at time *t*. At the individual level, **X** includes a set of age dummies, a categorical measure of educational attainment, sex, race and ethnicity, marital status, and full-time work status while working. At the state level, **X** includes the state-year unemployment rate, state-year controls for minimum monetary eligibility thresholds (adjusted for inflation and divided by 1,000 to scale the results to be interpretable, in units of $1,000), which may covary with ABP adoption, and state minimum wages (adjusted for inflation), which may affect the likelihood of monetary eligibility for a low-earning worker. We further conducted 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 the point estimates associated with ABP and was not included in the main

[5] We captured ABP use for the individual's state of residence; however, whether ABP is used to determine an individual's monetary eligibility is determined by the state of employment. In instances where a respondent's state of employment and state of residence differed, and one state used the ABP and the other did not, our measurement is incorrect. This is a limitation of our measure. Vroman (2001) found that 4 percent of claims in 1998 were interstate claims. Interstate claims were concentrated in states where interstate commutes are common, for example, in the District of Columbia and New England.

[6] We also used a lagged ABP variable, but found that the results were substantively similar to the nonlagged variable.

specification because of concerns that it was imprecisely measured.[7] Finally, we include state-fixed effects (γ) and year-fixed effects (θ). ε represents unexplained variation.

With state and year controls included, the ABP indicator represents the effect of adoption of the 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 affect UI receipt. We used CPS ASEC person-level probability weights and clustered our robust standard errors by state.

We first ran the model on the full sample ($N = 206{,}991$), and then stratified the sample to determine the effect of ABP on subpopulations. We first stratified by education attainment: less than high school ($n = 40{,}947$), high school only ($n = 77{,}875$), some college ($n = 58{,}293$), and bachelor's degree or higher ($n = 29{,}876$). Within the population with less than a high school degree, we stratified further to previously full-time ($n = 29{,}032$) and previously part-time ($n = 11{,}915$) unemployed workers.

Later, we estimated an interacted specification on the full sample that mirrored our stratified models, and our results proved quite consistent. In our main specifications, we employed stratified models rather than models with interaction effects because the stratified models allowed us to more clearly interpret the subgroup effects and more precisely model confounding variables for subpopulations that have vastly different experiences in the labor market, without requiring numerous interactions. This was important because we were focused particularly on the impact of the ABP on low-educated unemployed workers, as the policy is not likely to affect more highly educated unemployed workers, who typically easily meet monetary eligibility requirements.

RESULTS

Effect of the ABP by Education Level

In Table 3, column 1 reports the coefficients for our model run on the full population of unemployed workers, and the remaining columns stratify the sample into subgroups by education level and full-time or part-time status when working.

For the full unemployed population, we did not find a statistically significant association between use of the ABP and the probability that an unemployed worker would receive UI. Columns 2 through 5 in Table 3 show that there was also no significant ABP coefficient when the sample was stratified by education level only.

In columns 6 and 7, we show the association between ABP use and UI receipt for two subgroups: unemployed workers with less than a high school degree who were employed part-time when working, and those who were employed full-time when working. While there was no significant association between the ABP and the probability that a worker with less than a high school degree previously employed full-time would receive UI, there was a significant effect for those who were

[7] There is no official source of information on adoption of part-time work search requirements by state-year, prior to information provided by the Employment and Training Administration 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 is, 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.

Table 3. Receipt of Unemployment Insurance by education level and part-time/full-time status.

Independent variable	(1) All unemployed workers	(2) Bachelor's degree and higher	(3) Some college	(4) High school graduates	(5) Less than high school degree	(6) Less than high school school degree, part-time	(7) Less than high school degree, full-time
ABP	0.00352	−0.000955	−0.00648	0.00709	0.0109	0.0275**	0.00327
	(0.00504)	(0.0115)	(0.00714)	(0.00904)	(0.0128)	(0.0119)	(0.0149)
Ages 26 to 35	0.148***	0.177***	0.170***	0.143***	0.111***	0.0800***	0.135***
	(0.00586)	(0.00746)	(0.00617)	(0.00688)	(0.00744)	(0.00989)	(0.00863)
Ages 36 to 45	0.217***	0.272***	0.227***	0.206***	0.191***	0.115***	0.228***
	(0.00630)	(0.0102)	(0.0100)	(0.00680)	(0.0107)	(0.0128)	(0.0119)
Ages 46 to 55	0.262***	0.294***	0.257***	0.252***	0.268***	0.156***	0.314***
	(0.00574)	(0.0104)	(0.00667)	(0.00651)	(0.0101)	(0.0144)	(0.0117)
Ages 55 to 64	0.286***	0.277***	0.259***	0.293***	0.320***	0.183***	0.381***
	(0.00504)	(0.0136)	(0.00968)	(0.00882)	(0.0144)	(0.0213)	(0.0172)
High school graduate	0.0556***						
	(0.00567)						
Some college	0.0519***						
	(0.00628)						
College +	−0.00530						
	(0.00855)						
Female	−0.0304***	−0.0478***	−0.0323***	−0.0347***	−0.00864	0.00964	−0.0147
	(0.00690)	(0.00580)	(0.00869)	(0.00739)	(0.00833)	(0.00606)	(0.0106)
Black	−0.0444***	0.0138	−0.0268**	−0.0691***	−0.0475***	−0.0142	−0.0626***
	(0.00942)	(0.0138)	(0.0115)	(0.0118)	(0.00880)	(0.00972)	(0.0124)
Native American	−0.0246**	0.0295	−0.00998	−0.0250	−0.0504***	−0.0623***	−0.0488***
	(0.0108)	(0.0381)	(0.0197)	(0.0171)	(0.0127)	(0.0143)	(0.0169)
Asian/Pacific Islanders	−0.0688***	−0.0725***	−0.0675***	−0.0816***	−0.0419**	0.0360*	−0.0712***
	(0.0145)	(0.0149)	(0.0167)	(0.0181)	(0.0207)	(0.0196)	(0.0244)
Multiracial	−0.00367	−0.123*	0.0114	−0.00511	0.0203	0.0429	0.00977
	(0.0193)	(0.0633)	(0.0534)	(0.0552)	(0.0229)	(0.0333)	(0.0265)
Other	−0.0227**	−0.0222	0.000185	−0.0605***	−0.00969	0.0213	−0.0323
	(0.00934)	(0.0277)	(0.0211)	(0.0166)	(0.0190)	(0.0295)	(0.0317)
Hispanic	−0.0653***	−0.0470***	−0.0435***	−0.0738***	−0.0697***	−0.0364***	−0.0755***
	(0.0110)	(0.00924)	(0.0113)	(0.0110)	(0.0130)	(0.00708)	(0.0163)
Married	0.0608***	0.00866	0.0484***	0.0789***	0.0832***	0.0481***	0.0915***
	(0.00395)	(0.00555)	(0.00560)	(0.00552)	(0.00901)	(0.00806)	(0.0111)
Full-time	0.192***	0.209***	0.194***	0.195***	0.164***		
	(0.00664)	(0.00915)	(0.00774)	(0.00984)	(0.00827)		
State-year unemployment rate	0.0127***	0.0128***	0.0134***	0.0120***	0.0106***	0.00630**	0.0126***
	(0.00122)	(0.00310)	(0.00218)	(0.00230)	(0.00234)	(0.00272)	(0.00268)
R^2	0.146	0.113	0.147	0.152	0.169	0.089	0.143
N	206,991	29,876	58,293	77,875	40,947	11,915	29,032

Notes: Models are linear probability models weighted using CPS ASEC population weights. Robust standard errors are listed in parentheses. Sample includes unemployed workers aged 18 to 64. All models also include state-year UI monetary eligibility requirements, state-year minimum wage levels, and state- and year-fixed effects, not reported.

*$P < 0.10$; **$P < 0.05$; ***$P < 0.01$.

Source: Authors' analysis of Current Population Survey data, by King et al., 2010.

employed part-time. We estimate that ABP adoption is associated with a 2.8 percentage point increase in the probability that an unemployed worker with less than a high school degree and employed part-time when working will receive UI. This is the only significant effect that we found associated with ABP implementation.[8]

The other covariates included in the models reported in Table 3 provide a consistency check with other studies. We found that older workers and full-time workers were more likely to receive UI. Unemployed workers with a high school degree or some college were significantly more likely to receive UI than those with less than a high school degree, but having a college degree was not associated with a higher probability of UI receipt, after controlling for other factors in the model. In most models, blacks and Asian/Pacific Islanders were less likely to receive UI than non-Hispanic whites; Hispanics were less likely to receive UI than non-Hispanic whites; and women were less likely to receive UI than men. Across most subgroups, married unemployed workers were more likely to receive UI than unmarried unemployed workers. The point estimates on the demographic characteristics in column 6 of Table 3 are mostly smaller than in column 7. The state unemployment rate was associated with increased probability of UI receipt across all seven columns in Table 3; however, interestingly, the effect size was smallest for part-time workers without a high school degree.

While we prefer the specifications in Table 3, we replicated our estimates using an interacted specification, which allowed us to test the statistical significance of the differential effects of ABP by education and work hours subgroup (reported in Table 3). The results are reported in Table 4, where we modified equation (1) as follows for all unemployed workers:

$$\begin{aligned} UI_{i,j,t} = {} & \beta ABP_{j,t} + \delta \text{ education-work hours}_{i,j,t} \\ & + \varphi ABP_{j,t} \times \text{ education-work hours}_{i,j,t} + \lambda \mathbf{X}_{i,j,t} + \varepsilon_{i,j,t}. \end{aligned} \tag{2}$$

We included a set of eight mutually exclusive dummies categorized by educational attainment and work hours. The dummies were (a) full-time, bachelor's degree; (b) full-time, some college; (c) full-time, high school graduate; (d) full-time, less than high school degree; (e) part-time, bachelor's degree; (f) part-time, some college; (g) part-time, high school graduate; and (h) part-time, less than high school degree. We then interacted these dummies with ABP, using "full-time, some college" as the reference category, and this interaction term allowed us to estimate the subgroup-specific effect of the ABP. To further mirror our stratified models, rather than state- and year-fixed effects, we included state $\times$ education-work hours and year $\times$ education-work hours interactions in vector **X**. These interacted fixed effects variables captured more precisely the state- and time-specific trends for workers in the eight subcategories.

Table 4 presents selected point estimates. ABP adoption was not associated with an increase in the probability of UI receipt for a full-time worker with some college. In fact, mirroring the results from our stratified model, the point estimates for seven of the eight education-work hours subgroups were not statistically significant. Only the point estimates on the interaction term ABP $\times$ part-time, less than high school degree was statistically significant, with these workers being 2.52 percentage points ($3.59 + -1.07$) more likely to receive UI when states adopted the ABP. This point estimate was substantively similar to the results for this subgroup reported in Table 3 (2.75 percentage points).

[8] In many states, workers who seek reemployment at part-time hours are ineligible to receive UI. Despite this nonmonetary barrier to UI receipt, it appears that the ABP is most helpful to this population.

Table 4. Comparison of Alternative Base Period (ABP) effects using an interacted model.

Independent variable	Point estimate
ABP (full-time, some college)	−0.0107
	(0.0090)
ABP × full-time college graduate	0.0036
	(0.0165)
ABP × full-time high school graduate	0.0184
	(0.0115)
ABP × full-time less than high school degree	0.0111
	(0.0152)
ABP × part-time college graduate	0.0270
	(0.0188)
ABP × part-time some college	0.0189
	(0.0146)
ABP × part-time high school graduate	0.0152
	(0.0121)
ABP × part-time less than high school degree	0.0359*
	(0.0147)
Test of linear restriction (*P* value)	
ABP × full-time college graduate = ABP × part-time less than high school	0.094
ABP × full-time high school graduate = ABP × part-time less than high school degree	0.347
ABP × full-time less than high school degree = ABP × part-time less than high school degree	0.095
ABP × part-time college graduate = ABP × part-time less than high school degree	0.712
ABP × part-time some college = ABP × part-time less than high school degree	0.321
ABP × part-time high school graduate = ABP × part-time less than high school degree	0.162

Notes: Models are linear probability models weighted using CPS ASEC population weights. Robust standard errors are listed in parentheses. Sample includes unemployed workers aged 18 to 64. All models also include demographic characteristics, state-year unemployment rate, state-year UI monetary eligibility requirements, state-year minimum wage levels, and state × education-work hours and year × education-work hours fixed effects, not reported.

*$P < 0.05$.

Source: Authors' analysis of Current Population Survey data, by King et al., 2010.

Table 4 presents a test of linear restrictions to examine whether the ABP point estimate for this subgroup was statistically significantly different from those of the other subgroups. In addition to the full-time, some college group, the point estimate for the part-time, less than high school degree group differed significantly from the interaction terms for full-time, college graduates and for full-time, less than high school degree group, but not for the other subgroups.

Table 5 reports on three additional sensitivity analyses. Panel A assesses whether the ABP point estimate was different for states that were induced to adopt the ABP by the UI modernization provisions included in the ARRA, as compared to states that chose to adopt the ABP without the ARRA incentive. For this analysis, we added to equation (1) an interaction term for ABP × ARRA-induced. ARRA-induced is coded as 1 if the state was induced to adopt the ABP by the incentives included in the ARRA and 0 otherwise. We reported point estimates for ABP and also ABP × ARRA-induced for the full model that included all workers, and for the model restricted to part-time workers with less than a high school degree. In neither case was the

Table 5. Alternative Base Period (ABP) effects sensitivity analyses.

Independent variable	Point estimate
Panel A: Restricted model, ARRA UI modernization inducement	
All unemployed workers	
ABP	0.0026
	(0.0050)
ABP × induced	0.0068
	(0.0097)
Part-time, less than high school degree	
ABP	0.0303*
	(0.0133)
ABP × induced	−0.0183
	(0.0286)
Panel B: Restricted model, ABP × monetary eligibility interaction	
All unemployed workers	
ABP	0.0088
	(0.0106)
ABP × monetary eligibility threshold	−0.0022
	(0.0040)
Part-time, less than high school degree	
ABP	0.0211
	(0.0243)
ABP × monetary eligibility threshold	0.0026
	(0.0110)
Panel C: Reduced model (with only ABP, state- and year-fixed effects)	
Group	ABP Point Estimate
All unemployed workers	0.0054
	(0.0073)
Bachelor's degree and higher	−0.0013
	(0.0116)
Some college	−0.0057
	(0.0097)
High school graduates	0.0147
	(0.0113)
Less than high school degree	0.0034
	(0.0124)
Part-time, less than high school degree	0.0317*
	(0.0135)
Full-time, less than high school degree	−0.0047
	(0.0135)

Notes: Models are linear probability models weighted using CPS ASEC population weights. Robust standard errors are listed in parentheses. Sample includes unemployed workers aged 18 to 64. Models in panels A and B also include demographic characteristics, state-year unemployment rate, state-year UI monetary eligibility requirements, and state-year minimum wage levels, not reported. All models include state- and year-fixed effects, not reported.

*$P < 0.05$.

Source: Authors' analysis of Current Population Survey data, by King et al., 2010.

ABP × ARRA-induced interaction term statistically significant. (For the part-time, less than high school degree group, the point estimate was negative and not insubstantial, although not remotely significant.) In addition, when we added the ABP × ARRA-induced interaction, the point estimate on the stand-alone ABP variable remained similar to the main model without the ARRA interaction, which suggests that the pre- and post-ARRA behaviors do not change in basic interpretation.

Panel B assesses whether ABP adoption had a larger effect as a state's monetary eligibility rules became more stringent. We did this by adding an interaction term ABP × monetary eligibility threshold to equation (1), with state-year monetary eligibility thresholds adjusted for inflation and divided by 1,000, so a one-unit change in a state threshold corresponds with a $1,000 increase in the state's inflation-adjusted minimum monetary eligibility threshold. We reported point estimates for ABP and the new interaction term ABP × monetary eligibility threshold for the model including all workers, and the model restricted to part-time workers with less than a high school degree. In neither case was the new interaction term statistically significant, although for part-time workers with less than a high school degree, the point estimates seem to indicate the theorized direction, with the ABP being associated with a bigger effect as the minimum monetary eligibility threshold rises.

Finally, panel C compares the ABP point estimates from our equation (1) model to a reduced model that included 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 essentially unmoved by the exclusion of the other covariates. In fact, that was what we found. In all cases, the point estimates in the reduced model were substantively similar to the one with the full set of covariates. This offers additional evidence that our specification captured a natural experiment.

INTERPRETATION OF RESULTS

This study has a number of limitations. First among them is the underreporting of UI benefits, which may bias point estimates either upward or downward. However, the CPS ASEC is currently the best possible choice of data available for a study like this, and the quasi-experimental design somewhat mitigates the possibility that underreporting is biasing our estimates: Unless reporting levels vary with ABP adoption, the amount of bias our estimates suffer should be minimal. However, should underreporting vary with education level or full-time or part-time status, the differences between the effect of the ABP for these groups will be under- or overstated, relative to other groups. Unfortunately, to our knowledge there is no definitive work[9] on how underreporting of UI benefits varies by our demographic characteristics of interest, and we are unable to speculate as to whether and in which direction our estimates may be biased. This is a limitation of our study and should serve as an impetus for future research to examine differences in reporting rates by sociodemographic characteristics.

Another limitation is that the CPS ASEC does not allow us to parcel out the mechanical effects (in increasing eligibility among applicants) versus the behavioral effects (in increasing the number of applicants) of ABP. But it is also true that this study is the first that has examined both effects together in a nationally representative sample, so this study goes the furthest to date in assessing the efficacy of the ABP.

Our primary results suggest that ABP implementation is associated with an increase in the probability that low-educated, part-time unemployed workers will receive UI benefits of 2.8 percentage points, significant at the 0.05 level. Taking data from the most recent year in our sample, 2011, we estimate that universal adoption of the ABP would be associated with approximately 24,000 additional unemployed part-time workers without a high school degree receiving UI than would be true in

[9] Meyer and Goerge (2011) found that lower-income households reported Supplemental Nutrition Assistance Program (SNAP, formerly called food stamps) receipt at higher rates than higher-income households. However, it is unclear whether this finding, which examined reporting behavior among a low-income population, can be generalized to the broad population of unemployed workers.

the absence of the ABP. However, we estimate the total number of workers who experienced unemployment over the course of 2011 to be 16 million. Our estimates therefore 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 fifteen hundredths of a percent of the total unemployed population.

Although our results suggest that the ABP would extend UI coverage to an appreciable number of unemployed workers, we find an effect only for unemployed part-time 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 statistically significant 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 employed part-time, sees any significant change in levels of UI receipt associated with the policy change.[10] Because part-time workers work fewer hours, and because less-educated workers have lower hourly wages, this group is likely to have the most difficulty achieving monetary eligibility. 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 the 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 (2008), the study was conducted in a state with high monetary eligibility requirements. Our study examined the full population of unemployed workers, including nonapplicants; thus, our estimates should be expected to be substantially lower than these studies. Further, as O'Leary (2011) points out, Vroman and others did not take into account that workers who fail to achieve monetary eligibility at the point of job loss may be able to apply again a few months later when their standard base period includes earnings from the more recent quarter. This may have biased upward previous estimates of the ABP's impact.[11] O'Leary used longitudinal records and does account for this possibility. He found that in Kentucky, which has a lower minimum monetary eligibility threshold, the ABP would increase the UI receipt of *applicants* by 2.21 percent. While statistically insignificant, the point estimate we report for the association between the ABP and UI receipt among all unemployed workers is in line with the O'Leary estimate after accounting for the differences in the samples (O'Leary's estimates are for applicants, while ours are for all unemployed workers).

Our study goes beyond previous evaluations to offer an estimate of the effect of actual ABP implementation on UI receipt at the population level. Our findings suggest that adoption of the 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 nonmonetary eligibility requirements and rates of application may be important barriers to UI access for the broad group of low-earning unemployed workers (Gould-Werth & Shaefer, 2012; Holzer, 2000; O'Leary & Kline, 2008; Rangarajan, Razafindrakoto, & Corson, 2002; Shaefer & Wu, 2011). Our findings suggest further policy change would be necessary to substantially expand UI coverage for the broad group of low-educated unemployed workers.

[10] We also ran our analysis for part-time workers who had a high school degree or more education and found no significant effect.
[11] O'Leary (2011) points out that a worker's base period is determined from the time of UI application. So, if a worker applies in June and is not eligible under the standard base period, but would be under the ABP, that worker may be able to reapply in July when his or her standard base period has shifted.

CONCLUDING COMMENTS

We found that the ABP policy change would increase access to UI for 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 can lead to shorter average job tenures than has been common historically (Government Accountability Office, 2007; Kalleberg, 2009). Our study suggests that the ABP is a helpful intervention for low-educated workers who work fewer hours than most workers and who are likely to have a short job tenure. The combination of working fewer hours and having a shorter work tenure may make amassing enough base period earnings to qualify for UI particularly difficult. Thus, 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 UI coverage continue to exist among the broader population of low-educated workers. Our results thus indicate that ABP implementation 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 nonmonetary 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 could be effectively coupled with the ABP. The U.S. economy has changed dramatically since the UI 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.

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APPENDIX

Table A1. Unemployment Insurance dollar reporting rates: The ratio of weighted total of UI benefits reported in dollars from four household surveys to administrative totals.

Year	PSID	SIPP	CPS ASEC	CEX
1987	0.724	0.821	0.767	0.618
1988	0.816	0.785	0.759	0.649
1989	0.759	0.752	0.762	0.564
1990	0.785	0.828	0.824	0.587
1991	0.738	0.867	0.851	0.644
1992	0.681	0.829	0.746	0.538
1993	0.719	0.876	0.798	0.561
1994	0.872	0.836	0.942	0.699
1995	0.693	0.808	0.945	0.591
1996	0.806	0.697	0.862	0.437
1997		0.636	0.863	0.495
1998	1.089	0.573	0.865	0.483
1999		0.631	0.768	0.429
2000	0.801	0.766	0.753	0.413
2001		0.642	0.806	0.389
2002	0.709	0.566	0.742	0.380
2003	0.435	0.619	0.736	0.436
2004	0.924	0.758	0.748	0.475
2005		0.870	0.717	0.441
2006			0.691	0.416
2007			0.679	
Average	0.77	0.745	0.792	0.512

Notes: Averages computed over years are included in this reproduction. PSID = Panel Study of Income Dynamics. SIPP = Survey of Income and Program Participation. CPS ASEC = Current Population Survey Annual Social and Economic Supplement. CEX = Consumer Expenditure Survey.

Source: Reproduced from "The under-reporting of transfers in household surveys: its nature and consequences," by Meyer, Mok, & Sullivan, 2009.