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GOVERNMENT OLD-AGE SUPPORT AND LABOR SUPPLY:
EVIDENCE FROM THE OLD AGE ASSISTANCE PROGRAM

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ABSTRACT

Many major government programs transfer resources to older people and implicitly or explicitly tax their labor. In this paper, we shed new light on the labor supply and welfare effects of such programs by investigating the Old Age Assistance Program (OAA), a means-tested and state-administered pension program created by the Social Security Act of 1935. Using Census data on the entire US population in 1940, we exploit the large differences in OAA programs across states to estimate the labor supply effects of OAA. Our estimates imply that OAA reduced the labor force participation rate among men aged 65-74 by 8.5 percentage points, more than half of its 1930-40 decline. But both reduced-form evidence and an estimated structural model of labor supply suggest that the welfare cost to recipients of the high tax rates implicit in OAA's earnings test were small. The evidence also suggests that Social Security could account for at least half of the large decline in late-life work from 1940 to 1960.

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1 Introduction

Many of the most important government programs, including Social Security and Medicare, transfer resources to older people and tax their labor relative to that of younger people.[1] Standard economic theory predicts that such programs reduce late-life labor supply and that the implicit taxation reduces the ex-post value of the programs to recipients. Understanding the size and nature of such effects on labor supply and welfare is an increasingly important issue, as demographic trends have increased both the potential labor supply of the elderly and its aggregate importance, while simultaneously increasing the need for reforms to government old-age support programs. This raises three important questions. What are the effects of government old-age support programs on late-life labor supply? What is the relative importance of the two key features of these programs—the transfers to older people and the taxation of their labor—in determining these effects? And to the extent that taxation of labor is important, how large are the associated welfare costs to recipients?

We address these questions by investigating Old Age Assistance (OAA), a means-tested program introduced in the 1930s alongside Social Security that later became the Supplemental Security Income (SSI) program. OAA was large both in absolute terms—22 percent of people 65 and older received OAA in 1940—and relative to Social Security, which made no regular payments until 1940 and remained smaller than OAA until 1950. OAA transferred resources to older people and, through an earnings test, implicitly taxed their labor relative to that of younger people. Unlike Social Security and other social insurance programs that are national in scope and near-universal in coverage, OAA was state-administered and exhibited considerable variation across states in eligibility and benefit levels, from very small to substantial programs. The combination of wide variation across states and, compared to more recent periods, the relative paucity of private pensions and other government programs targeting the elderly provides a promising opportunity to learn about the effects of these programs.

The particular setting we study is of special interest because it marked the beginning of a large expansion of government old-age support through OAA and Social Security that coincided with large declines in labor force participation among older men. Figure 1(a) illustrates these trends during the early expansions of these programs, from 1920 through 1970. The striking time-series correlation between the expansion of the Social Security program after 1950 and declining labor force participation is often noted in discussions of Social Security and retirement (e.g., Feldstein and Liebman, 2002; Krueger and Meyer, 2002; Gruber, 2013; Coile, 2015). But, as the same authors note, there is still significant uncertainty about the causal relationship between the two trends. Moreover, during this

[1] See, for example, Goda, Shoven and Slavov (2009) and Gelber, Jones and Sacks (2017) on Social Security and Goda, Shoven and Slavov (2007) on Medicare.

period both OAA and Social Security implicitly taxed late-life earnings in a highly salient way. OAA payments were typically reduced dollar for dollar with earnings, and the Social Security earnings test withheld benefits from people earning more than a small amount, without any compensating increase in future benefits.[2] The extent to which the earnings tests reduced the value of OAA and Social Security is an unaddressed yet critical determinant of the welfare effects of the mid-century growth in government old-age support.

Our analysis combines large policy variation with recently-released data on the entire US population from the 1940 US Census. The large sample size of this dataset and its precise geographic information enable a wide range of empirical tests that would have been difficult or impossible with previously available data. Our main empirical tests make use of two sources of variation. The first is age eligibility requirements, which almost always limited eligibility for OAA to people 65 and older. Importantly, other modern-day programs that use age 65 as a cutoff, including Social Security, were either small or non-existent at the time, and private pensions were still relatively uncommon. Figure 1(b) shows that while labor force participation declined fairly continuously around age 65 in 1930, by 1940 the decline at age 65 was slightly sharper than at other ages and by 1960 the decline at age 65 was quite pronounced. The second source of variation is cross-state variation in payment and eligibility levels of OAA programs. The empirical analysis tests whether there is a greater reduction in labor force participation after age 65 in states with larger OAA programs relative to states with smaller programs.

Our estimates indicate that OAA significantly reduced labor force participation among older individuals. The basic patterns that we explore in the data are evident in Figure 2, which plots male labor force participation by age, separately for states with above- and below-median OAA payments per person 65 and older. Up to age 65, the age pattern of labor force participation was extremely similar in states with larger and smaller OAA programs. At age 65, however, there was a sharp divergence in labor force participation between states with larger OAA programs relative to those with smaller programs, and this divergence continued at older ages. Our regression results, which isolate variation in OAA program size due to state policy differences, imply that OAA can explain more than half of the large 1930–40 drop in labor force participation of men aged 65–74.

Although we find large effects of OAA on labor force participation, both reduced-form and structural results suggest that the welfare cost to recipients of OAA's implicit taxation of work was small. Combining our reduced-form results with an estimate of overall OAA recipiency for men aged 65–74 implies that 52 percent of OAA recipients in this group changed

[2] For example, between 1939 and 1950 Social Security withheld benefits from people earning more than $15 per month. In 2010 dollars, this amount corresponded to about $230 in 1939 and about $135 in 1950.

their labor force participation status in response to OAA. Assuming that the remaining 48 percent of recipients valued their benefits fully, a dollar of OAA benefits was worth at least \$0.48 of unconditional late-life income to the average recipient. Moreover, the effects of OAA were concentrated among men with low potential earnings. Labor force exit in response to OAA was greater for men with lower levels of education, and nearly half of the reduction in labor force participation from OAA was due to exit from unemployment or from employment in work relief programs that were targeted at individuals with poor labor market prospects.

In order to better understand the effects of OAA on labor supply and the value of OAA to recipients, we go beyond our reduced-form results by estimating a model of lifetime labor supply and retirement. Identification of the model comes from the pattern of bunching of retirements at the OAA eligibility age, at which OAA's earnings test creates a convex kink in the lifetime budget constraint, across different earnings groups. By comparing the extent of bunching among different earnings groups, we can jointly estimate eligibility, which we do not observe in the data, and preferences. The estimates suggest that 22 percent of the male population was eligible for OAA in 1940, and simulations indicate that, had all men aged 65–74 in 1940 been eligible for OAA, OAA would have reduced their labor force participation rate by 21 percentage points. Further simulations of the model indicate that, although OAA's earnings test explained close to half of its labor supply effects, recipients valued their benefits highly. The average dollar of OAA benefits was worth about \$0.95 of unconditional late-life income. The average value of OAA benefits is high for two reasons. First, the effects of OAA were concentrated among people with low potential earnings, for whom the cost of meeting an earnings test is smaller. Second, many people were inframarginal in the sense that they would have retired either before or relatively soon after the OAA eligibility age even without OAA, or even if OAA did not impose an earnings test.

In the final section of the paper, we ask what both the reduced-form results and the estimated model suggest about the role of government old-age support—and of Social Security in particular—in the growth of retirement over the mid-20th century. Although OAA was targeted at the poor elderly and had very different eligibility and payment determination rules from those of Social Security, OAA and Social Security shared the core features of transferring resources to older people while taxing their labor relative to that of younger people. Our results suggest that Social Security had the potential to drive at least half—and likely more—of the mid-century decline in late-life labor supply for men.

This paper relates to literatures on the labor supply and welfare effects of government old-age support and means-tested transfer programs. A large body of work has investigated the effects of Social Security and other government old-age support programs on labor supply and retirement (for reviews, see Diamond and Gruber, 1999; Feldstein and Liebman, 2002;

Krueger and Meyer, 2002; Coile, 2015). Our paper is especially related to a branch of this literature on the labor supply effects of OAA (Parsons, 1991; Friedberg, 1999) and to work on the role of government old-age support programs in the mid-20th century rise in retirement (Boskin, 1977; Moffitt, 1987; Parsons, 1991; Friedberg, 1999; Gelber, Isen and Song, 2016).[3] It is also especially related to a branch of this literature that has sought to decompose the labor supply effects into those due to income transfers and those due to changes in marginal incentives to work associated with the tax and benefit rules (e.g., Burtless and Moffitt, 1985; Friedberg, 2000; French, 2005; Gelber, Isen and Song, 2016; Gelber, Jones and Sacks, 2017). Our work on OAA, a means-tested program, also relates to other work on the labor supply effects of means-tested transfers. A large literature, reviewed by Moffitt (2003) and Ziliak (2015), has investigated the labor supply effects of Aid to Families with Dependent Children (AFDC) and Temporary Assistance to Needy Families (TANF). A much smaller literature (Neumark and Powers, 2000, 2006) has investigated the labor supply effects of Supplemental Security Income (SSI) for the aged, the successor program to OAA and the modern program that OAA most closely resembles.

Relative to the earlier literature, one important contribution of this paper is that the combination of our setting and approach allows us to credibly estimate the *full* labor supply effects of OAA programs, from essentially no program to the largest of those observed in 1940. Although the full labor supply effects of government old-age support and means-tested programs are crucial determinants of the welfare effects of such programs, most research on both old-age support programs and welfare programs is limited to the more modest aim of estimating effects of marginal changes in program parameters.[4] The existing literature has provided substantial evidence on the effects of marginal changes in Social Security benefits on retirement, for example by examining changes in retirement timing associated with changes in eligibility ages (e.g. Atalay and Barrett, 2015), by exploiting different treatment under program rules of otherwise similar people (e.g., Coile and Gruber, 2007; Liebman, Luttmer and Seif, 2009), and—less often—by using within-country differences in pension programs (Baker and Benjamin, 1999). But this literature has focused on data from the last several decades, a period in which benefit levels are high, multiple programs affect incentives to retire at specific ages, and there is significant bunching of retirements at certain ages. Any one of these factors alone would make extrapolations to zero inherently speculative.

[3] Although the time series relationship over the mid-20th century is striking, Costa (1998) and Lee (1998) note that attachment to the labor force among men 65 and older had already declined significantly between 1880 and 1910, before OAA and Social Security were established. Studying Union Army pensions and retirement in the first decade of the 20th century, Costa (1995) suggests that rising incomes could account for much of the rise in retirement over the 20th century.

[4] For example, in his review of research on the TANF program and its predecessor the AFDC program, Moffitt (2003) writes, "Probably the major methodological problem with these estimates is the obvious one that they are not based on any data in which AFDC was literally absent, but rather are extrapolations from estimated effects of the existing, positive level of AFDC benefits down to a benefit level of zero."

Another key contribution of this paper is to shed light on the welfare effects of government old-age support programs by using transparent sources of identification, and a variety of empirical methods, to estimate the cost to recipients of OAA's earnings test. Although the cost to recipients of earnings tests is a crucial determinant of the welfare effects of means-tested programs, relatively little work has sought to estimate this cost directly. To the best of our knowledge, the cost of the earnings test is completely unaddressed by papers on the early years of OAA and Social Security and is rarely addressed even in the vast literatures on government old-age support and means-tested programs, including Temporary Assistance to Needy Families and SSI. Those papers that do address this issue rely almost entirely on calibrated structural models, whereas our setting enables us to implement approaches based on both reduced-form methods and an estimated structural model.[5] Our finding that the cost to recipients of OAA's earnings test is small is notable in light of OAA's large effects on labor supply and its high implicit taxation of earnings.

2 Background on the Old Age Assistance Program

Old Age Assistance (OAA) was introduced alongside Old Age Insurance (OAI), which came to be known as Social Security, in the Social Security Act of 1935. OAA provided federal matching grants for state-administered, means-tested old age support programs for the low-income elderly. Social Security was initially small—it made no payments until 1940, and even then to less than two percent of the elderly. But the introduction of OAA led to a major and rapid expansion in government old-age support. In 1929, just seven states had laws providing for assistance to the elderly. By 1940, every state had an OAA program, and about 22 percent of people aged 65 and older received OAA payments. The average annual OAA payment was \$232 (\$3,615 in 2010 dollars), about 25 percent of 1939 median wage and salary earnings for 60–64-year-olds earning a wage, and slightly over half of the 25th percentile of wage earnings. OAA was by far the largest source of old-age support at this time, greatly exceeding both Social Security and employer pensions. Only in the 1950s did Social Security become the larger program, as cohorts "aged in" and legislation expanded eligibility and benefits.

[5]Friedberg (2000), using an estimated structural model, estimates the deadweight loss of the Social Security earnings test on the intensive margin for working beneficiaries during the period spanning the late 1970s and early 1990s. During this period, the taxation implicit in the earnings test was smaller than it was during the middle of the 20th century, the period on which we focus. Laitner and Silverman (2012) is another rare example of a welfare analysis of Social Security policies in an estimated model. Large literatures analyze the welfare effects of Social Security in calibrated computable general equilibrium models (e.g., Auerbach and Kotlikoff, 1987) and estimate the income and substitution effects of means-tested programs on labor supply using structural models (for a recent review of the literatures on several means-tested programs, see Moffitt, 2016).

States had a great deal of discretion in the design and administration of their OAA programs, and both the share of the elderly receiving OAA and average payments varied widely. Figure 3 shows county-level data from U.S. Social Security Board (1940*c*) on total OAA payments in the month of December 1939, scaled by the population 65 and older in the 1940 Census. Many state borders exhibit stark differences in payments per elderly person, which suggests that different state policies led to large differences in payments for individuals in similar circumstances.

In general, OAA programs were set up as either an income floor or a consumption floor (the latter of which takes into account all resources when determining payments), both of which implicitly tax recipients' income at a 100 percent rate, as benefits are phased out dollar-for-dollar with income. In practice, state or local OAA administrators evaluated the "needs" and "resources" of each applicant. The excess, if any, of needs over resources determined the size of the payment, up to a maximum level. In some states the level of "needs" could vary across people, while in others a common dollar amount was used.[6] In the analysis, we use measures of maximum payments to approximate variation across states in the level of the income or consumption floor. The statutory maximum monthly payment was $30 in most states ($470 in 2010 dollars), which was the federal matching cap, but ranged from $15 to $45, with eight states having no statutory maximum. The states that had no statutory maximum had a small number of very high payments, but the 99th percentile of payments were well in line with other states' legal maxima.[7]

State OAA laws specified a variety of eligibility criteria. All states required that an applicant have little income and be of at least a certain age, nearly always 65. Nearly all states had residency requirements. Many states also imposed asset tests and restricted eligibility to US citizens or long-term residents. Some states required that an applicant have no legally responsible relatives able to provide support. As was common in public assistance programs of the time, relief officials retained a significant amount of discretion in determining eligibility. For example, Lansdale et al. (1939) note some moralistic provisions in OAA laws (e.g., limiting eligibility to the "deserving") and emphasize the significant variation in how strictly certain eligibility requirements, such as relatives' responsibility laws, were enforced.[8] As a result, features of OAA other than statutory eligibility criteria affected recipiency rates (see, e.g., Fetter, 2017). These potential influences on recipiency will be relevant in the structural estimation in Section 6, in that they are a reason we estimate eligibility from behavior rather

[6] See Appendix Section A.2.1 for more detail. Appendix Figure A1, discussed in more detail there, shows distributions of payments in selected states to illustrate the differences across states in whether "needs" were set at a common level or varied across people.

[7] See Appendix Table A1, which reports basic features of each state's payments; also see Appendix Figures A2 and A3, which show maps of legal maximum payments and 99th percentile payments, respectively.

[8] Although not the focus of this paper, the discretion of OAA administrators left significant scope for discrimination (see, e.g., Quadagno, 1988, on racial discrimination in the South).

than observing it directly.

Table 1 shows summary statistics at the state level on recipiency and payments in December 1939. Variation in recipiency rates and benefits per recipient, which were not strongly correlated across states, generated wide variation in average OAA benefits: Payments per person 65 and older varied more than 13-fold across states.

3 Theoretical Predictions

The simplest model for understanding how OAA might affect the labor supply and welfare of recipients is a model of the lifetime budget constraint relating total lifetime consumption to the length of retirement, as illustrated in Figure 4.[9] OAA expands the set of consumption-leisure opportunities available to potential OAA recipients by paying recipients $\bar{y}$ for each period they do not work after the OAA eligibility age. OAA has an income effect that tends to hasten retirement and, for people who would retire after the OAA eligibility age if OAA benefits did not depend on earnings, a substitution effect that also tends to hasten retirement.

By reducing the private return to work after the OAA eligibility age but not before, OAA introduces a convex kink in the lifetime budget constraint at that age. For retirement ages younger than the OAA eligibility age, working an additional year increases total lifetime consumption by the full amount of earnings, w. For retirement ages older than the OAA eligibility age, working an additional year increases total lifetime consumption by the excess, if any, of earnings over the OAA benefit level, $\max\{0, w - \bar{y}\}$. OAA therefore imposes an implicit marginal tax on earnings after the OAA eligibility age, with implicit tax rate $\tau = \min\{1, \bar{y}/w\}$. With a smooth distribution of preferences for consumption versus leisure in the population, such a convex kink attracts more people than nearby allocations on the budget constraint. This prediction of a "bunching" of retirements at the eligibility age underlies both our reduced-form and structural estimation strategies.

The extent to which the earnings test affects labor supply and the value of the program to recipients depends critically on the distribution of people along the budget set. People who retire earlier are less affected by the earnings test. As long as leisure is non-inferior, someone who would retire before the OAA eligibility age even without OAA is unaffected

[9]This framework is better-suited to analyzing those OAA programs that provide income floors than those programs that provide consumption floors, since the latter might distort the timing of consumption. We assume throughout that OAA recipients do not bear any of the burden from the taxes necessary to finance the program. This is likely a good approximation for the cohorts we study, who finished most of their working years before OAA was introduced.

by the earnings test; his optimal retirement age with OAA, which is weakly earlier than it is without OAA, is before the earnings test reduces benefits. Someone who would work only a short time after the eligibility age without OAA is at most slightly affected by the earnings test.

4 Data and Empirical Approach

4.1 Data

The key data source that enables many of our empirical tests is the full-population microdata from the 1940 Census. In addition to the large size of the sample, an advantage relative to previously available datasets is precise geographic location, which enables empirical tests that would not otherwise be possible. We focus on men aged 55 to 74 in states in which the OAA eligibility age was 65 in 1939.[10] We restrict the sample to men with non-missing information on basic demographics (birthplace, race, citizenship status, marital status, and years of education). Our main analysis focuses on work behavior at the time of the 1940 Census. Some additional analyses use information on work and income outcomes in 1939. For each set of outcomes, we exclude from the sample those men with missing information on work (or income) outcomes in the relevant year.[11]

Table 2 describes the characteristics of the men in our sample. About 71 percent were in the labor force, and 65 percent were employed. An important component of overall employment in the late 1930s was "public emergency" employment—employment through one of the federal programs that provided work-based relief to the unemployed, such as the Work Projects Administration (WPA) (see, e.g., Fishback, 2007). About 62 percent of men in our sample were employed in either private or non-emergency government work and about 4 percent were employed in public emergency work. About half reported receiving any wage or salary income in 1939. (Those with no wage or salary income would include both those who did not work and those who worked but were self-employed.) Including those who reported zero wage and salary income in 1939, the average reported income was

[10] Three states—Missouri, New Hampshire, and Pennsylvania—had an OAA eligibility age of 70 in 1939 but reduced the eligibility age to 65 on January 1, 1940 to meet a requirement to continue receiving federal matching funds. The fact that the age requirement was changed just a few months before the 1940 Census complicates including them in the sample. We also exclude Colorado, in which long-term residents became eligible at age 60.

[11] The share of individuals missing data varies across states, but the variation in OAA policies that we use in the main specifications is not significantly associated with the probability of being dropped from the sample for 1940 outcomes; nor is it significantly associated with the probability of being dropped from the sample for 1939 outcomes, with the possible exception of older ages (70 and higher). These results are shown in Appendix Figure A4.

\$557 (corresponding to about \$8,672 in 2010 dollars). The only information on income from sources other than wages and salaries is whether such income totaled \$50 or more (about \$780 in 2010 dollars). Just over half of our sample reported receiving at least \$50 of such income. It may be helpful to note that vital statistics data suggest that in 1940, remaining life expectancy of men reaching age 65 was about 12 years (Grove and Hetzel, 1968).

As discussed in more detail below, our primary specifications limit comparisons to counties on either side of a state boundary. The "border county" sample excludes counties that did not border at least one other state in our sample. As shown in Table 2, a comparison of means across the full and border county samples indicates only small differences between the two, which suggests that inferences drawn from the border county sample likely apply to the population as a whole.

We use data on OAA payments and recipients from a variety of sources. We do not have individual-level data that includes OAA recipiency, although we show some results using the measure of non-wage income in the 1940 Census data, which includes OAA. As a summary policy measure, we use state OAA payments in December 1939 per person in the state aged 65 and older. The OAA data for this measure come from U.S. Social Security Board (1940*b*), which reports monthly data on total OAA dollar payments and the number of recipients at the state level. Geographically, the finest data we have on recipiency and payments is at the county level, from December 1939, as used in Figure 3. Neither of these sources contains detail on recipients' characteristics. We do, however, have state-level tabulations on new OAA recipients and the payments approved for these recipients for each fiscal year from 1936 through 1940, from research memoranda of the Social Security Board (U.S. Social Security Board, 1939*a*,*b*, 1941). We use this information for two purposes: first, to determine how high payments tended to be in states without statutory maximum payments, and second, to estimate an OAA recipiency rate for the demographic group we study (men aged 65–74) as opposed to the whole population 65 and older.

4.2 Empirical approach

We use two key sources of variation to investigate the effects of OAA. The first is the age-eligibility requirement, nearly always limiting assistance to people 65 and older.[12] The second key source of variation is the heterogeneity in state policy discussed in Section 2.

[12] Most states did not have mandatory birth certificates for the cohorts in our sample, so in addition to birth certificates a range of other records, such as marriage records and school records, were used to determine age-eligibility (Lansdale et al., 1939). Although misreporting or inaccurate knowledge of age is a potential concern (Ransom and Sutch, 1986; Elo and Preston, 1994), in Appendix Table A3, we find no evidence that our OAA variation is associated with the share of the population reporting they are 65 or older.

Combining variation in age-based eligibility and state policy, we test for differential changes in the age profile of labor force participation after the eligibility age, controlling flexibly for any age-specific effects common across states and for the possibility that state OAA policies were correlated with unobserved determinants of the level of labor force participation-age profiles.

A useful aspect of our context is that OAA was by far the largest source of old-age support for which 65 was a cutoff age in 1940. In more recent periods, changes in behavior at or around age 65 could be associated with any of a number of factors, such as eligibility for Social Security or Medicare. However, Medicare did not exist until 1965, and as noted in Section 2, Social Security made no monthly payments until 1940, and even then to a very small share of the elderly. Other public or private pensions at the time made payments to a significantly smaller share of the elderly than OAA did and were primarily relevant for people higher in the income distribution than OAA recipients. Appendix Section A.3.2 offers more detail.

Because our main specifications use age eligibility for identification, we do not directly identify any anticipatory effects of OAA on labor supply before the eligibility age; our estimates are net of any such effects. Differential trends across states in the age profile of labor force participation will provide some indication of the likely size of such anticipatory effects, however, and will also speak to the relative size of the net-of-anticipatory effects between the young elderly (those just turning 65) and older individuals.

We follow a simulated instruments strategy, in the spirit of Currie and Gruber (1996), to capture variation in observed levels of OAA driven by policy rather than by population characteristics. Using the earnings distribution for a national population of men aged 60–64, the oldest ineligible age group, we simulate payments per person 65 and older treating a state's maximum payment as an income floor and incorporating any earnings disregards.[13] The basic idea is that a state's maximum payment is correlated with its typical income floor and is not itself affected by labor market conditions or population characteristics. For the eight states with no legal maximum payment, we measure variation in income floors using the 99th percentile payment among recipients accepted in fiscal year 1938–39 in each state (based on information in U.S. Social Security Board, 1939*b*).[14] In all but a few cases, the

[13]The national population we use for each state omits the state itself, although in practice this makes little difference. Earnings disregards existed in only a few states and were always at low levels. For the purposes of the simulated instrument, we impute earnings for the self-employed by drawing from the distribution of earnings for wage earners with the same level of education and the same number of weeks worked.

[14]The idea is that with payments equal to the gap between "needs" and "resources," payments near the top of the distribution tend to reflect payments to individuals with virtually no resources—present in every state—and therefore likely reflect administrative norms or rules. We use the 99th percentile payment rather than the observed maximum payment because the latter could be driven by quite exceptional situations.

99th percentile payment is the same as the legal maximum in those states that had legal maxima (see Appendix Table A1). Appendix Figure A5 illustrates some of the variation used in the simulated instrument, showing the distribution of monthly earnings for the national population against the minimum, median, and maximum levels of the maximum payment (plus any earnings disregards) across states. Despite using only some of the eligibility and payment criteria, the resulting instrument is clearly predictive of realized payments per person 65 and older at the state level, as can be seen in Appendix Figure A6.[15] Appendix Section A.2.3 provides full details of the construction of the instrument.

The potential endogeneity of OAA policy is a concern if differences in policy were correlated with other, unobserved determinants of the shape of labor force participation-age profiles (any correlation between policies and the level of labor force participation-age profiles is differenced out by the estimation). We investigate the relationship between realized OAA payments and state-level demographics and income in Appendix Table A2. OAA payments per person 65 and older tended to be greater in higher-income states, suggesting that these states tended to have more generous policies. Greater elderly population shares, greater foreign-born shares, and lower non-white population shares also correlated with higher OAA levels. We discuss these results in more detail in Appendix Section A.3.1. To address potential policy endogeneity, in all of our main specifications we restrict comparisons to counties lying on either side of the same state border. The idea is to compare individuals who have similar characteristics and who are in the same labor market (and hence subject to similar shocks), but who face different OAA policies. Appendix Table A3 indicates that in unconditional comparisons using county-level data, similar correlations are evident between population characteristics and both realized and simulated payments. But when comparisons are limited to state boundaries by adding a fixed effect for the set of counties that touch each state border, these correlations disappear. This result supports the identification assumption that in the absence of OAA, the shape of labor force participation-age profiles would be similar in counties on either side of a state boundary.[16]

In our main specifications, for individual i living in state s and county c, we estimate equa-

[15] Appendix Figure A6 also shows, in some cases, considerable variation in observed payments for states with the same simulated payment. This is partly because assessed "needs," and hence the income floor, could vary across people in some states. Consistent with this, payments per recipient also exhibit some dispersion across states with the same maximum payment (see Appendix Figure A7). The variation is also partly due to important determinants of eligibility for OAA, such as those discussed in Section 2, that cannot be mapped into characteristics we observe in Census data. We discuss these issues at greater length in Appendix Sections A.2.2 and A.2.3.

[16] In principle, these comparisons could be refined further, for example by comparing only adjacent counties on either side of a state border. Because specifications using the entire state border pass covariate balance checks and placebo tests, we focus on these less computationally costly specifications. For all of the core results of the paper, however, using county pairs yields very similar results.

tions of the form

$$y_{iacsb} = \beta_c + \delta_{ba} + \sum_{a \neq \bar{a}} \gamma_a * \log(\text{payments per person 65+})_s + \Lambda' \mathbf{x}_{iacs} + \varepsilon_{iacsb}, \tag{1}$$

where a indexes age (either in single years or groups of years), $\bar{a}$ is a reference age, and $\mathbf{x}_{iacs}$ is a vector of controls. The variable of interest, $\log(\text{payments per person 65+})_s$, is the log of the December 1939 OAA payments per person 65 and older in state s. We instrument for this variable using the log simulated payment per person 65 and older. Appendix Table A4 reports first stage regressions.[17] We limit the sample to counties lying on state borders and define a border segment between two states (indexed by b) as the set of all counties in either state that touch the boundary between the two. Since some counties border two or more different states, a county, and hence all the individuals in it, will appear in the data as many times as there are states that it borders. The border segment-by-age fixed effects then limit comparisons of age profiles to men living on either side of the same border. Since our policy of interest varies at the state level, we cluster standard errors at the state level. This level of clustering also accounts for the duplication of observations in counties lying on multiple state boundaries.

5 The Effect of OAA on Labor Supply

5.1 Baseline results

Estimates of equation (1) indicate that larger OAA payments led to sharper declines in labor force participation after age 65. For our primary results we group ages into 5-year age bins, with ages 60–64 as a reference age. Table 3 reports both OLS and IV estimates. We focus our discussion on the (preferred) IV specifications; the OLS results follow similar patterns. All specifications indicate that states with larger OAA programs featured larger reductions in labor force participation after the OAA eligibility age. To demonstrate the importance of the border-by-age fixed effects, columns (1) and (2) show estimates from a variant on equation (1) replacing these with age fixed effects, in the full and border county samples. Interpreted at face value, these specifications suggest that larger OAA programs led to a greater reduction in labor force participation from ages 55–59 to ages 60–64. In principle, these reductions prior to eligibility could reflect anticipatory effects of OAA. But the fact that they disappear after including border-by-age fixed effects in column (3) suggests both

[17] All models are just-identified, so bias from weak instruments is unlikely to be a problem (see, e.g., Angrist and Pischke, 2009).

that these controls are important and that they are effective in controlling for underlying differences in age trends in labor force participation between states with larger and smaller OAA programs.[18] Our preferred specification, column (4), further controls for interactions of education and race with age. The coefficients suggest that a one standard deviation increase in log payments per person 65 and older (about 0.62 log points) was associated with a 3.9 percentage point reduction in labor force participation at ages 65–69, and a 4.7 percentage point reduction in labor force participation at ages 70–74. These results are broadly similar to those of Friedberg (1999) and Parsons (1991) (see Appendix Section A.5).

Panel (a) of Figure 5 plots the coefficients on the age-payment interactions when we estimate (1) using interactions with single years of age. There is little evidence of large differential age trends prior to age 65, but at age 65 states with larger OAA programs exhibit a sharp, differential decline in male labor force participation that levels out around age 69. Panel (b) shows estimates from the same specification using the closest measure to a "first stage" outcome available in the 1940 Census, an indicator for receipt of at least \$50 (about \$780 in 2010 dollars) of income from sources other than wages and salaries in 1939. Receipt of non-wage income follows a pattern analogous to that of labor force participation. It does not appear to trend differently prior to age 65 in states with greater expected OAA payments but shows a sharp, differential increase at age 65. Appendix Table A5 shows results for receipt of non-wage income analogous to those in Table 3.

To illustrate the relationship between exit from the labor force at age 65 and receipt of non-wage income, and the degree to which these outcomes varied across states, we take a simple approach to estimating "excess" exit from the labor force and receipt of non-wage income at age 65. We estimate the following model, separately by state:

$$y_i = \beta_0 + \beta_1 \mathbf{1}(\text{age}_i \geq 65) + \beta_2(\text{age}_i - 65) + \beta_3(\text{age}_i - 65)\mathbf{1}(\text{age}_i \geq 65) + \varepsilon_i \qquad (2)$$

where the outcome is either labor force participation or receipt of non-wage income.[19] In Figure 6 we plot the estimated breaks at age 65 from equation (2) for receipt of non-wage income against the estimated breaks in labor force participation. Declines in labor force participation line up strikingly well with increases in receipt of non-wage income. The results also illustrate the substantial variation across states in the overall drop in labor force

[18] It is also noteworthy that the first stage F-statistics indicate that our simulated instrument has a stronger relationship with observed payments when comparisons are restricted to state borders. This likely reflects greater predictive power of policy for observed outcomes when examining more similar populations.

[19] The 1940 Census has information only on age in completed years at the time of the Census, meaning that individuals who were 65 at the time of the Census may or may not have been eligible for OAA during 1939, the time period covered in the non-wage income question. Hence, in estimating the break in non-wage income we omit 65-year-olds. Note that for the primary outcomes in our analysis—work behavior during the week preceding the 1940 Census—age will be correctly observed for nearly all people.

participation at age 65, from nearly zero in Arkansas to 15 percentage points in Oklahoma.

In order to assess OAA's contribution to the large decline in labor force participation of the elderly over the 1930s, we re-estimate the state-border specification using the level rather than the log of OAA payments per person 65 and older. The results, reported in Appendix Table A6, are comparable to our main estimates at the mean level of OAA payments. The results imply that were it not for OAA, labor force participation among men aged 65–74 would have been 8.5 percentage points greater, about 17 percent of this group's observed labor force participation rate in 1940 of 51 percent. As another point of comparison, between 1930 and 1940 the labor force participation of men aged 65–74 fell from about 65 to 51 percent (see Appendix A.1 for details on comparability of labor force participation rates between these years). Our estimates suggest that OAA can explain about 60 percent of this 14 percentage-point decline.

Under slightly stronger assumptions, we can estimate what the age profile of labor force participation would have been in the absence of OAA in a way that allows for potential anticipatory effects. We compare labor force participation profiles across state boundaries, and assume that if OAA levels were the same in two states, the levels of labor force participation would be the same on either side of the boundary, rather than the weaker assumption in our main analysis of parallel age trends. Formally, we estimate

$$y_{iacsb} = \alpha_{ba} + \sum_{a} \gamma_a * (\text{payments per person 65+})_s + \varepsilon_{iacsb} \qquad (3)$$

with no age omitted from the summation. The counterfactual age profile of labor force participation with payments per person set to zero is shown in Figure 7. It is noteworthy that using this approach, we find reductions in labor force participation after age 65 similar to our main estimates, and any anticipatory effects of OAA on labor supply before age 65 appear to be quite small. As we discuss further in Section 6, counterfactual labor force participation in the absence of OAA is an important determinant of the cost of the implicit taxation of earnings by OAA's earnings test.

5.2 Heterogeneity by potential earnings

Also important for understanding the cost of OAA earnings tests are the potential earnings of the people OAA induced to exit the labor force, since people with lower potential earnings have lower costs from the earnings test. Panel (a) of Figure 8 shows the age 65–69 estimates from equation (1) separately by grouped years of education. The effects of OAA tended to be greater for men with less education, who tend to have lower potential earnings. Panel (b) shows the results of a similar exercise that uses a broader range of characteristics to predict earnings. In a sample of 45–54 year old men who were not self-employed, we regress wage and salary earnings on indicators for each number of years of education; for being non-white; for state of birth and for foreign birth; and for each of the 1,000 most common first names, exploiting the fact that first names contain information about socioeconomic status (Olivetti and Paserman, 2015). We then use the coefficients to predict earnings for men aged 55–74 and estimate equation (1) separately by decile of predicted earnings. The results line up well with those based on education alone, showing greater effects for lower deciles of predicted earnings.

Separating the effect of OAA on labor force participation into effects for employed and unemployed individuals provides further evidence that the effect of OAA on labor force participation was concentrated among people with poor labor market prospects. To the extent that the effects of OAA on labor force participation were driven by exit from unemployment or work-based relief programs such as the WPA, the costs of OAA work disincentives would be smaller. Table 4 shows estimates of equation (1) using overall employment (including work-based relief) as an outcome variable, as well as employment in private or public non-emergency work. (Results by single years of age are shown in Appendix Figure A8.) Comparison of the point estimates for different outcome variables suggests that at ages 65–69, about 21 percent of the reduction in labor force participation was associated with exit from unemployment and about 29 percent with exit from public emergency work. These results also suggest that the implicit taxation of work by OAA's earnings test may not have significantly reduced the value of OAA benefits to recipients.

5.3 Responses by local labor market conditions

An important question for assessing the generalizability of the results is whether the effect of OAA on labor force participation was larger in 1940 than it would have been in a context of lower unemployment. By the late 1930s unemployment had fallen significantly from its peak in the early 1930s, but it remained fairly high (Margo, 1993). In theory, the effects of a program like OAA on labor force participation could be larger or smaller in a context of high unemployment. On one hand, people might be more likely to exit the labor force in response to OAA if they are unemployed than if they are employed. On the other hand, high unemployment means that many people with weak attachments to the labor force might be out of the labor force regardless of OAA. To shed light on this question, we leverage the substantial geographic variation in labor market conditions during this time period (Wallis, 1989; Rosenbloom and Sundstrom, 1999). We calculate county unemployment rates for men aged 45–54 and estimate equation (1) separately for counties in the bottom quartile (which had an average of 5 percent unemployment) and the top quartile (an average of 20 percent unemployment). As shown in Appendix Figure A9, we find that reductions in labor force participation immediately after the eligibility age were somewhat greater in high-unemployment counties, and there are small reductions slightly before the eligibility age, around age 63, in these counties as well. The latter finding may reflect a greater propensity of men in high-unemployment counties to lose jobs shortly before reaching OAA eligibility and then to remain out of the labor force in anticipation of receiving OAA. Overall, however, it is striking how similar the effects of OAA are across widely varying labor market conditions.[20]

5.4 Further robustness checks

Given that relatively few states had OAA programs in 1930, a natural further check on the validity of our estimates is to test whether, conditional on comparisons across state borders, OAA policies in 1940 were systematically related to differential age patterns of labor force participation in 1930. If differential age trends in the underlying propensity to exit the labor force were driving our results, it is likely that we should see similar patterns in 1930. We estimate equation (1) using the 1930 complete-count Census data to test whether observed payments in 1940 predict labor force outcomes in 1930. We omit the nine states in our 1940 sample that had laws providing for old-age payments in 1930. Panel (a) of Figure 9 shows that there was no systematic difference in work behavior in 1930 after age 65 in states that had larger OAA programs in 1940. For comparison, Panel (b) shows estimates from 1940 using the same sample of states. The results are very similar to our main estimates, so the absence of differences in 1930 is not an artifact of using a different set of states.

We address various additional potential concerns in robustness checks reported in the appendix. Our main simulated IV estimates use the 99th percentile of observed payments in 1938–39 as a measure of "maximum" payments in states without legal maxima, based on the notion that these payments reflect typical standards of need. The typical concern with an observed-payments measure is that it could be contaminated by reverse causation. In practice, while population and labor market characteristics likely affect average OAA payments in a state, they are unlikely to affect payments as high up in the distribution as the 99th percentile, since it is extremely likely that there will be some people with no "resources" in every state. But as an alternative that does not use any information on observed payments, we have also tried assigning these eight states the highest legal maximum across states (45 dollars per month, in Colorado), based on the idea that the absence of a maximum payment

[20] Appendix Table A7 reports specifications with linear interactions between OAA benefits and the unemployment rate. The coefficients on the interaction terms are positive, which interpreted at face value suggest smaller effects in places with higher unemployment, and mostly statistically insignificant at standard levels.

means that payments in these states could have been as high as those in any other state. The results, reported in Appendix Table A8, are also in line with our main estimates. Given that there was significant variation across these eight states in the highest (and typical) payments actually granted, the first stage is weaker, with an F-statistic of 2.6 for the regressions describing labor market outcomes.

In principle, the comparison across state borders does not address concerns about differences in state policies other than OAA that relate to either old-age pensions or public assistance. In practice, introducing controls for other major types of pensions—railroad pensions or state and local government employee pensions—leads to very little change in the results, as documented in Appendix Table A9. Differences in state general assistance do not explain the results either, as shown in Appendix Table A10. These results are discussed in significantly more detail in Appendix Section A.3.3.

Finally, in Appendix Table A11 we test whether more generous OAA policies were associated with greater probability of having migrated between 1935 and 1940. As we discuss in more detail in Appendix Section A.6, the confidence intervals suggest that endogenous migration is at most an order of magnitude smaller than our labor force participation results. This result is unsurprising given the residency requirements for OAA in nearly all states.

6 The Cost to Recipients of OAA's Earnings Test

The overall welfare effect of OAA depends critically on the extent to which the large reductions in labor supply we have documented entail a social cost.[21] A key dimension of this question is the degree to which OAA's earnings test reduced its value to recipients. To the extent that a recipient "earned" his OAA benefit by reducing his earnings in response to the earnings test, his valuation of the benefit would be less than its full amount. This section presents evidence on the cost to recipients of OAA's earnings test based on our regression results and an estimated structural life cycle model.

6.1 Bounding the cost of the earnings test

Our regression results suggest a natural upper bound on the cost of the earnings test. Economic logic suggests that OAA benefits that would have been received even in the absence of

[21]Other important factors include the deadweight loss from raising taxes to finance the program, risk and the insurance benefits of OAA, and fiscal externalities (e.g., from people substituting to OAA from other government programs such as unemployment insurance and the WPA).

a behavioral response should be valued fully. The estimates reported in Section 5.1 suggest that about 8.5 percent of men aged 65–74 were out of the labor force because of OAA in 1940. No data directly measure the recipiency rate specifically for this group, but as detailed in Appendix Section A.4, a conservative estimate is about 16.5 percent. Hence, about 48 percent of recipients in this group were inframarginal, in the sense that they would have received OAA even without leaving the labor force in response to OAA. Under the assumption that these inframarginal recipients valued their OAA payments fully, this calculation suggests that the average recipient valued his OAA benefits at between 48 and 100 percent of their dollar amount.

Several considerations suggest that the average recipient valued his OAA payments significantly more than this lower bound. As shown in Section 5.2, a large fraction of marginal recipients were people who had low potential earnings or who had been unemployed or receiving work-based public assistance. The earnings test is not very costly to people with poor labor market prospects, in part because someone with lower potential earnings forgoes less consumption by not working.

Obtaining a point estimate requires making several assumptions, and in the next section we do so by estimating a life cycle model. To develop further intuition for the results, however, it is useful to consider an alternative approach to constructing a bound, in the spirit of Afriat (1967) and Varian (1982). This bound is based largely on the extent to which benefits are inframarginal over the life cycle for any particular individual. As an extreme example, someone who would have retired before the OAA eligibility age even in the absence of OAA is not affected by the earnings test as long as leisure is non-inferior, since he would retire before the eligibility age even if OAA did not have an earnings test. More generally, the earlier someone would leave the labor force in the absence of OAA, the lower the maximum possible cost of the earnings test. Furthermore, to the extent that income effects of OAA pull retirement forward, meaning that recipients would retire earlier even if OAA did not impose an earnings test, they tend to diminish the cost of the earnings test.

To construct the bound we use consumer theory to derive the minimum equivalent variation of OAA for someone with a given level of potential earnings, OAA benefit level, and retirement age in the absence of OAA. We introduce the assumptions that falling levels of labor force participation at older ages reflect retirement—that is, that exit from the labor force is an absorbing state—and that the counterfactual age profile of labor force participation shown in Figure 7 can be interpreted as reflecting the distribution of retirement ages that would arise in the absence of OAA. We combine this latent retirement age distribution with the observed joint distribution of state OAA benefit levels and earnings for men aged 48–52. Calculations making conservative assumptions about rates of take up across the earnings-benefit distribution suggest that within the class of preferences in which utility is quasilinear in retirement—the usual case in many applications of the life cycle model—the average \$1 of OAA was worth at least \$0.72 of unconditional late-life income. Details of this and related calculations are in Appendix Section A.7. Intuitively, a large fraction of OAA recipients would have retired either before or soon after the OAA eligibility age even without OAA, which limits the degree to which the earnings test can reduce the average value of OAA benefits to recipients.

6.2 Estimating the cost of the earnings test

In this section, we estimate the cost to recipients of the earnings test using an estimated life cycle model. The estimation targets empirical moments that provide transparent identification of the key parameters that govern the model's predictions about the counterfactuals of interest; in this sense, our general approach is broadly similar to that of, for example, Laitner and Silverman (2012). We adopt a standard model that is widely-used in diverse settings. This enables us to assess the extent to which the key determinants of behavior in other settings have similar effects in our setting as well.

6.2.1 Model

Consider a standard model of lifetime labor supply in which people choose how much to consume at each date and when to retire. Individual i at age t chooses whether to retire, if he has not already, and how much to consume in order to maximize the discounted sum of utility from age t forward,

$$U_{it} = \sum_{s=t}^{T} \beta^{s-t} \left(\frac{c_{is}^{1+\eta}}{1+\eta} - \delta_i \mathbb{1}(h_{is} = \bar{h}) \right), \ \eta \leq 0,$$

subject to a constraint on hours of work, $h_{is} \in \{0, \bar{h}\}$ (so there is only an extensive-margin labor supply decision), and a dynamic budget constraint,

$$a_{it+1} = (1+r)(a_{it} + N_{it} + \hat{w}_{it} h_{it} + b_{it} - c_{it}) \geq 0.$$

$\beta \in (0, 1]$ is the discount factor. The absolute value of η is the coefficient of relative risk aversion. δ_i is the disutility of work, which is allowed to vary across individuals, with cumulative density function $F(\delta)$. $c_{it} \geq 0$ is consumption, and $h_{it} \in \{0, \bar{h}\}$ is hours of work. a_{it} is assets, N_{it} is non-labor income, $\hat{w}_{it}$ is the wage, $\hat{w}_{it} h_{it}$ is labor earnings, and b_{it} represents OAA payments. The requirement that assets must be non-negative rules out borrowing.

We consider an OAA program that provides an income floor of $\bar{y}_{it}$ to individual i at age t:

$$b_{it} = \max\{0,\ \bar{y}_{it} - \hat{w}_{it} h_{it}\},$$

where $\bar{y}_{it}$ is the OAA benefit available to individual i in the period in which i is t years old. If individual i is not eligible for OAA at age t for any reason—due to being too young, to not meeting an asset test, or to having relatives who are able to support him—then $\bar{y}_{it} = 0$. This is a simplified version of a typical OAA program in 1940.[22]

Potential earnings, $w_{it} \equiv \hat{w}_{it}\bar{h}$, are constant in real terms over the life cycle. Everyone is assumed to have positive potential earnings, which tends to work against our main results. OAA benefits are fixed at their real values in 1940. People learn about OAA in 1936, a year in which many state OAA programs were introduced. Because assets and non-labor income are measured only coarsely in the data, we make the simplifying assumptions that initial assets when the individual enters the labor market at age 21 are zero and that OAA is the only source of non-labor income. Individuals live to age 75 with certainty ($T = 75$). Any assets the individual accumulates earn a constant real return of 3 percent per year, $r = 0.03$. The individual discounts future utility at this same rate, $\beta = \frac{1}{1+r} = \frac{1}{1.03} \approx 0.97$.

A challenge we face is that an individual's eligibility for OAA depended on many characteristics that are not available in the Census or alternative sources of data, which means we have to infer eligibility rather than measure it directly. In our baseline case, we assume that the probability that a randomly-chosen individual with potential earnings w_i is "eligible" for OAA—by which we mean that he would receive a positive OAA benefit if he met the age requirement and had low enough earnings—is a piecewise-linear function in the individual's potential earnings,

$$Pr(\text{eligible}_i | w_i) = \max\{0,\ \min\{1,\ \alpha_e + \beta_e w_i\}\}.$$

This measure of eligibility—which is exclusive of the minimum age restriction and the earnings test—is meant in part to approximate the many other restrictions that individuals must have met in order to qualify for OAA, including any requirements related to citizenship, residency, housing wealth, and relatives' characteristics. In addition to being exclusive of the age and earnings tests, this notion of eligibility further departs from the standard one in that it bundles together many things that are conceptually distinct and not necessarily

[22]The key simplification here is the assumption that there is a common income floor across people. As described in more detail in Appendix Section A.2.1, in many states there was significant variation in benefits even for people with no other source of earnings. This fact suggests that these states either had consumption floors rather than income floors or had income floors with heterogeneous income levels.

related to the usual meaning of the word "eligibility." For example, it includes un-modeled factors, such as incomplete information and stigma, that limit take up of OAA benefits. By definition, in the simulations take up is universal among people who are "eligible," old enough, and have low enough earnings.

Additional details about the model and estimation beyond those described in the text are in Appendix Section A.8.

6.2.2 Estimation strategy

The key tradeoff in the model is between consumption and the length of retirement. By retiring later, the individual can enjoy more consumption over his lifetime at the expense of less leisure (a shorter retirement). The marginal cost of retiring later is the disutility of work, δ_i, which is allowed to differ across individuals. The marginal benefit of retiring later is the extra consumption that retiring later buys, $\max\{0,\ w_{it} - \bar{y}_{it}\}$, valued at the marginal utility of consumption, c_{it}^{η}, which depends on the curvature of the utility function, η. OAA potentially affects both elements of the marginal benefit of retiring later. OAA reduces the consumption benefit of work after the eligibility age through its earnings test, and OAA reduces the marginal utility of consumption at any given retirement age through its income transfers. Both of these factors tend to lead people to retire earlier as a result of OAA.

These considerations guide our estimation strategy. We estimate the model using the Method of Simulated Moments, with the target moments measuring the extent of "bunching" of retirements at the OAA eligibility age. Intuitively, the estimation is based on comparing the extent of "bunching" of choices (retirement ages) at convex kinks in the budget set (the OAA eligibility age) that differ in their sharpness (due to different people facing different replacement rates from OAA, $\bar{y}/w$). As discussed in Section 3, OAA creates a convex kink in the lifetime budget constraint relating lifetime consumption to retirement length, and this kink is sharper for people who face higher replacement rates from OAA, $\bar{y}/w$. With a smooth distribution of disutility of work in the population, $F(\delta)$, such convex kinks lead to bunching of retirements at the OAA eligibility age, as some of the people who would have retired somewhat after the OAA eligibility age in the absence of the earnings test choose to hasten their retirements due to the substitution effects of the earnings test. The amount of bunching is informative about the level of eligibility for OAA. The greater the observed bunching, the greater the inferred eligibility. The speed with which the bunching "fades out" as the replacement rate declines is informative about the curvature of utility, η. The faster the fade out, the greater the curvature of utility, $|\eta|$.

This estimation strategy is related to the bunching strategy of Saez (2010). The most important difference is that some of the bunching in our context could be due to binding borrowing constraints, since the kink is in the intertemporal, not intratemporal, budget constraint. This complication prevents us from identifying structural parameters without estimating a full structural model. Although estimating a full structural model requires us to make stronger assumptions than must be made in simpler contexts, it has the advantage of enabling us to analyze a larger set of counterfactuals. Incomplete eligibility in our approach serves the same conceptual role as optimization frictions in the recent bunching literature (e.g., Chetty et al., 2011; Chetty, 2012; Kleven and Waseem, 2013; Gelber, Jones and Sacks, 2017). Although in general those who bunch may not be representative of the broader population of interest, in our context much of the labor supply response to OAA is driven by people who retire at the eligibility age, making them unusually representative of the broader set of people whose labor supply is affected by the program. Other research that analyzes the bunching of retirement ages includes Burtless and Moffitt (1984), Brown (2013), and Manoli and Weber (2016).

In principle, all of the key parameters of the model can be identified from the bunching of retirements at the OAA eligibility age. But intuition suggests and estimations confirm that the slope of the eligibility-potential earnings relationship, β_e, is not well identified using this information alone. Intuitively, the estimation has a hard time distinguishing between two potential sources of fadeout in the bunching of retirements in potential earnings: curvature in the utility function (η) and declining eligibility with potential earnings (β_e). This motivates a two-stage estimation procedure. In the first stage, we estimate the slope of the eligibility-potential earnings relationship, β_e, using the empirical relationship between earnings and housing wealth; we use housing wealth because it was the key non-income determinant of eligibility in many states that is measured in Census data. In the second stage, we estimate the key preference parameters, η and $F(\delta)$, and the level of the eligibility-potential earnings relationship, α_e, taking as given the first-stage estimate of β_e. The key results are robust to estimating all of the key parameters of the model in a single step, as shown in Appendix Table A12, but the two-stage procedure yields a more reasonable estimate of the relationship between eligibility and potential earnings.

The second stage of the estimation is based on the Method of Simulated Moments. The target moments are the proportional breaks at the OAA eligibility age in the labor force participation-age profile for each of 15 different earnings groups. We weight each moment by the inverse of its variance; more-precisely estimated moments receive greater weight in the estimation. Specifically, we estimate η and α_e by attempting to match the pattern of bunching of retirements at the OAA eligibility age, while requiring that the model also match the counterfactual distribution of retirement ages in the absence of OAA, which non-parametrically identifies $F(\delta)$. The key assumptions are that any heterogeneity in retirement behavior among people who face the same budget constraint is due to heterogeneity in the disutility of labor and that all earnings groups have the same counterfactual no-OAA retirement distribution.

6.2.3 Empirical inputs

"Bunching" of retirements by potential earnings level.— The main target moments in the estimation are measures of the bunching of retirements at the OAA eligibility age of 15 different earnings groups. Potential earnings, w, are not observed for those out of the labor force, so we approximate the bunching of retirements of different earnings groups using changes in the distribution of earnings at the OAA eligibility age. We continue to make the assumption, introduced in the second bound described in Section 6.1, that exit from the labor force at these ages is an absorbing state. We create separate indicator variables for reporting 1939 wage and salary income of \$1–100, of \$101–200, and so on in multiples of \$100 (\$1550 in 2010 dollars) for a total of 15 groups. We then estimate equation (2), reproduced below, with an indicator for each level of earnings as a separate dependent variable.[23] In order to focus on situations as similar as possible to the model, we estimate these regressions using data from Massachusetts only, a state whose OAA program appears to have closely approximated an income floor with a common level for all individuals. Our key results, as we discuss in Appendix Section A.8.1, are robust to estimating the model using data from other states or the full US.

$$y_i = \beta_0 + \beta_1 \mathbf{1}(\text{age}_i \geq 65) + \beta_2(\text{age}_i - 65) + \beta_3(\text{age}_i - 65)\mathbf{1}(\text{age}_i \geq 65) + \varepsilon_i \tag{2}$$

Under two assumptions, the share of men of a given potential earnings level who retire upon reaching 65, conditional on working up until age 65, equals β_1/β_0. First, we assume that for someone who worked in 1939, his actual earnings are his potential earnings, so that someone either earns his potential earnings or zero. Second, because the available measure of earnings does not include self-employment earnings, we also assume that self-employment is independent of earnings and responds in the same way to OAA as does wage and salaried employment. Note that this approach does not difference out factors other than OAA that may have induced retirement at age 65, but since these other sources were significantly smaller than OAA and relevant primarily for individuals higher in the income distribution (see Appendix Section A.3.2), this is likely of modest importance and would tend to work against our key findings.

[23] We use a uniform kernel and the Imbens and Kalyanaraman (2012) approach to select a bandwidth separately for each dependent variable. The results are robust to alternative choices on these dimensions. We omit 65 year olds from the regression since, observing a person's age only at the time of the Census (in April 1940), we cannot determine whether a given 65-year-old would have been eligible for OAA in 1939.

Estimation of equation (2) provides further evidence that the effects of OAA on labor supply were concentrated among those with low potential earnings. Appendix Figure A10 shows estimates of β_1/β_0 at each earnings level along with estimates of β_1. The point estimates suggest that at levels of potential earnings up to $800 per year, about 20 percent more men left the labor force at age 65 than would have been expected based on general trends in labor force participation by age. Appendix Figure A11 shows underlying shares by age for amounts up to $1,000. There are clear breaks in the underlying shares at age 65, and these breaks diminish at higher earnings levels.

Counterfactual retirement ages in the absence of OAA.— To identify $F(\delta)$, we require the estimated model to match the counterfactual distribution of retirement ages in the absence of OAA. We use the counterfactual age profile of labor force participation estimated in Section 5.1 together with the assumption that the observed cross-sectional relationship between labor force participation and age is a good proxy for the unobserved life-cycle relationship. Because the distribution of people along a non-linear budget constraint plays a key role in determining how policy changes affect behavior (e.g., Moffitt, 1986), the ability to identify the counterfactual distribution of people along the lifetime budget constraint on the eve of the major mid-20th-century expansions in Social Security greatly facilitates an understanding of the role of government old-age support programs in reducing late-life work at this time.

The relationship between earnings and housing wealth.— We use the relationship between earnings and housing wealth, the main determinant of eligibility other than income and age available in the Census data, to estimate the slope of the eligibility-potential earnings relationship in the first stage of the estimation. The estimate of the slope of the eligibility-potential earnings relationship, $\hat{\beta}_e$, is the slope of the empirical relationship between earnings and the fraction of people in Massachusetts with house values below the Massachusetts OAA eligibility threshold of $3,000. The underlying assumption is that the slope of the relationship between potential earnings and eligibility for OAA based on house value alone equals the slope of the relationship between potential earnings and eligibility for OAA based on all determinants of eligibility. Appendix Figure A12 shows the share of Massachusetts men aged 60–64 with less than $3,000 of house value as a function of wage and salary earnings. In addition to limiting OAA eligibility to people with less than $3,000 of equity in real property, Massachusetts imposed additional eligibility requirements as well, so the actual share eligible was less than the share that would be eligible based on the property test alone.

6.2.4 Estimation results and validation

The estimation is well-behaved and yields plausible results. Appendix Figure A13, the objective function, shows that the parameters are well-identified (see Appendix Section A.8.2 for details). The coefficient of relative risk aversion is 1.3 ($\hat{\eta} = -1.3$), indicating slightly greater risk aversion than the log utility benchmark ($\eta = -1$) at which income and substitution effects of wage changes exactly offset one another. This is within the usual range reported in the labor supply literature, despite the particular characteristics of our setting. Roughly 22 percent of the male population is estimated to be eligible for OAA, with eligibility declining from about 36 percent among those with the lowest potential earnings to about 13 percent among people with potential earnings of $2,000. Fitting a linear relationship suggests that no one with potential earnings greater than about $3,163 would have been eligible. Both the level of eligibility and its slope with potential earnings seem reasonable based on OAA eligibility rules and recipiency rates. The relatively low eligibility rate suggests that the effects of OAA would have been significantly larger had everyone been eligible for the program. If everyone had been eligible for OAA, the model predicts that OAA would have reduced the labor force participation rate in 1940 of men aged 65–74 by 21.5 percentage points, whereas our main reduced-form estimates imply a reduction of 8.5 percentage points. The results of this estimation and several robustness tests appear in Appendix Table A12.

The estimated model matches key features of the data well. Figure 10 compares the empirical and simulated moments, proportional breaks in labor force participation-age profiles at the OAA eligibility age for people with different potential earnings levels. The model matches the empirical pattern that increases in the probability of retiring at the OAA eligibility age are concentrated among groups with low potential earnings—primarily those between $0 and about $900 (about $14,000 in 2010 dollars), about 2.5 times the Massachusetts OAA benefit of $360.[24] The probabilities of retiring at the OAA eligibility age, especially the probabilities among groups with low potential earnings, pin down the level of the eligibility-potential earnings relationship. If everyone were eligible for OAA, nobody whose potential earnings were less than the OAA benefit level would work past the OAA eligibility age, since doing so would give up leisure for no gain in consumption. In this case, the probability of retiring at the OAA eligibility age conditional on not retiring before that age would be one—about four times the observed probabilities among groups with low potential earnings. The model infers from this pattern that even among groups with low potential earnings,

[24] The one wage group where the simulated and empirical moments match poorly is the lowest, those with potential earnings between $1 and $100 per year (about $1550 in 2010 dollars). Although we treat this as a valid wage group, it likely reflects incorrectly reported income or people with other sources of income in addition to a small amount of wage income. Given its large standard error, this moment has little effect on the results, and in any case works against the key finding of large responses at low earnings levels. The results are robust to dropping all of the low-earnings moments up to the OAA benefit of $360.

no more than about one-third of individuals were eligible for OAA in the sense we defined above.

As a validation exercise, we simulate labor force participation in 1940 of everyone aged 55–74. This requires an additional empirical input: the joint distribution of potential earnings and OAA benefits. In each state, we use the observed distribution of earnings in 1940 among people aged 48–52 with positive earnings together with the OAA benefit level in 1940. (Details are in Appendix Section A.9.) We find, in Figure A14, that the simulated labor force participation-age profile matches its empirical counterpart closely. Simulations of the model predict that OAA should have reduced labor force participation among 65–74-year-olds by 6.3 percentage points in 1940, not far from the main reduced-form estimate of 8.5 percentage points.[25] Overall, the results suggest that the model can provide a useful benchmark for understanding the value and labor supply effects of OAA in 1940 and for predicting the effects of Social Security during the middle of the twentieth century.

6.2.5 Effects of the earnings test on the labor supply and welfare of recipients

What was the role of OAA's earnings test in reducing labor supply, and how much did it reduce the value of the program to recipients? We address these questions by using the estimated model to simulate the behavior and outcomes of a particular cohort of the US population—that aged 55 in 1940—under a variety of budget constraints based on state OAA programs in existence in 1940. Details are in Appendix Section A.10.

Simulations of the model suggest that about half of the overall effect of OAA on labor force participation was due to its earnings test. Figure 11 shows the simulated age profile of labor force participation under three scenarios: no OAA, actual OAA, and a counterfactual unconditional OAA program, i.e., a program that pays the same fixed benefit regardless of the individual's current earnings. The results indicate that about 46 percent of the reduction in labor supply among men aged 65–74 was due to OAA's earnings test.

Despite the importance of the earnings test in reducing labor supply, further results suggest that the earnings test had little effect on the value of OAA benefits to recipients. Simulations of the model indicate that the average \$1 of benefits was worth about \$0.95 of unconditional late-life income, on the upper end of the bounds calculated using our regression results. The

[25] We also estimate an alternative version of the model with perfect capital markets and find that it is highly inconsistent with the pattern of bunching of retirements at the OAA eligibility age. With perfect capital markets, the simulated breaks in the earnings distribution at the OAA eligibility age among groups with very low potential earnings are zero, since everybody in these groups who is eligible for OAA retires strictly before age 64. The much better fit of the model with borrowing constraints is additional validation of the model given the poor functioning of household credit markets at the time (see, e.g., Rose, 2014).

average cost to recipients of the earnings test was small for two main reasons. First, as discussed in Section 6.1, a large fraction of OAA benefits were inframarginal; they would have been received even if recipients did not adjust their labor supply in response to OAA. The model implies that 49 percent of benefit-years were inframarginal without any behavioral response and 72 percent were inframarginal after income effects. Second, labor supply responses to OAA were highly concentrated among people with poor earnings prospects, mainly because they faced the highest replacement rates.

The conclusion that the cost to recipients of the earnings test is small is extremely robust to a wide range of alternative empirical inputs and assumptions. Appendix Table A12 reports results based on several alternative specifications, and Appendix Section A.8.1 discusses these and other results at length. For example, the results are highly robust to targeting the bunching of retirements observed in populations other than that of Massachusetts alone, including that of California, which also had a clear income floor with a common level for everyone, and of the United States as a whole.[26] Our main conclusions are also robust to making a wide range of alternative assumptions about non-labor income, lifespan, discounting, asset returns, and eligibility for OAA. The key role of inframarginal benefits in limiting the cost of the earnings test means that the main effect of many possible changes in the model would come through any effects on the counterfactual distribution of retirement ages in the absence of OAA. But because we force the estimation to match this distribution directly, the analysis is not very sensitive to changes in assumptions about the underlying determinants of retirement ages in the absence of OAA. Furthermore, our key conclusions are robust to larger-than-plausible changes in the distributions of counterfactual retirement ages and potential earnings, the primary determinants of the cost of the earnings test.

This robustness also speaks to the external validity of our conclusions. For example, while the earnings test was likely less costly because of the poor labor market conditions around 1940, using the (higher) 1950 earnings distribution and the (also higher) 1930 age profile of labor force participation increases the average cost only slightly, to $0.07.

7 Social Security and the Rise in Retirement

Government old-age support expanded dramatically from 1940 to 1960 in both recipiency rates and benefit levels. Combined OAA and Social Security payments per person 65 and

[26] Appendix Figure A15 shows empirical and simulated moments using the entire US (instead of only Massachusetts), and Appendix Figure A16 shows the underlying shares of men with earnings in each range. Appendix Figures A17 and A18 show the analogous results for California. Both the US and California exhibit the same basic patterns as Massachusetts.

older grew by a factor of more than six, from about \$850 to more than \$5,300 in 2010 dollars. This was partly due to an expansion of OAA in the late 1940s but was mainly due to the much greater growth of Social Security, which grew from \$41 per person 65 and older in 1940 to \$677 in 1950 and \$4,644 in 1960 (all in 2010 dollars). Social Security differed from OAA in many ways, but over this period Social Security imposed an earnings test not unlike those of OAA, only gradually liberalized over the 1950s. Although providing a definitive point estimate is beyond the scope of this paper, our results can provide some indication of the probable importance of this expansion.

The simplest approach to answering this question is an extrapolation based on our main regression results. This extrapolation suggests that the large growth in combined OAA and Social Security between 1940 and 1960 would be expected to decrease labor force participation among men aged 65–74 by 12.4 percentage points, or about 90 percent of the observed reduction of 13.5 percentage points. An obvious limitation of this approach is that it fails to account for other major changes over this period, perhaps most importantly rising wages and falling unemployment. It also fails to account for non-linearities in the relationship between log benefits per person and labor force participation, and for differences in the effects of changes in recipiency rates and changes in benefits per recipient.[27]

To overcome these limitations, we use our estimated life cycle model to try to obtain a lower-bound estimate of the effect of the growth in government old-age support from 1940 to 1960. To do so, we attempt to make consistently conservative assumptions that reduce the magnitude of the implied effect. Most important, we significantly understate mid-century Social Security. We simulate the effect of a Social Security program fixed at its 1939 characteristics, ignoring the large expansions in Social Security eligibility and benefits that occurred from 1950 onward. We also make assumptions that tend to reduce the implied effect of Social Security along several other key dimensions, including wage levels and growth rates, inflation, eligibility for Social Security supplemental benefits, and counterfactual labor supply in the absence of Social Security. The simulation compares the predicted behavior of a single cohort of early recipients—men aged 50 in 1940—under this relatively modest version of Social Security to its predicted behavior in the absence of government old-age support. Details are in Appendix Section A.10.2.

The results indicate that even this conservative representation of Social Security would reduce labor force participation among members of this cohort from age 65 to 74 by 7.3 percentage points, 54 percent of the 13.5 percentage point reduction observed from 1940 to 1960. As

[27]Simulations of our estimated model suggest that labor-supply effects are concave in the replacement rate, due to income effects shifting retirements earlier and thereby reducing exposure to the earnings test. This is a potential explanation for the sometimes weak relationship between Social Security benefits and labor supply in the aggregate time series (e.g., Moffitt, 1987).

shown in Appendix Table A12 and Appendix Section A.8.1, the prediction that Social Security significantly reduced labor supply from 1940 to 1960 is highly robust to a wide range of alternative assumptions and holds even in models that under-predict the effects of OAA in 1940. Within the standard lifecycle model framework, it is hard to imagine a model in which the rapid growth in government old age support from 1940 to 1960 would not reduce labor supply significantly.

We cannot provide a direct estimate of the cost to recipients of Social Security's earnings test over this period because the assumptions we make to be conservative in terms of labor supply generate bias of ambiguous sign. We can, however, offer speculation informed by our analysis in Section 6.2.5. The key determinants of the cost of the earnings test are replacement rates and counterfactual retirement ages. Rising replacement rates tend to decrease the cost of the earnings test, since greater income effects lead people to retire earlier and therefore have less "exposure" to the earnings test. Falling "latent" retirement ages also tend to decrease the cost of the earnings test by reducing "exposure" to the earnings test. Replacement rates increased significantly between 1940 and 1960 (Clingman, Burkhalter and Chaplain, 2014). It is less straightforward to know how latent retirement ages changed over this period, but it is notable that Figure 1(b) shows little apparent change in the labor force participation-age profile at ages below 65 from 1940 to 1960. Together with the small cost to recipients of OAA's earnings test, these considerations suggest that mid-century Social Security's earnings test was also unlikely to be very costly.

8 Conclusion

Many of the most important government programs transfer resources to older people and explicitly or implicitly tax their labor. In this paper, we investigate the labor-supply and welfare effects of the Old Age Assistance program in 1940. OAA was a large source of government old-age support at the time—nearly one quarter of all individuals 65 and older received OAA in 1940—and it helped pave the way for many of the important social insurance programs of the present day. Even independent of its historical importance, OAA presents a valuable opportunity for learning about the effects of government old-age support programs. Like many modern programs, it had both a transfer component and a high implicit tax on labor. But unlike many modern programs, it varied significantly across states and across otherwise-similar groups of people within states. The recent availability of Census data on the full US population in 1940 makes studying OAA a particularly fruitful way to shed light on the effects of these programs.

Our results suggest that OAA caused large reductions in labor supply in 1940, explaining more than half of the observed decline in labor force participation among men aged 65–74 over the 1930s. Yet both reduced-form regressions and an estimated life cycle model indicate that, while a significant share of this reduction in labor supply was due to substitution effects from the high implicit tax rates of OAA's earnings test, the reduction in the value of benefits to recipients associated with the earnings test was quite small. Predictions based on our regression estimates and our estimated life cycle model both suggest that Social Security accounted for at least half of the large mid-century decline in late-life labor supply. Taken as a whole, our results suggest that government old-age support programs can have large effects on labor supply, through both their transfer and taxation components, but that in the case of OAA circa 1940, the costs to recipients of the implicit taxation of work were quite small.

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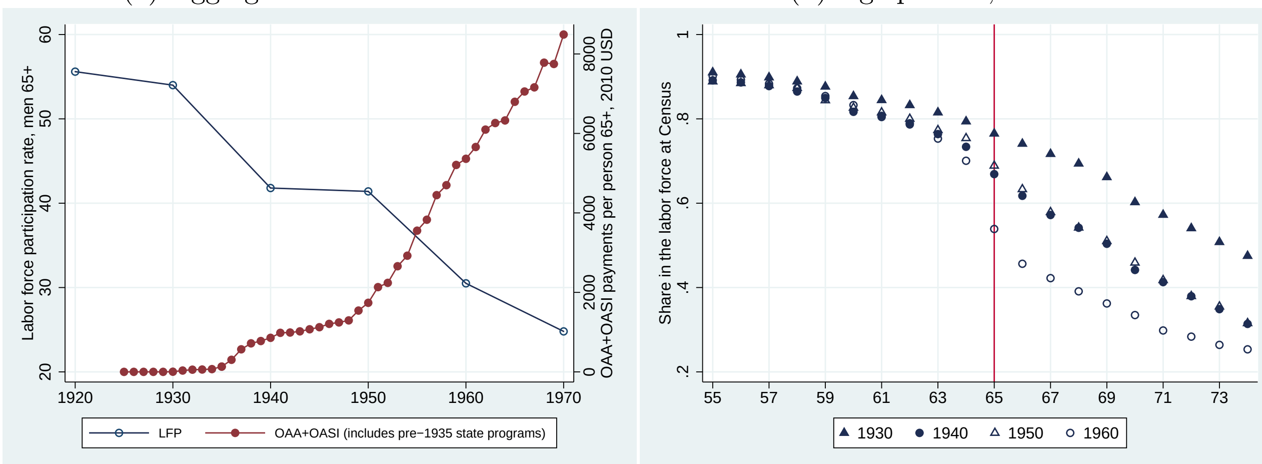
Tables and Figures

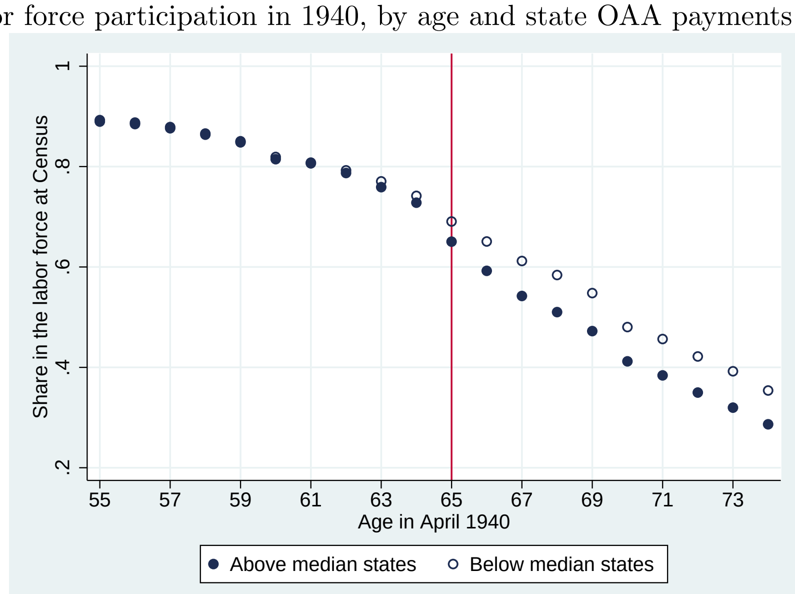
Figure 1: Government old-age support and retirement over the mid-20th century

(a) Aggregate trends 1920-70 (b) Age profiles, 1930-60



Notes: Panel (a) shows labor force participation rate of men 65 and older, from Series D35 of U.S. Bureau of the Census (1975), and payments under Old Age Assistance (OAA) and Old Age and Survivors Insurance (OASI) per person 65 and older, in 2010 US dollars. OAA payments data come from Parker (1936) for 1925 to 1935 and Series Bf621 of Carter et al. (2006) for 1936 onwards. OASI payments data come from Series BF396 of Carter et al. (2006). Panel (b) shows share of men in the labor force at each age, calculated from the 1930-60 Censuses. Rates prior to 1940 are adjusted for comparability: see Appendix Section A.1 for details.

Figure 2: Labor force participation in 1940, by age and state OAA payments per person 65+



Notes: Figure shows share of men in the labor force at the time of the 1940 Census, in states with above- and below-median OAA payments per person 65+ in 1939, for states with an eligibility age of 65 in 1939.

Figure 3: OAA payments per person 65 and above, December 1939

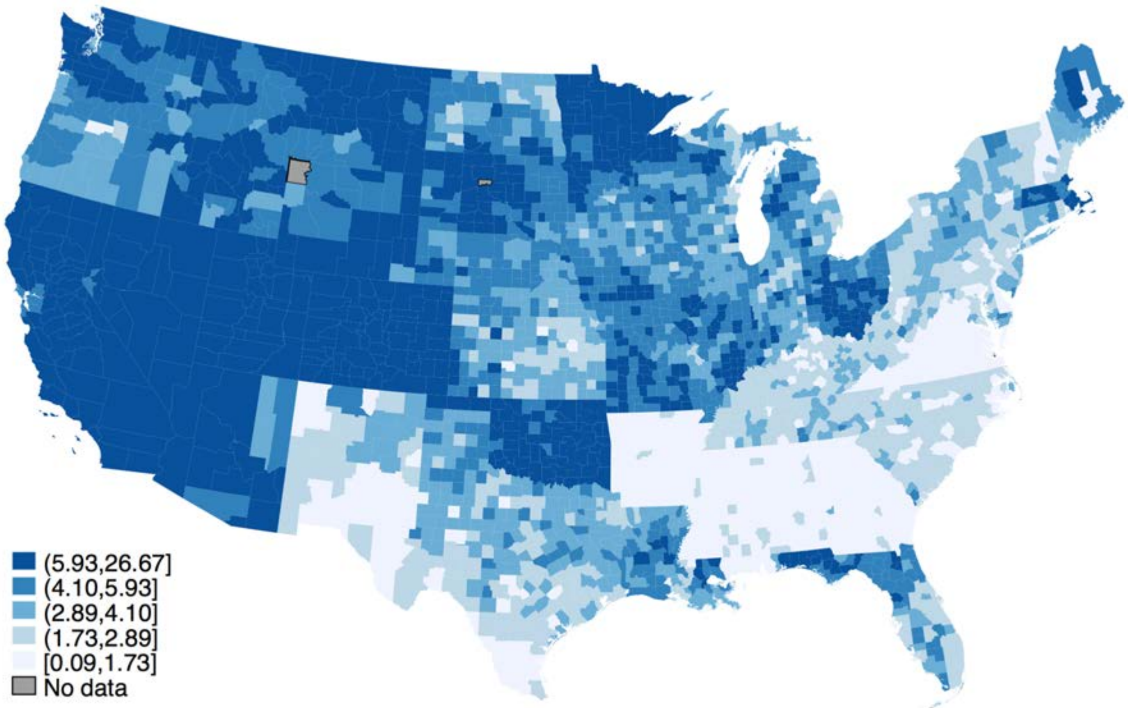
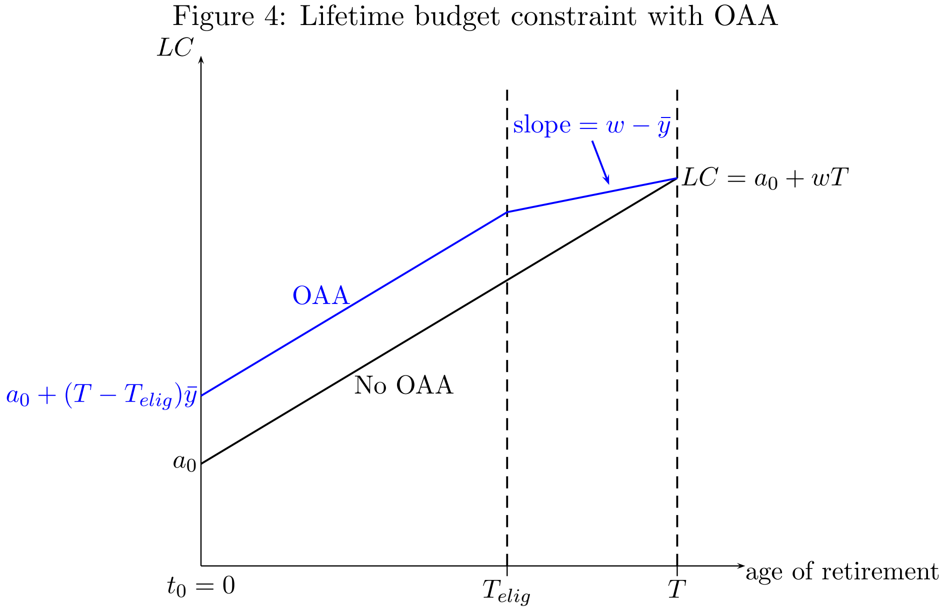


Figure shows total OAA payments for the month of December 1939 scaled by the population 65 and above in the 1940 Census. OAA payments data come from U.S. Social Security Board (1940*c*). In 2010 dollars, bottom quintile is approximately \$1.40 to \$27 and top quintile is \$93 to \$419.

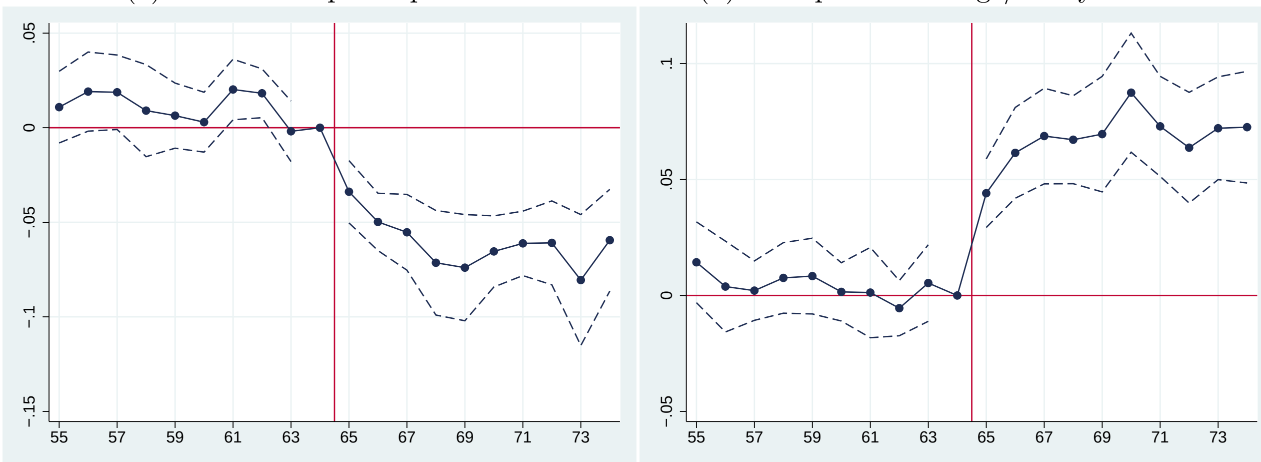
Figure 4: Lifetime budget constraint with OAA



Lifetime budget constraint relating the present value of lifetime consumption (LC) to age at retirement, with and without OAA. The OAA program depicted is an income-floor program with eligibility age T_{elig}, which implicitly taxes labor earnings at a 100 percent rate from the first dollar (by phasing out benefits dollar-for-dollar with labor income). For simplicity, the figure depicts the case in which the rate of return is zero, $r = 0$.

Figure 5: Differential changes in labor force participation and receipt of non-wage income

(a) Labor force participation (b) Receipt of non-wage/salary income

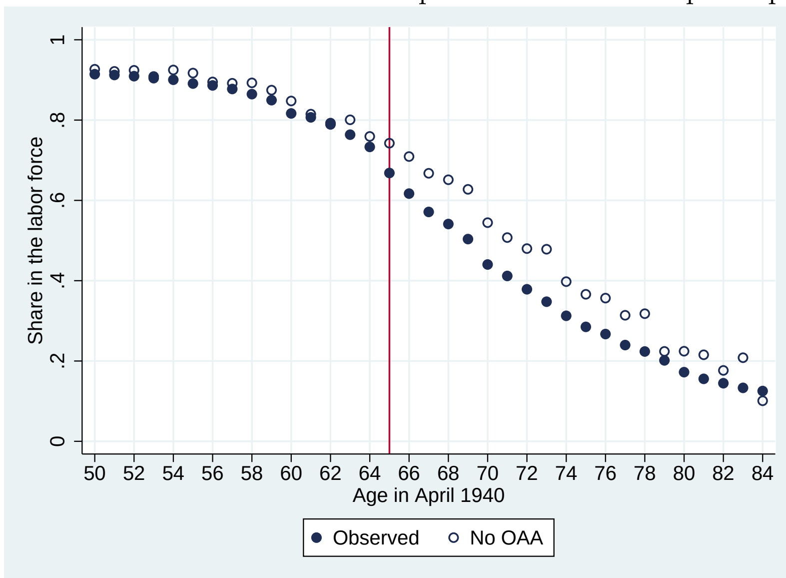


Notes: Figure shows point estimates and 95% confidence intervals on age-payment interactions from estimation of equation (1), using log simulated payment by age interactions as instruments for log per-65+ payment by age interactions and controlling for state border by age fixed effects. Sample restricted to counties on state boundaries, excluding counties on borders of states with age eligibility requirement other than 65 in 1939. Unit of observation is an individual-state border pair. Standard errors clustered at the state level. For Panel (a), $N = 2403915$ and Kleibergen-Paap rk Wald F-stat is 3.06. For Panel (b), $N = 2238476$ and Kleibergen-Paap rk Wald F-stat is 3.55.

Figure 6: Breaks in non-wage/salary income versus breaks in labor force participation, by state

Notes: Figure shows point estimates from estimation of equation (2) for receipt of non-wage income in 1939 against estimates for labor force participation in 1940, separately by state. Sample: men aged 56-64 or 66-73 at 1940 Census, in states with an eligibility age of 65 in 1939; breaks in receipt of non-wage income estimated on sample of men with non-missing 1939 income information ($N = 5277150$) and breaks in labor force participation estimated on sample of men with non-missing 1940 labor force participation information ($N = 5649733$).

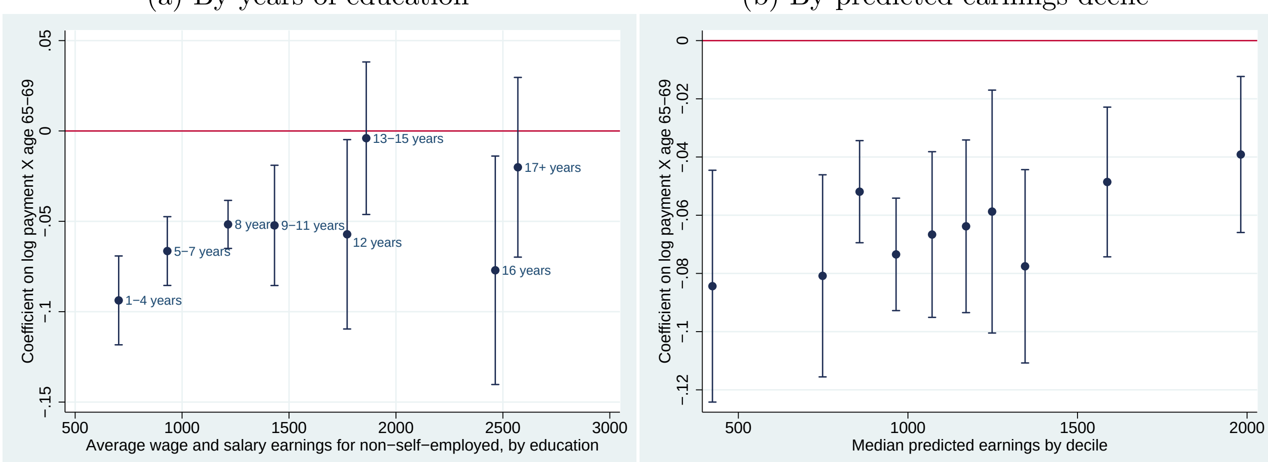
Figure 7: Actual and counterfactual profile of labor force participation



Notes: Figure shows observed rates of labor force participation by age and estimated counterfactual rates of labor force participation in the absence of OAA, based on IV estimates of equation (3).

Figure 8: OAA reduced labor supply more for men with lower potential earnings

(a) By years of education (b) By predicted earnings decile

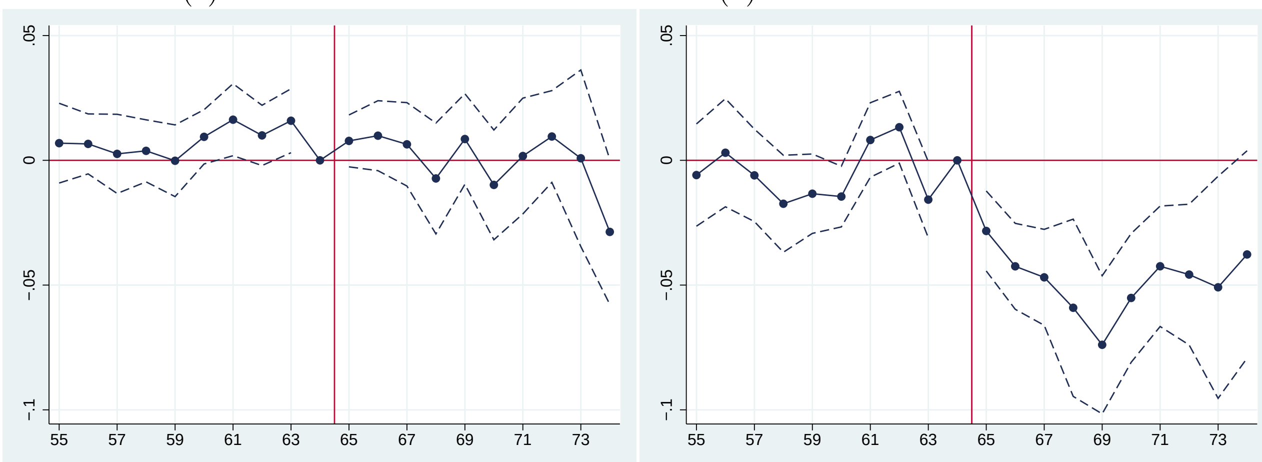


Notes: Figure shows estimates of age 65-69 interaction in equation (1) separately by grouped years of education (Panel (a)) and by decile of predicted earnings (Panel (b)), using log simulated payment by age interactions as instruments for log per-65+ payment by age interactions and controlling for state border by age fixed effects. Sample restricted to counties on state boundaries, excluding counties on borders of states with age eligibility requirement other than 65 in 1939. Unit of observation is an individual-state border pair. In Panel (a), education groups are arranged horizontally by average 1939 wage and salary earnings for 45-54 year old men who were not self-employed at the time of the Census.

Figure 9: 1940 OAA payments are not associated with differential age trends in 1930

(a) 1930 coefficients

(b) 1940 coefficients for same states

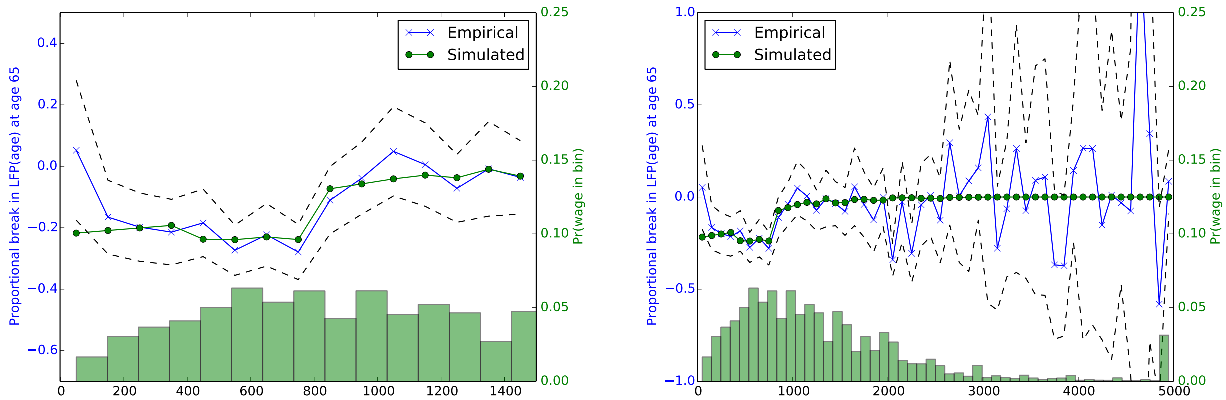


Notes: Panel (a) shows point estimates and 95% confidence intervals on age-payment interactions from estimation of equation (1) using "gainful employment" in 1930 as the outcome and December 1939 payments as the payment variable, with log simulated payment by age interactions used as instruments for log observed payment by age interactions and controlling for state border by age fixed effects. Sample includes only those states that had no old-age assistance program in 1930. For comparison, panel (b) shows analogous estimates for 1940 labor force participation for the same sample of states. In both, sample restricted to counties on state boundaries, excluding counties on borders of states with age eligibility requirement other than 65 in 1939. Standard errors clustered at the state level. For 1930 coefficients $N = 1602079$ and Kleibergen-Paap rk Wald F statistic is 1.36, for 1940 coefficients $N = 1930322$ and Kleibergen-Paap rk Wald F statistic is 1.44.

Figure 10: Empirical vs. simulated moments

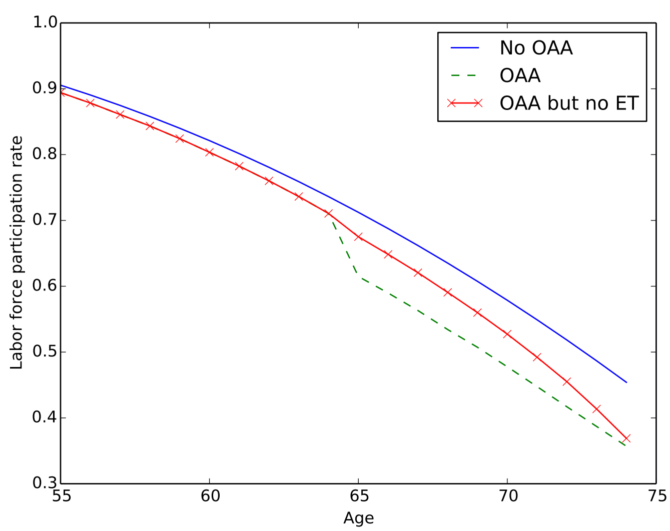
(a) Moments in the estimation

(b) All earnings levels



Notes: Empirical vs. simulated moments and annual earnings distribution for moments in the estimation (Panel (a)) and for all earnings levels, including those not in the estimation (Panel (b)). Dashed lines are 95 percent confidence intervals of the empirical moments. The moments are the proportional breaks in labor force participation-age profiles at age 65. Empirical moments correspond to the proportional breaks at age 65 in the share of men with the specified amount of wage/salary income in Massachusetts in 1939, relative to the predicted share at age 65 based on data from younger ages. The earnings distribution is the distribution of wage/salary income among men in Massachusetts aged 60–64 in 1939 who had any wage/salary income. For reference, the "income floor" in Massachusetts is \$360 per year (about \$5,600 in 2010 dollars). Earnings above \$5,000 are set to \$5,000.

Figure 11: Simulated effects of OAA on labor force participation



Notes: Simulated life cycle labor force participation profiles of the cohort of people aged 55 in 1940 in the US under different OAA programs. The policy underlying the "OAA but no ET" profile is a counterfactual OAA program that did not impose an earnings test.

Table 1: Basic features of state OAA programs

	Mean	Median	SD	Min	Max	N
OAA recipiency rate, December 1939	0.23	0.23	0.09	0.08	0.49	49
OAA payment per recipient, December 1939	17.93	18.90	6.49	6.01	32.97	49
OAA payment per person 65+, December 1939	4.16	3.59	2.59	1.01	13.17	49
Legal maximum payment	29.37	30	5.34	15	45	41
99th percentile payment	29.43	30	6.22	12	45	49
99th percentile payment, states with legal maximum	28.78	30	4.85	15	45	41

Includes the 48 states and the District of Columbia. '99th percentile payment' is for new recipients in fiscal year 1938-39. Eight states had no legal maximum payment. Recipiency rate and payments per person 65+ are normalized by state population from 1940 Census. Sources: data on OAA dollar payments and number of recipients from U.S. Social Security Board (1940*b*), data on legal maximum payments from U.S. Social Security Board (1940*a*), data on 99th percentile payment from U.S. Social Security Board (1939*b*).

Table 2: Summary statistics

	Full sample			Border county sample		
	Mean	SD	N	Mean	SD	N
Years of education	7.142	3.740	6722869	7.013	3.718	2055456
Completed primary school	0.547	0.498	6722869	0.528	0.499	2055456
Non-white	0.079	0.269	6722869	0.088	0.283	2055456
US citizen	0.946	0.227	6722869	0.953	0.211	2055456
Currently married	0.755	0.430	6722869	0.755	0.430	2055456
In the labor force in 1940	0.713	0.452	6722869	0.725	0.447	2055456
Employed in 1940	0.651	0.477	6722869	0.666	0.472	2055456
Employed, non-emergency work in 1940	0.616	0.486	6722869	0.630	0.483	2055456
Worked in 1939	0.720	0.449	6283146	0.730	0.444	1915313
Any wage/salary income in 1939	0.480	0.500	6283146	0.480	0.500	1915313
Wage/salary income in 1939	557	911	6283146	551	905	1915313
$\geq$\$50 in non-wage/salary income	0.516	0.500	6283146	0.518	0.500	1915313

Full sample: men aged 55-74 in states with 1939 eligibility age of 65 with non-missing demographic information (education, race, birthplace, citizenship, and marital status). For demographic variables and 1940 labor force and employment variables (reflecting labor force status in last week of March 1940), sample restricted to men with non-missing information on labor force status and non-missing demographic information. For 1939 employment and income variables, sample restricted to men with non-missing information for all 1939 employment and income variables and non-missing demographic information. State border county sample further limits to counties that border a state included in the sample.

Table 3: Labor force participation by state payments per person 65+ and age

Panel A. OLS results				
	(1)	(2)	(3)	(4)
Log per-65+ payment	0.018***	0.016***	0.000	0.001
× age 55-59	(0.005)	(0.004)	(0.004)	(0.003)
Log per-65+ payment	-0.061***	-0.057***	-0.049***	-0.049***
× age 65-69	(0.004)	(0.006)	(0.004)	(0.004)
Log per-65+ payment	-0.068***	-0.068***	-0.060***	-0.061***
× age 70-74	(0.008)	(0.009)	(0.006)	(0.006)
Observations	6722869	2403915	2403915	2403915
Sample	full	border	border	border
Border segment × age fixed effects	no	no	yes	yes
Education × age fixed effects	no	no	no	yes
Race × age fixed effects	no	no	no	yes
Panel B. IV results				
	(1)	(2)	(3)	(4)
Log per-65+ payment	0.048*	0.028***	0.006	0.006
× age 55-59	(0.019)	(0.007)	(0.005)	(0.005)
Log per-65+ payment	-0.070***	-0.060***	-0.063***	-0.063***
× age 65-69	(0.008)	(0.010)	(0.008)	(0.008)
Log per-65+ payment	-0.089***	-0.078***	-0.074***	-0.075***
× age 70-74	(0.016)	(0.013)	(0.010)	(0.010)
Observations	6722869	2403915	2403915	2403915
Kleibergen-Paap rk Wald F-stat	2.00	8.39	20.46	20.53
Sample	full	border	border	border
Border segment × age fixed effects	no	no	yes	yes
Education × age fixed effects	no	no	no	yes
Race × age fixed effects	no	no	no	yes

Dependent variable is indicator for being in the labor force at 1940 Census. In Panel B, log simulated payment by age interactions used as instruments for log per-65+ payment by age interactions. Sample for column (1): men aged 55-74 in states with 1939 eligibility age of 65. Columns (2)-(4) include only counties on state boundaries and exclude counties on borders of excluded states. Unit of observation in columns (2)-(4) is an individual-state border pair. All specifications include county fixed effects and 5-year age group fixed effects. All age-interactions are with 5-year age groups. Standard errors (in parentheses) are clustered at the state level. *: $p < 0.05$, **: $p < 0.01$, ***: $p < 0.001$

Table 4: Alternative labor force participation outcomes by state payments per person 65+ and age

	(1) In labor force	(2) Employed	(3) Non-emergency
Log per-65+ payment	0.006	0.008	0.001
$\times$ age 55-59	(0.005)	(0.005)	(0.004)
Log per-65+ payment	-0.063***	-0.050***	-0.032***
$\times$ age 65-69	(0.008)	(0.008)	(0.008)
Log per-65+ payment	-0.075***	-0.060***	-0.042***
$\times$ age 70-74	(0.010)	(0.009)	(0.008)
Observations	2403915	2403915	2403915
Kleibergen-Paap rk Wald F-stat	20.53	20.53	20.53
Sample	border	border	border
Border segment $\times$ age fixed effects	yes	yes	yes
Education $\times$ age fixed effects	yes	yes	yes
Race $\times$ age fixed effects	yes	yes	yes

Dependent variables: in labor force at 1940 Census (1), employed at 1940 Census (2), employed in private or non-emergency government work (3). Log simulated payment by age interactions used as instruments for log per recipient payment by age interactions. Sample restricted to counties on state boundaries, excluding counties on borders of states with age eligibility requirement other than 65 in 1939. Unit of observation is an individual-state border pair. All specifications include county fixed effects and 5-year age group fixed effects. All age-interactions are with 5-year age groups. Standard errors (in parentheses) are clustered at the state level. *: $p < 0.05$, **: $p < 0.01$, ***: $p < 0.001$

A Appendix

A.1 Comparability of labor force participation rates over time

The 1940 Census was the first Decennial Census to use the concept of "labor force participation," which was based on a person's employment or unemployment status in the last week of March 1940. Earlier Censuses, including the 1930 Census, provide information on the closely related but distinct concept of "gainful employment," measuring whether an individual reported having had an occupation in the previous year. Comparability of these concepts is not an issue for our main estimates, which are all based solely on the 1940 Census, but we do make adjustments when we make comparisons over time.

Durand (1948) reports adjustment factors to make data from 1930 and earlier comparable to measures of labor force participation from 1940 onwards. A separate adjustment factor is given for men in each 5-year age bin from ages 20 to 74 and for men aged 75 and older. We use these age-specific adjustment factors in all comparisons of 1940 data with data from earlier years. In Figure 1(b), which plots labor force participation in 1930 by single years of age, we linearly interpolate to obtain adjustment factors for each age.

On the general issue of constructing consistent measures of labor force activity over time, see Costa (1998) for further details and Moen (1988) for an extensive discussion.

A.2 State OAA programs and construction of the simulated instrument

A.2.1 The OAA budget constraint

Payments, and to some extent eligibility, under OAA programs were in general based on an income floor or a consumption floor. Lansdale et al. (1939) provides a contemporary overview of how these worked in practice. The most common method for determining a basis for a payment was a "budgetary deficiency" principle. This involved the determination of a basic standard of living and, based on this standard, an estimate of the "needs" of a particular applicant for a given length of time (which could vary across applicants) and an estimate of the applicant's "resources" (which would always include any regular income). The deficit determined the basis for an OAA payment.[28] This method had been common in the administration of relief to the poor prior to the growth of OAA. By the late 1930s, some states had also begun using a more explicit income floor, which amounted to specifying a standard amount for "needs" (such as 30 dollars per month). In both types of systems, increases in income would lead to a reduction in benefits.

It is relevant to both our regressions and the structural estimation that to the extent that "needs" varied across people according to unobserved characteristics, it need not have been the case that OAA payments in a state would have been the same to all individuals with

[28]Note that under this system, there is no necessary reason why retiring later would lead to higher OAA payments.

the same level of income. In practice, in many states payments varied substantially even across people with no other source of earnings. This issue is illustrated in Figure A1, which is based on data from U.S. Social Security Board (1939*b*). In Ohio, among new recipients in 1939, only about 10 percent of payments were at the legal maximum of $30 per month, even among recipients with no other source of income.[29] However, conforming with the contemporary literature (e.g., Lansdale et al., 1939), the data suggest that a few states did have programs that more closely resembled an income floor set at a common level across people. As examples, California and Massachusetts had legal minimum amounts for the sum of income and benefits. For recipients with no other source of income, these states saw payments cluster right around this minimum. For new recipients in Massachusetts in 1939, for example, close to 70 percent of recipients with no other source of income received payments of $30 per month. California specified both a maximum and minimum income plus benefit of $35 per month, and for recipients with no other income all payments clustered at this amount.

Lacking data on possible determinants of payments in states without a uniform income floor, our baseline estimation of the structural model utilizes data from Massachusetts only, a state whose OAA program is close to a uniform income floor for all individuals. The reason to prefer Massachusetts to California is that California had a $15 earnings disregard that slightly complicates the budget constraint, though we do not find any apparent effects of the earnings disregard. As discussed in Section A.8.1, the key results are robust to using moments estimated based on the entire US or California.

A.2.2 The relationship of maximum payments to the size of state OAA programs

The simulated instrument used in our analysis relies primarily on measures of state maximum payments. Its construction is described in more detail in Section A.2.3 of this appendix. This section provides more detail on variation in maximum payments across states, and how it was related to variation in the size of state OAA programs relative to the elderly population, to give background for its use in the simulated instrument.

Table A1 lists basic information for the 48 states and the District of Columbia, with states ordered by OAA per person 65 and older in December 1939. This table includes the four states excluded from our analysis sample—the three with an eligibility age of 70 in 1939, and Colorado, where long-term residents were eligible starting at age 60. The legal maximum, if present, indicates that the state OAA law limited monthly payments to that amount per individual. This information comes from the summary of state OAA laws in U.S. Social Security Board (1940*a*). A majority of states set legal maximum payments at $30 per month, the level beyond which the federal government would not match additional spending. Figure A2 shows a map of legal maximum payments by state, which illustrates the cross-state differences that can be used once comparisons are restricted to state boundaries. Some neighboring states that differ starkly in OAA payments, such as Texas and Oklahoma, do

[29] We do not directly observe payments to those with no other source of income, but rather the unconditional distribution of payments and the share of recipients with no other source of income. We assume that the recipients with other sources of income received the lowest payments.

not differ in maximum payments.

In our baseline simulated instrument, for the eight states that did not have legal maximum payments, we measure a *de facto* maximum as the 99th percentile payment in that state, and use this measure in place of a legal maximum. As noted in the main text, the idea is that with payments being set as the difference between "needs" and "resources," payments near the top of the distribution in a state would reflect payments to individuals with virtually no resources, and hence reflect differences across states in assessments of needs (the reason not to use the observed maximum payment is that it tended to reflect highly unusual situations). Table A1 reports our measure of the 99th percentile payment, as well as an observed maximum payment for each state. These variables are both based on grants to new recipients in fiscal year 1938-39, reported in U.S. Social Security Board (1939*b*). To estimate a 99th percentile payment, we use summary tables on the distribution of grant amounts by state. We have information on the share of payments in either 1- or 5-dollar bins, so we cannot always calculate the 99th percentile precisely. Instead, we identify the bin containing the 99th percentile and use the smaller value of the upper endpoint of the bin or the observed maximum payment (in two cases we also round 30.99 down to 30). For all but three of the states with statutory maximum payments, these 99th percentile payments were the same as the statutory maxima.[30] Figure A3 shows a map of 99th percentile payments by state.

Maximum payments are useful in that they tend to reflect, at least broadly, differences in the level of the income or consumption floor across states. Appendix Figure A7 plots average payments per recipient against our measure of state maximum payments. Especially at the modal maximum monthly payment of $30 there is considerable variation across states. Based on the description of payment determination in Lansdale et al. (1939), this variation likely reflects, in large part, differences between maximum payments and the typical administrative determination of "need," as well as variation in the level of "need" determined across different people. Despite the presence of some variation that maximum payments do not capture, differences in maximum payments across states are strongly predictive of average payments overall.

A.2.3 Construction of the simulated instrument

To simulate payments per person aged 65 and older in Section 5, we apply a measure of each state's maximum payment and any income disregards (which existed in five of the 45 states with an eligibility age of 65) to a national population of men and calculate a predicted payment per person in the sample, which is then used as an instrument for the observed total OAA payments per person in the state. The national population we use for each state omits the state itself, although in practice this has very small effects on the estimates. As described in the text and in Section A.2.2 of this appendix, for our main simulated IV exercise, we use statutory maximum payments in 1939 as a measure of maximum payments for all states that had statutory maxima, and the 99th percentile payment to new recipients in fiscal year 1938-39 in the eight states that had no statutory maximum in 1939. In Appendix Table A8

[30] Observed maximum payments were the same as the legal maxima in all but four states. In New Jersey and Connecticut the legal maxima were greater, and in Alabama and Utah the reported value of the observed maximum is greater. Alabama, unlike other states, reported the budget deficit approved rather than the actual payment approved. It is unclear why there is a discrepancy for Utah.

we show alternative results that use the highest legal maximum across states (45 dollars per month, in Colorado) for these eight states.

Given a measure of an individual i's earnings in 1939, for each state s we calculate a predicted OAA payment per person under that state's law as the mean over all individuals of

$$\text{payment}_{is} = \max\{0, \min\{(\text{max payment})_s, (\text{max payment})_s + (\text{income disregard})_s - (\text{income})_i\}\}$$

A typical application of a simulated IV strategy would use characteristics of a population in a base period to simulate the effects of policy changes over time. We are leveraging cross-sectional policy variation, not within-state policy variation over time, and hence address the potential for endogenous earnings responses by using an ineligible population to simulate payments. We use the population of men aged 60–64 in 1940 (those just under the eligibility age for OAA), in the 45 states that had an eligibility age of 65 (except for the state itself), to construct the instrument. Because self-employment earnings are not reported in the 1940 Census, for any person who reported being self-employed at the time of the Census and who worked a positive number of weeks in 1939, we impute earnings by randomly drawing 1939 earnings amounts from the population of non-self-employed men with the same number of years of education and the same number of weeks worked in 1939.

Figure A5 shows the distribution of monthly earnings (including these imputed values) for the population used in the construction of the simulated instrument, along with the minimum, median, and maximum values of the "income limit" (maximum payment plus income disregards, if any existed). A significant share, just over 20 percent, of men at these ages reported zero weeks worked in 1939 and zero earnings. Although many of these men would likely have been ineligible for OAA under its other eligibility criteria, this large share illustrates the potential for many OAA recipients to have been inframarginal non-participants in the labor force.

Appendix Figure A6 plots the actual level of OAA payments per person 65 and older in each state against the simulated value from this procedure. Particularly in this comparison across all states, there is considerable variation in the actual size of state OAA programs for states with the same simulated value—especially for those with the modal maximum payment of 30 dollars per month—but the simulated measure captures the positive relationship across states. In addition to the variation in assessments of "needs" within states or across states with the same maximum payment, the reasons for variation in observed payments per person for states with the same simulated payments per person include features of eligibility determination that are not included in our simulated instrument. For example, the Census lacks information on non-housing assets, which is necessary for determining eligibility in states with asset tests. Some data sources from this time period offer some potential for measuring relevant characteristics—the 1935–36 Survey of Consumer Purchases includes more detail on non-housing assets, for example—but in practice, maximum payments have provided most of the predictive power when we explored using these alternatives as well. For some other eligibility criteria, such as not having relatives able to provide support, it is unlikely that any realistic data source from the time would be sufficient. As noted in the text, it was also the case that state and local relief officials retained a significant amount of discretion in determining eligibility, so that features of OAA other than statutory eligibility criteria could have important effects on recipiency rates. Fetter (2017), for example, documents that holding other features of OAA laws fixed, the allocation of responsibility for funding OAA payments between local and state governments had significant impacts on recipiency rates, particularly in those states where local governments played a greater role in making decisions.

Table A4 reports first-stage regressions for the interactions used in our main results. Consistent with the positive relationship evident in Appendix Figure A6, for each of the interactions of log payments per person 65 and older with age, the corresponding interaction of the log simulated payment with age is highly statistically significant.

A.3 The relationship between OAA and other factors associated with retirement ages

A.3.1 Correlation of state OAA programs with state demographics and income

At the state level, OAA payments per person 65 and older were correlated with demographics and state incomes. Table A2 reports regressions at the state level, for the 45 states, including the District of Columbia, that had eligibility ages of 65 in 1939. States with larger elderly populations and states with larger foreign-born populations had larger OAA programs; states with larger non-white populations (almost entirely the South) had smaller OAA programs. (These variables are calculated using data from Haines (2010).) OAA programs tended to be larger in states that had higher levels of income. The two measures shown in Table A2 are median years of education for 25–54-year-olds and log median wage and salary earnings for 25–54-year-old men who were not self-employed (these are based on our calculations using individual-level Census data). Both were positively correlated with the size of state OAA programs. Note that, as emphasized by, e.g., Wallis (1987), the structure of the federal matching grants for public assistance meant that states with higher incomes not only tended to spend more on assistance, but also, because of this, received more in matching grants from the federal government as well.

The relationship between the size of state OAA programs and these characteristics is one reason why it is important to make narrower comparisons across states in our main specifications. Table A3 reports regressions on county-level data that relate these demographic and income measures to the two policy measures we use in the main specifications and robustness checks—rest-of-state payments per person and simulated payments per person. We limit the sample to the same counties that are included in the main analysis (border counties in states with an eligibility age of 65 in 1939). Panels A and C report regressions without fixed effects for state border groups, and show the same relationships observed in the cross-state comparison. Panels B and D, which add fixed effects for state borders, show that these systematic differences largely disappear once comparisons are made only across state borders (only one coefficient is statistically significant at conventional levels, and for only one of the two policy measures). These results support our assumption that once comparisons are limited to state borders, differences in OAA policies are not correlated with factors that lead to differential underlying age trends in labor force participation. They also provide support for the stronger assumption of equal levels of labor force participation used in estimation of the counterfactual age-labor force participation profile reported in Figure 7.

A.3.2 The prevalence of other types of pensions

In this section we provide further details on the prevalence of other sources of pensions circa 1940. As noted in the text, Social Security made no regular monthly payments until 1940, and even then they were quite small relative to OAA: less than 2 percent of the elderly received them in that year, compared to 22 percent of the elderly receiving OAA. Prior to 1940, Social Security made one-time payments to some workers. The original Social Security Act excluded work done after age 65 from coverage, and required a certain number of years of coverage in order to receive regular benefits. Hence, those who turned 65 between 1936 and 1939 received lump-sum payments to reimburse them for taxes collected before they reached 65. These payments ended after the 1939 Amendments to the Social Security Act extended coverage to work at ages older than 65. These payments would have been relevant only for workers who turned 65 in that year, however, and only about 7 percent of 65 year olds received them in 1939. These payments tended to be smaller than OAA: the average OASI lump-sum payment at age 65 in 1939 was about 77 dollars, whereas the average annual OAA payment per recipient was 232 dollars.

The other major sources of old-age pensions at the time were private pensions, state and local government pensions, federal civil service pensions, and railroad pensions. In 1940 there were about 160,000 monthly beneficiaries of private pensions (Carter et al., 2006, Series Bf848).[31] McCamman (1943) estimates that there were about 158,000 beneficiaries of state and local government pensions, but notes that a significant share of these were for police and firemen, who typically had retirement ages before 65. There were about 141,000 beneficiaries of railroad retirement benefits (Carter et al., 2006, Series Bf753) and about 32,000 beneficiaries of federal civil service pensions with a retirement age of 65 (Reticker, 1941). By way of comparison, slightly more than 9 million people were aged 65 and older in the 1940 Census. Hence, the total number of beneficiaries of these plans was only about 5 percent of the population 65 and older in 1940, and some of these plans had retirement ages other than 65. Average payments under these plans were also much larger than OAA (between 750 and 950 dollars per year) and were likely primarily relevant for people higher in the income distribution than OAA recipients. In Section A.3.3 we discuss variation across states in the prevalence of these other types of pensions.

A.3.3 OAA and other state policies potentially relating to work at older ages

A related set of potential concerns with our empirical approach is that states varied in other policies that may have differentially affected work after the OAA eligibility age. One class of potential confounders is other types of pensions. As noted in Section A.3.2, these pensions were certainly present in 1940, although they were substantially less prevalent than OAA. The evidence we present in this section, however, suggests that those pensions that did exist do not provide an alternative explanation for our results.

There is no comprehensive data source on private pensions by state, but there is little reason

[31] Much of the growth in private pensions took place during and after World War II, associated with expanded labor demand and wage controls during wartime, as well as the expansion of the income tax combined with preferential tax treatment of pensions (see, e.g., Sass, 1997).

to think that they varied systematically across state borders.[32] Despite the existence of some state policies that could have been relevant to the prevalence of private pensions, they do not appear to have been important. Latimer and Tufel (1940), in a study of industrial pensions, note that states sometimes had tax laws relating to private pensions, but mainly to point out that these taxes were too small or too uncommon to have been important determinants of private pension decisions. Latimer (1932) discusses state law relating to pensions more broadly and also suggests that these laws were not very important. Most often, the purpose of these laws was to specifically enable firms to establish pension systems, but by 1916, federal courts had already made decisions that implied providing a pension would be a legitimate function for any profit-making enterprise.

For the other two significant sources of pensions at the time—railroad pensions and pensions for state and local government employees—some state-level data are available. We digitized data on retirement payments under the Railroad Retirement System (which had assumed responsibility from private railroad pensions plans in 1937) in fiscal year 1939-40, by state, from the Annual Report of the Railroad Retirement Board (Railroad Retirement Board, 1941). We also digitized data on retirement payments under state and local government pension systems in 1941, from the first comprehensive survey of state and local government retirement systems (U.S. Bureau of the Census, 1943). In Table A9, we re-estimate our main specification adding these other types of pensions as controls. We scale total dollar payments under each type of pension by the state population 65 and older in 1940, and interact each measure with age group fixed effects. Across all specifications, controlling for these other types of pensions has little effect on the estimated effect of OAA. This is unsurprising, in that these other types of pensions were significantly smaller than OAA, primarily affected a higher-income population, and were at best weakly correlated with OAA across states.

Finally, we are not aware of state labor laws or characteristics of firms' hiring practices that would be a plausible explanation for our results. Many states had labor laws governing weekly work hours, for example, but for the most part these laws applied only to women (U.S. Department of Labor, 1940). We are not aware of other laws that would have applied to workers specifically after age 65. Some firms did report explicit age limits on employment, and there was some geographic variation in these age limits, but nearly all of these limits were at ages well below 65—typically between 40 and 50 (Latimer, 1932).

As for other forms of public assistance that may vary across states, we have also investigated whether differences across states in "general assistance" explains the results. "General assistance" comprised assistance payments provided by states or localities other than through the three categorical assistance programs that received federal matching funds under the Social Security Act (Aid to the Blind, Aid to Dependent Children, and Old Age Assistance). *A priori* there is no strong reason to expect that general assistance would drive the results, because states and/or localities would receive no federal matching funds for general assistance, and hence would have a strong fiscal incentive to provide aid to the elderly through OAA, at least up to the federal matching cap. We use state-level data on the total dollar value of general assistance payments in December 1939 from U.S. Social Security Board (1940*b*). This information was unavailable for four states. In a fashion parallel to our measure of OAA, for each county we calculate the per-capita general assistance payments in that state

[32]A standard source on private pensions over this period is Latimer (1932), which relies on a sample of firms and industries.

(however, we scale by the full population instead of the population 65 and older). In Table A10 we report estimates of our main specification with and without controls for general assistance (interacted with age fixed effects), limiting to a common sample. The results are quite similar.

A.4 The OAA recipiency rate among men aged 65–74

To calculate the fraction of OAA recipients among men aged 65–74 who adjusted their labor supply in response to OAA, we need to know the total number of OAA recipients among men aged 65–74. Unfortunately, while we know the total number of OAA recipients in the full population, we do not have a direct measure of the number of OAA recipients among men aged 65–74. *A priori* it is likely that the recipiency rate for this group would be below the overall 22 percent recipiency rate that includes both men and women as well as older individuals. We do have information on the age and sex of *new* recipients at various points in time, however, so to provide a rough measure of the relevant recipiency rate in 1940 (a stock), we add up flows into the program and adjust for mortality and for aging out of the 65–74 age group. From U.S. Social Security Board (1939*a*), U.S. Social Security Board (1939*b*), and U.S. Social Security Board (1941) we have the number of new male recipients in fiscal years 1937/38 through 1939/40 by age at the end of the fiscal year, where age is reported in two groups: 65–69 and 70–74. The annual reports of the Social Security Board for 1935/36 and 1936/37 (U.S. Social Security Board, 1937*a*,*b*) provide the same information for fiscal year 1936-37, although not all states collected data for the entire fiscal year, meaning that we understate inflows in that year (for the period from July 1, 1936 through September 30, 1936 we observe the age distribution but not separately by sex; we assume half of new recipients aged 65–69 and 70–74 were men, which is approximately true in subsequent years). We do not have data on (and hence exclude from our calculation) any individuals who started receiving OAA prior to July 1, 1936.

In adding up flows, we adjust for aging out of the 65–74 range and for mortality. All men who started receiving OAA between the ages of 65 and 69 from July 1, 1936 onwards would still have been aged 65–74 as of mid-1940. We make a conservative assumption about the ages of 70–74 year olds, which is that no new recipients aged 70–74 by mid-1937 would still be 74 or younger by mid-1940, one-third of those 70–74 in mid-1938 would be 74 or younger in mid-1940, and two-thirds of those 70–74 in mid-1939. We then assume that recipients' mortality rate was 5.5% per year, just above the mortality rate of 65–74 year old men in the second half of the 1930s (Grove and Hetzel, 1968). This calculation yields 523,987 male recipients aged 65–74 in mid-1940, compared to a male 65–74 population in the 1940 Census of 3,167,055, for a recipiency rate of about 16.5 percent.

A.5 Comparison of labor supply results to prior literature on OAA

To the extent that our estimates can be directly compared to those in the earlier literature, they are similar. Friedberg (1999) investigates the effects of OAA in a differences-in-differences analysis from 1940 to 1950 using Census samples, and estimates effects of payments per recipient that are similar to (and perhaps slightly larger than) our findings from 1940. In particular, she estimates a probit coefficient on log OAA payments per recipient of -0.264. Dividing by 2.5—following the rule-of-thumb comparison of probit coefficients to those from a linear probability model (e.g., Wooldridge, 2010, p. 573)—yields approximately -0.1, while estimating our main specification using payments per recipient yields coefficients between -0.080 and -0.095. The earlier study by Parsons (1991) uses state-by-year aggregate data from 1930 to 1950 and estimates that OAA could account for about half of the 1930–1950 decline in male labor force participation, in line with our results.

A.6 Testing for migration responses to OAA

A possible concern with the results is that individuals with high disutility of labor chose to move to states with more generous OAA programs when they became eligible, or migrated out of more generous states at a lower rate. In either case, our empirical test would overestimate the reduction in labor supply upon aging into eligibility. The minimum residency requirements imposed by almost all states makes the first type of migration less likely, but to address the possibility of higher in-migration and lower out-migration we test for such effects using information on state of residence in 1935. Appendix Table A11 reports estimates of the baseline specifications with the dependent variable indicating whether an individual lived in a different state in 1935. Point estimates are quite small, and the upper and lower bounds of the 95% confidence intervals are an order of magnitude smaller than our labor supply results.[33] Hence, net migration of individuals with lower baseline levels of labor supply to more generous states after aging into eligibility is unlikely to explain our results.

A.7 Bounding the cost to recipients of the earnings test based on counterfactual retirement ages in the absence of OAA

This section provides details of the calculations underlying the second bound of the cost of the earnings test reported in Section 6.1. The maximum cost the earnings test imposes on an individual is the maximum amount of benefits he loses by working past the eligibility age, $\min\{w, \bar{y}\}\phi(O)$, where $\phi(O)$ is the number of periods he works after the eligibility age when facing the OAA budget constraint.[34] If leisure is non-inferior, people work no more when facing the OAA budget constraint than when facing the no-OAA budget constraint,

[33] If migration prior to age 65 responds to OAA benefits but people continue to work while still ineligible, the baseline specification may not pick up such effects on migration. To assess the extent to which effects of this sort would influence our results, we have estimated an alternative specification that restricts comparisons to state borders and simply tests for differences in the probability of migration within each age group. The results of this alternative specification are similarly small in magnitude.

[34] $\phi(O)$ is based on the latent retirement distribution: It is the number of periods after the OAA eligibility age the individual would work if he were not constrained to spend a non-negative amount of time retired. Consider an individual whose latent retirement age with OAA exceeds the maximum age, T, and whose potential OAA benefit is no larger than potential earnings, $\bar{y} \leq w$. For such an individual, the maximum cost of the earnings test equals the maximum lifetime OAA benefits the individual could receive, so the minimum value of OAA to this individual is zero. (The maximum cost of the earnings test can be no larger than the maximum amount of OAA benefits the individual could receive by not working.)

$\phi(O) \leq \phi(N)$, where $\phi(N)$ is the number of periods he works after the eligibility age when facing the no-OAA budget constraint. So for any individual whose preferences are in the broad class in which leisure is not an inferior good, the maximum cost of the earnings test is $\min\{w, \bar{y}\}\phi(N)$.

We use the joint distribution of earnings and OAA benefits observed in the 1940 cross section. We use the counterfactual age profile of labor force participation estimated in Section 5.1 together with the assumption that the observed cross-sectional relationship between labor force participation and age is a good proxy for the unobserved life-cycle relationship. We make the conservative assumption that take up of OAA benefits is uniform across the joint earnings-OAA benefit distribution. This tends to bias upward the cost of the earnings test, since people with higher replacement rates, for whom the earnings test was less costly, were in reality more likely to take up benefits. The lack of bunching of retirements at any particular age in the no-OAA counterfactual tightens the bounds from this approach, since it means that everyone's marginal rate of substitution, a key input into these calculations, is point-identified rather than bounded. This is another advantage of the lack of bunching in our setting.

Within the class of preferences in which utility is quasilinear in retirement—the usual case in many applications of the life cycle model—the average \$1 of OAA was worth at least \$0.72 of unconditional late-life income. Under the opposite extreme (and non-standard) assumption that utility is quasilinear in consumption (and so borderline inferior in leisure/retirement), the average \$1 of OAA was worth at least \$0.57 of unconditional late-life income. Intuitively, the earnings test was not that costly because many people would have retired either before or relatively soon after the OAA eligibility age even without OAA or even if OAA did not impose an earnings test.

A.8 Estimation of the life cycle model

A.8.1 Estimation results and robustness

Table A12 reports results based on the baseline specification and several alternative specifications of the model. The parameter estimates are fairly stable across specifications, and the key conclusions are extremely robust. Additional robustness tests not reported in the table, and available upon request, include estimating the model based on data from California (instead of Massachusetts or the full US), setting the discount rate and interest rate to zero, doubling the slope of the counterfactual labor force participation-age profile absent OAA, dropping low-earnings moments, fixing the slope of the eligibility-potential earnings relationship to zero, and setting the maximum age to 80 and 85. Across all specifications, the cost to recipients of the earnings test is always less than 7 percent of benefits received, and the reduction in labor force participation from 1940–1960 due to Social Security is always at least 5.6 percentage points, 41 percent of the observed 13.5 percentage point decline from 1940 to 1960.

The poor labor market conditions in 1940 would tend to reduce the cost to recipients of the earnings test. As we discuss in Section 6.2.5, the key determinants of the cost of the earnings test are replacement rates and counterfactual retirement ages. Bad labor market conditions likely cause both of these to change in ways that reduce the cost of the earnings test. Bad labor market conditions reduce wages, which increases replacement rates (holding fixed benefit levels). This tends to decrease the cost of the earnings test, since higher replacement rates lead people to retire earlier due to income effects, which reduces their exposure to the earnings test. Bad labor market conditions also tend to reduce labor force participation, which reduces our inferred counterfactual retirement ages. This also tends to decrease the cost of the earnings test, since a greater fraction of benefits are inframarginal. To bound the likely effect of these issues, we estimate the cost of the earnings test under assumptions that are likely to overstate what its cost would have been had labor markets been "typical" in 1940. We assume that the labor force participation-age profile in the absence of OAA in 1940 matches the observed labor force participation-age profile in 1930 (in fact we assume it matches the "gainful employment" profile—which is to say, we do not apply the correction described in Section A.1—meaning that it is a slight overestimate of what "labor force participation" would likely have been had that concept been used in the 1930 Census). Given the trend reductions in late-life work, this may overstate the counterfactual retirement ages that would have arisen in 1940 had the labor market been better. We assume that the potential earnings distribution in 1940 matches the observed earnings distribution in 1950 of 45–54 year olds with positive earnings, which reflects the rapid growth in wages during the 1940s. Even with these assumptions, which bias upward the cost of the earnings test, we estimate that the earnings test reduces the value of the average dollar of OAA benefits to recipients by \$0.07. The robustness of the result about the low cost of the earnings test is driven largely by the substantial fraction of benefits that were inframarginal in the sense that people would have received them even without adjusting their labor supply. This is true even with the somewhat greater labor force participation in 1930.

The key results are also robust to using moments estimated based on the entire United States or California as opposed to Massachusetts.[35] Figures A16 and A18 plot the share of men earning each amount up to \$1,000 for the full US and California, respectively. The general patterns are the same in the US, California, and Massachusetts (Figure A11): At age 65 the probability of earning low amounts drops sharply, and the drops fade away by earnings levels of about \$900 or more per year. One wrinkle in estimating the empirical moments is the discreteness of observation of age, which technically violates the assumption of a continuous forcing variable. This becomes more relevant in the full-US case because of the larger sample. The pilot bandwidth proposed by Imbens and Kalyanaraman (2012), for example, is less than two years on either side (from which we cannot estimate a line). Hence, in the full-US estimation, we use the smallest feasible number of years of age on

[35] The disadvantage of estimating the model based on the full US is that eligibility requirements varied across states in hard-to-measure ways, and computation costs prevent us from estimating separate eligibility parameters for each state. We estimate the model based on California because it is one of two other states (the other was Nevada) that had a fairly unambiguous single income floor like Massachusetts, in that they set state-wide minimum values for the sum of income and payments clearly in the state OAA law. We do not estimate the model based on Nevada because of sample size issues given its small population and the data-intensive estimation procedure. Moreover, unlike California and Massachusetts, Nevada did not have an asset limit, so using Nevada would require us to use a different procedure from the one in the main analysis of Massachusetts. Based on distributions of payments, and consistent with the contemporaneous description by Lansdale et al. (1939) of how OAA programs operated, it appears that a number of other states were also implementing uniform income floors, but using de facto (as opposed to de jure) income floors would require a procedure for backing out the de facto level of the floor empirically.

either side of the eligibility cutoff, two years, to estimate all empirical moments. Figures A15 and A17 plot the empirical and simulated moments for the estimations based on the full US and California, respectively. The fit of the model is good in these cases as well, and the key conclusions are unchanged. In both cases, the average $1 of OAA is valued highly by recipients ($0.94 based on the full US and $0.95 based on California), and Social Security is predicted to reduce labor force participation among 65–74-year-olds significantly (11.0 percentage points based on the full US and 6.7 percentage points based on California).[36]

In addition to the robustness tests discussed above, it is useful to discuss the possible role that other assumptions might play in the results. Because of older workers' worse health and greater difficulty re-entering the labor force after adverse employment shocks (see, e.g., Costa, 1998), the assumption that potential earnings are constant over the life cycle likely overstates potential earnings at older ages. Overstating late-life potential earnings tends to bias us against our key findings, since it tends to increase the cost of the earnings test and decrease the labor-supply effects of Social Security. The assumption that OAA is the only source of non-labor income understates non-labor income among people eligible for OAA somewhat. Empirically, 72 percent of new OAA recipients in the 1939–1940 fiscal year had no source of income other than OAA (U.S. Social Security Board, 1941). The main effect of understating other sources of non-labor income is to reduce the estimated level of eligibility for OAA, which is a lower bound anyway. The assumption that individuals cannot borrow is consistent with the poor functioning of household credit markets at the time (see e.g. Rose, 2014) and is reinforced by our estimation results. An alternative version of the model with perfect capital markets is highly inconsistent with the pattern of bunching of retirements at the OAA eligibility age.

One reason the results are robust to a wide variety of possible changes in the model is the combination of two key aspects of our approach: We estimate the model based on breaks in labor force participation at the OAA eligibility age, and we require the model to match the distribution of retirement ages in the absence of OAA. The key determinants of the amount of bunching of retirement ages in response to OAA are the fraction of people who would retire soon after the OAA eligibility age in the absence of OAA and the curvature of the utility function. The main effect on the bunching of retirements of many possible changes in the model, e.g., a correlation between discount rates and the disutility of labor, would come through any effects on the counterfactual distribution of retirement ages in the absence of the program. But because we force the estimation to match this distribution directly, the analysis is not very sensitive to changes in assumptions about the underlying determinants of retirement ages in the absence of OAA.

Additional results suggest that the timing of information about OAA shaped the observed effects of the program. Our baseline assumption that people learned about OAA in 1936 (when many state OAA programs were introduced) means that people had relatively little time before 1940 to incorporate OAA into their plans. Simulations of the model indicate that OAA would have had significantly greater effects on labor supply in 1940 had people

[36] The only substantive difference across these three estimations is the estimate of the coefficient of relative risk aversion. In the estimation based on the US, the estimate of the coefficient of relative risk aversion (0.6) is significantly smaller than it is in the estimations based on Massachusetts (1.3), California (1.5), and indeed any other estimation. This might reflect the incorrect assumption in this particular estimation of a common eligibility-potential earnings relationship in all states.

had more time to build OAA into their plans.

A.8.2 Identification

Figure A13 plots the objective function. The figure reveals that the model is well-identified; moving away from the estimates along any dimension of the parameter vector increases the mismatch between the simulated and empirical moments, as measured by the classical minimum distance-type objective function. If instead of estimating the slope of the eligibility-potential earnings relationship using the observed relationship between earnings and house value (as we do in the baseline specification) we estimate the slope of the eligibility-potential earnings relationship together with the other key parameters in the second stage of the estimation, the model is not as well identified. In this case, the estimation has a hard time distinguishing the source of the fadeout in the bunching of retirements in potential earnings between curvature in the utility function (η) on the one hand and declining eligibility with potential earnings on the other (β_e). This is why we invoke other evidence (the observed relationship between earnings and house value) to estimate the slope of the eligibility-potential earnings relationship in our baseline specification. Fortunately, as shown in Table A12, the key results are not sensitive to this choice.

A.8.3 Estimating the latent retirement distribution

We estimate the curvature of utility from consumption, η, and the intercept of the eligibility-potential earnings relationship, α_e, by attempting to match the pattern of bunching of retirements at the OAA eligibility age, while also requiring that the distribution of the disutility of work, $F(\delta)$, be such that the model matches the counterfactual distribution of retirement ages in the absence of OAA. The key assumptions are that all heterogeneity in retirement behavior among people who face the same budget constraint is driven by heterogeneity in the disutility of labor and that all potential earnings groups have the same counterfactual no-OAA retirement distribution. We estimate the $F(\delta)$ distribution non-parametrically by using the model to invert the (counterfactual) distribution of retirements without OAA.

The Census data do not contain all of the information necessary to construct individuals' lifetime budget constraints. For example, the data contain only incomplete information about assets (just housing wealth) and non-labor income (just an indicator about whether it exceeds \$50 per year). This means that unobserved heterogeneity in assets or non-labor income could help explain the observed heterogeneity in labor supply among people who share the same observable components of lifetime budget constraints. Given OAA eligibility rules, however, assets and non-labor income are likely to be quite limited among the population of people potentially eligible for OAA. As noted earlier, this is consistent with evidence on the characteristics of new OAA recipients in the 1939–1940 fiscal year, which indicates that 72 percent of new recipients had no source of income other than OAA (U.S. Social Security Board, 1941). The main effect of understating non-OAA non-labor income is to reduce the estimated eligibility rate, which is a lower bound for other reasons as well.

In order to estimate the full distribution of the disutility of work, $F(\delta)$, we need to know the full latent retirement distribution, out to the maximum age at which the person with the lowest disutility of labor would work if he could. In the model, everyone lives to exactly age 75 and so cannot work beyond that age. So for any given budget constraint, there exists a range of δ values that lead the individual to work until age 75: from the threshold δ such that the individual is just indifferent between retiring at age 74 and 75 down to $\delta = 0$ (people to whom work provides no disutility and so would continue working as long as possible). People with low enough δ values would work longer if they could. They can be said to have a negative latent demand for retirement, where the latent demand for retirement is the number of years an individual would choose to enjoy leisure (not work) were it possible to consume negative amounts of leisure, i.e., to work longer than one's full lifetime. Working longer than one's lifetime has the benefit of increasing consumption through higher earnings and the cost of incurring the disutility of work in the "extra" periods. The latent retirement distribution is fundamentally unobservable, and the data become progressively less informative about this object at greater ages due to the small number of individuals at these ages and the bias induced by selective survival. We therefore use the estimated relationship between labor force participation and age from age 50 to 84 to fit a polynomial out to the age at which labor force participation becomes zero. This polynomial serves as our estimated distribution of latent retirement ages, from which we infer the distribution of the disutility of labor, $F(\delta)$. An important assumption implicit in this procedure is that the cross-sectional relationship between labor force participation and age is similar to what the age profile of retirements would have been for a single cohort (had government policies and other factors been held constant at their 1940 values).

A.8.4 Our application of the Method of Simulated Moments

The Method of Simulated Moments estimator is the parameter vector $\theta \equiv (\eta, \alpha_e, F(\delta))$ that minimizes the distance between the model-simulated moments and their empirical counterparts, where distance is measured by a classical minimum distance objective function. In the baseline specification we estimate η and α_e by attempting to match the pattern of bunching of retirements at the OAA eligibility age, while at the same time requiring that $F(\delta)$ be such that the model matches the counterfactual distribution of retirement ages in the absence of OAA.

Given a candidate parameter vector θ, we simulate the 15 moments—one for each of the 15 potential earnings groups whose probability of retiring at the OAA eligibility age we estimate—using the following procedure. First, we simulate the retirement ages of a large sample of individuals. This involves drawing an individual's potential earnings, disutility of work, and eligibility for OAA, and then calculating the individual's optimal retirement age. Second, we aggregate the simulated data into moments.

The moment for each potential earnings group is the proportional break in that group's (otherwise smooth) labor force participation-age profile at the OAA eligibility age. Formally, this is the probability of retiring immediately upon becoming eligible for OAA conditional on not yet having retired:

$$Pr(\text{Retire immediately upon becoming eligible for OAA} \mid \text{Not yet retired}).$$

This conditional probability can be written as the conditional expectation,

$$E(\mathbb{1}(\text{Retire immediately upon becoming eligible for OAA}) \mid \text{Not yet retired}).$$

The one wrinkle involved in implementing this in practice is that model time is discrete. In discrete time, estimating the "break" in the labor force participation-age profile at the eligibility age requires using information from other nearby parts of the profile, not just its level at the eligibility age itself. This is because, in discrete time, the fraction of people who retire "at the eligibility age," i.e., sometime during the discrete period (year in our case) in which they reach the eligibility age, is weakly greater than the "break" in labor force participation or "excess" retirements at that age, since it also includes retirements during the rest of that discrete period. We deal with this issue by following a procedure analogous to the one we use to estimate the empirical moments in the Census data, in the spirit of a regression discontinuity. We simulate predicted labor force participation at ages 63 and 64 (immediately before the eligibility age) and ages 66 and 67 (immediately after the eligibility age). We use these labor force participation rates to form two predictions of what labor force participation would have been at exactly the OAA eligibility age, age 65. One is based on participation before the OAA eligibility age (at ages 63 and 64). The other is based on participation after the OAA eligibility age (at ages 66 and 67). These predictions of participation at the OAA eligibility age are based on the assumption that, at least within two years of the OAA eligibility age, labor force participation declines linearly with age, except for any break at the OAA eligibility age. The estimated break, i.e., the probability of retiring immediately upon becoming eligible for OAA given that the individual is not already retired, is

$$\frac{LL(65) - RL(65)}{LL(65)},$$

where $LL(65)$ is predicted labor force participation at exactly age 65, the OAA eligibility age, based on labor force participation rates at younger ages ("left limit"), and $RL(65)$ is predicted labor force participation at exactly age 65 based on labor force participation rates at older ages ("right limit").

In practice, for computational feasibility we discretize both the potential earnings and disutility of work distributions. We assume that potential earnings take one of 15 values corresponding to the midpoint of the ranges that we use for estimating the empirical moments. For each candidate vector of parameter values, θ, and for each of the 15 possible potential earnings levels, w, we construct the simulated moment condition in the following way. First, we calculate the disutility of work distribution, $F(\delta; w, \eta)$. The $F(\delta; w, \eta)$ distribution is that which matches the counterfactual no-OAA retirement age distribution (predicted using variants on our main regressions), given potential earnings and the curvature of utility of consumption, w and η. Because time is discrete in the model, any given (discrete) retirement age is consistent with a range of δ values. We use the midpoint of these ranges. For each of these δ values, we calculate the optimal (discrete) retirement ages for people eligible and ineligible for OAA, $T_r^*(O; w, \bar{y}, \eta, \delta)$ and $T_r^*(N; w, \eta, \delta)$, respectively. We use these mappings from δ to optimal retirement ages with and without OAA together with the disutility of work distribution, $F(\delta; w, \eta)$, to calculate the full distributions of retirement ages with and without OAA for this potential earnings group, $F(T_r^*(O; w, \bar{y}, \eta, \delta))$ and $F(T_r^*(N; w, \eta, \delta))$, respectively. We use these distributions together with the fraction of people in this potential earnings group eligible for OAA, $Pr(\text{eligible}_i|w_i;\alpha_e,\beta_e)$, to calculate the overall retirement age distribution among this group, including both eligible and ineligible individuals, $F(T_r^*(w,\bar{y},\eta,\delta))$. Finally, we use this retirement-age distribution to calculate this potential earnings group's simulated moment based on the procedure detailed above.

The objective function is

$$g_N(\theta)'\ W\ g_N(\theta),$$

where $g_N(\theta)$ is the vector of moment conditions (whose elements are the differences between the empirical and simulated moments) and W is a positive definite weighting matrix. Pakes and Pollard (1989) and Duffie and Singleton (1993) show that the MSM estimator, $\hat{\theta}$, is consistent and asymptotically normally distributed under regularity conditions satisfied here. For our weighting matrix, we follow Pischke (1995) and use the inverse of the diagonal of the estimated variance-covariance matrix of the second-stage moment conditions.

A.9 Validation of the estimated life cycle model

A natural validation test of the model is to use it to simulate the cross-sectional relationship between labor force participation and age in 1940 and to compare the results with the observed empirical relationship. This requires an additional empirical input not used in the estimation: the joint distribution of potential earnings and OAA benefits. In each state, we use the observed distribution of earnings in 1940 among people aged 48–52 with positive earnings together with the OAA benefit level in 1940. Among other things, this tests the extent to which the model estimated based on Massachusetts data alone matches an important feature of the data on the full US. Figure A14 plots the results. The "No OAA" profile shows the counterfactual no-OAA profile predicted based on our regression results and presented in Figure 7. The "OAA" profile is the part that is relevant for testing the model. It is simulated based on the estimated model and can be compared to its empirical counterpart, also depicted in Figure 7. The model captures the key features of the data well and provides a fairly close fit quantitatively. The model predicts a roughly 6.3 percentage point reduction in average labor force participation over the ages 65–74, whereas our regression analysis indicated an 8.5 percentage point reduction. A relatively minor difference between the model and the data is in labor force participation at ages younger than the OAA eligibility age. The model predicts small but noticeable anticipatory effects in the years before OAA eligibility, whereas there is relatively little evidence of anticipatory effects based on our regression analysis. The close match between the model and the empirical evidence of the effects of OAA on labor supply, including the good fit of the simulated to the empirical moments, suggests that the model may be capturing some of the key factors that determine the labor-supply effects of OAA and so may be useful for understanding the effects of OAA and predicting the effects of the early Social Security program.

A.10 Simulations of the Effects of OAA and Social Security

This section presents details of the calculations underlying the simulations of the life cycle model discussed in Section 6 and Section 7. The goals of these calculations are to understand the observed effects of OAA—the value of OAA to recipients and the extent to which the labor-supply effects of OAA are due to income vs. substitution effects—and to predict the effects of Social Security. To this end, we simulate the model under various policies and calculate statistics of the simulated data. The key statistics concern the predicted effects of OAA and Social Security on retirement, the equivalent variation of OAA, and the income and substitution effects of OAA.

A.10.1 Simulating the effects of OAA

We simulate the effects of OAA as it existed in 1940 on the cohort aged 55 in 1940. The key ingredient of the simulation is the joint distribution of potential earnings and potential OAA benefit levels among this cohort. Each individual's potential OAA benefit is the 95th percentile OAA benefit in 1940 in his state. Due to a lack of data on assets other than housing, these calculations assume that all states implement "income-focused" OAA programs that do not limit benefits based on assets, other than any limitations that operate through our estimated model of eligibility. In the baseline specification, the probability that an individual is eligible for OAA is given by the eligibility-potential wage relationship estimated using data from Massachusetts only. Although the eligibility-potential wage relationship likely varies across states, the key results about the effects of the earnings test on labor supply and the value of OAA to recipients are not sensitive to the particular eligibility-potential wage relationship. For the distribution of potential earnings among individuals in a particular state, we use the observed distribution of earnings in 1940 among people aged 48–52 with positive earnings in that state. We assume that the unobservable distribution of self-employment earnings is the same as the observable distribution of wage and salary earnings. This is a strong assumption, but some broadly supportive evidence is that the education distribution of the self-employed in 1940 was quite similar to that of wage and salary workers. We further assume that potential earnings are constant over the life cycle. This assumption likely overstates late-life earnings (worse health or weaker labor demand for older workers likely lead potential earnings to decline with age, Costa, 1998, and in the cross section earnings fall with age), which tends to push against our key finding that the earnings test had little effect on the ex-post value of the program to recipients.

Given the subsequent changes in OAA over the 1940s, most of which increased OAA benefits, this simulation is not representative of the actual experience of any one cohort. Instead, it is meant to answer the question of what effects OAA would have been expected to have had it remained as it was in 1940.

The simulations reported in the text focus on the role of OAA's earnings test. Another important feature of OAA was its minimum age requirement, which meant that OAA payments were back-loaded to later ages. Given the evidence that borrowing constraints significantly affected the pattern of labor-supply responses to OAA, OAA's back-loaded payment structure may have reduced the value of OAA benefits to recipients relative to a cost-equivalent transfer made earlier in life. We find that the average OAA recipient values \$1 of present value worth of OAA benefits equally to \$0.75 in initial assets. Combined with our other results, this implies a non-negligible cost of OAA's back-loaded payment schedule, which is consistent with evidence of poorly-functioning household credit markets in this period (Rose, 2014).

A.10.2 Simulating the effects of Social Security

We simulate the effects of a counterfactually-modest Social Security program on the cohort of men aged 50 in 1940. The goal of this exercise is not to simulate the actual experience of this cohort. The goal is to simulate a simple counterfactual in which Social Security would be expected to have smaller effects than it actually did in order to estimate a lower bound of Social Security's likely effects.

This simulation requires three key inputs. One is Social Security program rules. We base our counterfactual Social Security program on the Social Security program as of the 1939 Amendments, which implied much lower eligibility and benefit levels than members of this cohort actually enjoyed due to subsequent expansions in Social Security. Total household benefits were the sum of primary benefits (for the worker) and supplementary benefits (for spouses and dependent children), up to a maximum of \$85 or 80 percent of the average monthly wage (AMW), whichever was smaller. The primary monthly benefit was the sum of (i) 40 percent of the first \$50 of the AMW plus 10 percent of the amount by which the AMW exceeds \$50 up to an AMW of \$250 and (ii) 1 percent of the amount in (i) multiplied by the number of years in which the individual earned at least \$200 in covered employment. The minimum primary benefit was \$10. Supplementary benefits for aged spouses and dependent children were one half of the primary benefit per person. We assume that only 50 percent of men qualify for supplemental benefits, whereas about 70 percent of 65–74-year-old men in 1940 were married. We assume that everyone had 15 years of covered employment regardless of when they retired. Taxes were 1 percent of covered earnings.

As of the 1939 Amendments, eligibility for Social Security was limited to workers in commerce and industry (except railroads), and excluded farm and domestic workers and non-farm self-employed, among others. We assume that only those individuals whose 1940 occupations were covered by Social Security as of the 1939 Amendments were eligible, thereby ignoring the large expansions in coverage during the 1950s and ruling out the possibility that more people worked in covered occupations after 1940. For men aged 48–52 in 1940 who had positive earnings, we estimate the share, by earnings level, who were in occupations in 1940 that were covered by Social Security as of 1939. A complication is that with only measures of wage and salary earnings, we do not observe earnings for the self-employed (who were ineligible for Social Security). We assume that self-employment status was independent of earnings, and within each earnings level we simply multiply the share of non-self-employed who were eligible by the share who were non-self-employed to estimate an overall share who were eligible. We follow Wendt (1938) to determine which workers were eligible for Social Security under its original provisions. When Census information on occupation and industry is too coarse we make assumptions that tend to reduce the estimated share eligible. These classifications imply that overall, about 42 percent of this cohort is eligible for our counterfactual Social Security program, whereas as of the end of 1959, 67 percent of men aged 65–74 were actually receiving benefits, based on our calculations from the Census and Social Security Administration (1960).

The second key input to the simulation is the wage histories of people eligible for Social Security. We assume that an individual's average nominal monthly wage over his entire career was 3.6 times its level in 1940. This is the nominal wage that the individual would have received in 1960—at the very end of his career—had he received the average rate of wage growth from 1940 to 1960 among production workers in manufacturing (Carter et al., 2006, Series Ba4362). To the extent that this rate of wage growth was high relative to wage growth overall during the "Great Compression" of the 1940s and 1950s (Goldin and Margo, 1992), it will tend to overstate wage growth of this cohort overall. More important, assuming that members of this cohort received their 1960 wages over their entire careers leads us to significantly overstate their lifetime wages. Overstating wages from 1939 until retirement understates the predicted effects of Social Security on labor supply by understating Social Security replacement rates.

The third key input to the simulation is the counterfactual retirement-age distribution that would have arisen in the absence of OAA and Social Security. We make the conservative assumption that the observed labor force participation-age profile in 1960 is the counterfactual retirement-age distribution that would have arisen in the absence of OAA and Social Security. This is conservative because by 1960 Social Security was a large program that likely had already reduced labor force participation substantially. Our assumption therefore understates counterfactual labor supply in the absence of the program, which tends to reduce the predicted effects of Social Security by reducing the amount of labor available to potentially be reduced by the program. This tends to reduce the effects of Social Security on the key statistic we simulate, labor force participation among people aged 65–74, since it reduces the fraction of people who would otherwise (in the absence of Social Security) retire after age 65. This ensures that our predictions about the likely effects of Social Security are conservative despite the various un-modeled factors that might have increased the demand for retirement, such as private pensions and changes in the prices of leisure goods.

Figure 1(b) shows the cross-sectional labor force participation-age profiles in 1930, 1940, 1950, and 1960. As documented by Costa (1998) and others, the labor force participation-age profile underwent major changes between 1930 and 1960. In 1930, there is no apparent change in the profile at age 65. By 1940, the profile drops slightly at age 65. Our analysis implies that this drop can be explained by the introduction and expansion of OAA during the 1930s. By 1960, it is apparent from the labor force participation-age profile that something special is going on around and after age 65. This is consistent with OAA and Social Security having a major impact on labor supply. Because of the large changes in the 1960 labor force participation-age profile around age 65, when we fit a polynomial to this profile to predict counterfactual labor force participation at ages beyond age 84, we fit it using ages 65 and older only. The resulting polynomial under-predicts labor force participation at ages younger than 65, but labor force participation at these ages is not relevant for the key statistic we wish to simulate, the reduction in labor force participation at ages 65–74.

Our simulations ignore OAA entirely. We do this to be conservative in terms of the total effect of government old-age support over this period, since OAA should have reduced labor supply still further. An important caveat, though, is that because our comparison is to a scenario with no old-age support, program substitution from OAA to Social Security would reduce the implied effect of Social Security relative to the *observed* level of labor force participation in 1940 (which was already lower because of OAA). The share of Social Security-eligible men who were also eligible for OAA is likely to be slightly lower than the overall OAA eligibility share (which we estimate to be 22 percent), since the earnings of men we classify as Social Security-eligible tend to be higher than those we estimate to be OAA-eligible. A rough correction would be to suppose that about 20 percent of men who left the labor force to take up Social Security would otherwise have taken up OAA, which would suggest multiplying our estimates by about 0.8. Although the model does not include some other factors that may have reduced late-life labor supply over this time period, such as private pensions (Stock and Wise, 1990; Samwick, 1998) and changes in relative prices, especially those of leisure substitutes and complements (Costa, 1998), we capture the combined effect of such factors on labor supply by using observed labor force participation in 1960 as our no-Social Security counterfactual level. This assumption tends to reduce the implied effect of Social Security.

We do not attempt to evaluate the welfare costs of the Social Security earnings test to recipients. In addition to our thought experiment being a policy experiment that was never actually realized, the set of assumptions we make to understate the overall impact of Social Security unfortunately makes it difficult to sign the bias in the cost of the earnings test. On the one hand, understating benefits reduces implied replacement rates, which tends to overstate the costs of the earnings test. On the other hand, understating counterfactual no-program labor supply in 1960 overstates the growth in the demand for retirement due to non-program factors like wage growth, which tends to understate the costs of the earnings test by making more years of retirement inframarginal.

A.10.3 Decomposition of the effects of OAA on retirement into income and substitution effects

We decompose the effects of OAA into income and substitution effects using the following method. We solve for the optimal retirement age under three budget constraints: OAA, No OAA, and "No OAA with Compensation."[37] We consider two different "No OAA with Compensation" budget constraints. Each is identical to the No OAA budget constraint except for one change. In one case, initial assets are increased exactly enough that the individual is able to achieve exactly the same utility that he would achieve under OAA. In the other case, non-labor income after the OAA eligibility age is increased exactly enough that the individual is able to achieve exactly the same utility that he would achieve under OAA. If capital markets were perfect, the individual would be indifferent between receiving an immediate transfer of assets and receiving a present value-equivalent increase in his future non-labor income. But with borrowing constraints, individuals weakly prefer an increase in initial assets to a present value-equivalent increase in late-life income. The estimated equivalent variation of OAA is therefore weakly greater under the late-life income compensation than it is under the initial assets compensation. In the text, we discuss the equivalent variation of OAA based on both measures, but for measuring income effects we use the late-life income compensation.

The income effect of OAA is the number of years earlier that people retire under the "No OAA with Compensation" budget constraint relative to the No OAA budget constraint due to being richer with OAA.[38] The substitution effect of OAA is the number of years earlier that people retire under the OAA budget constraint relative to the "No OAA with Compensation" budget constraint due to the taxation of late-life labor supply implicit in

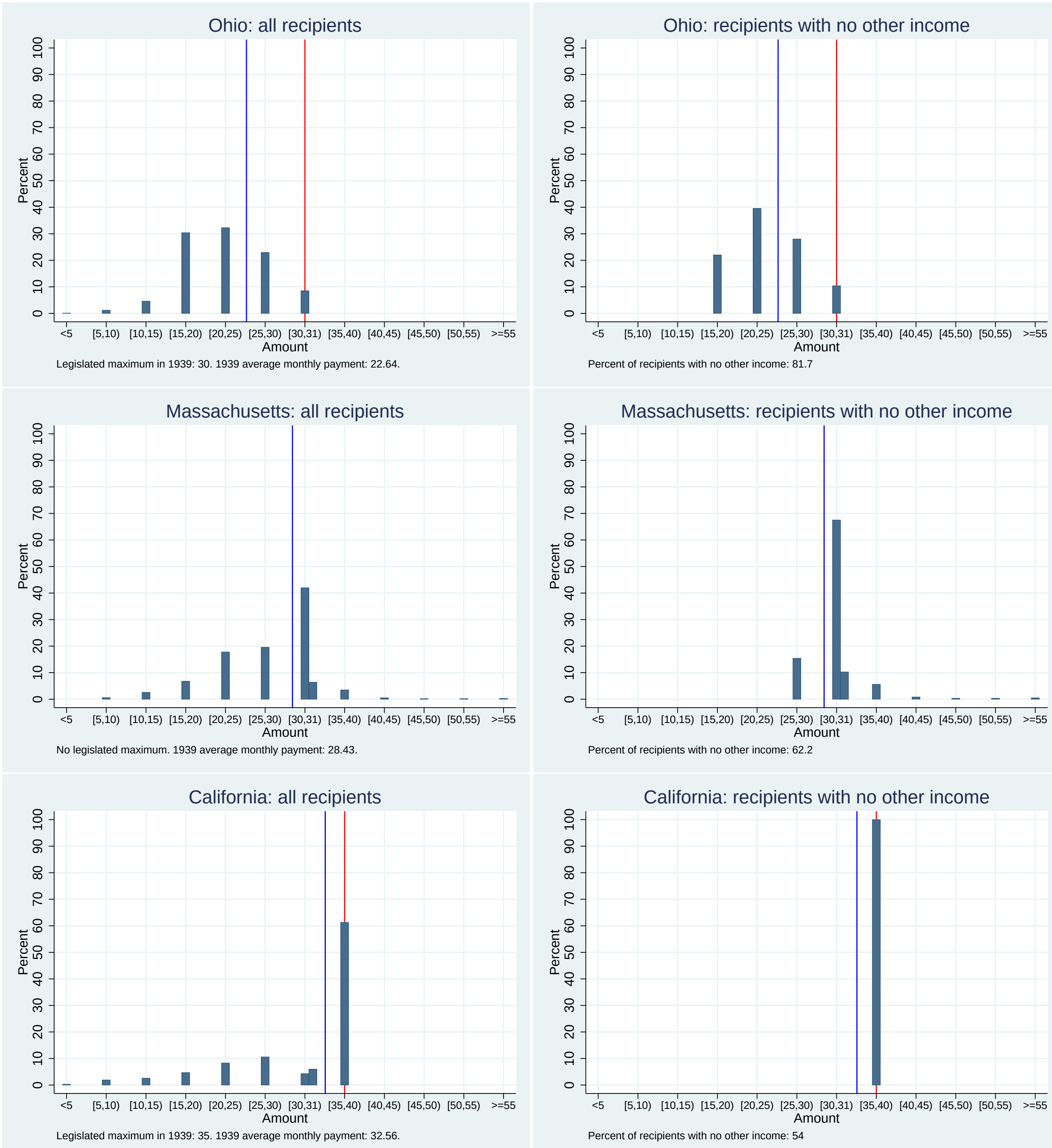
[37] We hold utility fixed at the level of utility the individual achieves with OAA in order to ensure invertibility in the presence of borrowing constraints.

[38] Recipients of OAA likely had their opportunity sets expanded by OAA since it was means-tested.

OAA's means tests.

Appendix Figures and Tables

Figure A1: Distributions of payments to new recipients in 1938-39, by state



Notes: Left figures show distributions of payment amounts to new recipients in 1938-39 by state, based on data from U.S. Social Security Board (1939*b*). Vertical lines correspond to average monthly payment and legislated maximum payment (if one existed) in 1939. Right figures show estimated distribution for recipients with no other source of income, under the assumption that those with other sources of income received the lowest payments.

Figure A2: OAA statutory maximum monthly payments, 1939

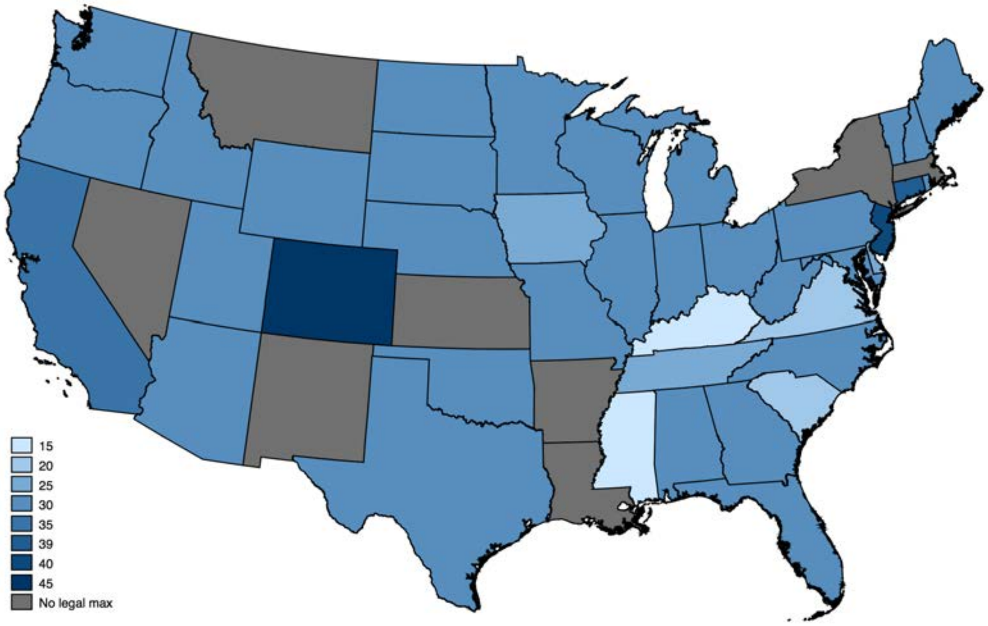


Figure shows statutory maximum monthly payment, from U.S. Social Security Board (1940*a*). Analysis sample excludes Colorado, Missouri, New Hampshire, and Pennsylvania.

Figure A3: OAA 99th percentile payments 1939

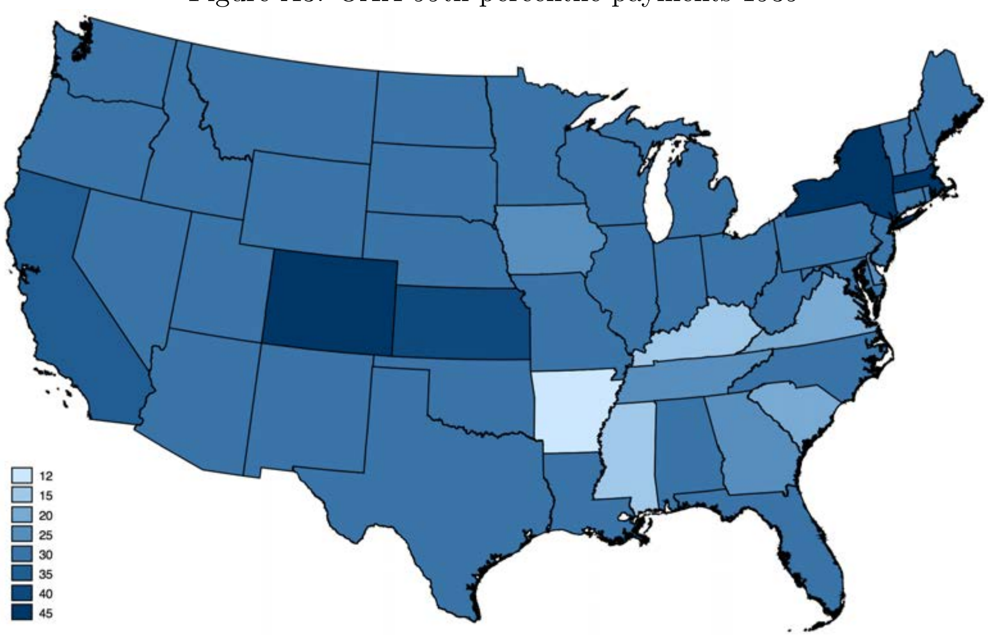


Figure shows estimate of 99th percentile payment, based on data from U.S. Social Security Board (1939*b*). For details on construction of this measure, see Appendix Section A.2.2. Analysis sample excludes Colorado, Missouri, New Hampshire, and Pennsylvania.

Figure A4: OAA variation and missing data

(a) Demographic and 1940 work (b) Demographic and 1939 work

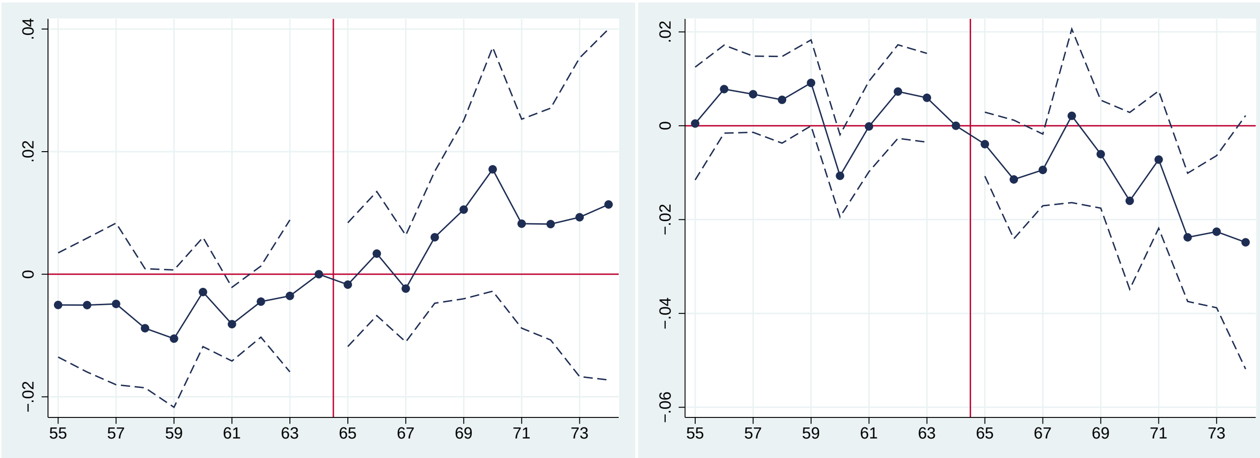


Figure shows point estimates and 95% confidence intervals on age-payment interactions from estimation of equation (1), using log simulated payment by age interactions as instruments for log observed payment by age interactions and controlling for state border by age fixed effects. Dependent variable in Panel (a) is missing (or allocated) information on demographics or 1940 labor force status; dependent variable for Panel (b) is missing (or allocated) information on demographics or 1939 work or income information. Demographic variables are sex, race, marital status, years of education, birthplace, and citizenship. Sample restricted to counties on state boundaries, excluding counties on borders of states with age eligibility requirement other than 65 in 1939. Unit of observation is an individual-state border pair. All specifications include county fixed effects and age fixed effects. Standard errors clustered at the state level. For both panels, $N = 2675836$ and Kleibergen-Paap rk Wald F-statistic is 3.06.

Figure A5: Distribution of monthly income in population used for simulated IV

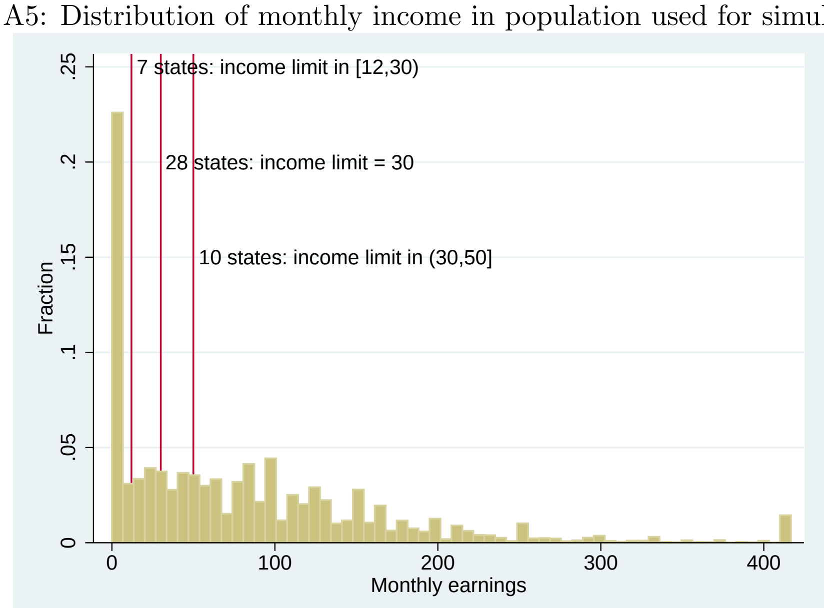


Figure shows distribution of monthly earnings in the population of men used for calculation of the simulated instrument (men aged 60-64 in states with an eligibility age of 65 in 1939). Monthly earnings is imputed for men reporting positive weeks worked in 1939 who were self-employed at the time of the Census, as described in Section A.2.3. Vertical lines show minimum, median, and maximum values of “income limit” (the sum of maximum payments and any income disregards) across states. These values are \$12, \$30, and \$50 per month.

Figure A6: Relationship between actual and simulated OAA payments

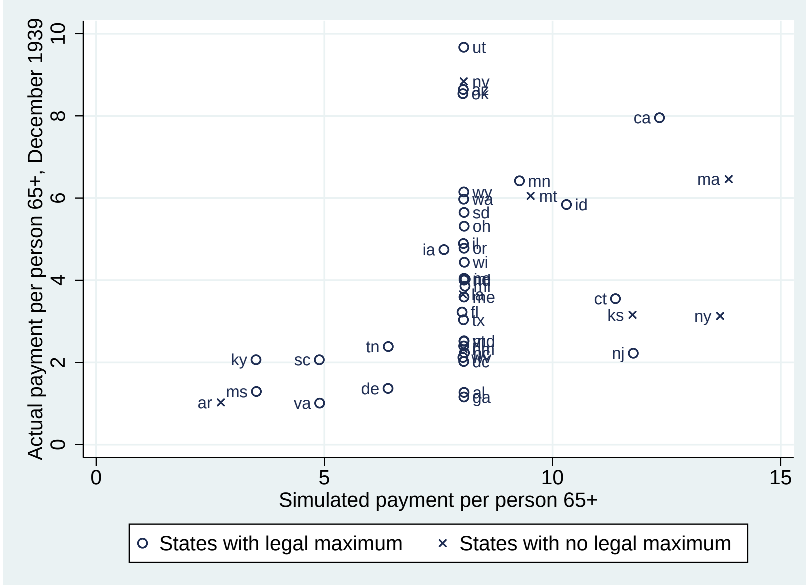


Figure shows relationship between observed state-level OAA payments per person 65 and older in December 1939 and simulated payments based on maximum payments and income disregards. "Maximum payment" is the statutory maximum monthly payment for those states with a statutory maximum and the 99th percentile payment for states without a statutory maximum. Only states with eligibility age of 65 in 1939 are included.

Figure A7: Relationship of maximum payments to payments per recipient

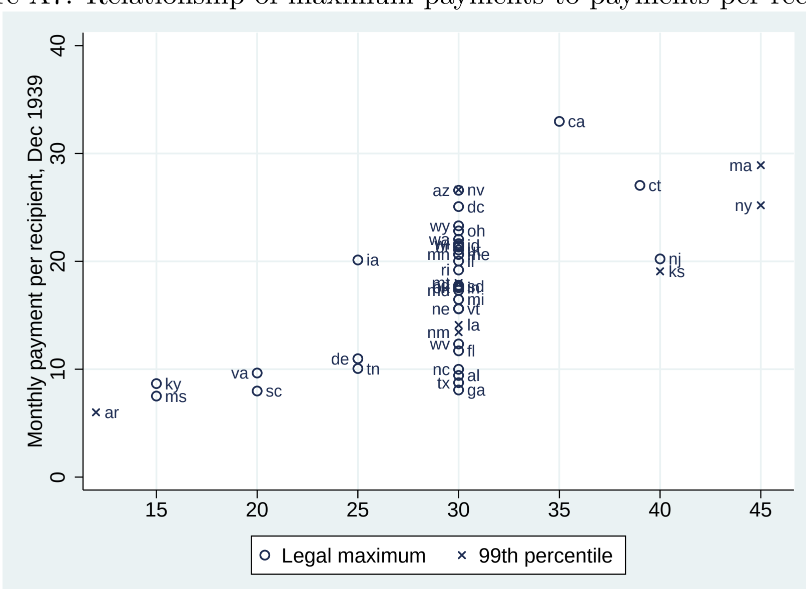
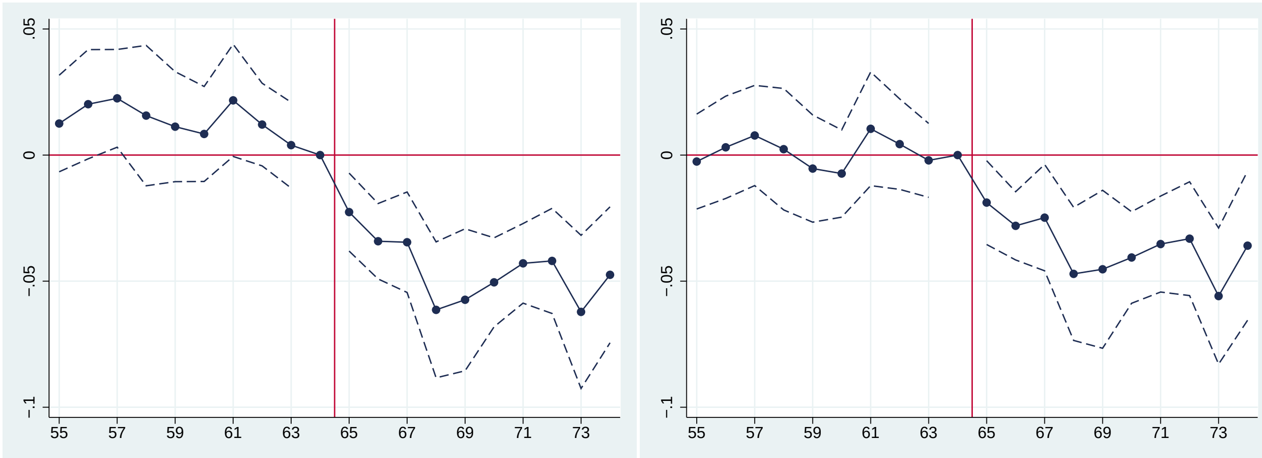


Figure shows relationship between average payment per recipient in December 1939 and statutory maximum monthly payment (for those states that had them) or 99th percentile payment (for states without a statutory maximum). Only states with eligibility age of 65 in 1939 are included. Sources: data on OAA dollar payments and number of recipients from U.S. Social Security Board (1940*b*), data on legal maximum payments from U.S. Social Security Board (1940*a*).

Figure A8: Alternative measures of labor force attachment

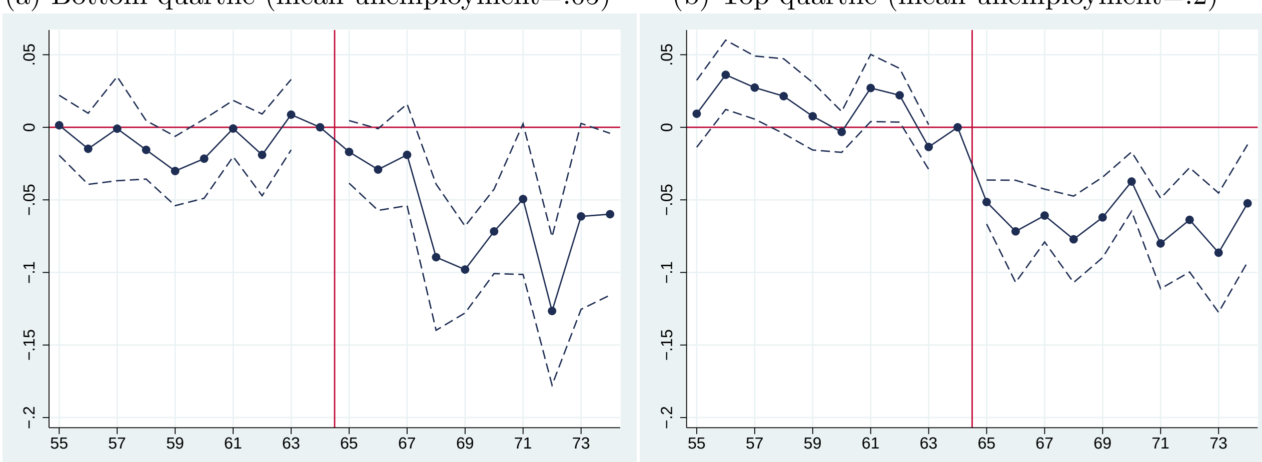
(a) employment (b) non-emergency employment



Notes: Figure shows point estimates and 95% confidence intervals on age-payment interactions from estimation of equation (1), using log simulated payment by age interactions as instruments for log per-65+ payment by age interactions and controlling for state border by age fixed effects. Standard errors clustered at the state level. For both panels, $N = 2403915$ and Kleibergen-Paap rk Wald F-stat is 3.06.

Figure A9: Effect of OAA on labor force participation by county unemployment rates

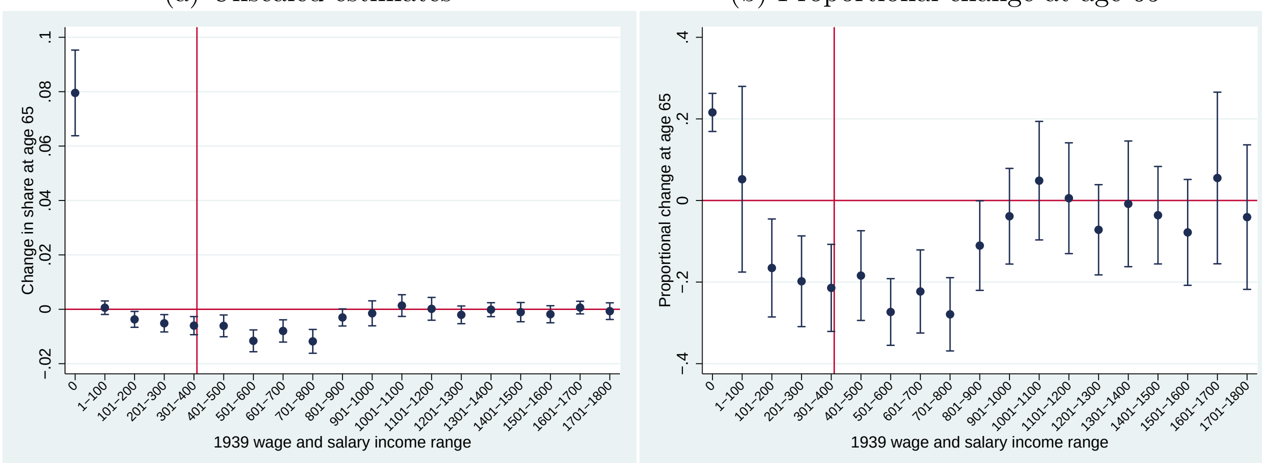
(a) Bottom quartile (mean unemployment=.05) (b) Top quartile (mean unemployment=.2)



Notes: Figures show point estimates and 95% confidence intervals on age-payment interactions from estimation of equation (1), using log simulated payment by age interactions as instruments for log per-65+ payment by age interactions and controlling for state border by age fixed effects. Panel (a) limits sample to counties in the bottom quartile of county unemployment rates, not weighting counties by population ($N = 306124$) and Panel (b) limits to counties in the top quartile of county unemployment rates ($N = 807613$). County unemployment rate is that of 45-54 year old men and includes work relief in unemployment. Standard errors clustered at the state level.

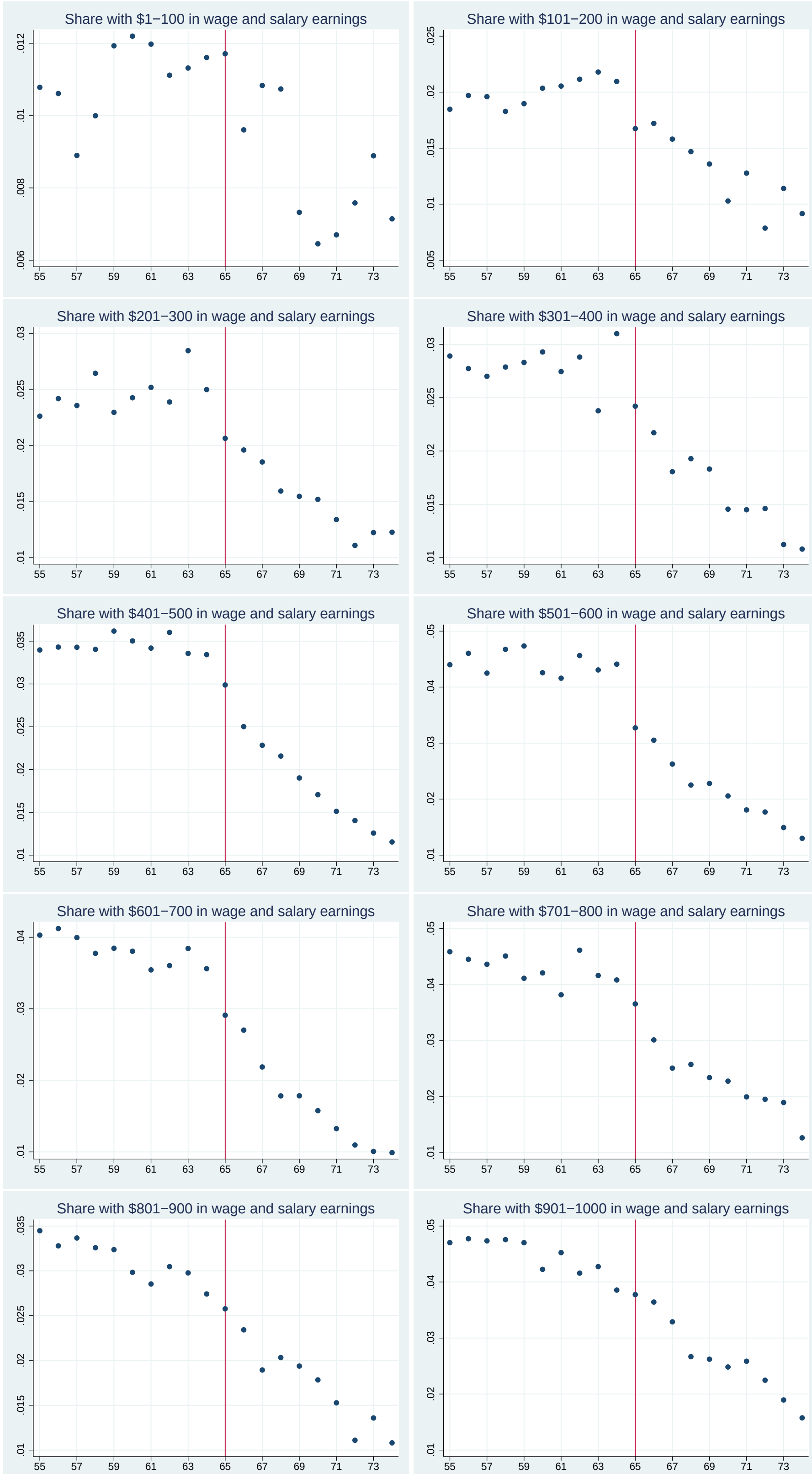
Figure A10: Change at 65 in share of men with specified amount of wage/salary income in 1939

(a) Unscaled estimates (b) Proportional change at age 65



Notes: Figures show point estimates and 95% confidence intervals from separate estimations of equation (2), with dependent variable indicating wage/salary earnings of each specified amount in 1939. Sample: men within IK bandwidth around age 65 at 1940 Census in Massachusetts. Vertical line denotes "income floor" of \$360 per year. Standard errors clustered by years of age. Panel (a) shows estimates of β_1; Panel (b) shows estimates of β_1/β_0 to measure proportional change at age 65 (with standard errors calculated using the delta method).

Figure A11: Share of Massachusetts men with specified 1939 wage and salary earnings

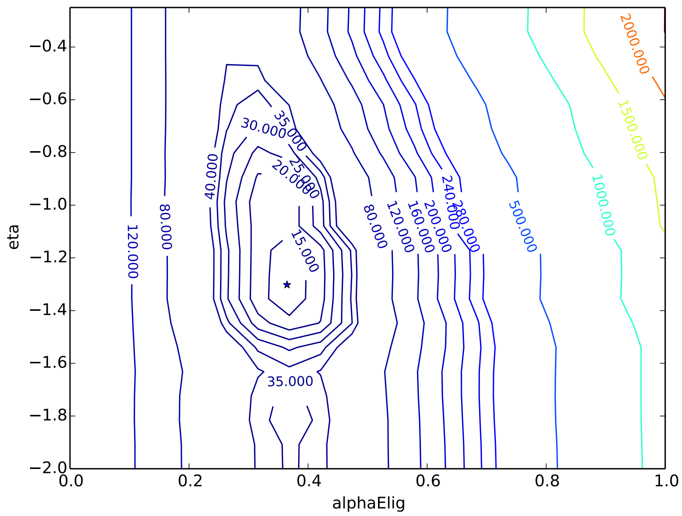


Notes: Figures show share of men reporting 1939 wage and salary earnings in specified range, by age at 1940 Census.

Figure A12: Share eligible for OAA based on their housing wealth

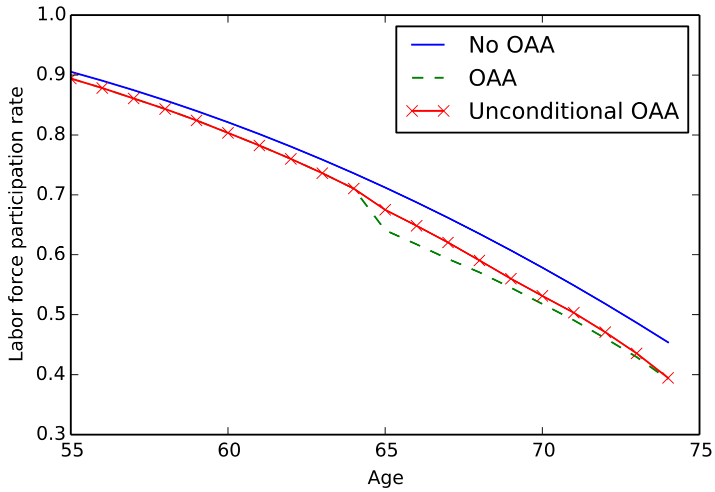
Notes: Share of Massachusetts men aged 60–64 who had less than $3,000 of house value, as a function of wage and salary income. Massachusetts limited eligibility for OAA to people with less than $3,000 in real property and did not have any home disregard. The figure therefore shows the share of people who were not ineligible for OAA on the basis of their house value alone.

Figure A13: Method of simulated moments objective function



Notes: Method of simulated moments objective as a function of η (“eta,” the negative of the coefficient of relative risk aversion) and α_e (“alphaElig,” the constant in the eligibility-potential earnings relationship). Higher contours indicate a worse fit of the model. The asterisk marks the estimated values.

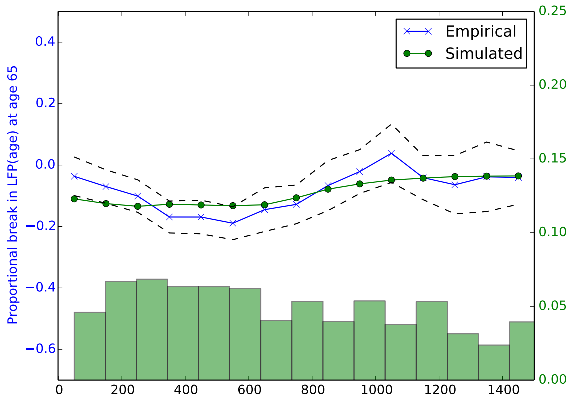
Figure A14: Simulated cross-sectional age-labor force participation profile in 1940



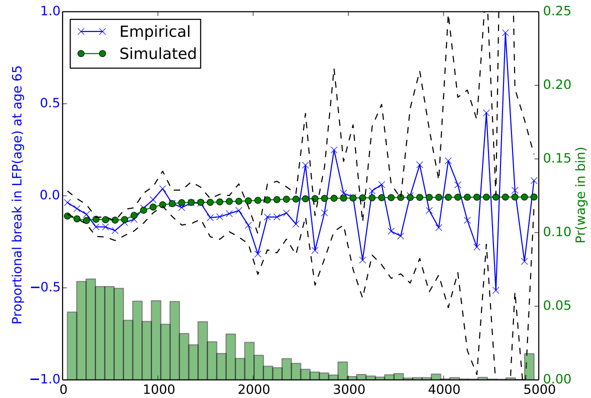
Notes: Simulated cross-sectional relationship between labor force participation and age in 1940 in the US. The "No OAA" profile is the counterfactual no-OAA profile predicted based on our regression results and presented in Figure 7. The "OAA" profile is simulated based on the estimated model. It can be compared to its empirical counterpart, also depicted in Figure 7. The "Unconditional OAA" profile is simulated based on the estimated model using a counterfactual OAA program that did not impose an earnings test. The difference between this figure and Figure 11 is that this figure focuses on the 1940 cross section, whereas Figure 11 focuses on the life cycle profiles of the cohort of men aged 55 in 1940. The predicted effects of OAA in the 1940 cross section are smaller than those over the life cycle of the cohort of men aged 55 in 1940 because in the latter case people had more time to build OAA into their plans.

Figure A15: Empirical vs. simulated moments for full-US estimation

(a) Moments in the estimation

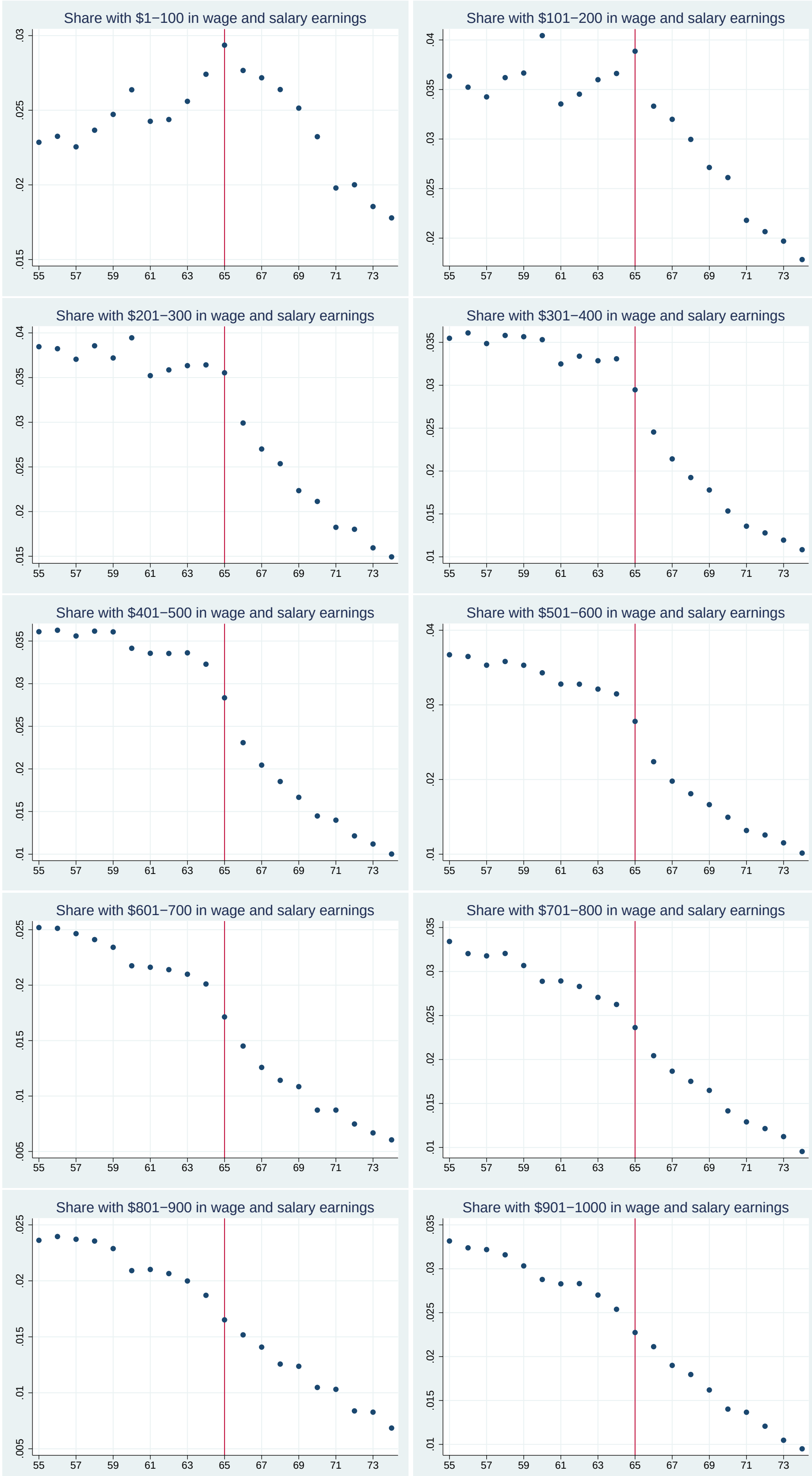


(b) All earnings levels



Notes: Empirical vs. simulated moments and annual earnings distribution for the full US for moments in the estimation (Panel (a)) and for all earnings levels, including those not in the estimation (Panel (b)). The moments are the proportional breaks in labor force participation-age profiles at age 65. Empirical moments correspond to the breaks at age 65 in the share of men with the specified amount of wage/salary income in 1939, relative to the predicted share at age 65 based on data from younger ages. The earnings distribution is the distribution of wage/salary income among men in the US aged 60–64 in 1939 who had any wage/salary income. For reference, the average OAA benefit in the US is $232 per year. Earnings above $5,000 are set to $5,000.

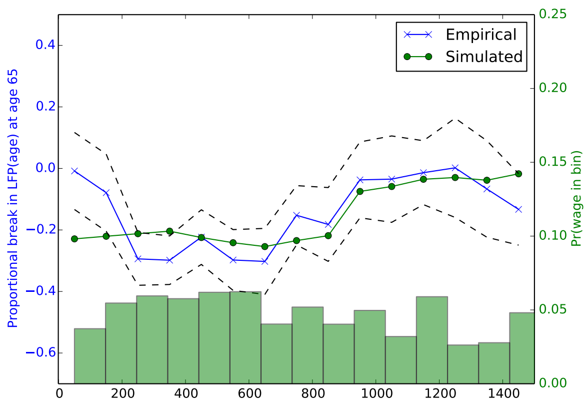
Figure A16: Share of all men with specified 1939 wage and salary earnings



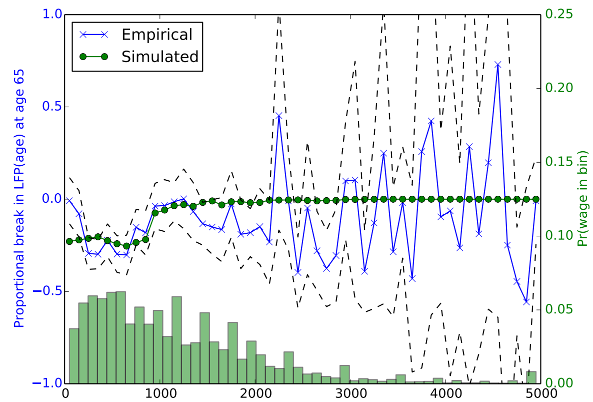
Notes: Figures show share of men reporting 1939 wage and salary earnings in specified range, by age at 1940 Census, for all states with an eligibility age of 65 in 1939.

Figure A17: Empirical vs. simulated moments for California estimation

(a) Moments in the estimation

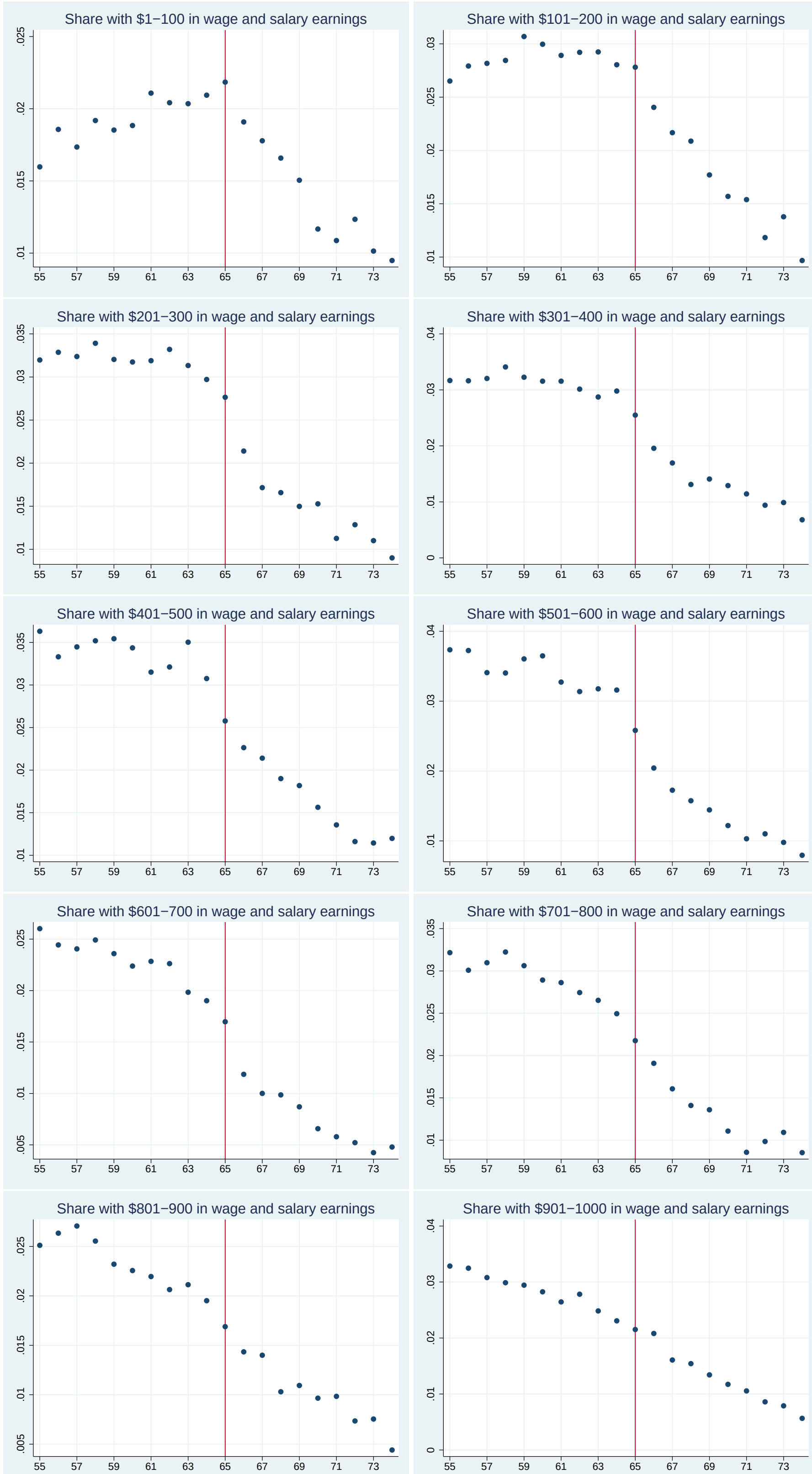


(b) All earnings levels



Notes: Empirical vs. simulated moments and annual earnings distribution for California for moments in the estimation (Panel (a)) and for all earnings levels, including those not in the estimation (Panel (b)). The moments are the proportional breaks in labor force participation-age profiles at age 65. Empirical moments correspond to the breaks at age 65 in the share of men with the specified amount of wage/salary income in 1939, relative to the predicted share at age 65 based on data from younger ages. The earnings distribution is the distribution of wage/salary income among men in California aged 60–64 in 1939 who had any wage/salary income. For reference, the "income floor" in California is \$420 per year. Earnings above \$5,000 are set to \$5,000.

Figure A18: Share of California men with specified 1939 wage and salary earnings



Notes: Figures show share of men in California reporting 1939 wage and salary earnings in specified range, by age at 1940 Census.

Table A1: Monthly payments in 1939

State	Age eligible	Legal max payment	Observed max payment	99th percentile payment	Payment per recipient	Payment per person 65+
Virginia	65	20	20	20	9.65	1.01
Arkansas	65	.	12	12	6.01	1.03
Georgia	65	30	30	25	8.07	1.16
Alabama	65	30	111	30	9.42	1.27
Mississippi	65	15	15	15	7.51	1.29
Delaware	65	25	25	25	10.98	1.37
New Hampshire	70	30	30	30	20.95	1.98
District of Columbia	65	30	39	30	25.08	2.02
South Carolina	65	20	20	20	7.98	2.07
Kentucky	65	15	15	15	8.66	2.07
West Virginia	65	30	30	30	12.34	2.12
New Jersey	65	40	30	30	20.22	2.22
North Carolina	65	30	30	30	9.99	2.23
New Mexico	65	.	42	30	13.43	2.33
Tennessee	65	25	25	25	10.06	2.39
Rhode Island	65	30	30	30	19.20	2.40
Maryland	65	30	30	30	17.31	2.52
Pennsylvania	70	30	30	30	21.77	2.52
Vermont	65	30	30	30	15.60	2.53
Texas	65	30	30	30	8.75	3.04
New York	65	.	86	45	25.20	3.13
Kansas	65	.	94	40	19.07	3.16
Florida	65	30	30	30	11.70	3.23
Connecticut	65	39	30	30	27.04	3.55
Maine	65	30	30	30	20.64	3.59
Louisiana	65	.	46	30	14.10	3.66
Michigan	65	30	30	30	16.47	3.86
North Dakota	65	30	30	30	17.78	4.00
Indiana	65	30	30	30	17.55	4.02
Nebraska	65	30	30	30	15.61	4.05
Wisconsin	65	30	30	30	21.65	4.44
Missouri	70	30	30	30	18.90	4.57
Iowa	65	25	25	25	20.13	4.75
Oregon	65	30	30	30	21.33	4.78
Illinois	65	30	30	30	20.03	4.89
Ohio	65	30	30	30	22.82	5.31
South Dakota	65	30	30	30	17.67	5.65
Idaho	65	30	30	30	21.47	5.84
Washington	65	30	30	30	22.04	5.97
Montana	65	.	30	30	17.99	6.05
Wyoming	65	30	30	30	23.29	6.15
Minnesota	65	30	30	30	20.64	6.42
Massachusetts	65	.	91	45	28.91	6.46
California	65	35	35	35	32.97	7.95
Oklahoma	65	30	30	30	17.59	8.54
Arizona	65	30	30	30	26.58	8.64
Nevada	65	.	30	30	26.64	8.84
Utah	65	30	47	30	21.06	9.67
Colorado	60 or 65	45	45	45	28.44	13.17

Notes: Includes the 48 states and the District of Columbia. '99th percentile payment' is for new recipients in fiscal year 1938-39. Eight states had no legal maximum payment. Recipiency rate and payments per person 65+ are for December 1939, and are normalized by state population from 1940 Census. Sources: data on OAA dollar payments and number of recipients from U.S. Social Security Board (1940*b*), data on legal maximum payments from U.S. Social Security Board (1940*a*), data on observed maximum payments and 99th percentile payment from U.S. Social Security Board (1939*b*).

Table A2: Variation in log OAA payments per person 65+ across states

	(1)	(2)	(3)	(4)	(5)
Share population 65 and above	11.895*				
	(5.759)				
Share population foreign born		4.103*			
		(1.555)			
Share population non-white			-2.878***		
			(0.491)		
Median years of education				0.364***	
				(0.076)	
Log median earnings					1.059***
					(0.244)
Observations	45	45	45	45	45

Dependent variable: log of OAA payments in December 1939 per person 65 and above. Sample includes states with 1939 eligibility age of 65. Median years of education is calculated for all people aged 25-54 in that state, median earnings is state median wage and salary earnings in 1939 for men aged 25-54 who were not self-employed. Heteroskedasticity-robust standard errors in parentheses. *: $p < 0.05$, **: $p < 0.01$, ***: $p < 0.001$

Table A3: Variation in Log OAA payments per person 65+ for border counties

Dependent variable	Share 65 and above	Share foreign born	Share nonwhite	Median years of schooling	Log median earnings
Panel A. Observed payments variable, no border fixed effects					
Log per-65+ payment	0.010**	0.033***	-0.127***	1.137***	0.291***
	(0.003)	(0.007)	(0.027)	(0.136)	(0.072)
Observations	1192	1192	1192	1183	1183
Panel B. Observed payments variable, border fixed effects					
Log per-65+ payment	0.002	0.003	0.015*	-0.045	-0.045
	(0.001)	(0.002)	(0.007)	(0.132)	(0.046)
Observations	1192	1192	1192	1183	1183
Panel C. Simulated payments variable, no border fixed effects					
Log simulated per-65+ payment	0.014*	0.058***	-0.175*	1.258***	0.534***
	(0.006)	(0.012)	(0.075)	(0.303)	(0.071)
Observations	1192	1192	1192	1183	1183
Panel D. Simulated payments variable, border fixed effects					
Log simulated per-65+ payment	0.002	0.001	-0.011	-0.205	-0.013
	(0.002)	(0.001)	(0.012)	(0.176)	(0.038)
Observations	1192	1192	1192	1183	1183

Sample: border counties in states with 1939 eligibility age of 65. Unit of observation is a county-state border pair. Smaller sample size for schooling and earnings is due to missing data in nine small border counties. Standard errors (in parentheses) are clustered at the state level. *: $p < 0.05$, **: $p < 0.01$, ***: $p < 0.001$

Table A4: Simulated IV first stage regressions

	(1) age 55-59	(2) age 65-69	(3) age 70-74
Log simulated per-65+ payment × age 55-59	0.897***	0.000	-0.000
	(0.113)	(0.001)	(0.000)
Log simulated per-65+ payment × age 65-69	0.002	0.892***	0.001
	(0.003)	(0.114)	(0.002)
Log simulated per-65+ payment × age 70-74	-0.004	-0.002	0.907***
	(0.003)	(0.002)	(0.110)
Observations	2403915	2403915	2403915
Sample	border	border	border
Border segment × age fixed effects	yes	yes	yes
Education × age fixed effects	yes	yes	yes
Race × age fixed effects	yes	yes	yes

Dependent variables: log state OAA payments per person 65+ in December 1939, interacted with indicator for specified age group. Sample restricted to counties on state boundaries, excluding counties on borders of states with age eligibility requirement other than 65 in 1939. Unit of observation is an individual-state border pair. All specifications include county fixed effects and 5-year age group fixed effects. All age-interactions are with 5-year age groups. Standard errors (in parentheses) are clustered at the state level. *: $p < 0.05$, **: $p < 0.01$, ***: $p < 0.001$

Table A5: Non-wage income by state payments per person 65+ and age

Panel A. OLS results				
	(1)	(2)	(3)	(4)
Log per-65+ payment	-0.010**	-0.005	0.006	0.006
× age 55-59	(0.003)	(0.004)	(0.005)	(0.005)
Log per-65+ payment	0.063***	0.062***	0.052***	0.053***
× age 65-69	(0.005)	(0.007)	(0.003)	(0.003)
Log per-65+ payment	0.093***	0.088***	0.073***	0.073***
× age 70-74	(0.009)	(0.011)	(0.005)	(0.006)
Observations	6283146	2238476	2238476	2238476
Sample	full	border	border	border
Border segment × age fixed effects	no	no	yes	yes
Education × age fixed effects	no	no	no	yes
Race × age fixed effects	no	no	no	yes
Panel B. IV results				
	(1)	(2)	(3)	(4)
Log per-65+ payment	-0.018**	-0.007	0.007	0.007
× age 55-59	(0.006)	(0.007)	(0.006)	(0.006)
Log per-65+ payment	0.066***	0.062***	0.061***	0.061***
× age 65-69	(0.008)	(0.009)	(0.006)	(0.006)
Log per-65+ payment	0.109***	0.100***	0.075***	0.075***
× age 70-74	(0.018)	(0.015)	(0.008)	(0.008)
Observations	6283145	2238476	2238476	2238476
Kleibergen-Paap rk Wald F-stat	1.98	8.79	21.60	21.66
Sample	full	border	border	border
Border segment × age fixed effects	no	no	yes	yes
Education × age fixed effects	no	no	no	yes
Race × age fixed effects	no	no	no	yes

Dependent variable is indicator for receipt of more than $50 in non-wage income in 1939. In Panel B, log simulated payment by age interactions used as instruments for log per-65+ payment by age interactions. Sample for column (1): men aged 55-74 in states with 1939 eligibility age of 65. Columns (2)-(4) include only counties on state boundaries and exclude counties on borders of excluded states. Unit of observation in columns (2)-(4) is an individual-state border pair. All specifications include county fixed effects and 5-year age group fixed effects. All age-interactions are with 5-year age groups. Standard errors (in parentheses) are clustered at the state level. *: $p < 0.05$, **: $p < 0.01$, ***: $p < 0.001$

Table A6: Main results using payments per person 65+ in levels

	(1) Non-wage income	(2) In labor force	(3) Employed	(4) Non-emergency
Per-65+ payment × age 55-59	-0.000	0.001	0.002	-0.001
	(0.002)	(0.002)	(0.002)	(0.001)
Per-65+ payment × age 65-69	0.019***	-0.020***	-0.014***	-0.009**
	(0.003)	(0.004)	(0.004)	(0.003)
Per-65+ payment × age 70-74	0.024***	-0.023***	-0.017***	-0.012***
	(0.004)	(0.004)	(0.004)	(0.002)
Observations	2238476	2403915	2403915	2403915
Kleibergen-Paap rk Wald F-stat	10.48	10.63	10.63	10.63
Sample	border	border	border	border
Border segment × age fixed effects	yes	yes	yes	yes
Education × age fixed effects	yes	yes	yes	yes
Race × age fixed effects	yes	yes	yes	yes

Dependent variables: receipt of non-wage income in 1939, in labor force at 1940 Census, employed at 1940 Census, employed in private or non-emergency government work at 1940 Census. Simulated payment by age interactions used as instruments for per-65+ payment by age interactions. Payments in 1940 dollars. Sample restricted to counties on state boundaries, excluding counties on borders of states with age eligibility requirement other than 65 in 1939. Unit of observation is an individual-state border pair. All specifications include county fixed effects and 5-year age group fixed effects. All age-interactions are with 5-year age groups. Standard errors (in parentheses) are clustered at the state level. *: $p < 0.05$, **: $p < 0.01$, ***: $p < 0.001$

Table A7: Test for heterogeneous labor force participation effects by county age 45-54 unemployment

	(1)	(2)
Unemployment rate × Log per-65+ payment × age 55-59	0.028	0.021
	(0.068)	(0.061)
Unemployment rate × Log per-65+ payment × age 65-69	0.059	0.106
	(0.141)	(0.129)
Unemployment rate × Log per-65+ payment × age 70-74	0.308	0.346*
	(0.170)	(0.156)
Log per-65+ payment × age 55-59	-0.000	0.001
	(0.010)	(0.009)
Log per-65+ payment × age 65-69	-0.063***	-0.068***
	(0.018)	(0.017)
Log per-65+ payment × age 70-74	-0.101***	-0.107***
	(0.023)	(0.022)
Observations	2402073	2402073
Kleibergen-Paap rk Wald F-stat	10.78	10.82
Sample	border	border
Border segment × age fixed effects	yes	yes
Education × age fixed effects	no	yes
Race × age fixed effects	no	yes

Dependent variable: in labor force at 1940 Census. Log simulated per-65+ payments used as instruments for observed log per-65+ payments. Sample: men aged 55-74 in states with 1939 eligibility age of 65, including only individuals in counties on state boundaries. All specifications include county fixed effects, 5-year age group fixed effects, interactions of age group effects with the unemployment rate, and border segment by age fixed effects. Unemployment rate is that of 45-54 year old men living in the individual's county and includes work relief in unemployment. Standard errors (in parentheses) are clustered at the state level. *: $p < 0.05$, **: $p < 0.01$, ***: $p < 0.001$

Table A8: Alternative simulated IV specifications

	(1) Non-wage income	(2) In labor force	(3) Employed	(4) Non-emergency
Log per-65+ payment	-0.016	0.007	0.012	0.004
× age 55-59	(0.018)	(0.008)	(0.009)	(0.008)
Log per-65+ payment	0.050***	-0.061***	-0.041***	-0.011
× age 65-69	(0.011)	(0.013)	(0.012)	(0.012)
Log per-65+ payment	0.054***	-0.093***	-0.070***	-0.036*
× age 70-74	(0.016)	(0.021)	(0.017)	(0.014)
Observations	2238476	2403915	2403915	2403915
Kleibergen-Paap rk Wald F-stat	2.09	2.64	2.64	2.64
Sample	border	border	border	border
Border segment × age fixed effects	yes	yes	yes	yes
Education × age fixed effects	yes	yes	yes	yes
Race × age fixed effects	yes	yes	yes	yes

Dependent variables: receipt of non-wage income in 1939, in labor force at 1940 Census, employed at 1940 Census, employed in private or non-emergency government work. Log simulated per-65+ payment by age interactions used as instruments for log per-65+ payment by age interactions. Simulated IV based on maximum payments (and any earnings disregards), assigning the highest legal maximum across states (45 dollars per month) to states with no legal maximum. Sample restricted to counties on state boundaries, excluding counties on borders of states with age eligibility requirement other than 65 in 1939. Unit of observation is an individual-state border pair. All specifications include county fixed effects and 5-year age group fixed effects. All age-interactions are with 5-year age groups. Standard errors (in parentheses) are clustered at the state level. *: $p < 0.05$, **: $p < 0.01$, ***: $p < 0.001$

Table A9: Controls for railroad pensions and state/local government pensions

	(1)	(2)	(3)	(4)
Log per-65+ payment	0.006	0.005	-0.009	-0.009
× age 55-59	(0.005)	(0.004)	(0.005)	(0.005)
Log per-65+ payment	-0.063***	-0.060***	-0.059***	-0.060***
× age 65-69	(0.008)	(0.007)	(0.010)	(0.010)
Log per-65+ payment	-0.075***	-0.069***	-0.069***	-0.070***
× age 70-74	(0.010)	(0.008)	(0.011)	(0.011)
Observations	2403915	2403915	2375865	2375865
Kleibergen-Paap rk Wald F-stat	20.53	23.11	11.43	13.20
Sample	border	border	border	border
Border segment × age fixed effects	yes	yes	yes	yes
Education × age fixed effects	yes	yes	yes	yes
Race × age fixed effects	yes	yes	yes	yes
Railroad × age fixed effects	no	yes	no	yes
State/local × age fixed effects	no	no	yes	yes

Dependent variable: in labor force at 1940 Census. Log simulated OAA payment by age interactions used as instruments for log per-65+ OAA payment by age interactions. Sample in all columns: men aged 55-74 in states with 1939 eligibility age of 65 and living in counties on state borders (columns 3 and 4 additionally omit states with missing information on state and local pensions). For definitions of railroad and state and local pension payments, see the text. Unit of observation is an individual-state border pair. All specifications include county fixed effects and 5-year age group fixed effects. All age-interactions are with 5-year age groups. Standard errors (in parentheses) are clustered at the state level. *: $p < 0.05$, **: $p < 0.01$, ***: $p < 0.001$

Table A10: Controls for general assistance

	(1)	(2)
Log per-65+ payment	0.005	-0.009
× age 55-59	(0.006)	(0.012)
Log per-65+ payment	-0.068***	-0.088*
× age 65-69	(0.015)	(0.039)
Log per-65+ payment	-0.067***	-0.094*
× age 70-74	(0.017)	(0.041)
Observations	2097968	2097968
Kleibergen-Paap rk Wald F-stat	7.34	1.54
Sample	border	border
Border segment × age fixed effects	yes	yes
Education × age fixed effects	yes	yes
Race × age fixed effects	yes	yes
Genl asst × age fixed effects	no	yes

Dependent variable: in labor force at 1940 Census. Log simulated OAA payment by age interactions used as instruments for log per-65+ OAA payment by age interactions. Sample in all columns: men aged 55-74 in states with 1939 eligibility age of 65 and non-missing general assistance data, and living in counties on state borders. Unit of observation is an individual-state border pair. All specifications include county fixed effects and 5-year age group fixed effects. All age-interactions are with 5-year age groups. Standard errors (in parentheses) are clustered at the state level. *: $p < 0.05$, **: $p < 0.01$, ***: $p < 0.001$

Table A11: Cross-state migration 1935-40 by state payments per person 65+ and age

	(1)	(2)	(3)	(4)
Log per-65+ payment	-0.0015	-0.0022	-0.0006	-0.0006
× age 55-59	(0.0022)	(0.0023)	(0.0021)	(0.0021)
Log per-65+ payment	0.0044	0.0005	0.0031*	0.0033*
× age 65-69	(0.0032)	(0.0027)	(0.0014)	(0.0014)
Log per-65+ payment	0.0059	0.0012	0.0016	0.0017
× age 70-74	(0.0054)	(0.0058)	(0.0037)	(0.0037)
Observations	6619726	2366217	2366217	2366217
Kleibergen-Paap rk Wald F-stat	1.98	8.30	20.56	20.63
Sample	full	border	border	border
Border segment × age fixed effects	no	no	yes	yes
Education × age fixed effects	no	no	no	yes
Race × age fixed effects	no	no	no	yes

Dependent variable: moved states between 1935 and 1940. Log simulated payment by age interactions used as instruments for log per-65+ payment by age interactions. Sample for column (1): men aged 55-74 in states with 1939 eligibility age of 65 and non-missing 1935 state of residence and 1940 employment information. Columns (2)-(4) include only counties on state boundaries and exclude counties on borders of excluded states. Unit of observation in columns (2)-(4) is an individual-state border pair. All specifications include county fixed effects and 5-year age group fixed effects. All age-interactions are with 5-year age groups. Standard errors (in parentheses) are clustered at the state level. *: $p < 0.05$, **: $p < 0.01$, ***: $p < 0.001$

Table A12: Estimation results and robustness

	Baseline	β_e in 2nd stage	Elig($w/\bar{y}$)	$\eta = -2$	$\eta = -1/2$	PVBC	Halve LFP(age) slope	Full US	Outside estimates
Parameter estimates									
$\hat{\alpha}_e$	0.36	0.30	0.36	0.37	0.29	0.38	0.31	0.20	
$1000 \times \hat{\beta}_e$	-0.12	-2.9E-10	-41.50	-0.12	-0.12	-0.12	-0.12	-5.9E-10	
$\hat{\eta}$	-1.3	-1.3	-1.3	-2.0	-0.5	-1.3	-1.0	-0.6	
Key implications									
Percentage of men eligible for OAA[a]	22.1	30.4	16.3	23.0	15.2	23.3	17.4	19.9	
Equivalent variation of OAA, %[b]	94.7	93.4	95.4	94.4	96.9	94.6	93.3	94.3	
Effect of OAA earnings test, % of total	45.9	50.7	42.4	37.0	53.6	51.2	53.3	62.0	
Reduc. in LFP(65–74) from Soc. Sec., p.p.[c]	7.3	7.4	7.3	5.6	12.1	7.3	8.7	11.0	
Validation tests									
Reduc. in LFP(65–74) from OAA in 1940	6.3	6.5	5.7	5.5	6.8	6.6	5.7	6.1	8.5[d]
OAA recip. rate among men 65–74, %	19.0	24.0	15.0	18.3	14.9	20.0	15.0	18.0	16.5[e]
Objective function value	11.1	8.9	11.1	33.9	37.3	24.4	14.9	25.0	

Notes:
Parameter estimates and implications of the estimated model under various assumptions. "β_e in 2nd stage" is an estimation in which the slope of the eligibility-potential earnings relationship is estimated jointly with the other parameters in the second stage. In all but one of the other estimations, this parameter is estimated separately in a first stage based on the slope of the relationship between earnings and house value in Massachusetts. The exception is the "Full US" estimation, in which estimating β_e based on housing wealth would be inappropriate given that many states did not condition eligibility on housing wealth. "Elig($w/\bar{y}$)" is based on the assumption that eligibility for OAA is linear in the ratio of the wage over the OAA benefit, whereas the baseline assumption is that it is linear in the wage. As a result, $\hat{\beta}_e$ from this estimation is not directly comparable to those from the other estimations, since it is in different units. "PVBC" is based on a model with perfect capital markets, in which individuals can borrow as much as they wish as long as they satisfy a present value budget constraint. "Halve LFP(age) slope" halves the slope of the counterfactual no-OAA labor force participation-age profle while holding fixed the level of the profile at age 65. This increases late-life labor supply, which tends to increase the cost of the earnings test. In all cases, the objective function is a standard classical minimum distance objective function, so lower values indicate a better fit of the model. See Appendix A.8.1 for a summary of the results from additional specifications as well.
(a) Percentage of men "eligible" for OAA is the percentage of men who would receive OAA benefits if they had no earnings and were 65 and older.
(b) Present value of the welfare-equivalent unconditional late-life income stream (received each year from age 65 on regardless of earnings) as a percentage of the present value of actual OAA benefits received.
(c) The observed reduction in LFP(65–74) from 1940 to 1960 (against which reductions from Social Security can be compared) was 13.5 percentage points.
(d) Extrapolation based on authors' instrumental variables regression results.
(e) Authors' calculations based on data on the characteristics of new recipients of OAA, 1936–1940.