Online Appendix
for
Government Old-Age Support and Labor Supply: Evidence from the Old Age Assistance Program

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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 inxterpolate 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.[1] This method had been common in the administration of relief to the poor prior to the growth of OAA. By the late 1930s,

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

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 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.[2] 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

[2] 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.

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 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.[3] 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

[3] 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.

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 IV, 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 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

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

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).[4] 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.

[4]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).

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 to think that they varied systematically across state borders.[5] 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

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

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 (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.[6] 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 V.A. 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.[7] If leisure is non-inferior, people work no more when facing the OAA budget constraint than when facing the no-OAA budget constraint, $\phi(O) \leq \phi(N)$, 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 $\min\{w, \bar{y}\}\phi(N)$.

[6] 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.

[7] $\phi(O)$ 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.)

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 IV.A 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 V.B, 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.[8] Figures A16 and A18 plot the share of

[8]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.

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 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).[9]

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

[9]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.

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 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 V and Section VI. 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."[10] 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

[10]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.

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.[11] 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 OAA's means tests.

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

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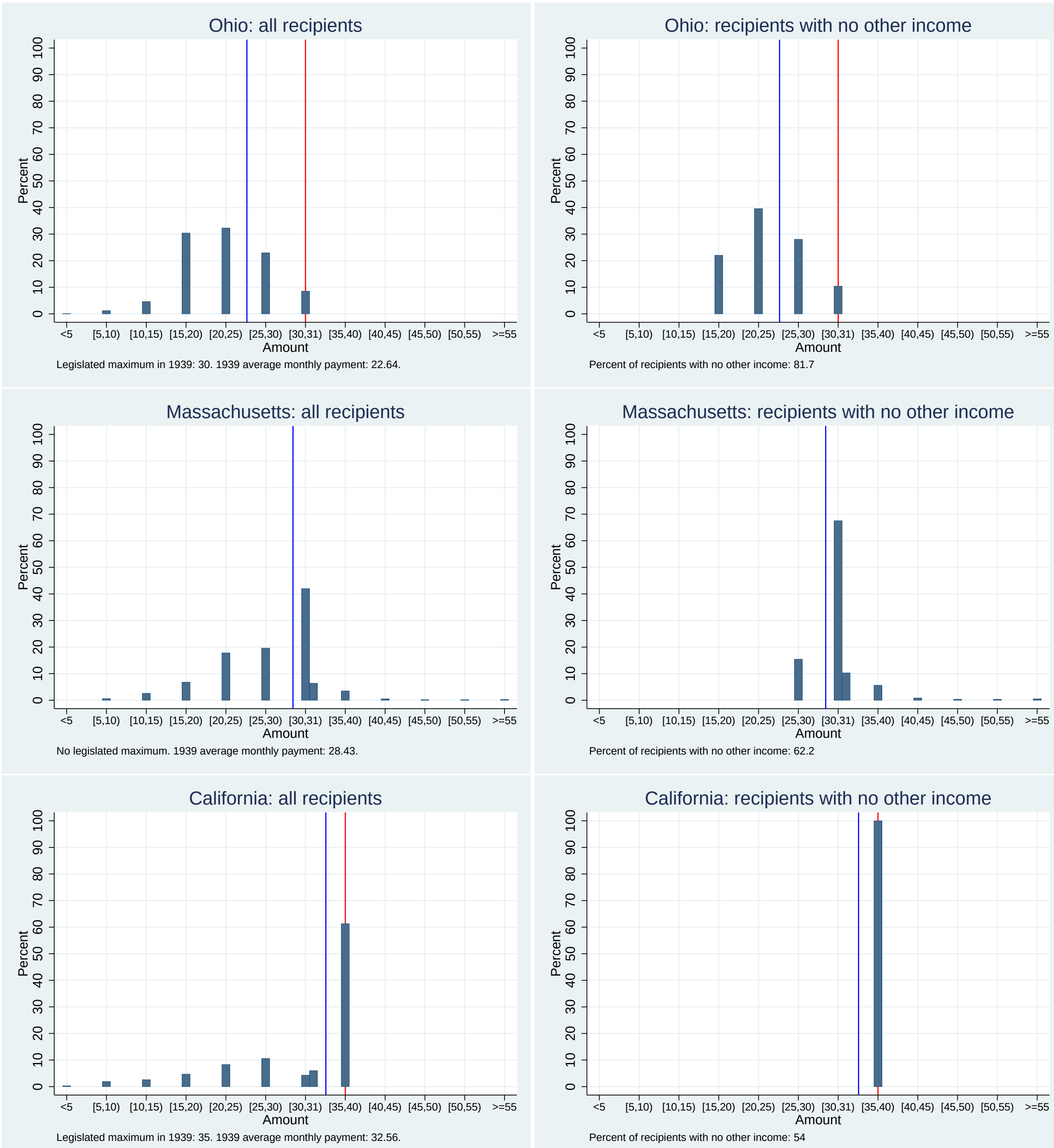
Appendix Figures and Tables

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

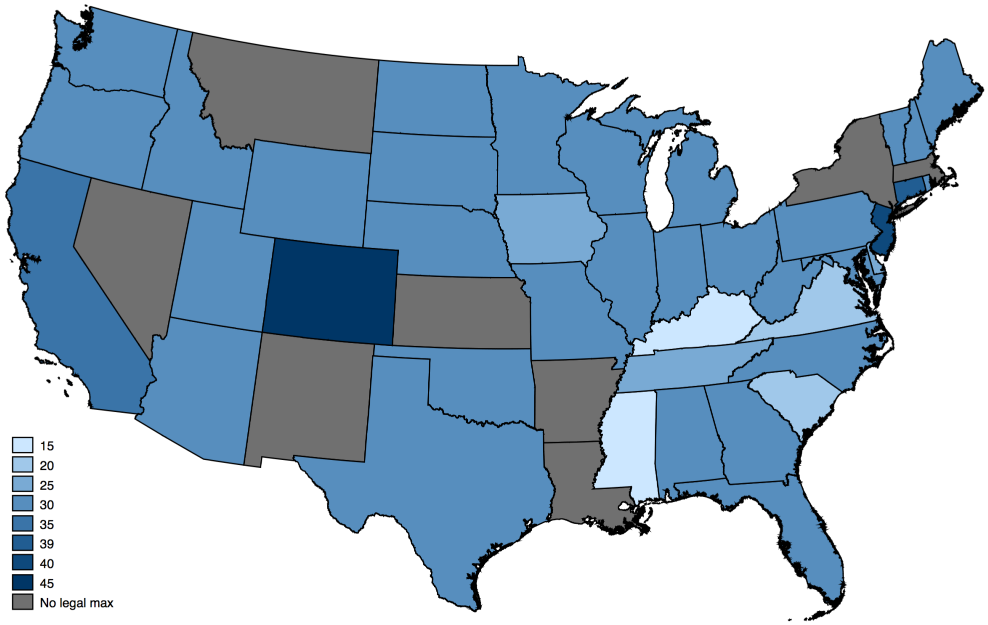


Figure shows statutory maximum monthly payment, from U.S. Social Security Board (1940*a*). Analysis sample excludes Colorado, Missouri, New Hampshire, and Pennsylvania.

Figure A3: OAA 99th percentile payments 1939

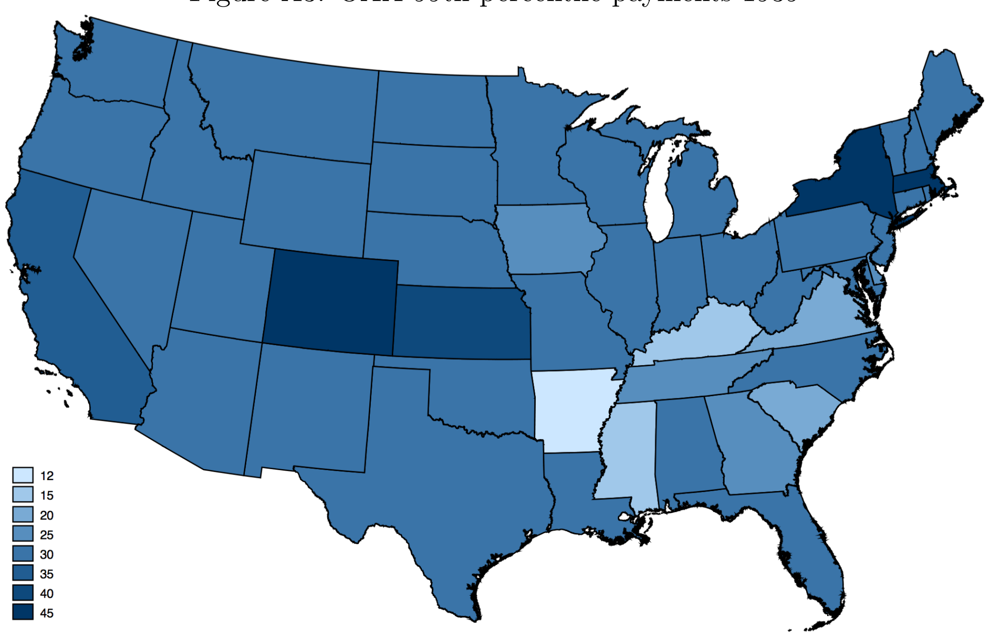


Figure shows estimate of 99th percentile payment, based on data from U.S. Social Security Board (1939*b*). For details on construction of this measure, see Appendix Section A.2.2. Analysis sample excludes Colorado, Missouri, New Hampshire, and Pennsylvania.

Figure A4: OAA variation and missing data

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

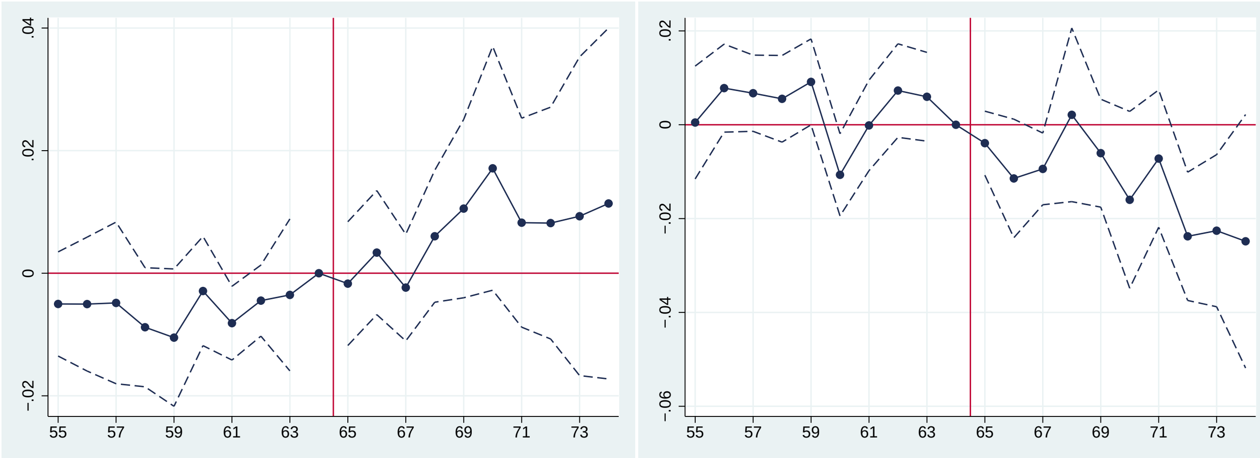


Figure shows point estimates and 95% confidence intervals on age-payment interactions from estimation of equation (1), using log simulated payment by age interactions as instruments for log observed payment by age interactions and controlling for state border by age fixed effects. Dependent variable 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

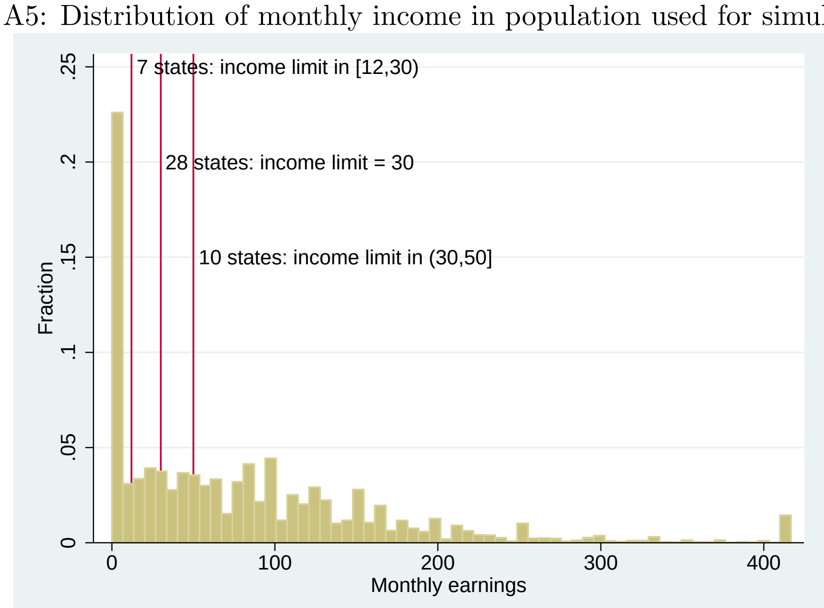


Figure shows distribution of monthly earnings in the population of men used for calculation of the simulated instrument (men aged 60-64 in states with an eligibility age of 65 in 1939). Monthly earnings is imputed for men reporting positive weeks worked in 1939 who were self-employed at the time of the 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

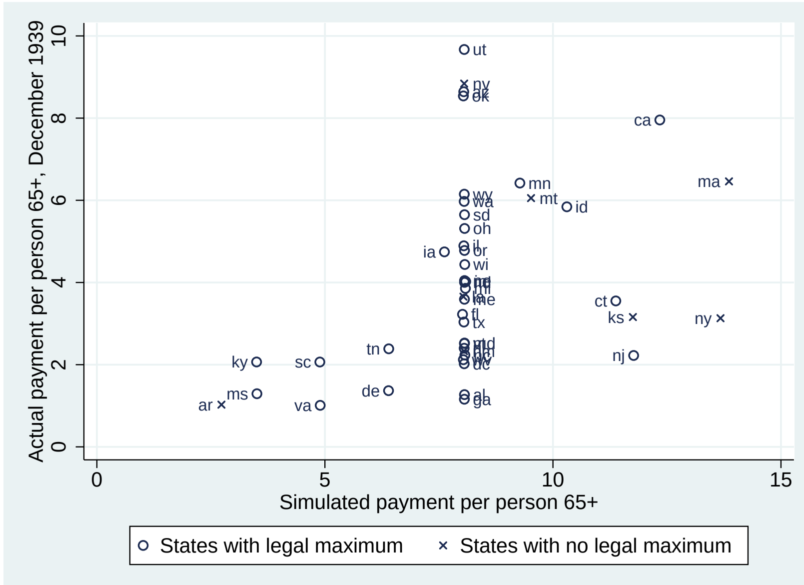


Figure shows relationship between observed state-level OAA payments per person 65 and older in December 1939 and simulated payments based on maximum payments and income disregards. "Maximum payment" is the statutory maximum monthly payment for those states with a statutory maximum and the 99th percentile 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

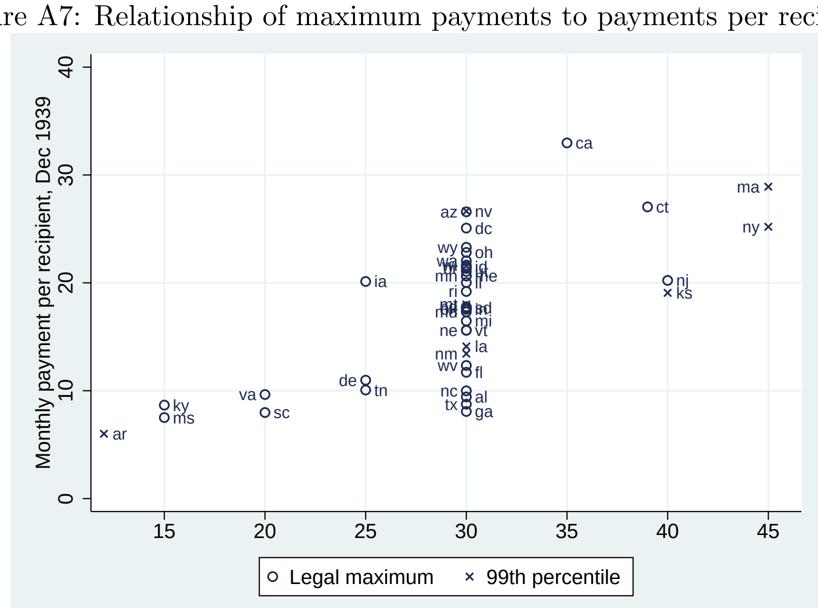
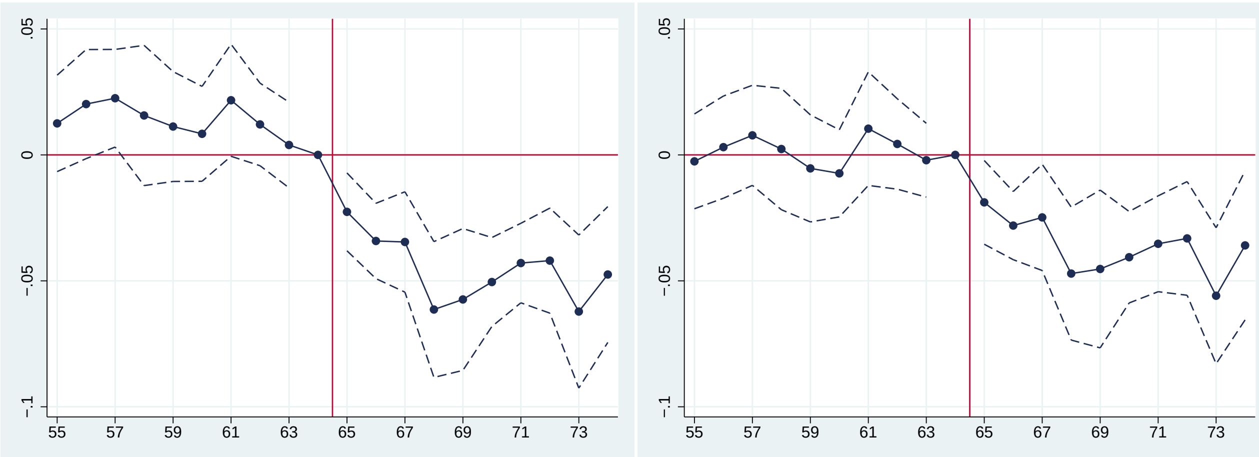


Figure shows relationship between average payment per recipient in December 1939 and statutory maximum monthly payment (for those states that had them) or 99th percentile payment (for states without a statutory maximum). Only states with eligibility age of 65 in 1939 are included. Sources: data on OAA 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

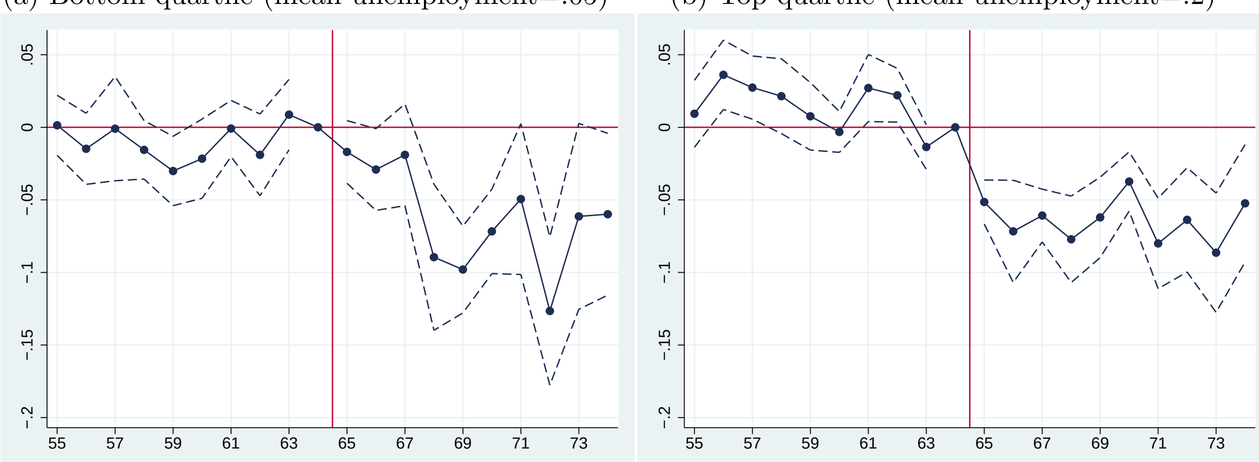
(a) employment (b) non-emergency employment



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

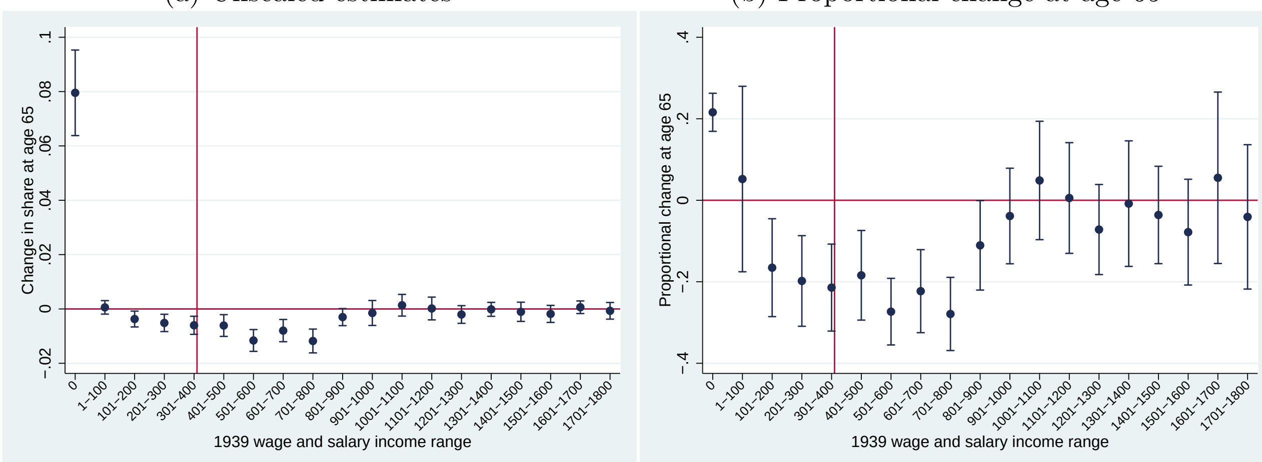
(a) Bottom quartile (mean unemployment=.05) (b) Top quartile (mean unemployment=.2)



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.

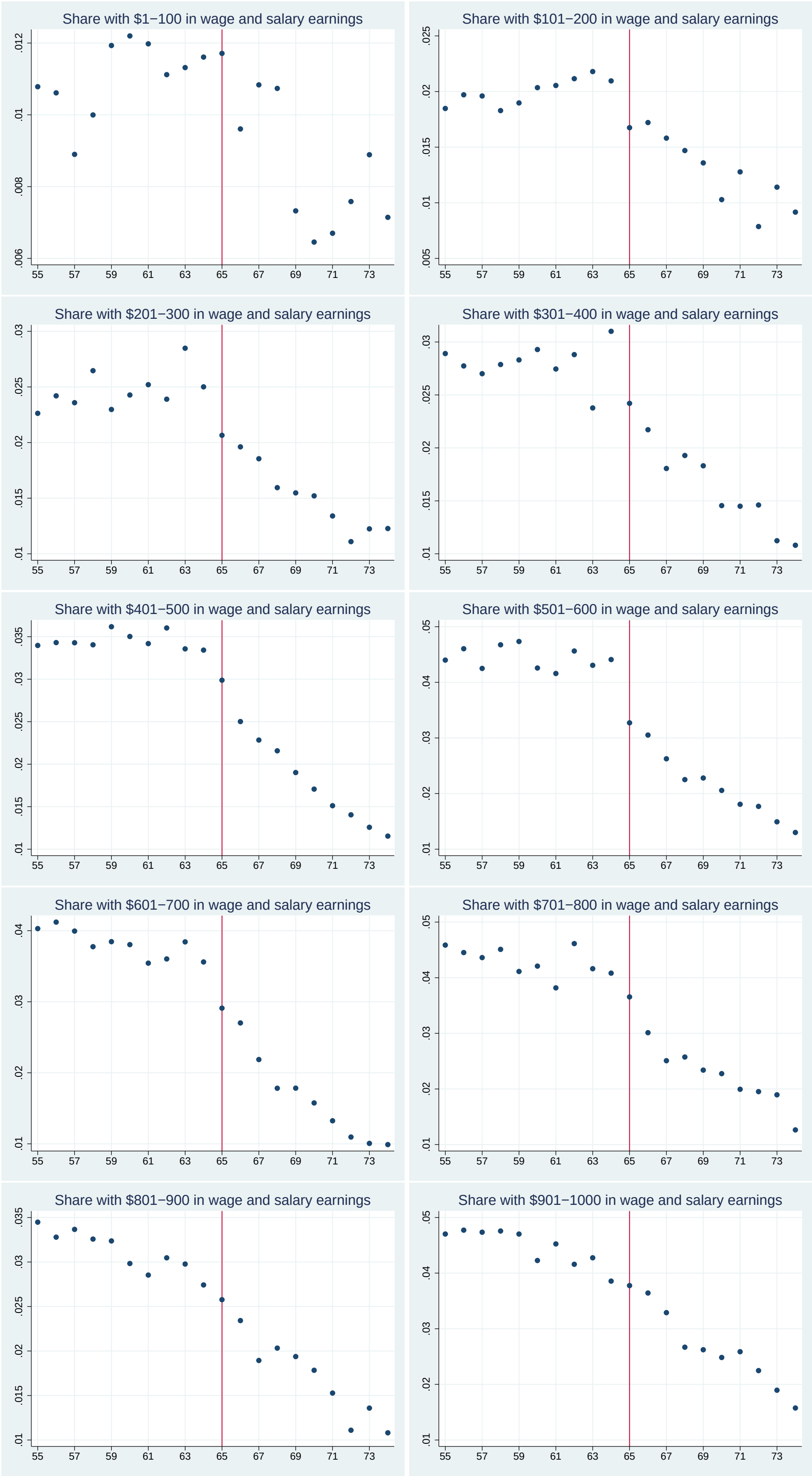
Figure A10: Change at 65 in share of men with specified amount of wage/salary income in 1939

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



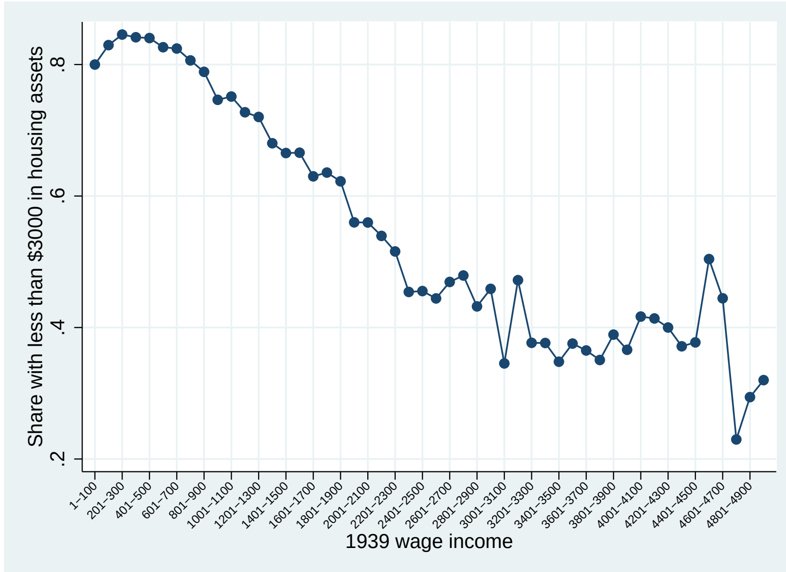
Notes: Figures show point estimates and 95% confidence intervals from separate estimations of equation (2), with dependent variable indicating wage/salary earnings of each specified amount in 1939. Sample: men within IK bandwidth around age 65 at 1940 Census in Massachusetts. Vertical line denotes "income 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



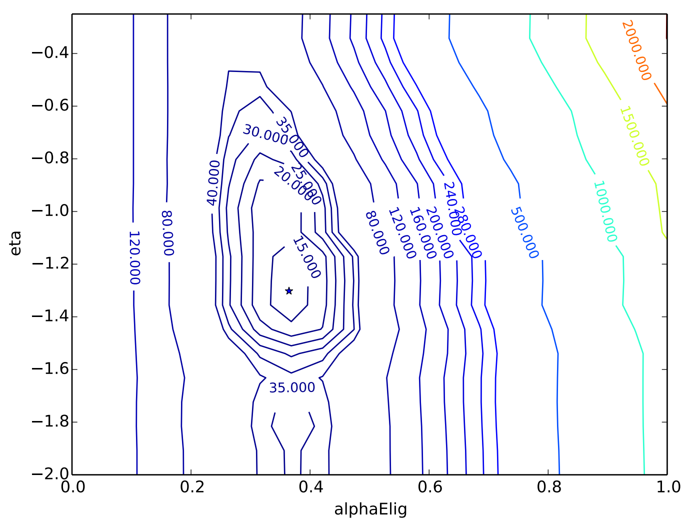
Notes: Figures show share of men reporting 1939 wage and salary earnings in specified range, by age at 1940 Census.

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



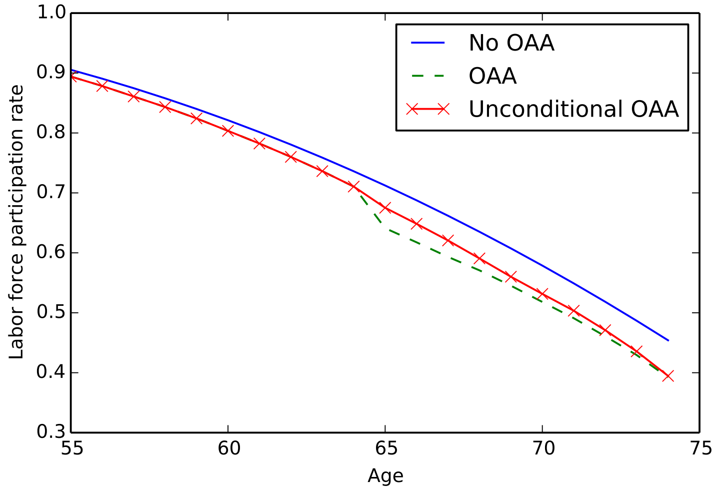
Notes: Share of Massachusetts men aged 60–64 who had less than \$3,000 of house value, as a function of wage and salary income. Massachusetts limited eligibility for OAA to people with less than \$3,000 in real property and did not have any home disregard. The figure therefore shows the share of people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

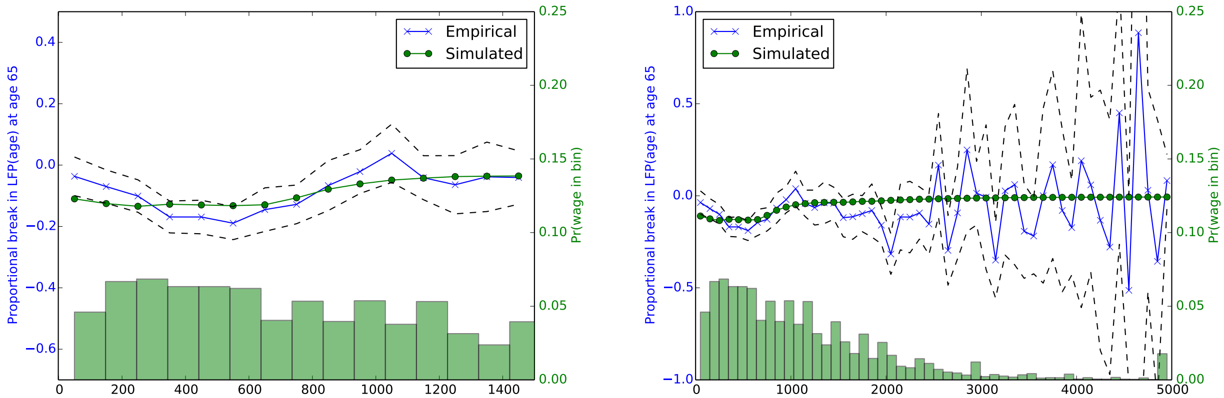
Figure A14: Simulated cross-sectional age-labor force participation profile in 1940



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.

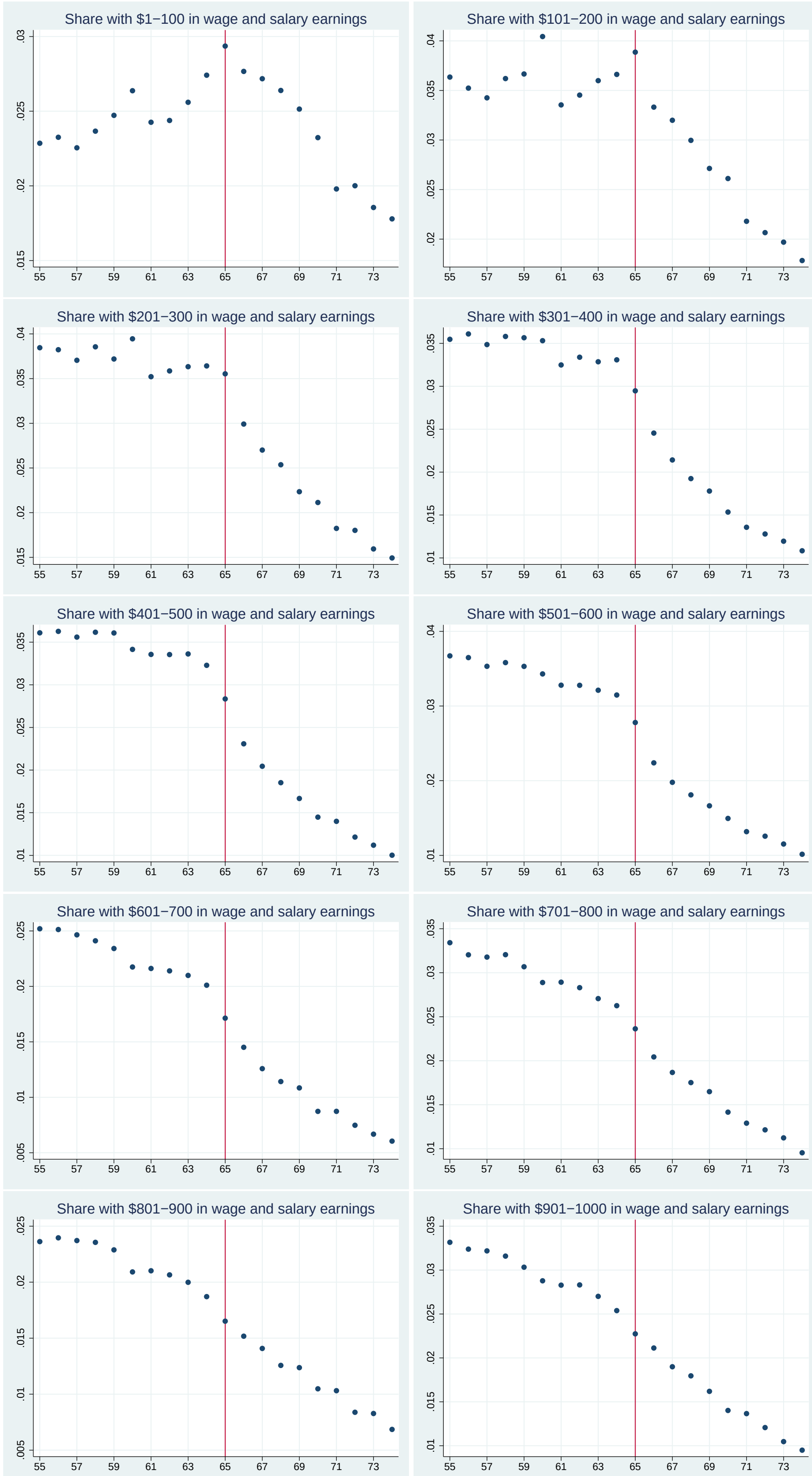
Figure A15: Empirical vs. simulated moments for full-US estimation

(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.

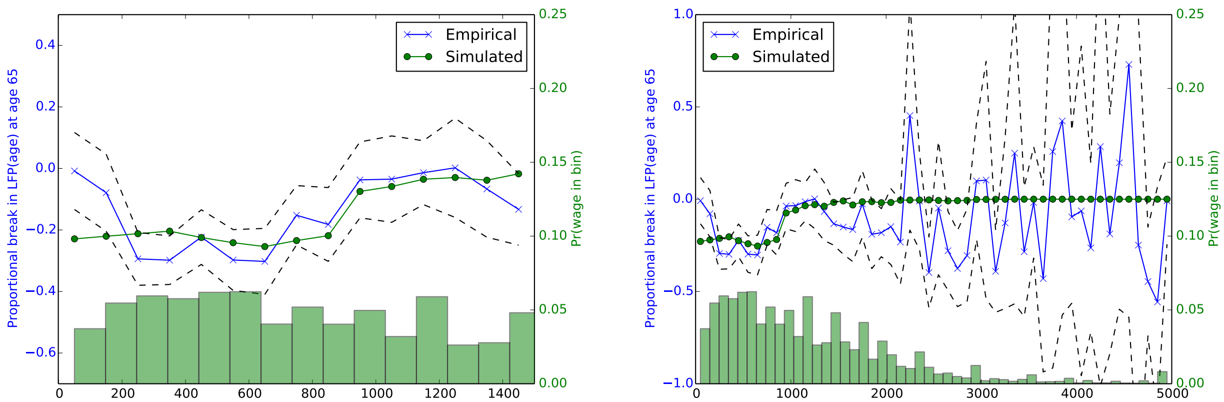
Figure A16: Share of all men with specified 1939 wage and salary earnings



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.

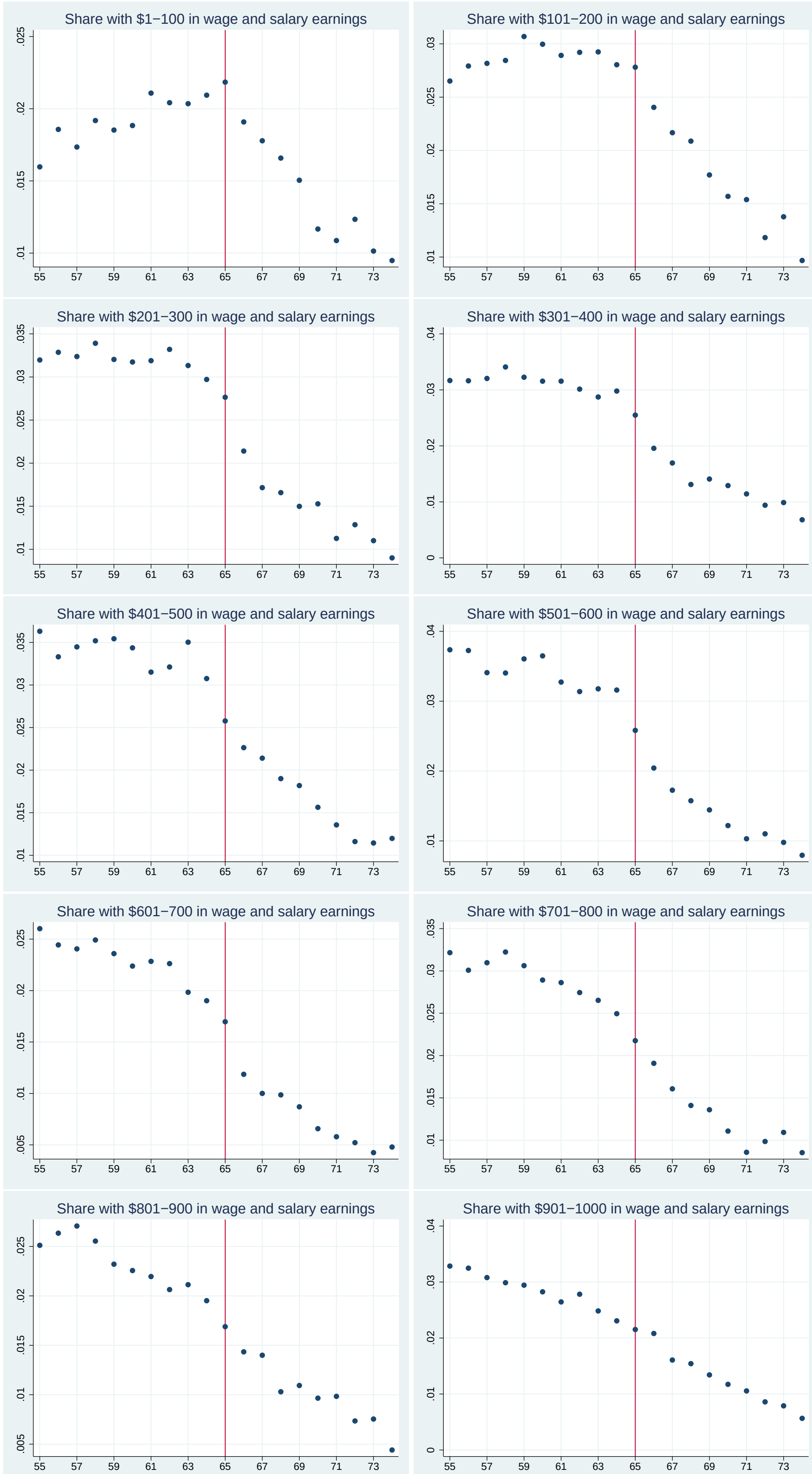
Figure A17: Empirical vs. simulated moments for California estimation

(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.

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.

Table A4: Simulated IV first stage regressions

	(1) age 55-59	(2) age 65-69	(3) age 70-74
Log simulated per-65+ payment $\times$ age 55-59	0.897	0.000	-0.000
	(0.113)	(0.001)	(0.000)
Log simulated per-65+ payment $\times$ age 65-69	0.002	0.892	0.001
	(0.003)	(0.114)	(0.002)
Log simulated per-65+ payment $\times$ 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 $\times$ age fixed effects	yes	yes	yes
Education $\times$ age fixed effects	yes	yes	yes
Race $\times$ age fixed effects	yes	yes	yes

Dependent variables: 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.

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.

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 (0.002)	0.001 (0.002)	0.002 (0.002)	-0.001 (0.001)
Per-65+ payment × age 65-69	0.019 (0.003)	-0.020 (0.004)	-0.014 (0.004)	-0.009 (0.003)
Per-65+ payment × age 70-74	0.024 (0.004)	-0.023 (0.004)	-0.017 (0.004)	-0.012 (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.

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.

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.

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.

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.

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.

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.