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Matthias Giesecke and Philipp Jäger[1]

Pension Incentives and Labor Supply: Evidence from the Introduction of Universal Old-Age Assistance in the UK

Abstract

We study the labor supply effects and welfare implications of introducing a universal means-tested old-age assistance program in times of very limited social protection. We take advantage of a unique historical reform: The Old-Age Pension Act (OPA) of 1908, which, for the first time, provided pensions to older people in the UK. Using recently released full-count census data covering the entire population, we exploit variation at the newly created age-based eligibility threshold. Our results show a considerable and abrupt decline in labor force participation of 6.0 percentage points (13%) when older workers reach the eligibility age of 70. This sudden drop only occurs at the age cutoff and only after the OPA was implemented. Despite the considerable labor supply decline, the overall efficiency loss from the OPA was limited and most likely outweighed by equity gains.

JEL-Code: D61, H21, H55, J14, J22, J26

Keywords: Old-age assistance; labor supply; retirement; regression discontinuity design; equity-efficiency trade-off

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1 Matthias Giesecke, RWI, IZA and Federal Ministry of Finance; Philipp Jäger, RWI. *– We thank Ronald Bachmann, Stefan Bauernschuster, Sebastian Braun, David Card, Fabian Dehos, Erik Hornung, Anica Kramer, Christian Proano, Laura Salisbury and Han Ye for helpful comments and suggestions. We also thank the participants of several conferences and seminars for insightful discussions. Johanna Muffert provided excellent research assistance.* *– All correspondence to: Philipp Jäger, RWI, Hohenzollernstr. 1-3, 45128 Essen, Germany, e-mail: philipp.jaeger@rwi-essen.de*

1 Introduction

Many government old-age assistance programs intend to prevent poverty through means-tested income transfers to the elderly. To assess the welfare implications of these programs, the equity gains from redistribution have to be traded off against efficiency losses that arise from behavioral changes such as reductions in late-life labor supply. So far, the equity-efficiency trade-off has received little attention in the analysis of old-age assistance. A deeper understanding of this trade-off, however, is becoming increasingly important as aging populations challenge social security systems around the world.

This article assesses the labor supply and welfare implications of the Old-Age Pension Act (OPA) of 1908 that established universal means-tested pensions in the UK at the expense of higher taxes for top income earners. In contrast to most of the existing literature, the unique historical setting allows us to estimate the labor supply response when moving from essentially no program to an old-age assistance scheme that covered a large share of the elderly population.[1] Using recently released full-count population data from the UK census in 1911, we isolate the causal effect of the program along the lines of an age-based eligibility threshold that has been introduced by the OPA at the age of 70.[2] This threshold induces a discontinuity in the retirement probability that we use to identify changes in the labor force participation (LFP) rate by adopting a regression discontinuity (RD) design. We finally translate our reduced form estimate into a welfare statement, comparing efficiency losses of the policy due to early retirement with equity gains from redistribution during times of high income inequality[3] and little social protection.

The historical setting, besides being interesting in its own right, also involves several other features that strengthen the empirical analysis. Enacted in 1908 and effectively

[1] In 1911, 57.5% of people who reached the eligibility age of 70 did receive a pension that was administered through the OPA.

[2] Despite the low expectancy at birth, the UK had a sizable elderly population at that time (more than one million people were 70 and older in England and Wales alone). Adjusting for child mortality shows that about 35% of the birth cohort 1841 reached age 70 in 1911 (life tables Human Mortality Database, 2018). Conditional on having celebrated the 70th birthday, this cohort had a remaining lifetime of 9 years on average.

[3] In 1911, 8.4% of total income (before taxes) in the UK went to the top 0.05% compared to 3.4% in 2000 (Atkinson, 2007).

implemented in 1909, the OPA was the largest means-tested old-age assistance program at that time[4] and served as an archetype for public old-age assistance in the U.S. In contrast to the U.S. program that was introduced in the 1930s during the turbulent times of the Great Depression, the OPA was established in a period of stable economic conditions which ensures that our results are not confounded by macroeconomic trends. Although the UK was a large industrialized economy when the OPA was introduced, it still lacked a comprehensive welfare system. Private pension schemes and other government programs that addressed older workers were either small and uncommon or did not exist. Furthermore, program substitution towards unemployment or health insurance was impossible because these programs became effective in 1912 and thus after the implementation of the OPA (1909) and data collection (1911).

Previous studies have suggested that the introduction of the OPA had little impact on labor supply (Johnson, 1994; Costa, 1998).[5] This conjecture is based on aggregate labor force participation (LFP) rates from British censuses, suggesting that the downward trend in LFP rates among older men in the UK did not accelerate after the introduction of the OPA. Using more disaggregated data, we provide striking evidence that LFP rates of older men and women did decline promptly and considerably as a direct consequence of the introduction of the OPA.

Our main findings indicate that, in absolute terms, the LFP decline amounts to 6.0 percentage points at the age of 70 just after the implementation of the OPA. Relative to a participation rate of 46% at age 69, the LFP rate thus declines by 13% when moving marginally above the age-based eligibility threshold.[6] The LFP drop is predominantly driven by reductions in work activity and much less by unemployment. Our estimates are

[4]Germany already introduced a public pension scheme in 1891, but pensions were based on contributions and not means-tested. Belgium, France, Denmark, New Zealand and parts of Australia had means-tested old-age assistance programs in place before 1909. However, these programs were smaller with respect to the absolute number of recipients or the benefit level.

[5]In contrast, the expansion of old-age assistance in the U.S. has been shown to coincide with a strong decline in labor force participation of older men (see e.g. Feldstein and Liebman, 2002; Krueger and Meyer, 2002; Coile, 2015; Fetter and Lockwood, 2018).

[6]In absolute terms, the effects are larger for men (7.9 percentage points) compared to women (3.2 percentage points). In relative terms, however, the reactions are smaller for men (10%) as compared to women (16%) when taking into account that LFP rates marginally below the eligibility age (at 69) are large among men (76%) and rather low among women (20%).

consistent with a static lifetime labor supply model that predicts bunching of retirements at the kink in the lifetime budget constraint that occurs at the eligibility age.

Labor supply responses are considerably larger at the bottom of the earnings distribution than at the top. We document that the LFP reduction of 15.4% in the lowest earnings quartile more than doubles the LFP decline of 7.2% in the highest quartile. This finding is consistent with the theoretical prediction that people with low earnings should respond more strongly to means-tested cash transfers because they have to forgo less earnings to pass the means test than people with high earnings. Our results also indicate that the LFP decline is stronger for older workers with close family ties in comparison to individuals without. This finding is in line with the hypothesis that private transfers from children, spouses, or other household members served as old-age insurance when public pensions were absent. We further provide evidence for joint retirement decisions of couples. We document a significant LFP drop of 3 percentage points among eligible individuals (age 70 or older) when their spouse also reaches the eligibility age. This result strongly suggests spousal correlations in the taste for leisure and corresponding labor supply decisions, in line with previous findings on joint retirement decisions of couples (Blau and Riphahn, 1999; Gustman and Steinmeier, 2000; Baker, 2002; Atalay et al., 2019).

Old-age assistance programs generate efficiency losses if people retire early to pass the means test. In this case, recipients value £1 of old-age assistance by less than £1. We approximate the willingness to pay for the policy by combining our causal estimate, the recipience rate of the OPA, and pre-existing parameters from a life-cycle model estimated by Fetter and Lockwood (2018). We find that, despite the considerable labor supply decline, OPA benefits were valued highly, by at least £0.93 of unconditional old-age income for each pound of OPA benefits. Comparing the willingness to pay to the actual costs of the program, we conclude that the program created only modest inefficiencies. When trading off the moderate efficiency loss against welfare gains through redistribution from the top to the bottom of the earnings distribution, we conclude that the overall impact of the OPA was arguably welfare enhancing.

This paper makes three major contributions to the literature. First, it adds to earlier studies on labor supply responses to marginal changes in existing pension systems (e.g. Krueger and Pischke, 1992; Börsch-Supan, 2000; Mastrobuoni, 2009; Liebman et al., 2009; Brown, 2013; Atalay and Barrett, 2015; Manoli and Weber, 2016). We extend this strand of the literature by studying a first-time introduction of an old-age assistance system, thus investigating the change from no program to universal coverage. Our estimates quantify the immediate and full labor supply effects of the OPA, starting from a benefit level of zero without having to rely on extrapolations. Such estimates are informative for the introduction of old-age assistance systems in today's developing countries and also serve as a benchmark for marginal changes in contemporaneous pension programs.

Second, the institutional setting allows for highly credible and transparent identification. The reform introduced an age-based eligibility threshold at the age of 70, where every UK citizen became eligible conditional on having passed the means test. Using unique census data with a large number of observations permits us to precisely quantify the labor supply effects of the program in the local environment of the age cutoff, based on an RD design. This research design builds on many recent studies that have used age-based eligibility thresholds to identify policy-relevant effects (see e.g. Card et al., 2008; Battistin et al., 2009; Carpenter and Dobkin, 2009; Card et al., 2009; Anderson et al., 2012, 2014; Carpenter and Dobkin, 2015, 2017; Fitzpatrick and Moore, 2018). Our RD design adds to previous research that used cross-state variation in benefit levels to identify the labor supply effects of introducing old-age assistance in the U.S. (Fetter and Lockwood, 2018). In contrast to this approach, we can rule out that policy endogeneity drives our results. This would be the case if unobserved factors that correlate to policy preferences and old-age labor supply were systematically tied to the identifying variation. However, our research design is solely based on the age-based eligibility threshold which makes identification highly transparent and visible from graphical evidence.

Third, the transparent redistributive design of the OPA allows us to make a welfare statement based on our causal estimates. Our welfare analysis is based on a simple three-stage framework: first, we translate our reduced form estimate into the willingness to pay for the program. Second, we approximate the efficiency loss by relating the willingness to pay to the costs of the program. Third, we make a welfare statement by comparing the efficiency loss to the equity gain from redistribution. To the best of our knowledge, we are the first to study the equity-efficiency trade-off for the introduction of an old-age assistance program. In the literature on incremental changes within contemporaneous pension systems, welfare analyses are rarely built on credible reduced form estimates.[7] Therefore, this paper contributes to a better understanding of the overall welfare impact of old-age assistance programs.

The remainder of this paper is structured as follows. Section 2 provides historical and institutional details on the old-age assistance program in the UK and how its introduction creates exogenous variation that we use for identification. Section 3 outlines theoretical predictions on labor supply responses that are generated from a static model of lifetime labor supply. Section 4 outlines the research design and describes the unique data source from the UK census. Section 5 presents results, sensitivity checks and falsification tests. Section 6 quantifies the welfare impact of the program. Section 7 concludes.

2 Historical Background and Institutional Details

The Old-Age Pension Act of 1908 (OPA) introduced means-tested, non-contributory minimum pensions for British citizens financed by the central government.[8] The OPA was a major social policy intervention and the first one to specifically target the elderly in a time of very limited social protection. The law was debated in the British Parliament in May 1908, passed through in August 1908 and the first pensions were eventually paid out in January 1909. At that time, neither unemployment nor health insurance existed because both of these programs only became effective in 1912.

Given that pensions were means-tested, the coverage of the OPA was astonishingly high. In 1911, almost 60% of people who had reached the eligibility age were granted a

[7]A recent paper by Andersen et al. (2020) is a notable exception. The authors study the efficiency-equity trade-off in the case of a Norwegian pension reform that removed an early retirement subsidy.

[8]See Casson (1908) for details.

pension in England and Wales (613,873 out of 1,068,486 according to the Department of Labour Statistics, 1915, p. 184). The vast majority of pension recipients (about 93% in 1911) also received the maximum pension of 5 *shillings* per week. According to Feinstein (1990), this amounted to approximately 22% of average wages.

The OPA was a response to the perceived inadequacy of the the existing poor relief system that provided only very basic protection and involved considerable sanctions such as the loss of voting rights and the requirement of working in a workhouse unless the person could prove to be sufficiently unfit. The newly introduced pensions were not only less restrictive but also involved more generous benefits[9], and thus considerably more older people applied for them (Thane, 2000). In contrast to the poor law, which was administered and financed at the local level giving local authorities a lot of discretion in the assignment of financial aid, the OPA was enacted as a nation-wide right for older workers who met the specified eligibility criteria for receiving a pension.

Pension eligibility was mainly based on two criteria: age and inadequate means.[10] First of all, older workers only became eligible when reaching the age of 70. The original proposal for the reform, dating back to 1899, recommended a retirement age of 65, which would have been more in line with the retirement rules in the few pre-existing pension schemes that typically specified an age between 60 and 65.[11] However, the original suggestion was considered too expensive. Given the low life expectancy at birth (below 50 for males in 1911), a retirement age of 70 seems high by today's standards. The low life expectancy, however, was mainly driven by high infant mortality. Once reaching the age

[9]At the time of the introduction of the OPA, poor relief was about 3 *shillings* per week and thus amounted to 60% of the new pensions that were legislated by the OPA (Casson, 1908).

[10]Moreover, the law included a residency requirement of 20 years. Even if the claimant had satisfied all of the the eligibility criteria, he could still be disqualified due to the following reasons. First, receiving poor relief or having received poor relief any time between January 1908 and December 1910. Second, habitually failing to work according to her ability. And third, being detained in a lunatic asylum, or in any place as a recipient of poor relief or a criminal lunatic or being in jail (or ordered to be imprisoned) less than ten years ago.

[11]Small-sized pension schemes within the boundaries of the respective firm existed previous to the large-scale introduction of the OPA. However, these pensions were rather informal and discretionary. More formalized schemes only existed in very few larger firms, for example in the railway industry, or the public sector. By far, these pensions did not reach the universal coverage rate that was introduced through the OPA (see Thane, 2000, for details on pre-existing pensions in the UK).

of 70 in 1911 (birth cohort 1841), people could expect to live another 9 years on average (men: 8.5 yrs., women: 9.5 yrs. Human Mortality Database, 2018).

Second, eligibility was conditional on a means test. Claimants had to prove to the pension authority that their annual means were below 630 *shillings* (54% of average annual wages at that time) to receive any pension. To become eligible for the maximum pension, an income of less than 420 *shillings* (36% of average annual wages) was required. Annual income was calculated based on the actual income received in the previous year (including family transfers) and augmented, if applicable, by the rental value of living in one's own house as well as a hypothetical return on property that so far had not been used commercially even though it could have been. Spousal income was considered explicitly. [12] The law also prescribed that if individuals intentionally deprived themselves of resources, the value of these resources would still be included in the calculation of the annual means.

Pensions were granted based on a gliding scale as depicted in figure 1 and table 1. However, the overwhelming majority received the maximum pension of 260 *shillings* per year. For simplicity, the pensions did not account for differences in the costs of living across regions, even though this has been considered during the legislation process.

Applications for pensions were typically made at the local post office. They were checked by a pension officer appointed by the treasury and a local pension committee. Most applications were approved with a rejection rate of only 10% (The Times, 1909; Old Age Pensions Committee, 1919). The main reasons for disapproval were the inability to prove the individual age, a failed means test or receiving poor relief. Proving age was easy for pension claimants in England and Wales because birth registers (and thus birth certificates) existed at least since 1837. Verification was more difficult in Scotland and especially challenging in Ireland, making rejections based on the age criterion far more frequent in Ireland (Old Age Pensions Committee, 1919).

[12] Spouses living together in one household had to fulfill two criteria: 1) Their own income and 2) the per person income of the couple had to be below the threshold. An example provided in Casson (1908) clarifies this rule: imagine a married couple. The man earns 800 *shillings* and the woman 200 *shillings*. In this case, the woman would be eligible for a pension and has to claim an income of 500 *shillings* ((800+200)/2). The man, however, would not be eligible as his own income of 800 *shillings* exceeds the threshold.

Despite the relatively high retirement age, the costs of the newly introduced pension payments were considerable. In the budget year 1911, £6.3 million were spent on old-age pensions in England and Wales. For the entire UK, old-age pension expenditures amounted to £9.8 million, making pensions one of the largest single spending item (5.7% of overall expenditures in the budget of 1911,(House of Commons, 1911)). Pensions were financed almost exclusively by taxes on high income earners (roughly the top 1% in the UK).[13] For this purpose, new progressive tax brackets were introduced (House of Commons, 1911): The marginal tax rate for people with annual earnings of £2,000 - £3,000 increased by around 1.2 percentage points (to 5.0%), for £3,000 - £5,000 by around 2 percentage points (to 5.8%) and above £5,000[14] by around 4.5 percentage points (to 8.3%). The marginal tax rate for the majority of tax payers earning between £160 and £2000 was not increased and remained at 3.8%.[15] Most people did not pay any income tax because their earnings were below the taxation threshold of £160 (2.7 times the mean earnings) and were also not affected by income tax changes.[16]

3 Theoretical Predictions

We use a static labor supply model to generate predictions on how people respond to the introduction of means-tested old-age assistance. The model captures the financial incentive from the OPA by relating the present value of lifetime consumption to the retirement age when old-age assistance is available. Figure 2 depicts the lifetime budget constraint both with and without old-age assistance, assuming certain lifetimes (T) and that the individual earns w for each year of work. Without the OPA, individuals who never work have lifetime consumption possibilities of α (e.g. from family transfers). Individuals who work over the entire lifespan have lifetime consumption of $\alpha + wT$. With OPA available, the set of lifetime consumption opportunities expands, introducing a parallel

[13] Other less important components of the financing system were death duties as a type of inheritance tax and stamp duties on investments. For details on the financing sources of the OPA, see Murray (2009).

[14] Earnings above £5,000 constituted the top 0.05% of earnings, according to Atkinson (2005).

[15] This group included about 75% of tax payers, e.g. 750,000 out of 1,000,000 (Murray, 2009).

[16] At that time, only one million people paid income taxes in the UK (Murray, 2009) compared to more than 36 million people living in England and Wales alone.

upward shift of the lifetime budget constraint. This transfer component induces non-distorting income effects[17], such that individuals optimize over leisure and consumption, conditional on the availability of the pension.

Starting at the eligibility age R_{ELIG}, the availability of OPA induces a convex kink in the budget set due to the implicit tax through the means test. Since benefits b are conditional on insufficient means (recall figure 1), labor earnings are taxed by an implicit rate of $\tau = min\{1, b/w\}$. Assuming continuously distributed preferences of consumption and leisure in the population and that leisure is a normal good, this feature of the budget constraint implies that some individuals above the kink will reduce their retirement age. For these individuals, consuming marginally more leisure is compensated by the income from pension benefits. Due to the convexity of the kink, the model predicts bunching of retirements just at the eligibility age (R_{ELIG}) for those who would retire at ages slightly above R_{ELIG} if pensions were not available. In the region above R_{ELIG}, the price of leisure declines relative to the price of other consumption goods. This relative price change, induced by the minimum pension, introduces distorting substitution of labor into leisure. In our empirical analysis we will show an immediate drop in labor supply at the age-based eligibility threshold ($R_{ELIG} = 70$). This is exactly what the static labor supply model predicts and reflects the distorting substitution effect from the implicit tax through the means test.

To further illustrate how labor earnings are taxed implicitly, figure 3 relates annual earnings to the sum of annual earnings and old-age assistance. The figure shows how the pension schedule creates jumps in the total individual income from working and through government transfers, inducing disincentives to increase labor supply at those points where a marginal increase in earnings will lead to a reduction in total income. For example, a person who earns 420 *shillings* per year receives annual benefits of 260 *shillings* while a person who earns 421 *shillings* per year only receives 208 *shillings*. On aggregate, this extends to the entire lower region of the earnings distribution, where a person who earns

[17] Pure income effects from doubling the legal minimum pension in the Ukraine have been studied by Danzer (2013). In this case, pension eligibility rules are not subject to a means test such that the benefit design does not distort labor supply decisions.

680 *shillings* per year obtains just the same amount of money as a person who earns only 420 *shillings* but qualifies for the full amount of old-age assistance of 260 *shillings*.

4 Empirical Strategy and Data

4.1 Exogenous Variation and Research Design

The identifying variation that we use to estimate the causal effect of pension availability on LFP is based on the age cutoff that was introduced by the OPA. Pension eligibility at age 70 creates a discontinuity in the local environment between age 69 and 70. Along the lines of this age threshold, we adopt an RD with the age as assignment variable. The identifying assumption is that the outcome of interest, LFP, would evolve smoothly between age 69 and 70 if the OPA had not been introduced. Any discontinuous jump of the outcome at the eligibility cutoff can be attributed to the availability of the pension if other programs did not affect LFP at the respective age.

Reaching eligibility does not necessarily mean that people retire instantaneously and claim pensions. At the eligibility threshold, however, the probability of retiring exhibits a discontinuous jump due to the fact that a substantial share of older workers become eligible for the OPA while claiming the pension was not possible below the age cutoff.[18] Since pension eligibility also depends on other criteria such as the means test, there is imperfect compliance and hence the retirement probability does not jump from zero to one. This setting can be referred to as a *fuzzy* RD[19], where treatment is not fully determined by the age cutoff.

4.2 Estimation

The observable outcome y_a is an indicator of LFP that takes the value one if the individual is in the labor force at age a and zero otherwise. We thus estimate the equation

[18] For a similar setting where an age-based eligibility threshold of retirement is used to study consumption outcomes, see Battistin et al. (2009).

[19] See Lee and Lemieux (2010) for an overview on RDs.

$$y_a = \beta_0 + \beta_1 \mathbb{1}(a \geq 70) + \beta_2 f(a) + \varepsilon_a \tag{1}$$

where the coefficient of primary interest, β_1, measures the percentage point difference in LFP, comparing the share of people in the labor force marginally above the age cutoff (age 70) to the respective share marginally below the age cutoff (age 69). To account for the possibility of a functional relationship between the outcome LFP and the assignment variable age, the function $f(a)$, which is allowed to vary on either side of the age cutoff, not only includes age linearly but also as a second order polynomial. However, graphical evidence suggests that the age-LFP relationship is essentially linear close to the age-cutoff.

To show that our results are exceptionally robust against changes in the specification, we implement several alternative estimation procedures that are common in the RD literature. We extend the baseline estimation framework with uniform weighting to more flexible local non-parametric estimates that put more weight on observations close to the cutoff (triangular kernel weighting). We also present bias-corrected point estimates with robust standard errors as suggested by Calonico et al. (2014a,b) and provide detailed results on how the estimates differ by bandwidth choice and the order of the polynomial (see section 5.6).

4.3 Data and Summary Statistics

The analysis relies on full-count individual level census data for three decennial UK census waves collected in the spring of 1891, 1901 and 1911. The dataset is a recent release by the Integrated Census Microdata (I-CeM) project (Higgs et al., 2013), distributed by Integrated Public Use Microdata Series International (IPUMS International Minnesota Population Center, 2018).[20] We use information for England and Wales, thus excluding

[20] The I-CeM project collaborated with the website findmypast.org to transcribe and harmonize several historical British censuses, encompassing data collected in the years 1851, 1861, 1881, 1891, 1901, and 1911. Recent economic studies that have used selected waves are, for example, Arthi et al. (2019) and Beach and Hanlon (2019).

Scotland, Ireland and the Channel Islands because data is not available for the other regions at all points in time. Moreover, birth certificates, which substantially reduce age-misreporting, only existed in England and Wales for a sufficiently long period. Finally, we exclude persons with unknown gender (less than 0.1 % of the population) or age (0.2 % of the population).

4.3.1 Dependent Variable

Our definition of labor force status is based on the gainful employment concept which was used before the UK adopted the current labor force definition in 1961. In contrast to the current definition, which categorizes people based on their activity status (working or seeking work) in a specific reference week, the gainful employment concept derives the labor market status from the occupation of the respondents. In particular, we include people in the labor force (LFP = 1) if they specify an occupation or report to be unemployed.[21] Individuals are considered out of the labor force (LFP = 0) if they report no occupation or that they have retired from a specific occupation.[22] Both the current definition of LFP and the gainful employment concept are closely related. Costa (1998) constructs participation rates based on the gainful employment concept for the U.S. until the 1990s, showing that the patterns of both series match. Similarly, Johnson (1994) argues that the change of the definition in 1961 did "appear to have had little effect on the enumeration of older workers" (Johnson, 1994, p. 109). Based on this evidence, the two concepts arguably yield very similar patterns of LFP over time. For our empirical analysis, potential differences between the two concepts are of little relevance because the gainful employment concept did not change throughout the time under study (1891 - 1911). Differences only need to be kept in mind when comparing the results to the current LFP concept.

[21]We also include people that report to be formerly employed in the labor force. Recoding this subgroup as not in the labor force does not affect our results.

[22]Following the census in 1891, retirement was explicitly recognized as a separate category and retirees were not considered economically active anymore (Johnson, 1994), which is arguably consistent with being out of the labor force. We adjust the labor force variable constructed by IPUMS International by defining individuals to be out of the labor force if they state an occupation but add that they have retired already.

4.3.2 Summary Statistics

Using full-count census data enables us to zoom in directly at the age cutoff. Figure 4 shows the distribution of observations over age for the 1911 census, including 150,293 individuals at age 69 and 140,288 individuals at age 70. The drop in sample size from age 69 to 70 is natural as sample size declines steadily with age. While the 1911 census counts more than 400,000 individuals at age 50, the number of individuals drops below 5,000 at age 90.

Summary statistics in table 2 (upper panel) report a considerable decline in LFP between age 69 and 70. The drop totals to 7 percentage points (from 46% to 39%), while differences in other observable characteristics are fairly small. At age 70, individuals are less often married and the share of foreign born individuals is slightly higher. The lower panel in table 2 reports summary statistics for the baseline estimation sample with five age-years below the cutoff (65 - 69, N: 803,208) and above the cutoff (70 - 74, N: 551,100). Including additional age-years naturally leads to a larger differential in mean LFP rates of 48% below the eligibility age and 34% above. Differences in means of other socio-economic background variables are very similar when using five-year age brackets. Again, the share of married individuals declines over age, both the shares of foreign born individuals and single person households slightly increase over age and the rest remains largely unchanged.

Throughout this study, we examine the labor supply responses jointly for men and women. Since the participation rates of women were generally low, we also present the main results separately for men and women.[23] In general and for both men and women, the decline of LFP rates over age evolved less steeply than today. We will argue in this paper that one major explanation for this phenomenon was the low coverage by social security and old-age pensions in the early 20th century.

[23]Table 2 indicates that the average LFP rate marginally below the eligibility threshold (age 69) was 46%. However, while only 20% of women were still economically active at that age, the LFP rates of older men were relatively high and totaled to 76% (figure 12). See section 5.4 for separate results regarding men and women.

The majority of older workers (almost one third) was active in sectors such as crafts or related trades (table 3). The second largest sector included service workers, followed by skilled agricultural and fishery work and elementary occupations. In contrast, only few older workers earned a living as senior officials or managers, technicians or other professionals. We use occupational information later-on to construct a measure of labor earnings. This allows for estimates of the labor supply response to pension availability in different regions of the earnings distribution and is insightful regarding the functioning of the means test.

5 The Labor Supply Effect of the OPA

5.1 Baseline Results

Table 4 reports the decline in labor supply due to the OPA for our preferred RD specification.[24] The estimate shows a precisely measured drop in LFP of 6.0 percentage points when people reach the age-based eligibility threshold. Contrasting the LFP distribution over age separately for the 1901 and the 1911 censuses, figure 5 reveals the striking difference in age-specific LFP patterns before and after the OPA became effective in 1909. Without universal coverage by old-age assistance in 1901, participation rates decline gradually over age (figure 5, panel a). In contrast, we observe an abrupt drop of LFP in 1911 (figure 5, panel b) exactly at the age-based eligibility cutoff. The sudden drop only appears at the eligibility age and only after the OPA has been introduced. Since no other social security programs existed that would induce LFP changes between age 69 and 70, the LFP reduction is caused by the availability of the OPA.

The estimated decline is sizable both in absolute and relative terms. Departing from a participation rate of 46% at age 69, the estimated absolute decline of 6.0 percentage

[24] Our preferred specification uses a bandwidth of 5 age-years to the left (65 - 69) and to the right (70 - 74) of the age cutoff, a linear polynomial in age, and uniform kernel weighting putting the same weight on all observations. Graphical evidence in figure 7 suggests that linear and quadratic age polynomials reproduce the LFP-age relationship very similarly and thus, for brevity, we report the linear specification for all subsequent RD specifications including placebo tests and robustness checks. Throughout, the results are very similar for the quadratic polynomial and are available from the authors upon request.

points translates into a relative decline of 13%. Given the scale of the program and the limited social security net at that time, the substantial decline in participation rates is not surprising.

The considerable negative impact of the OPA on LFP rates of older workers, however, can be easily overlooked in aggregated labor market trends. As noted by Johnson (1994) and Costa (1998), the decline of LFP rates for the aggregate of older men in the UK has not accelerated after the OPA was implemented in 1909. Using a simple difference-in-differences (DiD) representation[25], we argue that simply extrapolating past aggregate LFP trends, however, does not produce a valid counterfactual scenario. Figure 6 shows that participation rates of people slightly below the eligibility age decreased between 1891 and 1901 but increased between 1901 and 1911. Thus, assuming that these age-groups are a valid counterfactual for the age-groups (slightly) above the eligibility age, the pace of the LFP decline for older workers would have slowed down between 1901 and 1911 without the OPA.

Our results are consistent with those of Fetter and Lockwood (2018), who study the introduction of the old-age assistance program in the U.S. They find that the labor supply of men aged 65 to 74 declined by 8.5 percentage points between 1930 and 1940 as a consequence of the introduction of old-age assistance in 1935. Our estimated LFP decline of 6.0 percentage points is slightly smaller in magnitude, which is mostly due to the inclusion of women in our sample (for more information on gender differences see section 5.4). In contrast to the U.S. case, where almost half of the decline was explained by exits

[25] For this exercise, we add the census wave of 1891 in addition to 1901 and 1911, which allows us to graphically test the common trend assumption. For the two age groups at the eligibility threshold set by the OPA, the graph indicates a common trend in the pre-treatment period between 1891 and 1901, showing that LFP rates of the two groups move in tandem between 1891 and 1901. The common trend is stable for neighboring ages below the eligibility threshold (65-69) and above (70-74), as is evident from figure B.1 in the appendix. After that, between 1901 and 1911, the two age groups diverge in terms of LFP. The DiD estimate quantifies the LFP differential to range between 4.0 to 4.5 percentage points difference between older workers aged 69 and 70 (table 5). It eliminates any time-constant unoberservables that would confound estimating the impact of the introduction of the OPA on LFP. The DiD is robust against including any available observable characteristics such as number of own children in the household, household size, and indicators for being married, foreign born, disabled, and region (covering the 53 counties in England and Wales).

from work relief programs and unemployment, direct transitions from work to retirement were more prevalent in the UK.

To examine the labor supply response of the population that retires directly from active work, we now disregard the un-employed and deviate from the main LFP measure by setting LFP = 1 only for those who are actually working (and LFP = 0 for everyone else). We assume that those individuals in the data who report an occupation are actively working and that those who report to be unemployed are not working. This exercise shows that almost the entire decline in LFP is driven by people who stop working (figure 8). The LFP rate among active workers declines significantly by 5.7 percentage points at the age-based eligibility threshold (table 6, column 1). When setting LFP = 1 for those in the labor force who are unemployed, the drop is virtually non-existent (table 6, column 2). The declining LFP rates are hence strongly driven by older workers who directly retire from active work.

5.2 Labor Supply Effects Across the Earnings Distribution

Looking at labor supply responses across the earnings[26] distribution allows us to test the key implication of the earnings test that people with higher earnings potential have to forgo more income to become eligible for old-age assistance. Figure 9 depicts the earnings distribution in 1911 and indicates the quartiles of the distribution (mean: 58.6£).

Figure 10 shows that the negative labor supply response is more pronounced at the lower end of the earnings distribution than at the top. In the first quartile, LFP declines by 15.4% and thus more than twice compared to the fourth quartile (7.2%). This finding is consistent with the theoretical prediction that substituting labor earnings with old-age assistance is less likely when earnings potential is high. High-earnings individuals have to trade-off a considerable drop in consumption opportunities against leisure gains from retiring.

[26]The census data do not directly measure earnings, but as detailed in appendix C, we construct earnings based on occupational information that we match to more recent U.S. census data that do report personal earnings. Thus, our earnings variable rather measures earnings potential instead of actual earnings.

5.3 Family Background and Old-Age Insurance

The family can function as old-age insurance, especially when individuals are credit constrained and do not have access to social security. The insurance argument has mainly been pointed out with regard to children (Leibenstein, 1957; Caldwell, 1982; Cain, 1983; Boldrin and Jones, 2002), but public old-age assistance can replace any type of family-related transfers (e.g. from spouses or other related household members). This mechanism was recognized during the legislation process of the OPA and, after some debates, it was finally decided that voluntary family transfers must be included in the calculation of the annual means of a pension claimant (Casson, 1908).

Given that private transfers are considered in the means test, family ties might affect eligibility and the corresponding labor supply response to the OPA. Older workers who receive private transfers (e.g. from family members) might either not pass the means test or have already used the family transfers to retire before the age of 70. Consequently, their labor supply response at the eligibility age is expected to be less pronounced. In contrast, individuals who do not receive private transfers have to work until pension benefits become available (if they lack savings) and thus are expected to react more strongly when they reach the eligibility age.

Although we do not have information on within-family cash transfers, we can test three dimensions of family ties based on existing living arrangements and marital status. First, we proxy inter-generational transfers from children by distinguishing older workers with and without own children in the household.[27] Second, we use marital status to proxy transfers in spousal relationships. And third, we compare single versus multiple person households as a more general measure for transfers from closely related individuals.

Table 7 reports RD estimates from samples that are stratified by individuals with and without own children in the household. These estimates indicate that older workers without own children in the household show stronger declines in LFP rates: at the age-based eligibility threshold, LFP falls by 6.9 percentage points. In contrast, LFP declines

[27] Note that given the age of their parents almost all children are in working-age already. More than 95% of children are 15 years or older, more than 90% are 20 years or older

by only 5.1 percentage points among older workers with children. Graphical evidence in figure 11 supports the estimation results. The finding of stronger LFP reactions among older workers without children is consistent with the hypothesis that children served as a type of old-age insurance before social security systems existed. It must be noted, however, that the point estimates for individuals with and without children only weakly differ from each other (z-statistic: 1.8).

The insurance argument is particularly salient when looking at solitaire individuals. Consistent with this view, the labor supply reductions among non-married individuals are larger compared to the sub-sample of married individuals (table 7) although these estimates do not significantly differ (z-statistic = 0.6). Individuals living completely on their own (single households) respond very strongly to the pension incentive. Individuals living alone show large LFP reductions of 9.5 percentage points, while those living with one or more other persons react significantly less (5.7 percentage points, z-statistic: 3.2).

One concern is that living arrangements correlate with earnings. In fact, we do find that earnings are lower in each of the three groups that show more pronounced labor supply reductions (no children, non-married, and living alone). However, we argue that the results are not entirely driven by earnings gaps for two reasons. First, the earnings gaps are rather small (£6 to £8 on average). Second, the earnings gaps are not correlated with the labor supply response gaps. For instance, we see the biggest earnings gap between married and non-married people, but the smallest LFP response gap.

5.4 Labor Supply Effects by Gender

A particularly robust finding documented in the literature on labor supply is that women have much larger labor supply elasticities than men, especially on the extensive margin of labor supply (see Keane, 2011, for a review). One important difference between men and women, in particular when using data from the early 20th century, is that a large share of women was not part of the labor force based on the prevalent definition. This also holds in the cohorts under study here and is apparent from comparisons of LFP rates between older men and women (figure 12). The figure indicates that LFP rates at the age of 69 were much larger among men (76%) compared to women (20%). RD estimates on separate male and female samples (table 8) show that, in absolute terms, the labor supply responses are larger among men (7.9 percentage points) as compared to women (3.2 percentage points). When placing the smaller female labor supply reduction into the context of their smaller LFP rate, their relative reduction of 16% is much larger compared to men (10%). This result is consistent with larger female labor supply elasticities as documented in the literature.

RD estimates across the earnings distribution also differ between men and women (figure 13). Especially in the second earnings quartile, women show dramatic labor supply reductions of up to 27%. Due to their relatively low earnings[28], women also had low adjustment costs in terms of forgone earnings to pass the means test. This is largely in line with the observation that the female labor supply response is particularly large at the lower end of the female earnings distribution.

5.5 Within-Family Spillovers: Joint Retirement of Couples

Many empirical studies suggest that within-family spillovers play an important role in retirement decisions (Blau and Riphahn, 1999; Gustman and Steinmeier, 2000; Baker, 2002; Atalay et al., 2019). Any correlation of tastes for leisure between partners may lead to joint retirement behavior of couples.

The UK census allows us to link the labor force status of spouses. We use this information to study whether people who have already reached the eligibility age (70) respond to the pension eligibility of their spouse. Therefore, we plot LFP rates of people above the age of 70 against the age of their corresponding spouse. Figure 14 shows that there is a sizable jump at the age cutoff in 1911 (OPA available), indicating that LFP rates drop among individuals whose spouse reaches age 70 and thus also becomes eligible. This graphical evidence is further supported by RD estimates in table 9 that report a significantly measured LFP drop of 3.1 percentage points of eligible individuals whenever

[28] The mean earnings of women in 1911 are 41£, which amounts to two thirds of mean earnings among men (60£).

their spouse becomes eligible. The significant drop in LFP is mainly driven by eligible men (3.3 percentage points) in the moment when their wife becomes eligible but not vice versa. These results are strongly suggestive for within-family spillovers and are in line with previous findings on joint retirement decisions of couples.

5.6 Validity and Robustness of the RD

Several sensitivity checks and falsification tests support the validity of our RD design. Based on these exercises, the estimates presented above can be given a causal interpretation.

5.6.1 Smoothness Analysis

We start verifying the validity of the RD by testing whether pre-determined covariates exhibit a discontinuity at the threshold. This smoothness analysis includes observable characteristics on the number of own children in the household, the share of individuals living in urban areas, the share of married, foreign born and disabled persons, and the share of individuals living alone (similar to the summary statistics reported in table 2). These variables mainly evolve smoothly around the age-based eligibility threshold, as suggested by the estimates in table 10 and figure B.2. Although discontinuities arise for some variables (share of married, foreign born, disabled and single individuals), these differences are negligibly small. Further estimates on respective sub-samples indicate that these groups do not drive the results (table 11 and figure B.3). We conclude that measurable discontinuities in observables at the age-based eligibility threshold are small and uncorrelated to LFP and thus unproblematic for the identification. In summary, the smoothness analysis suggests that no relevant changes other than the outcome of interest (LFP) take place at the age cutoff, thus supporting the validity of the RD.

5.6.2 Placebo Tests

Next, we conduct two placebo tests to show that LFP effects do not appear at any arbitrarily chosen age cutoff. First, we run the baseline RD specification on the 1901 census to ensure that the abrupt LFP decline measured after the introduction of the OPA in 1911 does not occur in 1901. Table 12 (upper panel) indicates that there is no sizable LFP decline in 1901, consistent with graphical evidence in figure 5 (panel a) that indicates a smooth LFP decline over age before the pension was introduced. Table 12 also shows that there is no substantial drop in LFP at age 60 in the 1901 census. Since the birth cohort that reached age 60 in 1901 eventually reached age 70 in 1911, this robustness check rules out that individuals who play a key role for the identifying variation in the RD design already exhibited an LFP decline at earlier ages before the OPA was introduced.

Second, we repeat the analysis for arbitrarily chosen placebo age cutoffs in 1911. RD estimates in table 12 (lower panel) show that we do not observe a similar decline in LFP rates at any hypothetical age cutoff other than the true eligibility threshold at age 70.[29]

5.6.3 Bandwidth Choice, Polynomials, and Bias Correction

We finally present several adjustments to the baseline specification that have been proposed in the RD literature to verify robustness of the estimates (table 13). First, the estimates are not sensitive to bandwidth choice and only change little when making the bandwidth arbitrarily large. Second, changing the polynomial degree when modeling LFP as a function of age does not indicate sizable changes in the coefficients. Alternative estimation techniques also indicate that the baseline estimates are robust when estimating local non-parametric and bias-corrected point estimates with robust standard errors as suggested by Calonico et al. (2014a,b). Overall changes in estimates are only moderate even when undertaking considerable changes in bandwidth, age polynomials, weighting scheme (uniform vs. triangular kernel) or estimation procedure.

[29] This result holds for any other hypothetical age cutoff. Additional placebo tests on arbitrarily chosen age cutoffs are available from the authors upon request.

6 Welfare Implications

We now translate the causal estimate of the labor supply decline into a statement on the welfare impact of the OPA. We first approximate inefficiencies that arise from the program by relating the willingness to pay for the OPA to the actual costs of the policy. In a second step, we compare this efficiency loss to the equity gain and then conclude that the program was most likely welfare enhancing. Throughout, we distinguish two classes of pension recipients. The first group consists of infra-marginal individuals who are eligible for pension payments from the OPA irrespective of their behavior. Most importantly, infra-marginal individuals pass the means test without reducing their labor supply. The second group consists of marginal individuals who adjust their labor supply in response to the OPA.

6.1 Willingness to Pay

We obtain the bounds of the willingness to pay by relating the causal estimate of the labor supply reduction (6 percentage points) to the pension recipience rate in close distance to the eligibility threshold. In the age-group 70-74 (the same as in the baseline RD) at least 17.5% received a pension.[30] Among recipients, this yields a share of 34% marginal individuals who reduce their labor supply in response to the reform (0.06/0.175 = 34%). Reversely, this also implies that 66% of the recipients do not adjust their labor supply and thus can be classified as infra-marginal individuals.[31]

Assuming that infra-marginal individuals valued their pension benefits fully[32], the average recipient valued £1 of pension benefits by at least £0.66. The average valuation of

[30]The recipience rate of 17.5% is an absolute lower bound when assuming that every individual above age 74 receives a pension. Since the UK census does not provide direct information on recipients, we use the absolute number of recipients in 1911 from historical documents and then compute the recipience rate. For details and alternative assumptions on computing the recipience rate, see appendix D.

[31]The share of infra-marginal individuals was higher among women compared to men for at least two reasons. First, receiving a pension was more common for women compared to men (64% of pension recipients were women, but only 58% of the population aged 70 or older). Second, the behavioral response to the policy was smaller in absolute terms because women were less often part of the labor force.

[32]This assumption is plausible since infra-marginal individuals would receive the pension anyway because they have sufficiently low means to pass the means test.

benefits thus ranges somewhere between 66% and 100% because it is likely that marginal individuals value their pension benefits by more than zero. Especially for marginal individuals with relatively low earnings or for those who leave the labor force from unemployment, the valuation is likely to be high and perhaps close to 100%.

We use the estimates by Fetter and Lockwood (2018), generated by a lifetime labor supply model for the U.S., to approximate the average valuation for these marginal individuals in the UK. According to their results, each marginal worker who retired from active work (and not unemployment or work relief programs) in the U.S. valued one dollar of old-age assistance by at least 80 cents. [33] In what follows, we also assume that marginal recipients in the UK valued their OPA benefits by at least £0.8 for each pound of the program. We believe this value serves as a natural lower bound as the design of the old-age assistance system in the U.S. was similar to the one in the UK (e.g. in terms of benefit levels).[34] However, the eligibility age was higher in the UK (70 vs. 65 in most US states). Assuming that the dis-utility from work increases with age when work becomes more strenuous and burdensome, marginal recipients in the UK would arguably gain more utility from retiring than their younger counterparts in the U.S.

For the lower bound of infra-marginal individuals (66%), this implies that the average pound of OPA benefits was worth at least £0.93 of unconditional old-age income ($0.66 \times 1 + 0.34 \times 0.8 = 0.93$). We thus conclude that the willingness to pay for the program (i.e. its average valuation) was high despite its considerable labor supply impact. These results are in line with those of Fetter and Lockwood (2018) but they develop differently. While the share of infra-marginal individuals among OPA recipients is higher (66% or

[33] Fetter and Lockwood (2018) report that 48% of the individuals are infra-marginal (assuming full valuation) and that the remaining 52% are marginal individuals (valuation between zero and one). Based on their final simulation result of an average valuation of 0.95 dollars for each dollar from the program, we have that $0.48 \times 1 + 0.52 \times x = 0.95$ and thus the average valuation among marginal individuals is $x = 0.9$. Further, Fetter and Lockwood (2018) report that about 50% among marginal individuals retire from employment, so that the rest is only substituting from pre-existing work relief programs or unemployment. If we assume that the latter group values the assistance fully, the average valuation among marginal individuals who retire from active work is $y = 0.8$ since $0.5 \times 1 + 0.5 \times y = 0.9$

[34] Fetter and Lockwood (2018) report that the benefit level in the U.S. was 25% of median earnings of wage-earners aged 60-64. The OPA benefits amounted to 22% of mean total wage earnings.

more[35] compared to 48%), marginal individuals in the UK predominantly retire from active work. Hence, it seems plausible to assume that the average valuation of one pound of OPA benefits is lower (0.8 vs. 0.9) among marginal individuals in the UK compared to the US.

6.2 Efficiency Loss

We calculate the efficiency loss by dividing the recipients willingness to pay for the policy (£0.93 per pound of old-age assistance) by the the cost of transferring £1 to a pension recipient. Without behavioral changes or administrative costs, this benefit-to-cost ratio would be one, so that deviations from one measure the efficiency loss of the transfer program.

Costs for the government arise due to cash transfers to pension recipients (mechanical costs)[36] and due to behavioral responses that affect government revenues. In the case of the OPA, mechanical costs amount to £1.03, which is the sum of £1 for each pound paid out to pension recipients plus £0.03 of administrative costs. Administrative costs of the OPA were low and only amounted to 3% (Old Age Pensions Committee, 1919), because the OPA mainly relied on existing institutions (post-offices etc., see Murray, 2009).

Government revenues were hardly affected by behavioral changes of older workers because their tax burden was low in the first place. Individuals earning below £160 did not pay any income tax (House of Commons, 1911), which was more than 2.7 times the mean earnings in 1911 (£58.6, Feinstein, 1990). Thus, only few individuals were subject to income taxes in the UK at that time.[37] The tax burden among the elderly was even lower given that they earned much less than the average citizen (Spender, 1892).

[35] In a more realistic scenario the share of infra-marginal individuals is 85% (see appendix D and table A.1 for details). Under this scenario, the valuation of the average pound of OPA benefits would be even closer to one and amount to £0.97 ($0.85 \times 1 + 0.15 \times 0.8 = 0.97$).

[36] Note that our definition of mechanical costs differs from the one in Finkelstein (2019) because we scale the benefit and costs by the number of recipients and not the infra-marginal recipients. Since the scaling factor cancels out eventually, our benefit-to-cost ratio is alike to the one suggested in Finkelstein (2019).

[37] There were only about 1 million income tax payers in the UK in 1909/1910 compared to 36,353,455 inhabitants in England and Wales alone in 1911.

Assuming no impact on government revenues, the benefit-to-cost ratio amounts to $\frac{0.93}{1.03} = 0.90$. Given that the ratio is numerically close to one, each pound spent on old-age assistance is valued almost as high as the costs accruing from the program. Thus, the efficiency loss of the program was relatively low. In the next section, we also discuss potential labor supply responses of top income earners to increased taxation that have been disregarded so far.

6.3 Equity-Efficiency Trade-off

If the policy was ultimately welfare enhancing depends on whether the efficiency loss is outweighed by equity gains. Under the widespread assumption of diminishing marginal utility of consumption, income transfers from the top to the bottom of the income distribution are welfare enhancing in absence of behavioral changes. The OPA is a prime example of a policy that redistributed from top income earners (roughly the top 1%) to people at the bottom of the earnings distribution and thus increased equity.

Whether these equity gains are outweighed by potential efficiency losses depends on three factors. First, the size of inefficiencies created by behavioral responses of older workers, second, the size of inefficiencies created by behavioral responses of high income earners to increased taxation, and third, the social welfare function. While the efficiency loss due to behavioral changes of older workers is low (see section 6.2), we can neither directly measure responses to taxation of high income earners nor observe the social welfare function. However, if the social welfare function is logarithmic in consumption [38], we can calculate the social marginal utility of consumption at the top of the earnings distribution, where taxes are levied. This computation suggests that, among tax payers, the social marginal utility of an additional pound is only 1.5% of the social marginal utility at the lower margin where the means test phases in.[39] Thus, assuming a logarithmic

[38] This assumption is similar to the one by Diamond and Saez (2011) (p. 169).

[39] We compare an additional pound in the lower earnings region where the means test phases in (moving from £30 to £31), to higher earnings at the bottom of the first progressive tax bracket (moving from £2000 to £2001).

welfare function[40], the efficiency loss from higher taxation must have been (unrealistically) huge in order to outweigh the equity gain. Given the modest marginal tax rates even at the top, such losses are very unlikely. Against this background we conclude that the program was most likely welfare enhancing.

7 Conclusions

In this paper, we take advantage of a unique historical reform – the introduction of means-tested pensions in the UK in 1909 – to study the effect of old-age assistance on labor supply and welfare. Using recently digitized full-count census data, we estimate the causal impact of the policy on the labor supply of older workers by exploiting variation around the newly created age-based eligibility threshold.

We find a considerable reduction of labor supply in the local environment of the eligibility threshold after the pension was implemented. When reaching the eligibility age of 70, the labor force participation rate declines by 6 percentage points or 13%. This decline was strongly driven by older workers who directly retire from active work (and not unemployment). Labor supply reductions are larger when earnings potential is low and when family ties, as a proxy for private transfers, are weak. Our results are also suggestive for joint retirement decisions of couples.

The transparent redistributive design of the pension system allows us to translate our reduced form estimate into a welfare statement. We argue that the introduction of old-age assistance in the UK was welfare enhancing despite the considerable labor supply decline for two reasons. First, the policy created only small inefficiencies because, relative to the immense size of the program, the labor supply decline was moderate. Second, equity gains were large as the policy redistributed from the top to the bottom of the earnings distribution in a time of relatively high income inequality and only limited social security.

[40]Using an iso-elastic utility function with a relative risk aversion parameter above 1, as used in Finkelstein (2019) (p.7), yields even larger gains from redistribution.

The historical setting allows us to credibly identify the full labor supply effect and welfare impact of an old-age assistance program and thus extends a large body of literature that quantifies the incentive effects of marginal changes in existing pension schemes. Our results are policy-relevant, not only for developing countries without universal coverage by old-age assistance (e.g. in Sub-Sahara Africa).[41] The substantial and immediate response to the creation of a completely new social security program is also knowledgeable for future policy changes in developed countries.

[41] An overview on old-age assistance coverage worldwide is provided by www.pension-watch.net.

References

Abramitzky, R., L. P. Boustan, and K. Eriksson (2014). A Nation of Immigrants: Assimilation and Economic Outcomes in the Age of Mass Migration. *Journal of Political Economy 122*(3), 467–506.

Andersen, A. G., S. Markussen, and K. Roed (2020). Pension Reform and the Efficieny-Equity Trade-Off: Impacts of Removing an Early Retirement Subsidy. *IZA Discussion Paper No. 12918*.

Anderson, M., C. Dobkin, and T. Gross (2012). The Effect of Health Insurance Coverage on the Use of Medical Services. *American Economic Journal: Economic Policy 4*(1), 1–27.

Anderson, M. L., C. Dobkin, and T. Gross (2014). The Effect of Health Insurance on Emergency Department Visits: Evidence from an Age-Based Eligibility Threshold. *Review of Economics and Statistics 96*(1), 189–195.

Arthi, V., B. Beach, and W. W. Hanlon (2019). Recessions, Mortality, and Migration Bias: Evidence from the Lancashire Cotton Famine. *Discussion Paper* (building on NBER Working Paper No. 23507).

Atalay, K. and G. F. Barrett (2015). The Impact of Age Pension Eligibility Age on Retirement and Program Dependence: Evidence from an Australian Experiment. *The Review of Economics and Statistics 97*(1), 71–87.

Atalay, K., G. F. Barrett, and P. Siminski (2019). Pension incentives and the joint retirement of couples: Evidence from two natural experiments. *Journal of Population Economics 32*(3), 735–767.

Atkinson, A. B. (2005). Top Incomes in the UK Over the 20th Century. *Journal of the Royal Statistical Society: Series A (Statistics in Society) 168*(2), 325–343.

Atkinson, A. B. (2007). The Distribution of Top Incomes in the United Kingdom 1908-2000 (Chapter 4). In A. B. Atkinson and T. Piketty (Eds.), *Top Incomes over the*

Twentieth Century. A Contrast Between Continental European and English-Speaking Countries, pp. 82 – 140. Oxford University Press.

Baker, M. (2002). The Retirement Behavior of Married Couples: Evidence from the Spouse's Allowance. *The Journal of Human Resources 37*(1), 1–34.

Battistin, E., A. Brugiavini, E. Rettore, and G. Weber (2009). The Retirement Consumption Puzzle: Evidence from a Regression Discontinuity Approach. *American Economic Review 99*(5), 2209–26.

Beach, B. and W. W. Hanlon (2019). Censorhip, Familiy Planning, and the Historical Fertility Transition. *NBER Working Paper No. 25752*.

Blau, D. M. and R. T. Riphahn (1999). Labor Force Transitions of Older Married Couples in Germany. *Labour Economics 6*(2), 229–252.

Boldrin, M. and L. E. Jones (2002). Mortality, Fertility, and Saving in a Malthusian Economy. *Review of Economic Dynamics 5*(4), 775–814.

Brown, K. M. (2013). The Link Between Pensions and Retirement Timing: Lessons from California Teachers. *Journal of Public Economics 98*, 1 – 14.

Börsch-Supan, A. (2000). Incentive Effects of Social Security on Labor Force Participation: Evidence in Germany and Across Europe. *Journal of Public Economics 78*(1-2), 25–49.

Cain, M. (1983). Fertility as an Adjustment to Risk. *Population and Development Review 9*(4), 688–702.

Caldwell, J. C. (1982). *Theory of Fertility Decline.* London: Academic Press.

Calonico, S., M. D. Cattaneo, and R. Titiunik (2014a). Robust Data-Driven Inference in the Regression-Discontinuity Design. *Stata Journal 14*(4), 909–946.

Calonico, S., M. D. Cattaneo, and R. Titiunik (2014b). Robust Nonparametric Confidence Intervals for Regression-Discontinuity Designs. *Econometrica 82*(6), 2295–2326.

Card, D., C. Dobkin, and N. Maestas (2008). The Impact of Nearly Universal Insurance Coverage on Health Care Utilization: Evidence from Medicare. *American Economic Review 98*(5), 2242–58.

Card, D., C. Dobkin, and N. Maestas (2009). Does Medicare Save Lives? *The Quarterly Journal of Economics 124*(2), 597 – 636.

Carpenter, C. and C. Dobkin (2009). The Effect of Alcohol Consumption on Mortality: Regression Discontinuity Evidence from the Minimum Drinking Age. *American Economic Journal: Applied Economics 1*(1), 164–182.

Carpenter, C. and C. Dobkin (2015). The Minimum Legal Drinking Age and Crime. *The Review of Economics and Statistics 97*(2), 521–524.

Carpenter, C. and C. Dobkin (2017). The Minimum Legal Drinking Age and Morbidity in the United States. *The Review of Economics and Statistics 99*(1), 95–104.

Casson, W. A. (1908). *Old-Age Pensions Act, 1908: Together with the Text of the Regulations Made Thereunder Dated 15th October, 1908, and Official Circulars and Instructions for the Guidance of Pension Authorities by the Local Government Boards of England, Scotland, and Ireland; Annotated and Explained, with Historical Introduction.* C. Knight & Company, Limited.

Coile, C. C. (2015). Economic Determinants of Workers' Retirement Decisions. *Journal of Economic Surveys 29*(4), 830 – 853.

Costa, D. L. (1998). The Evolution of Retirement. In *The Evolution of Retirement: An American Economic History, 1880-1990*, pp. 6–31. NBER and University of Chicago Press.

Danzer, A. M. (2013). Benefit Generosity and the Income Effect on Labour Supply: Quasi-Experimental Evidence. *The Economic Journal 123*(571), 1059–1084.

Department of Labour Statistics (1915). Seventeenth Abstract of Labour Statistics of the United Kingdom. Abstract of Labour Statistics.

Diamond, P. and E. Saez (2011). The Case for a Progressive Tax: From Basic Research to Policy Recommendations. *Journal of Economic Perspectives 25*(4), 165–90.

Feinstein, C. (1990). New Estimates of Average Earnings in the United Kingdom, 1880-1913. *The Economic History Review 43*(4), 595–632.

Feldstein, M. and J. B. Liebman (2002). Social Security (Chapter 32). Volume 4 of *Handbook of Public Economics*, pp. 2245 – 2324. Elsevier.

Fetter, D. K. and L. M. Lockwood (2018). Government Old-Age Support and Labor Supply: Evidence from the Old Age Assistance Program. *American Economic Review 108*(8), 2174–2211.

Finkelstein, A. (2019). Welfare Analysis Meets Causal Inference: A Suggested Interpretation of Hendren. *Note, MIT Economics*.

Fitzpatrick, M. D. and T. J. Moore (2018). The Mortality Effects of Retirement: Evidence from Social Security Eligibility at Age 62. *Journal of Public Economics 157*, 121 – 137.

Gustman, A. L. and T. L. Steinmeier (2000). Retirement in Dual-Career Families: A Structural Model. *Journal of Labor Economics 18*(3), 503 – 545.

Higgs, E., C. Jones, K. Schürer, and A. Wilkinson (2013). Integrated Census Microdata (I-CeM) Guide; 1851-1911. *University of Essex, Department of History*.

House of Commons (1911). Revenue and expenditure (England, Scotland, and Ireland). *House of Commons Papers 220*(45), 16.

Human Mortality Database (2018). University of California, Berkeley and Max Planck Institute for Demographic Research, Rostock. *American Economic Journal: Economic Policy available at www.mortality.org or www.humanmortality.de*(data downloaded on 01.10.2018).

Johnson, P. (1994). The Employment and Retirement of Older Men in England and Wales, 1881 – 1981. *The Economic History Review 47*(1), 106–128.

Keane, M. P. (2011). Labor Supply and Taxes: A Survey. *Journal of Economic Literature 49*(4), 961–1075.

Krueger, A. B. and B. D. Meyer (2002). Labor Supply Effects of Social Insurance (Chapter 33). Volume 4 of *Handbook of Public Economics*, pp. 2327 – 2392. Elsevier.

Krueger, A. B. and J.-S. Pischke (1992). The Effect of Social Security on Labor Supply: A Cohort Analysis of the Notch Generation. *Journal of Labor Economics 10*(4), 412–37.

Lee, D. S. and T. Lemieux (2010). Regression Discontinuity Designs in Economics. *Journal of Economic Literature 48*(2), 281–355.

Leibenstein, H. (1957). *Economic Backwardness and Economic Growth : Studies in the Theory of Economic Development.* John Wiley and Sons.

Liebman, J. B., E. F. P. Luttmer, and D. G. Seif (2009). Labor Supply Responses to Marginal Social Security Benefits: Evidence from Discontinuities. *Journal of Public Economics 93*(11 - 12), 1208 – 1223.

Manoli, D. and A. Weber (2016). Nonparametric Evidence on the Effects of Financial Incentives on Retirement Decisions. *American Economic Journal: Economic Policy 8*(4), 160–82.

Mastrobuoni, G. (2009). Labor supply effects of the recent social security benefit cuts: Empirical estimates using cohort discontinuities. *Journal of Public Economics 93*(11 - 12), 1224 – 1233.

Minnesota Population Center (2018). Integrated Public Use Microdata Series, International: Version 7.0. Minneapolis, MN: IPUMS.

Murray, B. (2009). The ‘Peoples Budget´ A Century On. *Journal of Liberal History Liberal Democrat History Group*(64), 4 – 13.

Old Age Pensions Committee (1919). Old Age Pensions: Appendix to the Report of the Departmental Committee on Old Age Pensions Including Minutes of Evidence. Technical report, London.

Spender, J. A. (1892). The State and Pensions in Old Age. *London, SWAN SONNENSCHEIN & CO.*.

Thane, P. (2000). *Old Age in English History: Past Experiences, Present Issues* (1 edition ed.). Oxford UK ; New York: Oxford University Press.

The Times (1909). Old-Age Pensions. *The Times*, page 9.

Appendix

Tables

Table 1: Pension Schedule

Annual Means X	Weekly Pension Entitlement
X $\leq$ £21	5 Shillings
£21 $<$ X $\leq$ £23, 12s and 6d	4 Shillings
£23, 12s and 6d $<$ X $\leq$ £26 and 5s	3 Shillings
£26 and 5s $\leq$ $<$ X £28, 17s and 6d	2 Shillings
£28, 17s and 6d $<$ X $\leq$ £31, 10s and 6d	1 Shilling
X $>$ £31 and 10s	–

Source: UK legislation (1908). *Note:* X denotes annual means, s denotes shillings and d denotes pence. £1 corresponds to 20s or 240d.

Table 2: Summary Statistics by Age (Census 1911)

	69		70	
	Mean	S.D.	Mean	S.D.
Share Female	0.55	0.50	0.56	0.50
Share Urban	0.10	0.30	0.10	0.30
Share Married	0.50	0.50	0.47	0.50
Share Foreign Born	0.04	0.20	0.06	0.24
Share Disabled	0.02	0.13	0.02	0.14
N Children in the Household	0.8	1.1	0.7	1.0
Share Single Households	0.07	0.26	0.08	0.28
Labor Force Participation Rate	0.46	0.50	0.39	0.49
Observations	150,293		140,288	

	65 - 69		70 - 74	
	Mean	S.D.	Mean	S.D.
Share Female	0.55	0.50	0.57	0.50
Share Urban	0.10	0.30	0.10	0.29
Share Married	0.54	0.50	0.43	0.50
Share Foreign Born	0.04	0.20	0.05	0.22
Share Disabled	0.02	0.13	0.02	0.14
N Children in the Household	0.9	1.2	0.7	1.0
Share Single Households	0.07	0.25	0.09	0.29
Labor Force Participation Rate	0.48	0.50	0.34	0.47
Observations	803,208		551,100	

Source: UK Census wave 1911 and IPUMS. *Note:* Reported values for men and women. The upper panel reports values marginally below and above the age cutoff (69 and 70). The lower panel reports values for the baseline estimation sample with 5 age-years below the cutoff (65 - 69) and above the cutoff (70 - 74).

Table 3: Occupational Composition

	Census 1911	
Occupation	Frequency	Percent
Legislators and Managers	10,781	0.8
Professionals	17,262	1.3
Technicians	9,292	0.7
Clerks	14,500	1.1
Service Workers	151,528	11.2
Agriculture and Fishery	105,472	7.8
Crafts and Related Trades	170,959	12.6
Machine Operators and Assemblers	39,143	2.9
Elementary Occupations	54,348	4.0
Armed Forces	1,881	0.1
Active	575,166	42.5
Inactive	779,142	57.5
Total Observations	1,354,308	100

Source: UK Census wave 1911 and IPUMS. *Note:* Reported values on occupations based on ISCO classification at the 1-digit level for individuals aged 65 - 74.

Table 4: RD Estimates of Labor Force Participation at the Age Cutoff (Baseline)

	Baseline RD Estimate
	-0.060** (0.006)
Observations	1,354,308

Source: UK Census wave 1911 and IPUMS. *Note:* RD estimate of the labor force participation rate (dependent variable) using an indicator for the age cutoff (= 0 if age < 70; = 1 if age >= 70). Linear polynomials in age. Specification uses a bandwidth of 5 age-years to the left (age 65-69, N: 803,208) and to the right (age 70-74, N: 551,100) of the age cutoff and uniform weighting on all observations. **, * denotes significance at the 1% and 5% level respectively. Standard errors (in parentheses) are clustered at age in years.

Table 5: Difference-in-Differences Estimates of Labor Force Participation

	(1)	(2)	(3)
DiD Estimate	-0.045** (0.003)	-0.040** (0.002)	-0.040** (0.002)
Controls	–	✓	✓
Controls X Census Year	–	–	✓
Observations	715,714	698,921	698,921

Source: UK Census wave 1891, 1901, 1911 and IPUMS. *Note:* Estimates for men and women aged 69 and 70. **, * denotes significance at the 1% and 5% level respectively. Standard errors in parentheses. The dependent variable is the labor force participation rate.

Table 6: RD Estimates of Labor Supply Response

	Actively Working (1)	Unemployed (2)
	-0.057** (0.006)	-0.002* (0.001)
Observations	1,354,308	

Source: UK Census wave 1911 and IPUMS. *Note:* RD estimates of the labor force participation rate (dependent variable) using an indicator for the age cutoff (= 0 if age < 70; = 1 if age >= 70). Estimates based on separate definitions of labor force status: LFP = 1 for individuals who are actively working (column 1) or LFP = 1 for individuals who are unemployed/formerly employed (column 2). All regressions use a bandwidth of 5 age-years to the left (age 65-69) and to the right (age 70-74) of the cutoff, a linear polynomial in age, and uniform weighting on all observations. **, * denotes significance at the 1% and 5% level respectively. Standard errors (in parentheses) are clustered at age in years.

Table 7: RD Estimates Stratified by Family Background

	RD Coefficient	Observations
	Children	
Children in Household	-0.051** (0.006)	627,733
No Children in Household	-0.069** (0.008)	726,575
z-Statistic (Significant Differences)	1.8	
	Marital Status	
Married	-0.056** (0.006)	679,145
Non-Married	-0.062** (0.008)	675,163
z-Statistic (Significant Differences)	0.6	
	Household Size	
Single Person Households	-0.095** (0.008)	102,938
Multiple Person Households	-0.057** (0.007)	1,251,370
z-Statistic (Significant Differences)	3.2**	

Source: UK Census wave 1911 and IPUMS. *Note:* RD estimates of the labor force participation rate (dependent variable) using an indicator for the age cutoff (= 0 if age < 70; = 1 if age >= 70). Estimates for sub-samples along the lines of three types of family background variables. All regressions use a bandwidth of 5 age-years to the left (age 65-69) and to the right (age 70-74) of the cutoff, a linear polynomial in age, and uniform weighting on all observations. **, * denotes significance at the 1% and 5% level respectively. Standard errors (in parentheses) are clustered at age in years. Testing for differences between two coefficients with known variances yields a z-test statistic (assumed to be normally distributed), where β is the estimated coefficient for the respective sub-sample and SE denotes the corresponding standard errors. This yields $z = \frac{\beta_C - \beta_{NC}}{\sqrt{SE^2_{\beta_C} + SE^2_{\beta_{NC}}}} = 1.8$ for testing between children/no children, $z = \frac{\beta_M - \beta_{NM}}{\sqrt{SE^2_{\beta_M} + SE^2_{\beta_{NM}}}} = 0.6$ for testing between married/non-married, and $z = \frac{\beta_{SP} - \beta_{MP}}{\sqrt{SE^2_{\beta_{SP}} + SE^2_{\beta_{MP}}}} = 3.2$ for testing between single/multiple person households.

Table 8: RD Estimates by Gender

	Men	Women
	(1)	(2)
	-0.079** (0.012)	-0.032** (0.004)
Observations	602,458	751,850

Source: UK Census wave 1911 and IPUMS. *Note:* RD estimates of the labor force participation rate (dependent variable) using an indicator for the age cutoff (= 0 if age < 70; = 1 if age >= 70). Estimates for the sub-samples of men (column 1) and women (column 2). All regressions use a bandwidth of 5 age-years to the left (age 65-69) and to the right (age 70-74) of the cutoff, a linear polynomial in age, and uniform weighting on all observations. **, * denotes significance at the 1% and 5% level respectively. Standard errors (in parentheses) are clustered at age in years.

Table 9: RD Estimates depending on Spousal Age (conditional on having reached 70)

Own Age	Total	Men	Women
	True Age Cutoff		
	Age of Spouse: 70, Census 1911		
	-0.031** (0.010)	-0.033** (0.009)	0.004 (0.003)
Observations	142,263	90,815	51,448
	Placebo Cutoffs		
	Age of Spouse: 60, Census 1911		
	-0.008 (0.008)	-0.005 (0.015)	0.020* (0.010)
Observations	33,669	28,085	5,584
	Age of Spouse: 65, Census 1911		
	0.012 (0.005)	0.015 (0.005)	0.001 (0.003)
Observations	76,804	57,616	19,188
	Age of Spouse: 60, Census 1901		
	0.020** (0.005)	0.018** (0.006)	-0.003 (0.009)
Observations	31,163	26,894	4,269
	Age of Spouse: 70, Census 1901		
	-0.006* (0.003)	0.012* (0.007)	0.002 (0.001)
Observations	112,066	73,370	38,696

Source: UK Census wave 1901 and 1911 and IPUMS. *Note:* RD estimates of the labor force participation rate (dependent variable) using an indicator for the age cutoff (= 0 if age < 70; = 1 if age >= 70) of the respective spouse. Estimates conditional on having reached the eligibility age of 70, separately for the full sample (1), men (2), and women (3). All regressions use a bandwidth of 5 age-years to the left (age 65-69) and to the right (age 70-74) of the cutoff, a linear polynomial in age, and uniform weighting on all observations. **, * denotes significance at the 1% and 5% level respectively. Standard errors (in parentheses) are clustered at age in years.

Table 10: Smoothness Analysis

Share Urban	0.003 (0.003)
Share Married	-0.012* (0.005)
Share Foreign Born	0.013** (0.004)
Share Disabled	0.001* (0.000)
N Children	-0.007 (0.004)
Share Single Households	0.005** (0.001)
Observations	1,354,308

Source: UK Census wave 1911 and IPUMS. *Note:* RD estimates of the respective observable (dependent variable) using an indicator for the age cutoff (= 0 if age < 70; = 1 if age >= 70). All regressions use a bandwidth of 5 age-years to the left (age 65-69) and to the right (age 70-74) of the cutoff, a linear polynomial in age, and uniform weighting on all observations. **, * denotes significance at the 1% and 5% level respectively. Standard errors (in parentheses) are clustered at age in years.

Table 11: RD Estimates on Sub-Samples

Sub-Sample	RD Coefficient	Observations
Married	-0.056** (0.006)	679,145
Non-Married	-0.062** (0.008)	675,163
Native-Born	-0.061** (0.007)	1,247,484
Non-Disabled	-0.061** (0.007)	1,329,510
Non-Single Person Households	-0.057** (0.007)	1,251,370
Single Person Households	-0.095** (0.008)	102,938

Source: UK Census wave 1911 and IPUMS. *Note:* RD estimates of the labor force participation rate (dependent variable) using an indicator for the age cutoff (= 0 if age < 70; = 1 if age >= 70). Estimates for sub-samples as indicated. All regressions use a bandwidth of 5 age-years to the left (age 65-69) and to the right (age 70-74) of the cutoff, a linear polynomial in age, and uniform weighting on all observations. **, * denotes significance at the 1% and 5% level respectively. Standard errors (in parentheses) are clustered at age in years.

Table 12: Placebo Tests

Age Cutoff	RD Coefficient	Observations
	Census 1901	
70	-0.009* (0.004)	1,041,111
60	0.001 (0.003)	1,889,492
	Census 1911	
69	-0.017 (0.013)	1,446,757
65	-0.004 (0.002)	1,820,152
60	0.001 (0.002)	2,293,676

Source: UK Census wave 1901 and 1911 and IPUMS. *Note:* RD estimates of the labor force participation rate (dependent variable) using an indicator for different placebo age cutoffs (as indicated). All regressions use a bandwidth of 5 age-years to the left (age 65-69) and to the right (age 70-74) of the cutoff, a linear polynomial in age, and uniform weighting on all observations. **, * denotes significance at the 1% and 5% level respectively. Standard errors (in parentheses) are clustered at age in years.

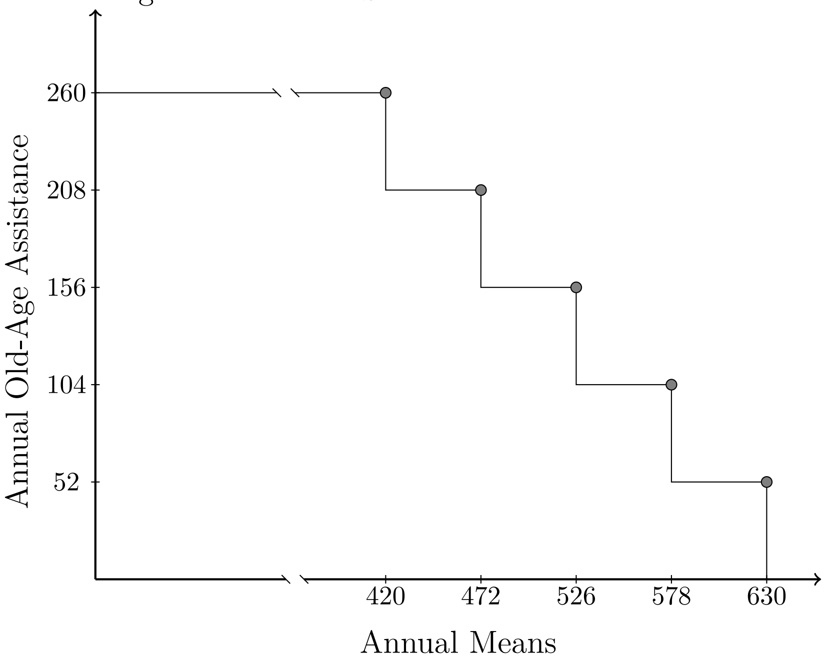
Table 13: Alternative Specifications of the Baseline RD - Bandwidth Choice, Local Polynomials, and Bias-Corrected Confidence Bands

	Parametric Linear	Parametric Quadratic	Local Non-Parametric: Conventional	Local Non-Parametric: Bias-Corrected
Bandwidth (Age-Years)				
2	-0.048** (0.000)	-0.036** (0.000)	–	–
3	-0.056** (0.004)	-0.038** (0.000)	–	–
4	-0.058** (0.005)	-0.053** (0.005)	-0.054** (0.003)	-0.038** (0.001)
5	-0.060** (0.006)	-0.053** (0.003)	-0.056** (0.004)	-0.050** (0.005)
6	-0.063** (0.007)	-0.055** (0.003)	-0.058** (0.005)	-0.052** (0.003)
7	-0.064** (0.008)	-0.056** (0.004)	-0.059** (0.006)	-0.053** (0.003)
8	-0.066** (0.008)	-0.058** (0.004)	-0.061** (0.007)	-0.055** (0.003)
9	-0.070** (0.009)	-0.055** (0.005)	-0.062** (0.007)	-0.056** (0.003)
10	-0.072** (0.009)	-0.057** (0.005)	-0.064** (0.008)	-0.054** (0.003)
20	-0.110** (0.012)	-0.061** (0.007)	-0.090** (0.011)	-0.055** (0.007)

Source: UK Census wave 1911 and IPUMS. *Note:* RD estimates of the labor force participation rate (dependent variable) using an indicator for the age cutoff (= 0 if age < 70; = 1 if age >= 70). OLS estimates use uniform weighting of all observations. Local non-parametric estimates use triangular kernel weighting, putting more weight on observations closer to the cutoff. Bias-corrected estimates use the bias correction proposed by Calonico et al. (2014a,b) including robust standard errors and triangular kernel weighting. **, * denotes significance at the 1% and 5% level respectively. Standard errors (in parentheses) are clustered at age in years, unless otherwise specified.

Figures

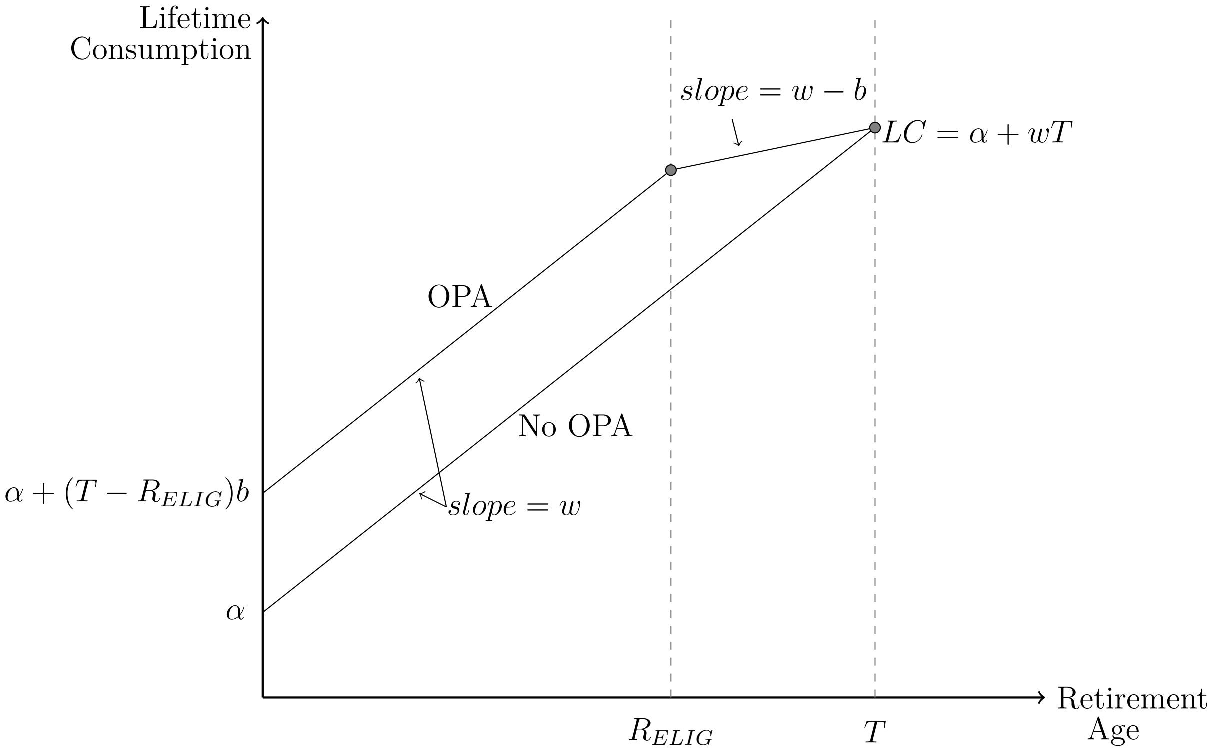
Figure 1: Pension Schedule and Means Test



Source: Own graph.

Note: The figure depicts critical values of the means test. 20 Shillings = 1 Pound. The maximum pension of 260 Shillings per year (5 Shillings per week) was granted to individuals with annual means of no more than 420 Shillings (36% of the average annual wage in 1911). Individuals with annual means of more than 630 Shillings (54% of the average annual wage in 1911) did not qualify for the pension.

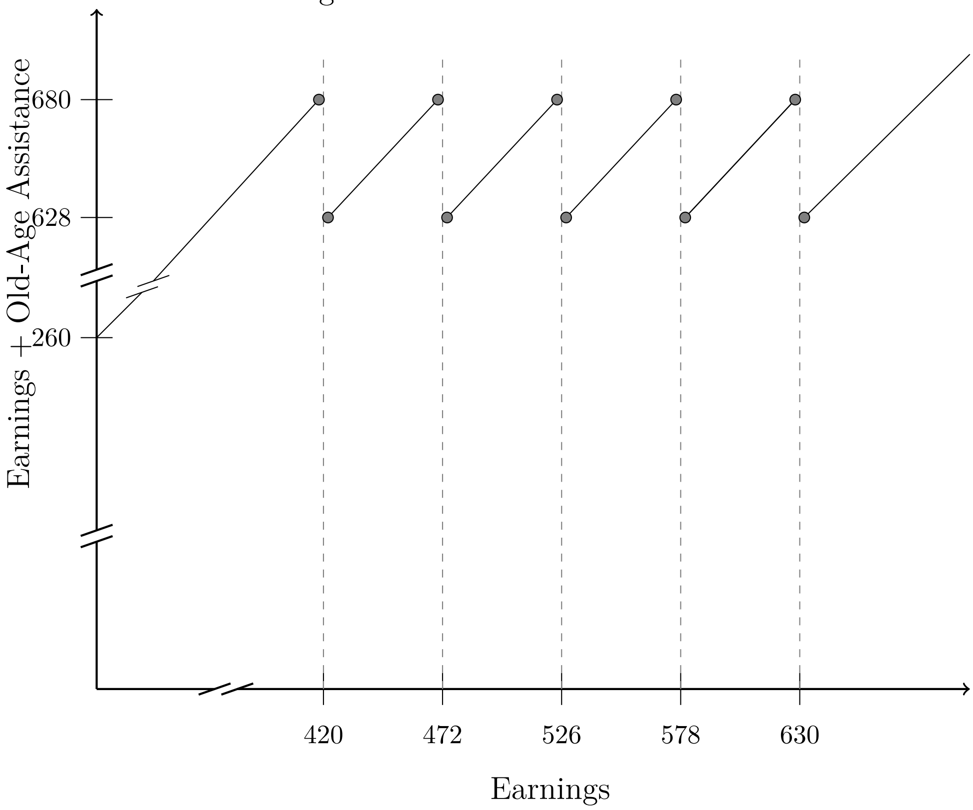
Figure 2: Lifetime Budget Constraint



Source: Own graph.

Note: The figure relates the present value of lifetime consumption to the retirement age.

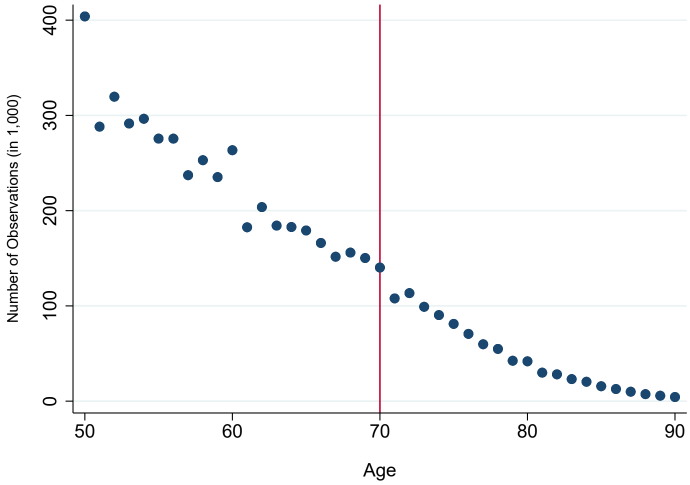
Figure 3: Pension Incentives



Source: Own graph.

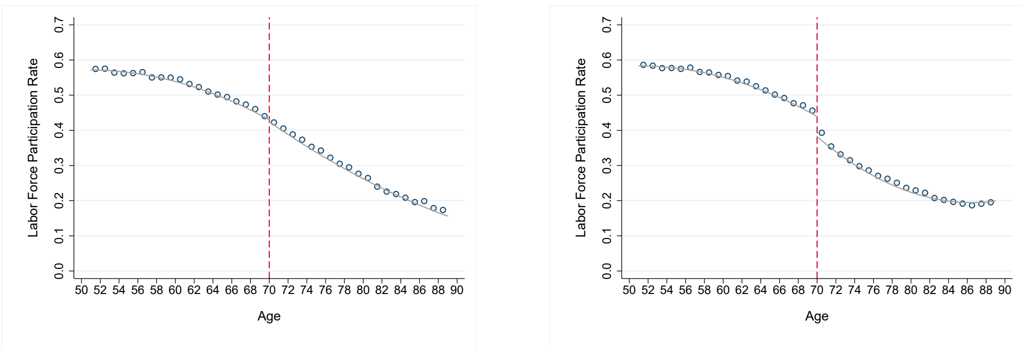
Note: The figure depicts the piece-wise constant regions of pension benefits by relating earnings to the total of earnings and old-age assistance.

Figure 4: Number of Observations by Age (Census 1911)



Source: Own calculations based on UK Census (wave 1911) and IPUMS. *Note:* The vertical line indicates the age-based eligibility threshold at age 70 that was introduced by the OPA in 1909.

Figure 5: Labor Force Participation by Age

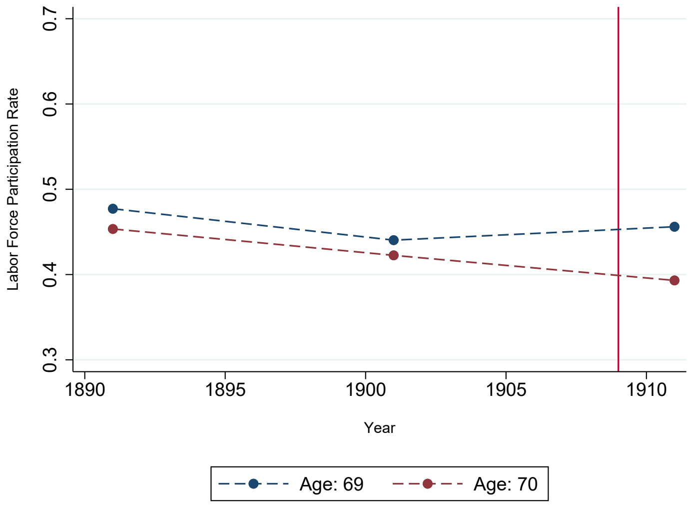


(a) Census 1901

(b) Census 1911

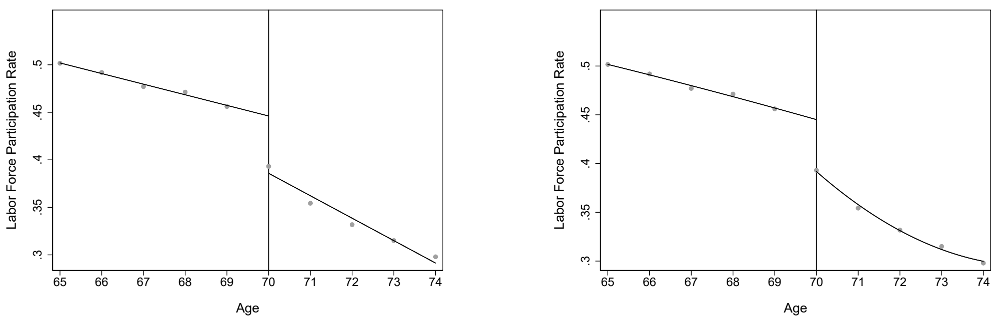
Source: Own calculations based on UK Census (waves 1901 and 1911) and IPUMS. *Note:* The vertical line indicates the age-based eligibility threshold at age 70 that was introduced by the OPA in 1909.

Figure 6: Labor Force Participation Over Time by Age-Based Eligibility



Source: Own calculations based on UK Census (waves 1891, 1901 and 1911) and IPUMS. *Note:* The vertical line indicates the introduction of old-age assistance by the OPA in 1909.

Figure 7: Functional Form of Labor Force Participation and Age Around the Cutoff

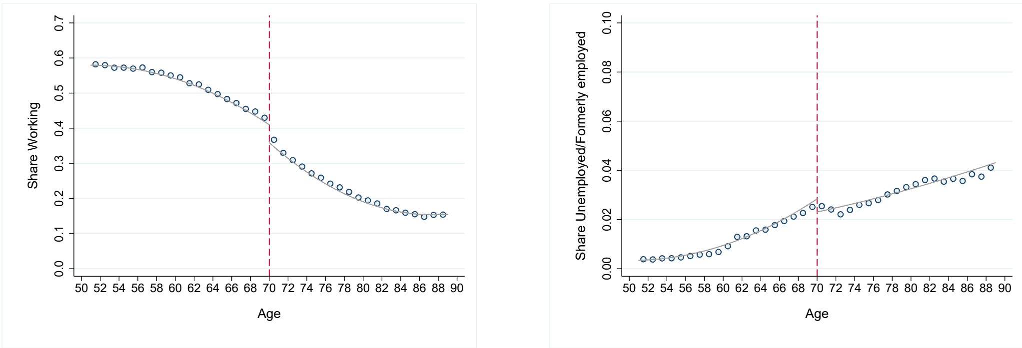


(a) Linear

(b) Quadratic

Source: Own calculations based on UK Census (wave 1911) and IPUMS. *Note:* The vertical line indicates the age-based eligibility threshold at age 70 that was introduced by the OPA in 1909.

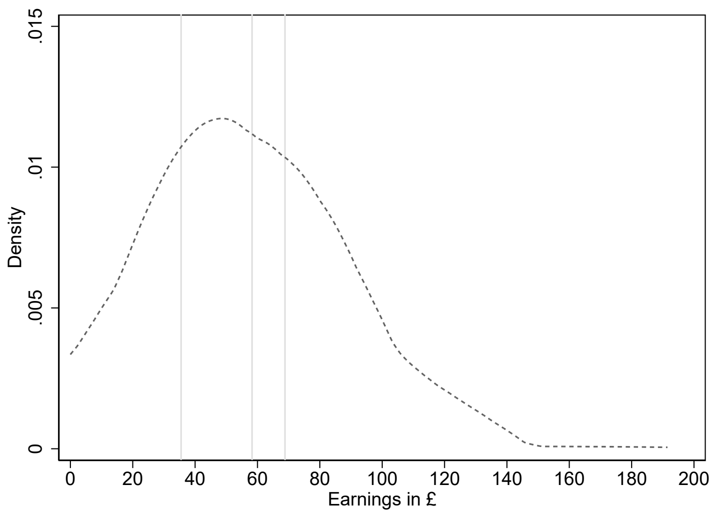
Figure 8: Work Status (1911)



(a) Actively Working (b) Unemployed

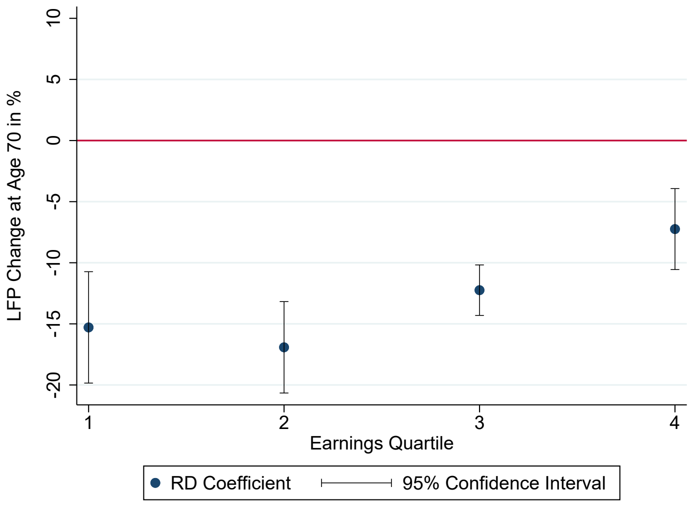
Source: Own calculations based on UK Census (wave 1911) and IPUMS. *Note:* The vertical line indicates the age-based eligibility threshold at age 70 that was introduced by the OPA in 1909.

Figure 9: The Distribution of Earnings in 1911



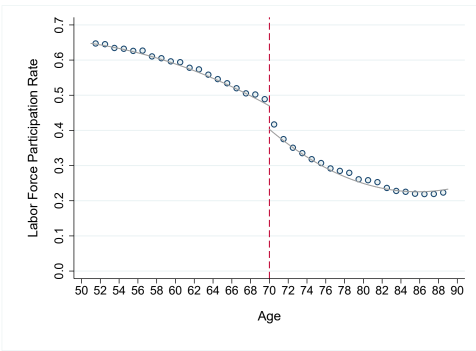
Source: Own calculations based on UK Census (wave 1911) and IPUMS. *Note:* The graph reports the earnings distribution as approximated from 3-digit occupational scores (for details, see appendix C). Vertical lines indicate the earnings quartiles.

Figure 10: RD Estimates Across the Earnings Distribution

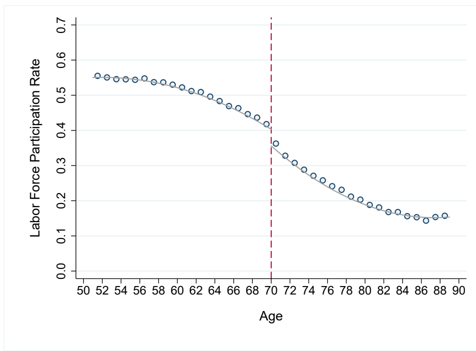


Source: Own calculations based on UK Census (wave 1911) and IPUMS. *Note:* RD estimates of LFP by earnings quartile in percent.

Figure 11: Labor Force Participation by Age and Children (1911)



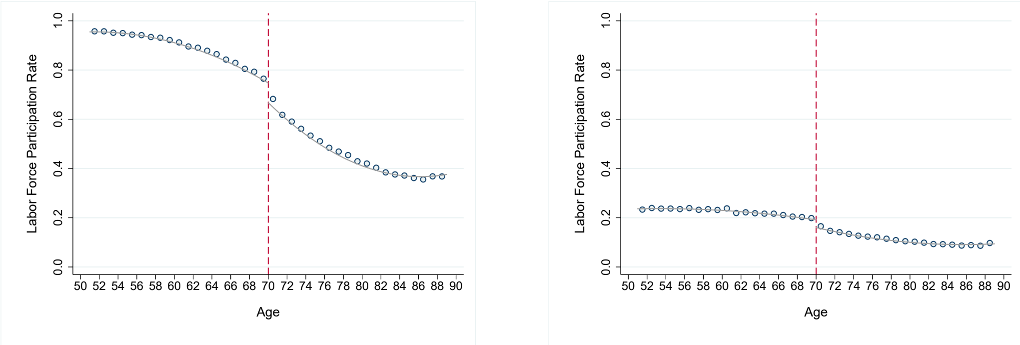
(a) Without Own Children



(b) With Own Children

Source: Own calculations based on UK Census (wave 1911) and IPUMS. *Note:* Reported values are labor force participation rates over age, separately for individuals without own children in the household (panel a) and with own children in the household (panel b). The vertical line indicates the age-based eligibility threshold at age 70 that was introduced by the OPA in 1909.

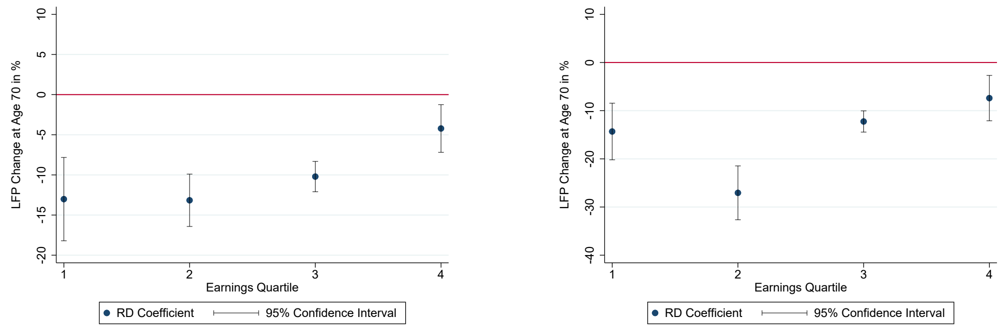
Figure 12: Labor Force Participation by Gender (1911)



(a) Men (b) Women

Source: Own calculations based on UK Census (wave 1911) and IPUMS. *Note:* Reported values are labor force participation rates over age, separately for men (panel a) and women (panel b). The vertical line indicates the age-based eligibility threshold at age 70 that was introduced by the OPA in 1909.

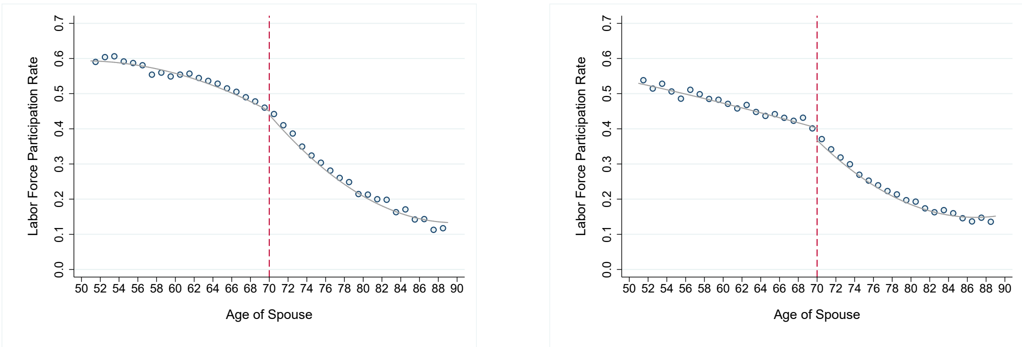
Figure 13: RD Estimates Across the Earnings Distribution by Gender



(a) Men (b) Women

Source: Own calculations based on UK Census (wave 1911) and IPUMS. *Note:* RD estimates of LFP by earnings quartile in percent, separately for men and women.

Figure 14: Labor Force Participation by Age of Spouse



(a) Census 1901 (b) Census 1911

Source: Own calculations based on UK Census (waves 1901 and 1911) and IPUMS. *Note:* LFP rates (vertical axis) are conditional on having reached the eligibility age (70) and are plotted by age of the respective spouse (horizontal axis). The vertical line indicates the age-based eligibility threshold at age 70 that was introduced by the OPA in 1909.

A Additional Tables

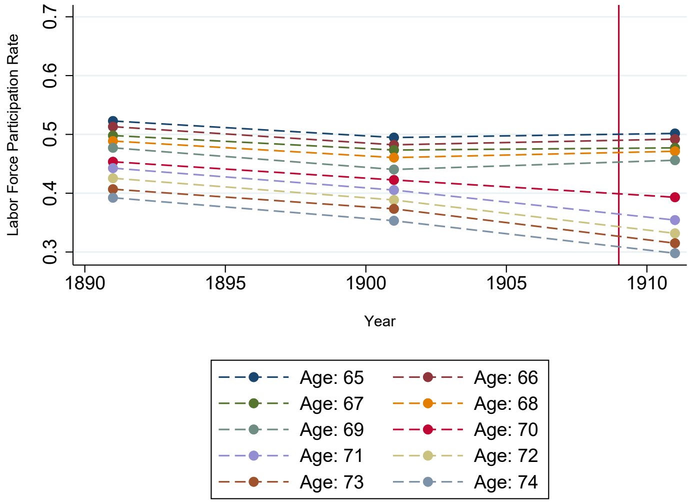
Table A.1: Bounding the Willingness to Pay for the OPA

Assumed Recipience Rate 75+	Residual Recipience Rate 70 - 74	Share Marginal Individuals	Share Infra-Marginal Individuals	Minimum Average Value of £1 from OPA
100%	17.5%	34%	66%	£0.66
90%	26.9%	22%	78%	£0.78
75%	41.0%	15%	85%	£0.85
57.5% (average)	57.5% (average)	10%	90%	£0.90

Source: Own calculation. *Note:* The table bounds the willingness to pay for the OPA. Reported values refer to four alternative scenarios, based on assumptions about the recipience rate at age 75+ (out of the known total of 613,873 recipients in 1911). Assumed recipience rates yield the respective residual recipience rates at age 70 - 74 and the corresponding shares of marginal and infra-marginal individuals.

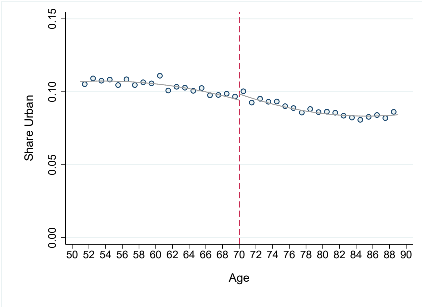
B Additional Figures

Figure B.1: Labor Force Participation Over Time by Age-Based Eligibility: More Ages

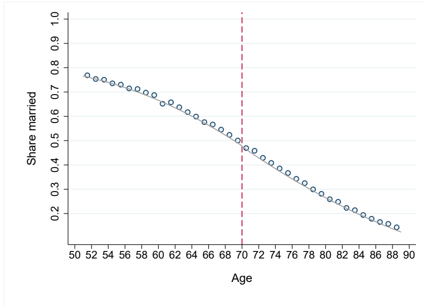


Source: Own calculations based on UK Census (waves 1891, 1901 and 1911) and IPUMS. *Note:* The vertical line indicates the introduction of old-age assistance by the OPA in 1909.

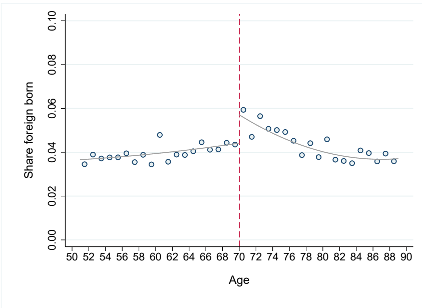
Figure B.2: Continuity of Observable Characteristics at Age Cutoff



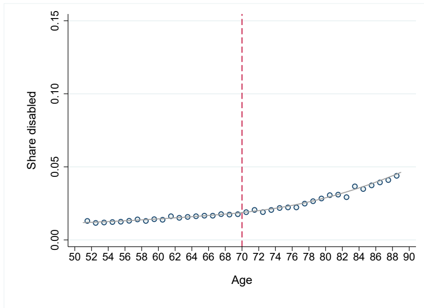
(a) Urban



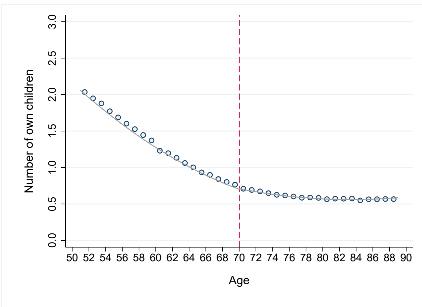
(b) Married



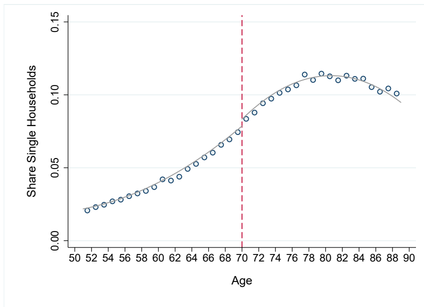
(c) Foreign born



(d) Disabled



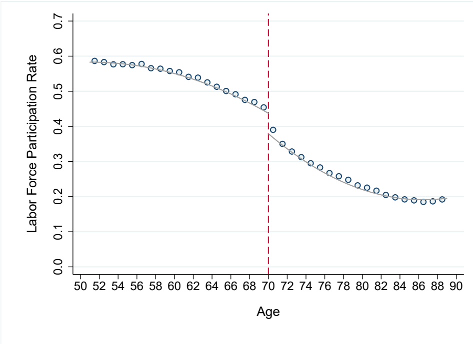
(e) Own Children in Household



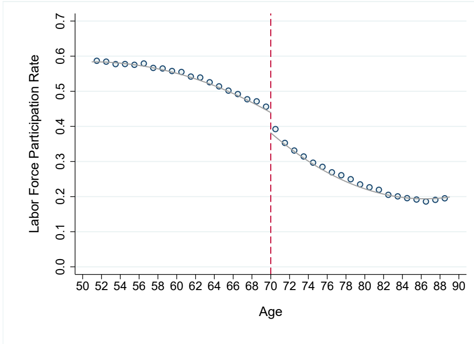
(f) Share Single Households

Source: Own calculations based on UK Census (wave 1911) and IPUMS. *Note:* The vertical line indicates the age-based eligibility threshold at age 70 that was introduced by the OPA in 1909.

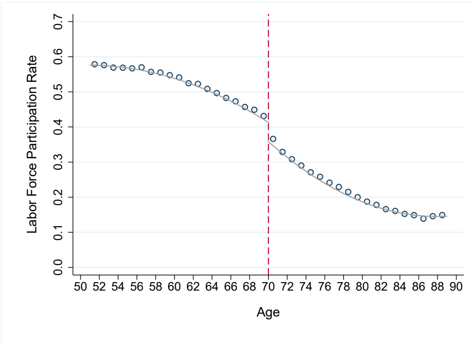
Figure B.3: Further Sensitivity Checks: Labor Force Participation by Sub-Samples



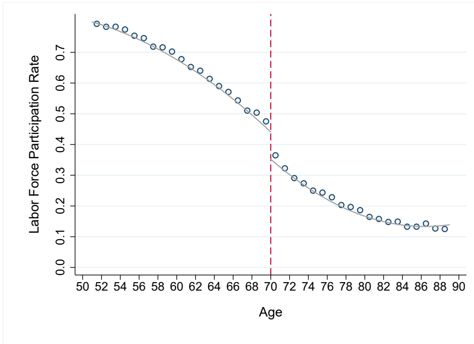
(a) Native-Born



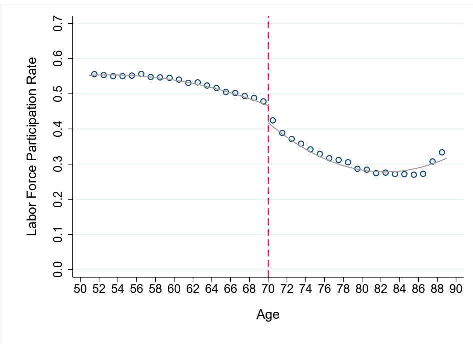
(b) No Disabled



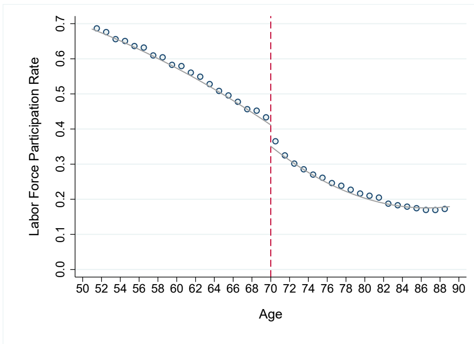
(c) Large HH (>= 12 Persons)



(d) Single HH



(e) Married



(f) Non-Married

Source: Own calculations based on UK Census (wave 1911) and IPUMS. *Note:* The vertical line indicates the age-based eligibility threshold at age 70 that was introduced by the OPA in 1909.

C Approximating Individual Labor Earnings

Earnings are neither reported directly in the UK census of 1911, 1901 nor 1891. Therefore, we approximate earnings based on occupational status as has been done in the previous literature (e.g. Abramitzky et al., 2014). For this purpose, we match the occupational income score (occscore) that is generated from the three-digit occupational level (occ1950) from IPUMS for the U.S. census in 1950 to the UK census in 1911 (see details below). We use the U.S. 1950 census because individual level census data that include earnings are not available for the UK before 1991. Moreover, comprehensive earnings information for the U.S. has not been collected before 1950.

Based on this procedure we can match an earnings score to 97% of individuals who state an occupation in the UK census of 1911. The implicit assumption behind this procedure is that the earnings distribution, conditional on occupation, is similar in the UK in 1911 and the U.S. in 1950. Even if this assumption is partly violated, our results would remain valid since we are only interested in a broad approximation of the earnings distribution rather than exact earnings levels.

C.1 Further Details on Matching the Income Score from the U.S. 1950 Census to the UK 1911 Census

The occupational classification in IMPUS differs between the U.S. and UK census. In particular, the three-digit occupational system used in the U.S. (1950 Census Bureau Occupational Classification system: occ1950) does not directly match with the five-digit occupational level used in the UK (Historical International Standard Classification of Occupations HISCO: occhisco). To solve this problem, we make use of the fact that the U.S. census of 1880 from IPUMS includes both occupational coding systems (occ1950 and occhisco), which means that every individuals' occupation is coded both in the occ1950 and occhisco coding scheme. We thus match the U.S. census from 1880 that also includes the occupational income score (occscore), derived from the incomes of the U.S. census 1950, to the UK census of 1911. In case that multiple occ1950 codes are matched to one occhisco code, we assign the average occscore for each occhisco code (except for "occupational title unclassifiable", "ambiguous responses", "other non-occupational response" and "no occupation and not in universe/not applicable").

To circumvent several adjustments with respect to aggregate price changes (deflating from 1950 to 1911), differential trends in GDP growth between the U.S. and the UK, and the overall conversion of U.S. dollars to British pounds, we normalize the mean value of earnings (measured in 1950 U.S. dollars) to match the UK mean earnings in 1911 that are documented in historical earnings data. This simple and transparent procedure preserves the ranks of the earnings distribution that is needed for the analysis (depicted in figure 9).

Occupation information is available for all individuals in the work force, but only for a subset of retirees who reported their former occupations. Restricting the sample to those who reported their occupation may provide flawed results due to selectivity.[42] To avoid selectivity and to be able to nevertheless use the universe of older workers for studying labor supply effects across the earnings distribution, we proceed as follows. First, we construct a dependent variable for each earnings quartile (obtained from occupation information, see above) that indicates whether the individual is in the labor force *and* in the specific earnings quartile (LFP = 1) or not (LFP = 0). The estimates obtained from this exercise represent the LFP decline, separately for each earnings quartile, relative to the total size of the respective age group. The advantage of this approach is that we can include all retired individuals irrespective of their reporting status because we only need occupational data for older workers who are still active. Since the interpretation of the estimates changes, figures 10 and 13 report the relative activity decline for each earnings quartile, dividing the estimate by the percentage of individuals in the labor force in the respective quartile (LFP = 1) at the age of 69.

[42]Consider the extreme case where only individuals with prestigious occupations report their original occupation. In this case, we would observe no decline in LFP by age for non-prestigious occupations because men who retire from these occupations simply drop from the universe.

D Computing the OPA Recipience Rate

The UK census does not directly report information on pension recipients. However, we do have information on the total number of recipients in 1911 (Department of Labour Statistics, 1915) which allows us to compute a lower bound of the recipience rate in the local environment of the age cutoff (age 70 to 74). In particular, we know that the total number of recipients in 1911 was 613,873, while 551,100 people were 70 to 74 years old and 517,386 were 75 or older. Assuming that every individual 75 and older received the pension, there would be 96,487 pensions left for the age group 70 to 74 (613,873 - 517,386 = 96,487). Relating this remaining number of pensions to the population in the age range 70 to 74 thus yields the lower bound for the recipience rate of 17.5% (96,487/551,100).

Table A.1 provides further scenarios based on different assumptions of the recipience rate for people 75 and older. While assuming that *every* individual above age 74 receives the pension yields the minimum recipience rate, it is more realistic to assume that less than 100% of the population aged 75 or older actually received the pension. Assuming that only 90% (75%) of individuals above age 74 receive the pension, the "residual recipience rate" in the age bracket 70 - 74 rises to 26.9% (41%) respectively. This also implies lower (higher) shares of marginal (infra-marginal) individuals. The lower bounds for valuing one pound from the OPA are then much closer to one (78% and 85% respectively). When assuming the average recipience rate of 57.5% uniformly at all ages (613,873 total recipients / 1,068,486 individuals alive above age 70 = 0.575), one pound from the OPA is valued almost fully (90%).