Large-scale social transfer and labor market outcomes: The case of the South African pension program

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Selected Paper prepared for presentation at the 2016 Agricultural & Applied Economics Association Annual Meeting, Boston, Massachusetts, July 31–August 2

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July 2016

Abstract

This paper assesses the effects of the South African old age pension program, the largest cash transfer program in the country, on labor force participation and employment of the elderly and prime-aged individuals. During 2008–2010, the minimum eligible age for males gradually decreased from 65 to 60. Exploiting this change as a natural experiment, the paper finds that the pension significantly discouraged the elderly to work. The intention-to-treat effects estimated using three different independent datasets imply that the labor force participation rate of men aged 60–64 significantly decreased by 5.8, 11.3 and 8.9 percentage points, depending on the datasets used. Correspondingly, the probability of being employed decreased by 4.1–11.8 percentage points. The local average treatment effects estimated suggest that once the elderly started to receive the benefit, the probability of participating in the labor force and being employed decreased by 29.4 and 31.6 percentage points respectively although these estimates are not statistically significant. The estimation of the effects for prime-aged individuals, on the other hand, is in progress.

JEL classification: H55, I38, J08, J21, J26

Keywords: Social transfers, Pension, Labor supply, Retirement, South Africa

1 Introduction

Social transfer programs in low- and middle-income countries have been increasing. According to World Bank (2015), there are about 20 social safety net programs in an average

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developing country, and among various types of safety net programs, cash transfers are particularly becoming more prevalent. In Africa, for example, 40 countries, out of 48, offered unconditional cash transfers in 2014. While transfer programs have been proved to contribute to poverty reductions (Fiszbein and Schady, 2009; Haushofer and Shapiro, 2016), it has often been said that these transfers may discourage work. For better policy designing, assessment of this unintended disincentive effect is required. Yet, there are few rigorous studies such as Alzúa et al. (2013) and Banerjee et al. (2015), and reliable quantification of the effect is lacking.

In this paper, I evaluate effects of the South African old age pension program on labor supply and employment outcomes of pensioners and their (grand)children. The pension program is the largest social assistance program in the country in terms of budget size, and serves nearly three million people. Also, its benefit is remarkably generous. The amount of the benefit is almost double the national median income and sufficient enough to support both recipients and their children. Besides, since it is noncontributory, and its eligibility depends primarily on ages, the pension program is virtually a universal cash transfer. Thus, the pension program is a good case to assess effects of large-scale social transfers on labor supply.

In the context of South Africa, too, quantifying the disincentive effects of the pension program is required from the following two viewpoints. First, good quantification of them leads to accurate assessment of its fiscal sustainability. Since the end of Apartheid, the country has developed generous social grant programs, and the government spending on social safety net programs amounts to 3.51% of the GDP, higher than the average of the OECD countries, 2.92% (World Bank, 2015). Its fiscal sustainability, however, is not certain (Woolard and Leibbrandt, 2010). Second, it helps understand causes of the persistently high unemployment rate in the country, which has long been standing around 20%. Although the literature has attempted to reveal its causes (Banerjee et al., 2008; Rodrik, 2008), the causes still remain unclear. It may be the case that generous large-scale social grant programs cause it by allowing dependency of people on the programs.

To estimate effects of the pension program, I exploit a change in the eligible ages for males as a natural experiment. The male threshold age had been 65 until 2008 but was gradually reduced to 60 during the period 2008 and 2010. Thus, men aged 60–64 were eligible after the change while the same-age men prior to it had not. My estimation strategy, therefore, is to compare outcome variables of men aged 60–64 before and after it. To the best of my knowledge, this paper is the first study using this policy change.

Empirical results using three different independent datasets show that the pension program has a significant disincentive effect for elderly men. The estimates of the intention-to-treat effect suggest that because of the pension, the labor force participation rate of 60–64 years old men decreased by 5.77%, 11.29%, and 8.90% points, depending on the dataset used. The local average treatment effect on the labor force participation rate is estimated at 29.35% points although this point estimate is not significant. On the other hand, the estimation of the effects on prime-aged labor market outcomes is still a work in progress, and I do not have conclusive results yet.

I contribute to three strands of the literature. First, I add to the literature about effects of unconditional cash transfer on labor supply and employment. Although the literature on unconditional cash transfer is extensive and still growing (Baird et al., 2011; Case et al., 2005; Edmonds and Schady, 2012; Handa et al., 2015), there are a limited number of papers examining its effects on labor market outcomes.

Second, I contribute to the literature on pensions and elderly retirement. I identify and quantify how much, if any, the elderly retirement rate is decreased by old age pensions. This area of the literature is large and growing, and past papers have confirmed that pension programs increase retirement rates (Gruber and Wise, 2004). Yet, the literature as well as policy makers need good quantification of the effects because better quantification enables more accurate assessment of fiscal burden and overall impacts on labor market and economy. Since pension programs are, by their nature, usually not embedded with exogenous variations exploitable for rigorous estimation and hard to run experiments, the literature has had a relatively small number of papers using exogenous variations compared with those relying on descriptive analyses (Krueger and Meyer, 2002) although the number is recently growing. Besides, the magnitude of the effects is expected to vary between countries, we need evidence from different countries. Most previous papers study pensions in developed countries (Krueger and Pischke, 1992; Staubli and Zweimüller, 2013; Fetter and Lockwood, 2016), and few examine ones in developing countries. Among few papers studying pensions in low- and middle-income countries are Danzer (2013) for Ukraine, Kaushal (2014) for India, Juarez and Pfutze (2015) for Mexico, and de Carvalho (2008) for Brazil. There is no rigorous study that estimate effects of the South African pension program, exploiting exogenous variations.

Third, I contribute to the whole literature on the South African pension program. The literature has explored its effects on various outcomes such as labor supply of the elderly (Ranchhod, 2006), employment of the prime-aged (Bertrand et al., 2003; Posel et al., 2006; Ardington et al., 2009), education (Edmonds, 2006), health and nutrition (Duflo, 2003; Gordon and Miller, 2012), and household composition (Edmonds et al., 2005; Hamoudi and Thomas, 2014). In a descriptive manner using labor force survey data in 2000–2001, Ranchhod (2006) shows that the pensions decrease the labor force participation and employment of the elderly. Regarding the prime-aged labor supply and employment, the existing evidence is mixed. Bertrand et al. (2003) find that the pensions, particularly the pensions paid to women rather than men, decrease the labor force participation rate and hours worked of prime-aged individuals residing with pensioners. Posel et al. (2006) and Ardington et al. (2009), on the other hand, find that the pensions increased labor supply of the prime-aged by leading them to migrating for work.

My contribution to this literature on the South African pension is that I provide the first, or second, quasi-experimental evidence. None of the previous papers relies on quasi-experiments or other exogenous variations except Duflo (2003), who exploits an expansion of the pension program to Black Africans in early 1990s. All the other empirical strategies are essentially to compare outcome variables between elderly individuals who have reached eligible ages for the pensions and relatively young elderly who are not yet eligible, or to compare between households with and without pensionable-age individuals. A critical concern in this approach is that these two groups may not be comparable. In other words, it is hard to distinguish effects of the pensions from effects of being old. For example, households with persons aged 60 and over would be different from households without such persons even if the pension program was not in place. Most of the past studies, however, do not control for ages of elderly people due to collinearity between ages and age-eligibility. There are few papers that run regression discontinuity estimation across the eligible age thresholds (Edmonds et al., 2005; Gordon and Miller, 2012). Although they better control for the age effects by comparing between those slightly younger than the threshold age and those slightly older, their estimates still sorely rely on discontinuous differences across the thresholds, which may not reflect effects of the pension. Given that the minimum eligible age had been 65 for males and 60 for females for a long time and that these ages could be also important in other institutions, the discontinuity, if any, could represent something else other than effects of the pension program. My empirical strategy, by contrast, overcomes this limitation. Relying on the natural experimental variation, I am able to throughly control for age-fixed effects and hence disentangle the effects of the pensions from any other effects that are correlated with ages.

2 The old-age pension program

The current scheme of the old-age pension program has been established since early 1990s. It has four remarkable features. First, the benefit is substantial. The maximum amount is 1080 Rand per month ($140 or $230 PPP adjusted) in 2010, which is approximately 75% higher than the median per capita income. The amount has been adjusted to inflation, so the real value has been basically the same over the last two decades.

Second, the maximum amount is paid almost universally in practice regardless of recipients' income and assets although the amount is supposed to vary with their income and assets.

Third, the pension is noncontributory and has generous eligibility criteria. The eligibility depends on age and a means-test. It is not conditional on being retired. The means-test is based on income and assets of applicants and their spouses, but not their children. In 2015, if applicants are single, annual income and the assets value of applicants should be not more than 64,680 Rand (approximately 5,000 USD) and 930,600 Rand (72,000 USD) respectively. If applicants are married, the double cutoff values apply for the combined income and assets of applicants and spouses. These cutoffs are so high for Black Africans that most of them pass it.[1,2] Thus, for blacks, the eligibility is determined primarily by age. Consequently, most of age-eligible Africans receive the pension grant. The pension recipient rate is extensively presented in the section of data.

Fourth, there was a policy change that reduced the male minimum eligible age from 65 to 60 during 2008–2010 whereas the female age remained the same at 60. Until March 2008 the male eligible ages had been 65 and above, but the threshold age decreased to 63 in April 2008, to 61 in April 2009, and to 60 in April 2010 according to the Social Assistance Amendment Bill dated on April 22, 2008. The means-test for the newly included group was identical to that for the other group. This change actually increased the recipient rate among men aged 60–64 as the following section shows. I exploit the change for my identification strategy.

The second and third features ease typical econometric issues. Because the fixed amount is almost universally paid to age-eligible blacks, the pension is virtually unconditional cash transfers. Hence, the selection issue does not arise. Also, pensioners have no incentive to adjust their income, assets or household composition in order to become eligible or increase the amount of the benefit.

3 Data

I use three different independent groups of datasets. The first group is the Community Survey 2007 and a 10% sample of the Population Census 2011 in South Africa. They were collected in February-March 2007 and October-December 2011, respectively. The community survey is a nationally representative, large-scale household survey, and the types of information collected is very similar to those of the census. Since both of the two data include fairly

[1] The medium income of all the *employed* black is 26,004 Rand in 2010 while that of the employed white is 114,000 Rand. Note that these medians are for those employed. Given the high unemployment rate, the median including the unemployed is much smaller.

[2] The assets are not examined in practice due to difficulty in the valuation of assets (McEwen et al., 2009).

detailed information of employment, I am able to construct variables of employment status. An advantage of using the datasets is their large sample sizes.

The second dataset is the Quarterly Labor Force Survey (QLFS) in 2008–2015.[3] Compared with the first datasets, the QLFS includes more detailed information of employment such as hours worked. Another advantage of the data is that it is available in every year between 2008 and 2015. Since the male threshold age was reduced sequentially in 2008–2010, I can examine how effects of the pension emerged over the period.

The last dataset is the National Income Dynamics Study (NIDS) in wave 1 (2008), wave 2 (2010–11), and wave 3 (2012). They are longitudinal. Since they tracked respondents across the country. Thus, the attrition is considered to not seriously correlated with migration. The NIDS has a distinctive feature. They asked information about "non-resident household" members. In most surveys including the Community Survey 2007, the Census 2011, and the QLFS, a household member is defined to be an individual who stays in a household for at least four nights per week on average during reference weeks and shares meals and other essentials for living. Thus, seasonal migrants are not classified as household members, and their information is not collected. On the other hand, the NIDS uses more inclusive definition of a household. It regards individuals as household members if they lived in households for at least 15 days during the last 12 months no matter whether they stayed in reference weeks. Then, it further classifies household members into resident household members and non-resident ones. That is, household members who satisfy the standard residence condition, i.e., staying for at least four nights per week on average, are resident household members, and the other household members are non-resident. (Note that the definition of a *resident* household member is identical to that of a household member in usual surveys.) Since the NIDS collected information of both resident and non-resident members, I am able to conduct analysis on non-resident members, who are missing in the other datasets. To estimate effects of the pension especially in the context of South Africa, it is important to take into account non-resident household members because the pension may affect residency status as suggested by the past studies (Edmonds et al., 2005; Posel et al., 2006; Ardington et al., 2009; Hamoudi and Thomas, 2014). Ardington et al. (2009), for example, find that the pension induces labor migration. Thus, a resident household member who decides not to migrate may be negatively selected in terms of employability, so estimation relying only on a dataset consisting of resident household members may be confounded.[4] This possible confounder,

[3]Slightly different quarterly labor force surveys before 2008 are available although the current version of the paper does not use them.

[4]When effects on pensioners are estimated, migration is not an issue. This is because pensioners were surveyed in pre- and post-periods of the policy change no matter whether they were migrants or not. However, it is an issue when effects on prime-aged children of pensioners are estimated.

however, is able to be dealt with by the unique feature of the NIDS.

Importantly, not all the datasets have information of pension receipt status. The community census 2007 and the NIDS include it while the census 2011 does not. The QLFS has it only for those who are not employed. Thus, the effects I estimate in most analyses are intension-to-treat (ITT) of the pension for men aged 60–64. Since age-eligibility is a major policy instrument in pension programs, these ITT estimates are of general interest. I additionally estimate the local average treatment effects (LATE) using the NIDS, which include the information of pension receipt status of all observations.

Pension recipient rates

Using the Community Survey 2007, the upper panel in figure 1 presents pension recipient rates by races in 2007, when the threshold age was 65 for men and 60 for women. The rates among men are zero at ages 64 and below for all races, but they suddenly rise at age 65 particularly for Black Africans and Others. At age 68 and above, about 85% of African males are recipients while only about 22% of whites are. For women, a similar pattern is found. These results suggest that the pension is important particularly for elderly Africans whereas it is not so for whites. Also suggested is that the age eligibility primarily determines pension receipt status of Africans.

The same facts are found in the lower panel of figure 1. The panel presents recipient rates of all races in 2008 using the NIDS. Although it is noisy,[5] it shows that the rate is high among age-eligible Africans and Others but low among whites. Also, shown is that most Africans receive the pension once they become age-eligible. These facts support that pension receipt status is largely determined by the age eligibility.

Using the NIDS, figure 2 compares the recipient rates among Africans before and after the policy change. The figure implies that the reductions in the male eligible ages indeed increased the recipient rate among the affected age group. According to the upper panel, which compares the rates between 2008 and 2010–2012, there are more recipients in 2010–12 than in 2008 among men aged 60–64, who had not been age-eligible until March 2008 but were age-eligible in 2010–2012.[6,7] By contrast, the recipient rates among women are similar

[5]Note also that the recipient rates at ages under the thresholds are not necessarily zero. Suspected is that this is due to measurement errors. In the same data, pension recipients are found even among the aged 20s and 30s.

[6]To be precise, men aged 60 was not age-eligible in January–March 2010 because the threshold decreased from 61 to 60 in April 2010. In the datasets used, however, all the observations in 2010 were interviewed in May 2010 and afterward. Hence, all of men aged 60 interviewed in 2010 were age-eligible.

[7]At ages 64 and 65, the recipient rate in 2010–12 is fairly comparable to, and not higher than, that in 2008. This may be due to the fact that some observations in the 2008 data were interviewed after April 2008 so that they were already age-eligible as of the interview dates.

between 2008 and 2010–12 at all ages including 60–64. The same patterns are found in the lower panel, which compares the pension recipient rates between 2008, 2010–2011, and 2012. An interesting observation in it is that the recipient rate among men aged 60–64 gradually increased during the period between 2008 and 2010–11 and between 2010–11 and 2012. This may indicate that it took time for elderly people to respond to the reductions in the eligible ages and also that the application process took time.

The QLFS also shows that the recipient rate among men aged 60–64 increased after the policy change. Figure 3 presents recipient rates of pension recipients among *not-employed* persons. Note that interpretation of the figure requires careful attention because the changes in the rates shown may reflect changes in the likelihood of being not-employed rather than changes in that of being a pension recipient. According to the upper panel, which shows the rates in the pre-priod (2008) and the post-period (2011–2015), the rate among men aged 60–64 is higher in 2011-15 by about 30% points than in 2008 whereas that among women look similar between years.[8] The lower panel, which shows the rates separately in 2008, 2009, 2010, 2011 and 2015, gives more interesting results. In the period from 2008 Q1 to 2009 Q1, during which the male minimum eligible age decreased from 65 to 63, the recipient rates increased among men aged 63 and 64 but did not among the other age group. In the period from 2009 Q1 to 2010 Q1, during which the threshold decreased from 63 to 61, the rates increased among men aged 61 and 62 but not among the others. From 2010 Q1 to 2011 Q1, during which the minimum age decreased from 61 to 60, the rates increased particularly among men aged 60. Finally, from 2011 Q1 to 2015 Q1, during which the minimum age did not change, there is no clear increase in the rates. All of these findings imply that people actually responded to the policy change.

Labor market outcomes

Using the three groups of datasets, figure 4 demonstrates the labor force participation rates among Africans before and after the reductions in the male minimum eligible ages.[9] On the whole, the rates among men aged 60–64 are lower in the post-period than in pre-period whereas those among men at the other ages do not look systematically different between the the periods. By contrast, however, the rates among women are similar between the periods at all ages including 60–64. These results may imply that the reductions in the eligible ages induced retirement.

[8]The reason that the proportion of pension recipients under the eligible ages is relatively high is that the proportion also includes disability grant recipients. See the note in the figure.

[9]Note that the labor force participation rates in the three panels are shown to be comparable to each other in the figure despite that the datasets are different. This comparability is also found in figures 5 and 6 and suggests reliability of the datasets to some extent.

Similar results are found in the proportions of the employed among Africans (figure 5).[10] The proportions for men at ages 60–64 are higher in the post-period than in the pre-period, but those at the other ages are not so different between the periods. For women, on the other hand, the proportions are similar between the two periods at all ages. These results support the supposition that the expansion of the pension eligibility increased retirement.

Lastly, figure 6 displays the proportions of the unemployed among Africans. The differences in the proportions between the pre- and post-periods across the ages are not as clear as is the case for the labor force participation rates and the proportions of the employed. Nevertheless, the figure seems to suggest that the proportions of the unemployed among men aged 60–64 decreased. In the upper panel, which uses the community survey 2007 and the population census 2011, the proportions among men aged 60–64 do not differ between 2007 and 2011 whereas the corresponding proportions among women increase in 2007–2011. This might be that although there was an upward underlying trend in unemployment among elderly men and women, the upward trend among men aged 60–64 was suppressed by the expansion in the eligibility because the expansion induced unemployed men aged 60–64 to leave the labor force. In the middle panel, which is based on the QLFS, the proportions of the unemployed indeed decrease among men aged 60–64 while it does not change among men at the other ages and among women. The bottom panel, which uses the NIDS, is not very informative due to its noisiness.

Note that the decrease in unemployment among men aged 60–64, if any, looks smaller in absolute terms than the decrease in employment. For example, in the middle panels in figures 5 and 6, both of which are based on the QLFS, the decrease in unemployment is at most 5% points while that in employment is roughly 10–15%. Hence, the observed decrease in the labor force participation rates may be driven mainly by employed people quitting jobs and getting retired.

All the above descriptive information suggests that the reductions in the male eligible ages induced men aged 60–64 to retire. I more rigorously investigate these descriptive facts in the following sections.

4 Empirical strategy

The source of the exogenous variation I exploit is the reductions in the threshold age for men. The threshold had been 65 and above until March 2008, and was lowered to 63 in

[10] The proportion of the employed explained here is different from the so-called employment rate. The former is the proportion of the employed among *all people* while the latter is that of the employed among the *labor force*.

April 2008, to 61 in April 2009, and to 60 in April 2010. Thus, men aged 60 and 64 were not eligible for the benefit prior to the reductions but eligible after them. My empirical strategy is essentially to compare those "treated" age-groups with the other age-groups in a difference-in-difference manner and to interpret a difference-in-differences in means between ages and years as a causal effect of the pension.

A basic estimation equation is

$$y_{ijt} = \beta_j + \eta_t + \delta_j Post_t + \alpha' X_{ijt} + \epsilon_{ijt}, \tag{1}$$

where y_{ijt} is a labor market outcome variable of male i aged j in year t; β_j is the age-fixed effect; η_t is the year-fixed effect; $Post_t$ is the dummy indicating that year t is after the policy change; X_{ijt} is controls; δ_j is age j-specific time trend from the pre- to post-periods. The parameter of interest is δ_j. As I show in the following paragraphs, if the pension has an effect on y_{ijt}, then δ_j exhibits a certain pattern across j, and such a pattern tells about the effect.

To understand what δ_j represents, consider the following structural relationship:

$$y_{ijt} = \beta_j + \eta_t + f(\cdot) + X_{ijt} + \epsilon_{ijt}, \tag{2}$$

where the function $f(\cdot)$ is the effect of the pension on y_{ijt}. The overall goal of the paper is to specify, identify, and quantify $f(\cdot)$. It is reasonable to specify $f(\cdot)$ such that its arguments are the amount of the pension that individual i receives in year t, the total amount he has received by year t, the lifetime amount of the pension that he expects in year t, and individual characteristics, i.e., $f(CurrentPen_{ijt}, PastPen_{ijt}, E[LifePen]_{ijt}, W_{ijt})$.[11] This specification is quite flexible.

The function can be written as $f(CurrentPen_{ijt}, PastPen_{ijt}, E[LifePen]_{ijt}, W_{it}) = f(j, t, W_{ijt})$,[12] given that the monthly amount of the pension is identical to every recipient and that whether an individual receives the pension is largely determined by his age. In other words, $CurrentPen_{ijt}$, $PastPen_{ijt}$, and $E[LifePen]_{ijt}$ are functions of j and t. To understand this, let's take a look at individuals in 2007 and 2011. The amount that individuals currently receive ($CurrentPen_{ijt}$) is easy. (For now, the fixed annual amount of the pension is denoted R.) In 2007, it is zero for men aged at and below 64 and R for men at and over 65, i.e., $CurrentPen_{ij2007}$ is 0 for $j \leq 64$ and R for $j \geq 65$. In 2011, $CurrentPen_{ij2007}$ is 0 for $j \leq 59$

[11] As individual characteristics, I use W_{ijt} here instead of X_{ijt} to be explicit in that the characteristics in $f(\cdot)$, which determine the pension effect and thereby affect an labor market outcome, might be different from X_{ijt}, which directly affect an labor market outcome.

[12] I assume that the expected remaining lifetime is the same for everyone conditional on his age.

and R for $j \geq 60$. The shapes of $CurrentPen_{ij2007}$ and $CurrentPen_{ij2011}$ are illustrated in figure 7. Next, $PastPen_{ij2007}$ and $PastPen_{ij2011}$ (the total amount that had been paid as of 2007 and 2011, respectively) are proportional to how many months a male i aged j in 2007 and 2011 could have received the pension, i.e., how long he had been age-eligible.[13] This is because the monthly amount had been the same in the real value and because most black men actually received the pension once they became age-eligible. Their shapes are given in figure 8. Lastly, figure 9 depicts the shapes of $E[LifePen]_{ijt}$ for $t = 2007, 2011$, which are derived in a analogous reasoning. As shown in figures 7, 8 and 9, it is now clear that $CurrentPen_{ijt}$, $PastPen_{ijt}$, and $E[LifePen]_{ijt}$ are functions of j and t.

I now assume that $f(j, t, W_{ijt}) = f(j, t)$, i.e., $f(\cdot)$ does not depend on W_{ijt}, individual characteristics excluding age. In other words, there does not exist heterogeneity in effects of the pension between individuals. This assumption is probably not the case even though the pension amount is identical across all recipients regardless of their past employment histories and current earnings. In an extreme case, for example, the pension may have zero effect for a person who would not work anyway because he is seriously ill or has abundant income sources. Nonetheless, I impose this assumption for the sake of clarity in this section. Actually, the assumption does not threat my identification strategy unless uncontrolled characteristics across which the heterogeneity occurs systematically differ between years and between ages. This is because my empirical strategy is a difference-in-difference approach exploiting the policy change. Whether the assumption holds or not, it is able to identify average effects of the pension.

I explore three different plausible cases regarding how the pension affects an employment outcome, and their corresponding functional forms of $f(j, t)$.

<u>Case 1:</u> *The effect appears once an individual starts receiving the pension, and its magnitude is binary, i.e., depends only on whether he receives it or not.*

This case is plausible if an individual faces credit constraints and cannot borrow money against a future pension grant. Thus, the effect appears only after he reaches the eligible age. For men in 2007 and 2011, case 1 implies $f(j, 2007) = \gamma_0 \mathbf{1}(j \geq 65)$ and $f(j, 2011) = \gamma_0 \mathbf{1}(j \geq 60)$, where $\mathbf{1}(\cdot)$ is the indicator function. This function follows the same shape as in figure 7. In 2007, the effect is zero until age 64 and suddenly emerges at ages 65 and over with the same size. In 2011, it is zero until age 59 and appears at ages 60 and over with the

[13] For example, in the case of the 2007 community survey data and the 2011 population census, the data collection was in February–March 2007 and October 2011. Thus, as of the data collection, men aged 61 years and six months in March 2007 had never been eligible whereas their counterparts in October 2011 had been age-eligible for 18 months.

identical size. Then, the structural relationship becomes

$$y_{ij2007} = \beta_j + \eta_{2007} + \gamma_0 \mathbf{1}(j \geq 65) + X_{ij2007} + \epsilon_{ij2007}, \text{ and}$$
$$y_{i'j2011} = \beta_j + \eta_{2011} + \gamma_0 \mathbf{1}(j \geq 60) + X_{i'j2011} + \epsilon_{i'j2011}.$$

Looking at these equations, we now know that δ_j in the empirical equation 1, i.e., the age j-specific difference between pre- and post-periods, is constant at zero for $j = 55, \ldots, 59, 65, \ldots, 70$ and constant at a fixed value for $j = 60, \ldots, 64$ in case 1. δ_j's look like the lower panel in figure 10. The size of the difference in δ_j represents the effect of the pension.

Case 2: *The effect emerges once an individual starts receiving it. The effect consists of a binary part, i.e., whether he receives it, and a cumulative part, i.e., how long he has received it.*

In this case, the pension not only immediately affects an outcome variable in the binary manner but also cumulatively affects it. This is plausible if individuals only a current cash-flow of the pension but also savings of the pension matters to a labor supply decision. For example, while some individuals retire immediately after starting to receive the pension, others retire once a carry-over of the pension reaches an enough amount. The effect, therefore, initially emerges in a discontinuous manner and thereafter grows gradually. This profile of the effect resembles the upper panel in figure 11. The panel visualizes $f(j,t)$ for $t = 2007, 2011$.[14] The lower panel takes the difference between the two years. The coefficients δ_j's should have a similar shape to the lower panel. The effect of the pension is indicated in differences in δ_j between $j = 60, \ldots, 68$ and $j = 55, \ldots, 59, 69, 70$. The sizes of δ_j $j = 60, \ldots, 64$ imply the sum of the binary and cumulative parts, and those for $j = 65, \ldots, 68$ do the cumulative part.

Case 3: *The effect emerges even before an individual starts receiving it, and its magnitude depends only on the lifetime benefit he expects to receive.*

This case is different from the earlier ones. It is plausible if an individual is not bounded by credit constraints so that he is able to make forward-looking decisions by solving a lifetime utility maximization problem. The pension affects him only through affecting his lifetime budget. Thus, the effect does not depend on whether the pension is currently being paid or how long it has been paid. With the additional assumption that the effect is proportional to the lifetime benefit, the effect $f(j,t)$ for $t = 2007, 2011$ looks like figure 9. The empirical regression coefficients δ_j's represent the differences in the effect between years. They are flat

[14] To draw the figure, I additionally assume that the cumulative part is proportional to the length of the pension receipt period.

until age 60, decreasing at ages 61–68, and constant at zero from age 69.

In summary, the coefficients δ_j reflect effects of the pension, and their patterns across j tell which of the three cases is the most plausible. In the beginning of the empirical results section, I show patterns found and discuss the most plausible case.

A key assumption for δ to represent causal effects is that unobserved time variant characteristics do not systematically differ between ages, or in other words that the time trend in an outcome variable between pre- and post-periods would have been identical for males at each age of 55–70 even if the eligible ages had not been changed. I argue that this assumption is valid considering that confounding time trends that are consistent with the predicted patterns discussed above are not likely. In addition, I empirically examine this assumption by checking prior trends from 2001 to 2007.

5 Effects on elderly labor market outcomes

5.1 Between-periods difference: How the effects of the pension look

Pension recipient rates

To verify that the reductions in the male eligible ages increased pension recipients, I start by estimating equation 1 with the dependent variable being a dummy for receiving the pension. The estimation is run separately for men and women aged 55–70.

Figure 13 plots the coefficients δ_j, i.e., the difference in recipient rates at age j between the pre- and post-periods. In the both panels of figure 13, the coefficients for men stay around zero until age 59. At ages 60–64, they leap to about 0.4, and all coefficients except for that at age 61 in the lower panel are significantly different from zero. At ages 65 and over, they return to 0–0.2. For women, on the other hand, the coefficients basically fluctuate around zero and do not possess any particular patterns. These results reconfirm that the reductions in the male eligible ages increased the recipient rates among men aged 60–64 but did not affect the pension receipt status of the other men.

Labor force participation, employment, unemployment

I conduct similar exercises where the dependent variables are dummies for being in the labor force, employed, and unemployed. Figures 14, 15 and 16 plot the coefficients δ_j. According to the upper panel of figure 14, the male labor force participation rates decreased during the considered periods by about 0.1 at ages 55–59 and by 0.05–0.1 at ages 65 and above. At ages 60–63, by contrast, the rate decreased sharply by 0.15. The middle panel of figure 14 finds similar results although its estimates are not as precise as in the upper panel. At the ages 55–59 and 65–70, the decrease ranges 0.05–0.2, but at the ages 60–64 it is large in the range of 0.2–0.3. For women, the labor force participation also decreased, but its magnitude is almost flat across all ages. These results indicate that although during the relevant period there was an overall downward trend in the labor force participation across all ages, the policy change in the male age-eligibility further pushed down the participation rates of men aged 60–64.[15]

The proportions of the employed basically follow the same pattern as the labor force participation rates (figure 15). For men there was an overall downward trend in employment, but the decrease is more striking at ages 60–64. For women the size of the changes in the proportion is rather flat across ages. The results on unemployment are subtle but still indicative (16). In the upper panel, the coefficients for men at ages 60–64 are smaller than at 65–70 although they are similar to those at 55–59. For women the coefficients are about the same level across all ages. These results may show that if there had not been the policy change, the proportion of the unemployed among ages 60–64 would not have decreased as much as it actually did. In the middle panel too, the point estimates are smaller at ages 60–64 than at the other ages.

All the results above suggest that the pension has negative effects on labor supply. Now, I discuss which of cases 1, 2 and 3 explained in the preceding section is the most plausible. Although the confidence intervals of the estimated age-specific differences between pre- and post-periods are not tight enough to have a firm conclusion, case 1 seems to be the most plausible. According to the tightest estimates using the Community Survey 2007 and the Population Census 2011 (the upper panel, figure 14), the result that the magnitudes of the decrease in the labor force participation rate are similar at ages 60–62 and small at age 63 is not consistent with the prediction from case 2, and the result that the decreases at ages 55–59 are smaller than those at 60–64 is not consistent with the prediction based on case 3. By contrast, the relatively flat size of the decreases at ages 60–64 is consistent with the prediction of case 1. In the rest of the paper I assume case 1. That is, the effects of the pension are assumed to depend only on whether the pension grant is currently paid. Although this assumption may not perfectly hold, the results obtained so far indicate that the assumption is a good approximation. With this assumption I quantify the size of the effects.

[15] The bottom panels in all figures, which are based on the NIDS, have large confidence intervals and seem to be noisy probably because of the relatively small number of observations in the NIDS. Thus, they are not informative.

5.2 Difference-in-difference estimates: The size of the effects

Presenting the figures of between-periods differences, I have shown that the reductions in the minimum eligible age for males increased the recipient rates and decreased labor supply and employment of 60–64 years old African males. In this section I quantify these effects of the reductions.

5.2.1 Intention to treat estimates

An estimation equation is:

$$y_{ijt} = \beta_j + \eta_t + \gamma AgeEligible_{jt} + \alpha' X_{ijt} + \epsilon_{ijt}, \tag{3}$$

where y_{ijt} is an outcome variable such as pension receipt status, labor force participation, and employment; β_j is the age fixed effect; η_t is the year fixed effect; and X_{ijt} is controls. The key explanatory variable is $AgeEligible_{jt}$, the dummy indicating that a j year olds individual is age-eligible in year t. The parameter of interest is γ. It represents an effect of being age-eligible, or an intention-to-treat (ITT) effect, for African men aged 60–64. Since the eligibility threshold varies over years due to the policy change, γ is identifiable.

Table 1 shows estimates for the effect of the reductions in the minimum eligible age on the take-up of the pension. It shows that the reductions increased the recipient rate among 60–64 years old men by 27–34% points. This magnitude may seem to be smaller than one might expect. This may be because it took time for people to know about the change, apply for the pension, and start receiving it. Actually, the increase in the take-up rate during the period from 2008 to 2012, 38% points (column 4), is higher than that from 2008 to 2010–11, 27% points (column 3).

Table 2 reports estimates for the effects on labor market outcomes. It clearly shows that the expansion in the eligibility significantly decreased the labor force participation, employment, and unemployment. The labor force participation rates decreased by 5.77% points based on the Community Survey and the Population Census data (column 2), by 11.29% points based on the NIDS (column 4) and by 8.90% points based on the QLFS (column 8). According to the same columns 2, 4 and 7, the probability of being employed decreased by 4.09%, 11.84%, and 4.99% points respectively; and the probability of being unemployed decreased by 1.68%, −0.55%, and 3.91% points, respectively. Although these point estimates are different from each other, the differences are not statistically significant. Whether the additional controls are included (columns 2, 4 and 8) or not (columns 1, 3 and 7) does not much change the estimates. Across years of the data used (columns 4–6 for the NIDS and columns 8–13 for the QLFS) the estimates differ, but the differences are probably within a reasonable range. As a final note, the sizes of the standard errors substantially vary depending on the data used simply because of the different sample sizes. The estimates using the NIDS have large standard errors whereas those using the Community Survey and the Population Census are fairly precise. For example, the 95% confidence interval of the effect on labor force participation (column 2) is $[-7.38, -4.15]$ based on the Community Survey and the Population Census.

5.2.2 Local average treatment effect estimates

Since the take-up of the pension largely depends on one's age-eligibility, and the age-eligibility criterion has the exogenous variation across periods, it is feasible to run instrumental variable estimation to estimate the local average treatment effects (LATE) of the pension grants for men aged 60–64 who would receive the transfer only if they are age-eligible. I estimate the following equation using the dummy for being age-eligible as an instrument:

$$y_{ijt} = \beta_j + \eta_t + \kappa Pension_{ijt} + \alpha' X_{ijt} + \epsilon_{ijt}, \tag{4}$$

where $Pension_{ijt}$ is the dummy indicating that individual i receives the pension. The parameter, κ, is the LATE. Since only the NIDS includes the information of pension recipient status every year, I conduct the 2SLS estimation only with the NIDS.

Table 3 shows results of the estimation. The estimates for the LATE of the pension on labor force participation is large. They are between -29.35% to -34.32% points. The LATE on employment and unemployment is estimated at -31.35% to -45.40% points and 1.63% to 11.09% points, respectively. Although most of these estimates are not significantly different from zero probably due to the small sample size, the point estimates suggest that the size of the LATE is economically substantial.

5.3 Falsification tests

5.3.1 The common trend assumption

My empirical strategy relies on the common trend assumption that the time trend in an outcome variable between pre- and post-periods would have been identical for males at each age of 55–70 if the reductions in the minimum eligible age had not occurred. To examine this assumption, I check prior time trends using the Population Census 2001 and the Community Study 2007. Figure 17 plots point estimates and confidence intervals for the differences in the outcome variables between 2001 and 2007. In contrast with the differences between 2007 and 2011 in figures 14, 15 and 16, there is neither hike or drop at ages 60–64. Besides, the trends between 2001 and 2007 for men and women look similar to each other. Table 4 reports estimation results of equation 3 using the 2001 and 2007 data. In this placebo test, the eligibility ($AgeEligible_{jt}$) is defined as if the minimum eligible age decreased from 65 to 60 during the period between 2001 and 2007. The estimated effect on labor force participation is positive at 1.15–1.72% points; that on employment is insignificant; that on unemployment is positive at 1.23–1.24% points. All these estimates have the different sign than the estimates obtained earlier for the period 2007–2015. Thus, no prior trends which could have extended to the period 2007–2015 and driven the main estimation results are found. This result supports that the common trend assumption holds.

5.3.2 Effects for elderly women

While the minimum eligible age for men decreased, that for women did not change from 60. Thus, if the observed effects of the pension for men were really driven by the pension, similar effects for women should not appear. I have already shown by figures 14–16 that there are no noticeable signs of certain changes in the labor market outcomes of elderly women at any particular ages. To more formally check if elderly women experienced similar changes, I conduct the difference-in-difference estimation by equation 3 for the subsample of women aged 55–70. In the estimation I construct the eligibility dummy as if the female minimum eligible age was reduced in the same manner as the male one was. The estimation result does not detect any placebo effects for women at all (table 5).

5.3.3 Effects for white elderly men

The pension recipient rates among whites are much smaller than those among Africans as shown in figure 1. For example, in 2007 the rates among 66 years old African males and whites were over 73% and 20% respectively according to the Community Survey 2007. Give this fact that the white elderly are less dependent on the pension, if the observed changes in labor market outcomes of African males were really caused by the policy change, similar changes should not be observed among the whites. To check this, I run the difference-in-difference estimation (equation 3) for the white males aged 55–70. The estimation does not find similar results (table 6).[16]

[16] The point estimates using the NIDS are highly negative although they are not significant. I suspect that these large point estimates are due to the small sample size. The number of observations used are less than 300.

6 Effects on prime-aged labor market outcomes

This section investigates effects of the pension on prime-aged individuals who live with pensioners. While previous papers find positive effects of the South African pension program on prime-aged labor supply and employment (Posel et al., 2006; Ardington et al., 2009) and negative effects on them (Bertrand et al., 2003), none of them has exogenous variations to identify the effects, so their estimates may be biased by elderly's age effects. Their estimations are essentially comparison between prime-aged persons living with age-eligible persons and ones living without them elderly men aged 60–64, but the comparison does not control for ages of coresiding elderly persons. Thus, if living with relatively young elderly persons such as men aged 60–64 is different from living with men aged 70–74, their estimates are counfounded by such differences. Even if the comparison is narrowly undertaken such as between living with men aged 64 and living with men aged 65, its results may not be sorely driven by the pension effects if the age 65 is an important timing for other institutions or any. What I aim here is to provide better estimates for the *pure* effects of monetary transfer to coresiding elderly on prime-aged adults. To do so, I exploit the reductions of the male minimum eligible age and control for elderly's age effects.[17]

The estimation equation is

$$y_{it} = \eta_t + \sum_{j=55}^{71} \lambda^{Mj} Old_{it}^{Mj} + \sum_{j=55}^{71} \lambda^{Fj} Old_{it}^{Fj} + \rho AgeEligibleOld_{it}^{M} + \alpha' X_{it} + \epsilon_{it}, \qquad (5)$$

where $AgeEligible_{it}^{M}$ is a dummy for living with an age-eligible *male*, and Old_{it}^{Mj} and Old_{it}^{Fj} are dummies indicating that individual i in period t lives with j years old male and female respectively.[18] Since Old_{it}^{Mj} and Old_{it}^{Fj} control for elderly's age fixed effects.

This estimation of the effects on prime-aged labor market outcomes is still a work in progress, and I do not have conclusive results yet. (Ongoing estimation results are not reported in this paper.)

7 Conclusion

Social protection programs in the developing world have been expanding. Though these programs aim to provide safety nets, they may cause unintended effects such as the disin-

[17] A downside of my approach is that what I am able to estimate at best is the effects of pension grants paid to males age 60–64. That is, my estimates do not directly tell the effects of grants to males aged 65 and above or females at any ages.

[18] To be accurate, Old_{it}^{M71} and Old_{it}^{F71} are dummies for living with a male and female aged 71 and *over*.

centive effect for work. Given the recent rapid growth in those programs, assessment of the disincentive effect is increasingly important for better policy making. Among various types of social programs, old age pension programs need rigorous quantification of the disincentive effect because the developing world will turn into the aging stage, or at least the mature stage, in the not too distant future. Better quantification allows more reliable prediction of the future fiscal burden and reviews of the financial sustainability.

Exploiting the policy change in the South African old age pension program, this paper analyzes the effects of the pension grants on labor supply and employment of elderly and prime-aged people. The pension program has offered sizable benefit under generous eligible criteria since 1990s to overcome the inequality inherited from the Apartheid era. While the program had had a stable structure since early 1990s, the minimum eligible age for males was reduced from 65 to 60 in 2008–2010.

Using three different datasets, my empirical analysis finds that the reductions of the eligible ages significantly discourage elderly men aged 60–64 to work. In those affected group, the labor force participation rate decreased by 5.77%, 11.29%, and 8.90% points depending on the different datasets used. Corresponding to this result, the probability of being employed decreased by 4.09%, 11.84%, and 4.99% points; and the probability of being unemployed decreased by 1.68%, −0.55%, and 3.91% points. Besides, the LATEs on labor force participation, employment, and unemployment are estimated, respectively, at 29.35%, 31.58%, and −2.23% points.

A policy implication from the analysis is that the government needs to take into account the substantial disincentive effect on the elderly and to consider an appropriate design of the program that offers necessary social assistance to the elderly while minimizing the disincentive effect. To do so, however, we need to also assess the effectiveness of the program in producing the intended effects, i.e., how much the program assists the elderly.

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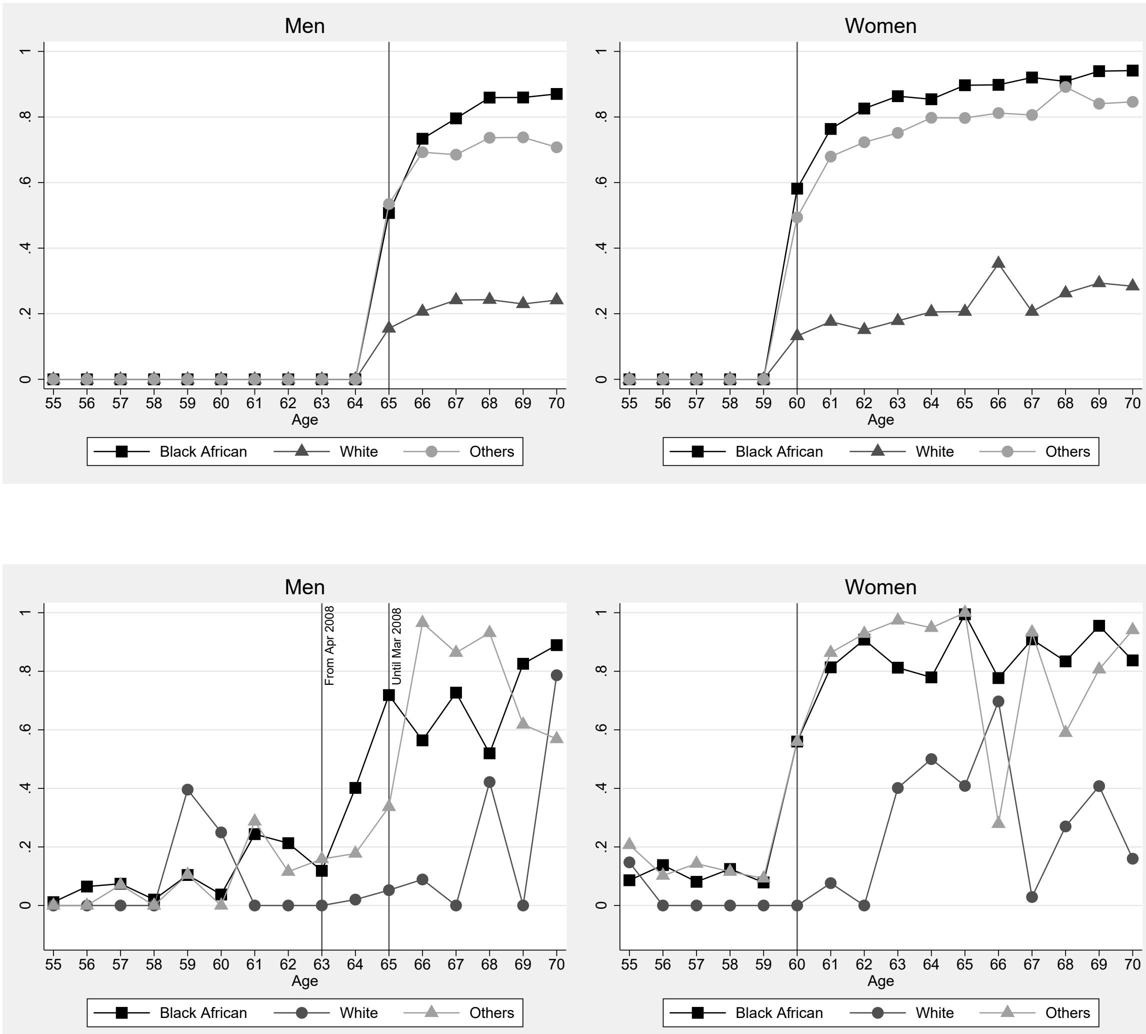


Figure 1: Pension recipient rate by races in 2007 (upper) and 2008 (lower)

Note. *Upper figure*: The data used is the community survey 2007. Pension receipt status is based on a response to a question asking whether an individual receives old age pension. The sampling weight is taken into account. The race group, Others, in the figure consists of Colored and Asian. The vertical lines correspond to the threshold ages in 2007.

Lower figure: The data used is the National Income Dynamics Study (NIDS) in wave 1. Although the survey was conducted throughout 2008, including the period after the first reduction in the male threshold age from 65 to 63 in April 2008, the survey period was concentrated in early 2008. 38% of all the observations were surveyed by the end of March; 62% by the end of April; 91% by the end of June. Pension receipt status is based on a response to a question asking whether an individual receives state old age pension. The sampling weight is taken into account. The race group, Others, in the figure consists of Colored and Asian. The vertical lines correspond to the threshold ages. The somewhat noisy rates among whites and others is due to a small number of observations in single age-gender group.

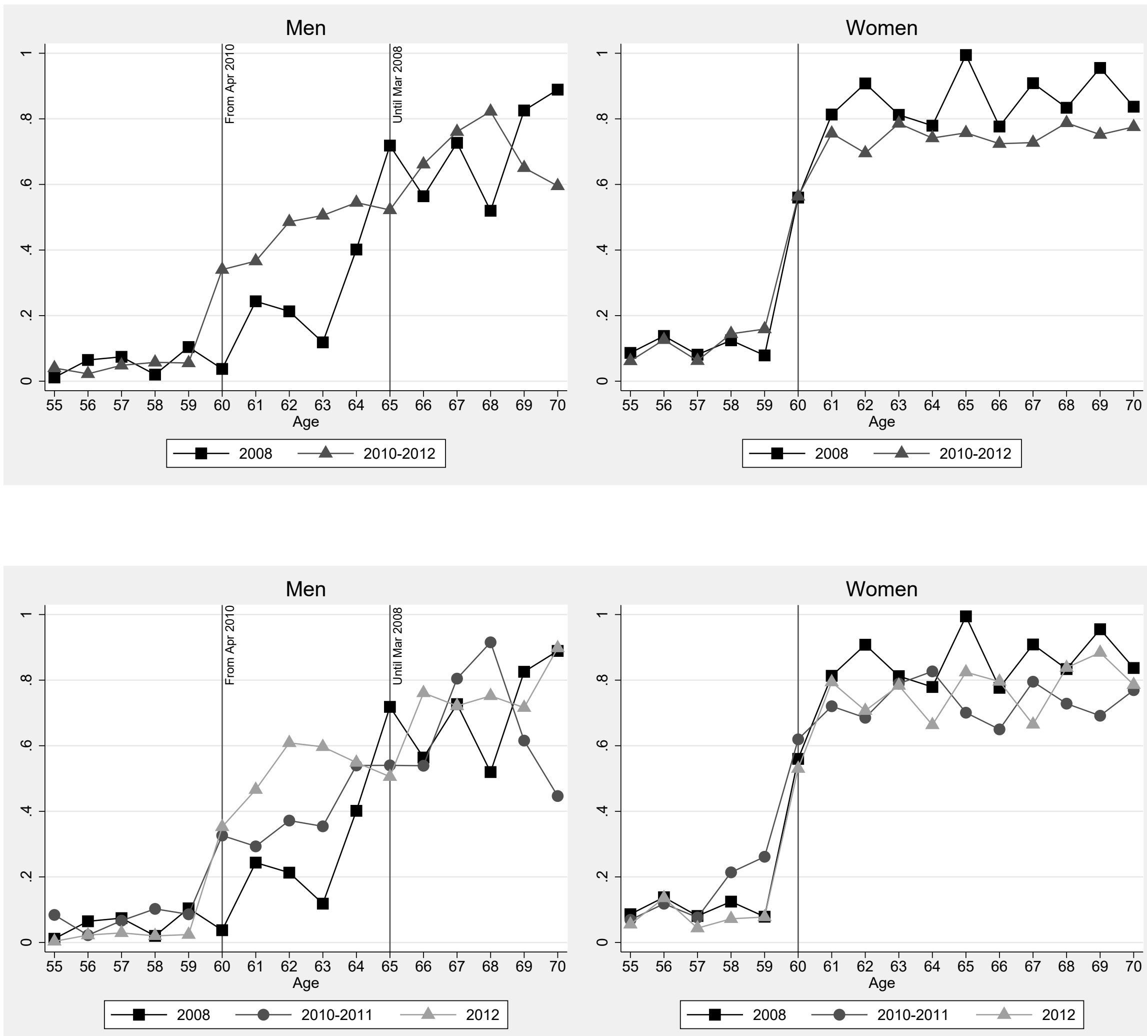


Figure 2: Pension recipient rate among Africans in 2008, 2010-2011, and 2012

Note. Both of the upper and lower figures are based on the National Income Dynamics Study (NIDS) waves 1, 2 and 3. The upper one presents pension recipient rates in 2008 (wave 1) and 2010–12 (waves 2 and 3) while the lower shows 2010–2011 and 2012 separately. Although the data collection of the wave 1 was conducted throughout 2008, including the period after the first reduction in the male threshold age from 65 to 63 in April 2008, the survey period was concentrated in early 2008. 38% of all the observations were surveyed by the end of March; 62% by the end of April; 91% by the end of June. The pension receipt status is based on a response to a question asking whether an individual receives state old age pension. The sampling weight is taken into account.

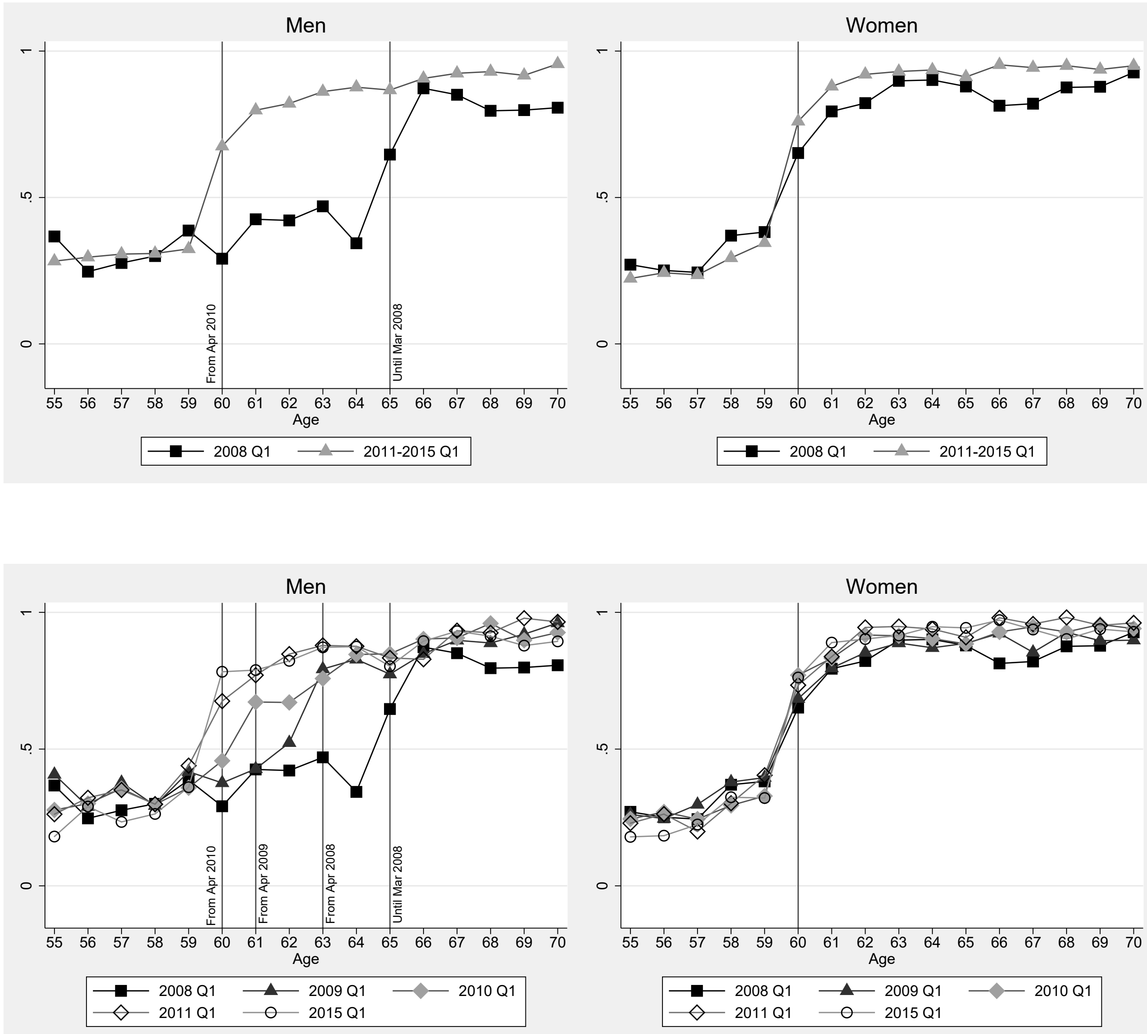


Figure 3: Pension recipient rate among not-employed Africans in 2008 to 2015

Note. The data used is the Quarterly Labor Force Surveys (QLFS) in the first quarter of 2008–2015. Only the subsample of not-employed Africans is used because information of pension receipt status is available only for those not employed. The sampling weight is taken into account. Pension receipt status is determined based on a response to the question "How do you support yourself? Do you receive old age or disability pension?" Although it includes the disability pension, it is not an issue because the disability pension, formally named as the disability grant, is equivalent to the pension grant for disabled people who are age-eligible. The amount of the grants are equal to each other, and once those who have received the disability grant reach the pension-eligible age, the disability grant is taken over by the old age pension grant. Actually, in the data of the community survey 2007, no disability grant recipients are found at and above the pension eligible ages. In the National Income Dynamics Survey, the rate of disability grant recipients suddenly drops at the pension eligible age thresholds. Therefore, among those who answered the question, I am able to interpret that those under the threshold ages are disability grant recipients and that those at and above them are pension recipients.

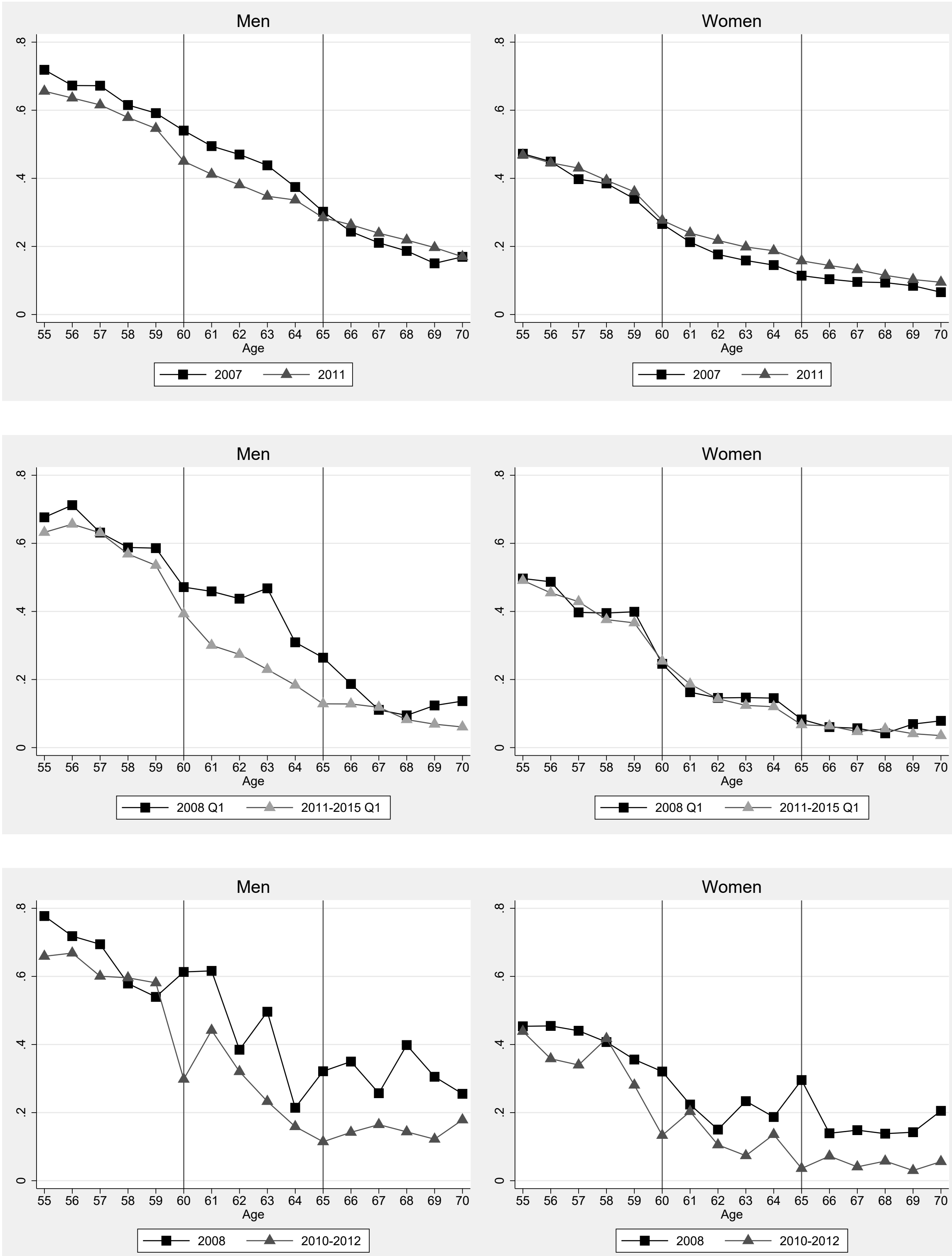


Figure 4: Labor force participation rates among Africans before and after the policy change

Note. The upper panel uses the community survey 2007 and the population census 2011; the middle uses the Quarterly Labor Force Surveys; the lower uses the National Income Dynamics Studies. The sampling weights are taken into account.

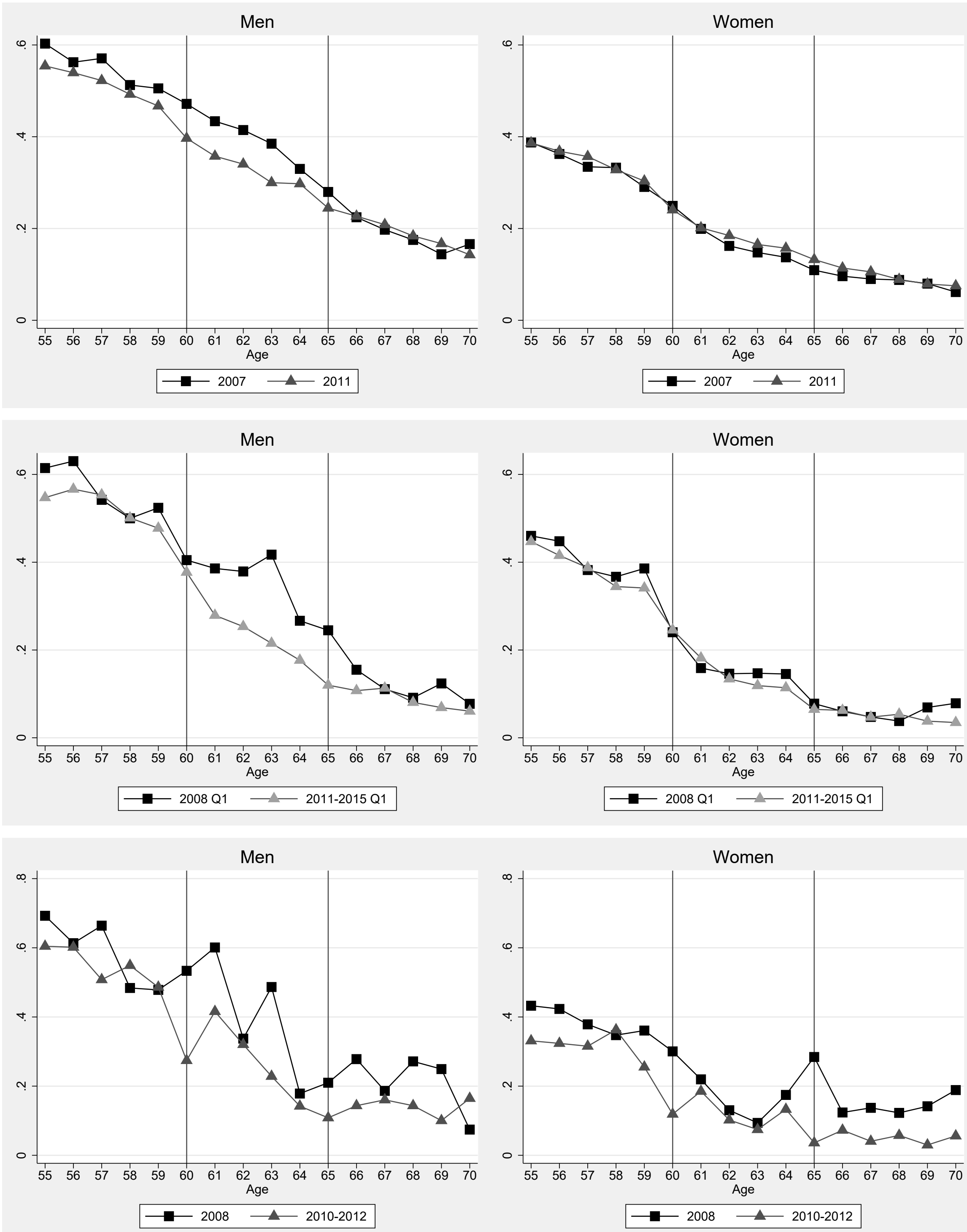


Figure 5: Proportions of the employed among Africans before and after the policy change

Note. The proportions of the employed presented here are different, by definition, from the so-called employment rates. The former is the proportion of the unemployed among all people while the latter is the proportion of them among the labor force. The upper panel uses the community survey 2007 and the population census 2011; the middle uses the Quarterly Labor Force Surveys; the lower uses the National Income Dynamics Studies. The sampling weights are taken into account.

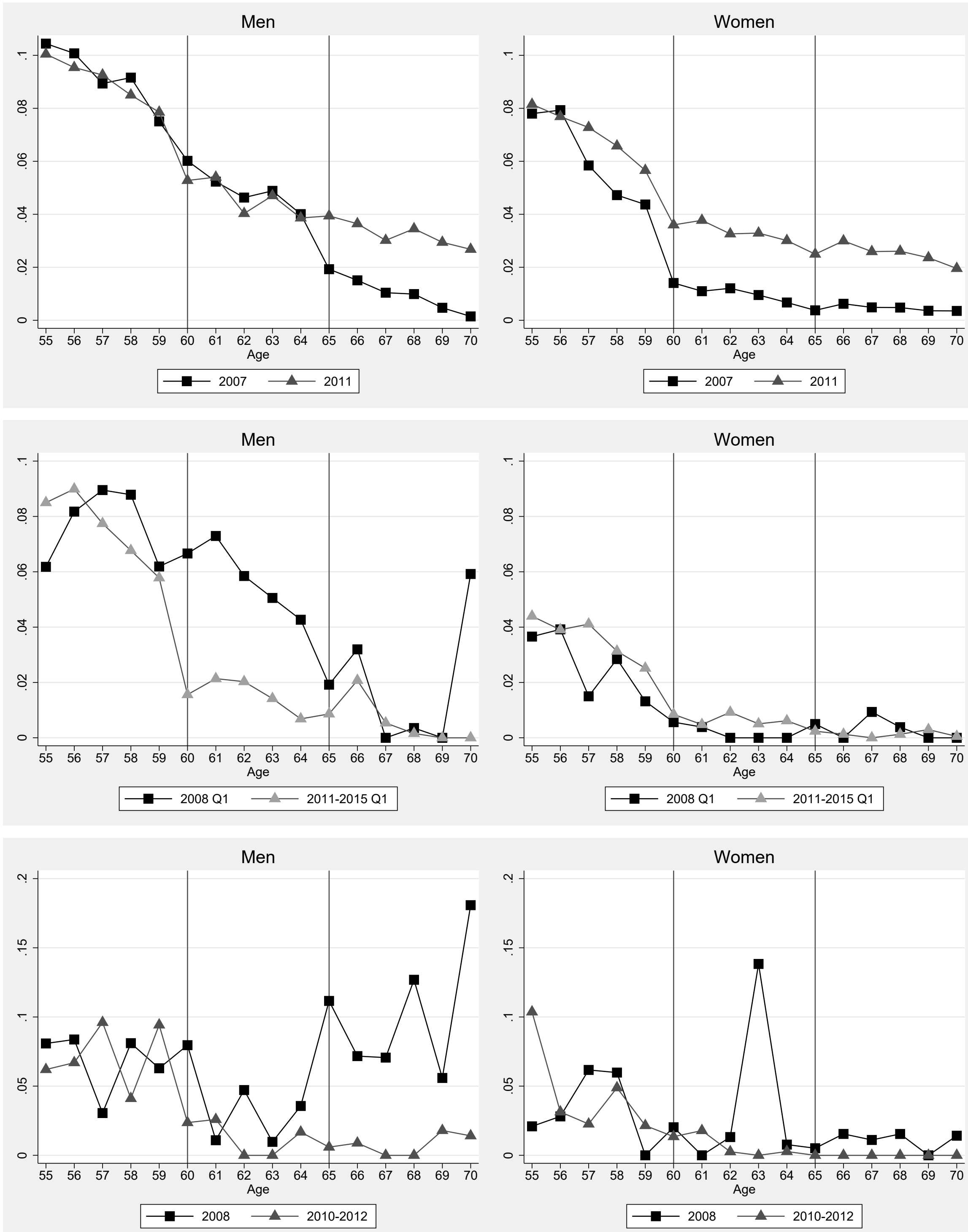


Figure 6: Proportions of the unemployed among Africans before and after the policy change

Note. The proportions of the unemployed presented here are different, by definition, from the so-called unemployment rates. The former is the proportion of the unemployed among all people while the latter is the proportion of them among the labor force. The upper panel uses the community survey 2007 and the population census 2011; the middle uses the Quarterly Labor Force Surveys; the lower uses the National Income Dynamics Studies. The sampling weights are taken into account.

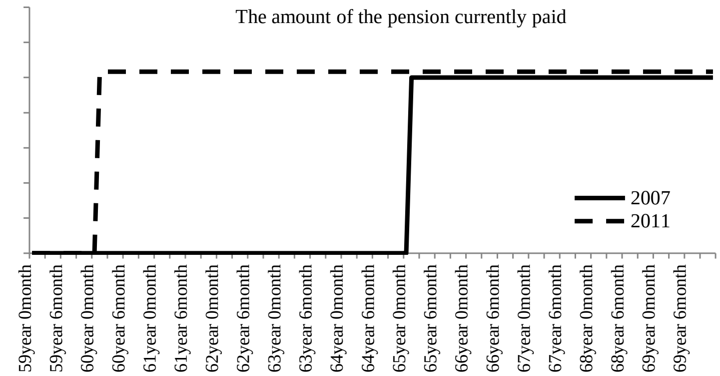


Figure 7: The amount of the pension that men at different ages received in 2007 and 2011

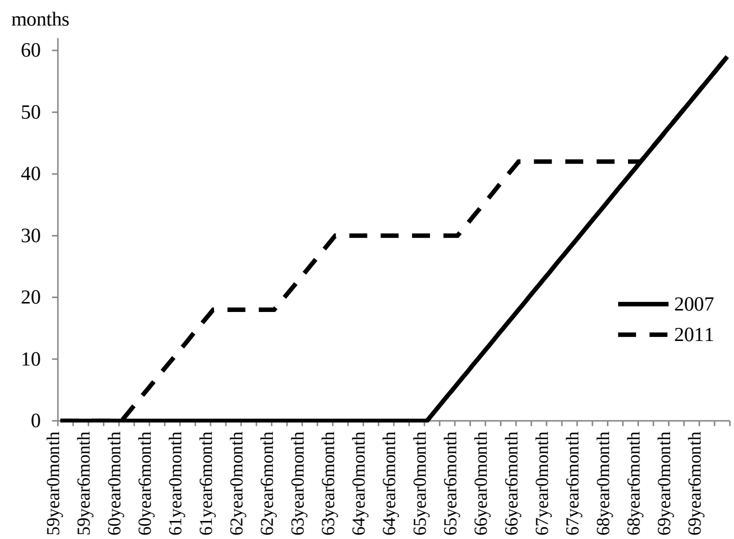


Figure 8: For how many months males at different ages could have received the pension as of 2007 and 2011

Note. The reference times considered here are March 2007 and October 2011, when the Community Survey 2007 and the Population Census 2011 were collected. Thus, for example, 62 years and 0 month old men in March 2007 and October 2011 could have received the pension for 0 month and 18 months, respectively.

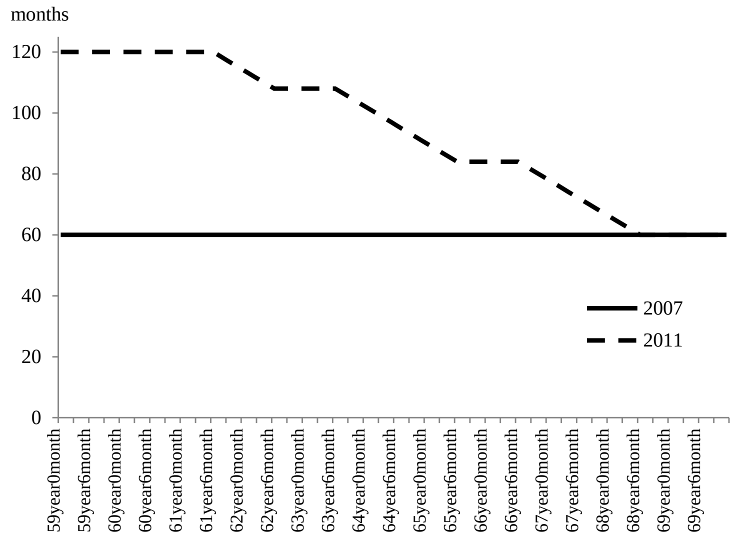


Figure 9: Lifetime pension benefit: the total number of months for which men expect to receive the pension until age 70

Note. Shown is the number of months for which men in Feb–Mar 2007 and Oct–Dec 2011 expect to receive the pension until age 70. In 2007 they did not know the eligible ages would be lowered and expected to start to receive it from age 65. In 2011 they expect the total number of months according to the new eligible ages.

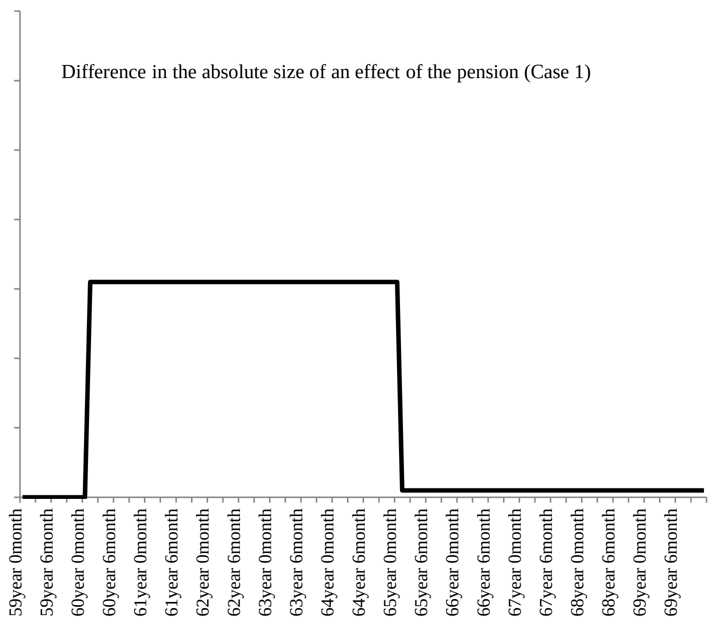
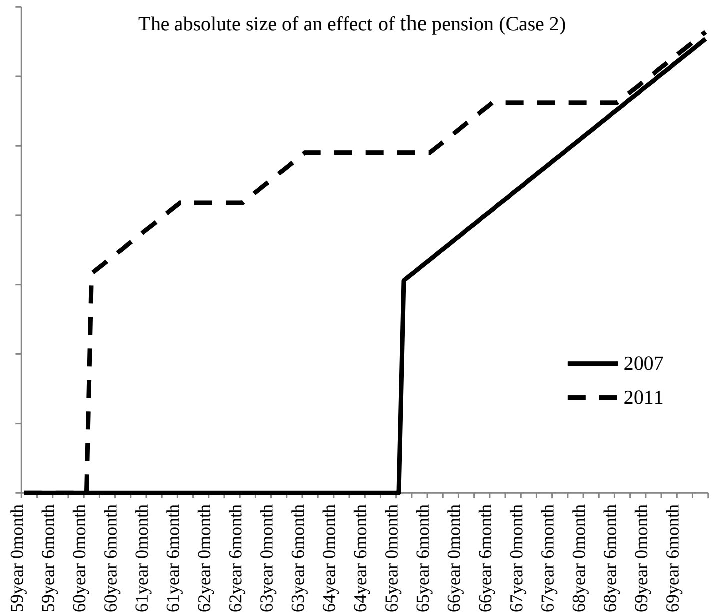


Figure 10: Difference in the predicted effect at each age between 2007 and 2011 (Case 1: Credit constraints exist, and the effect is binary.)

Note. Shown is the difference in the predicted size of an effect of the pension for men between Feb–Mar 2007 and Oct–Dec 2011. The size is assumed to be binary depending on whether the pension is currently paid.



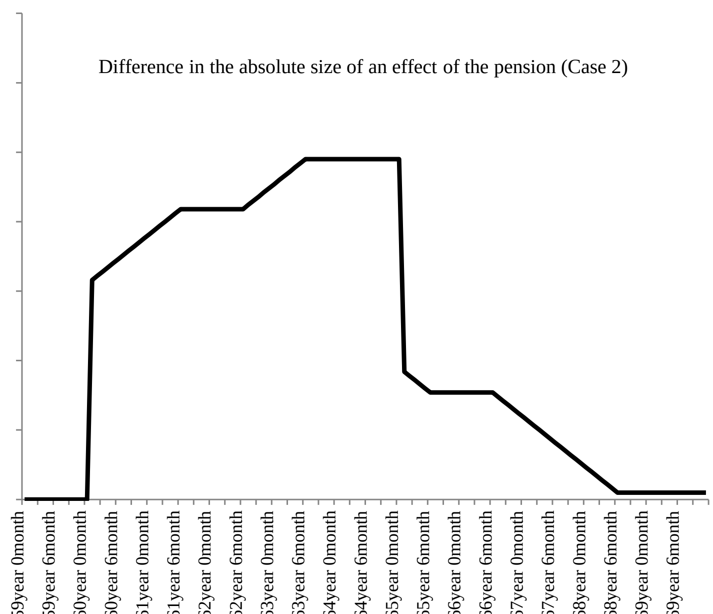


Figure 11: Predicted effect at each age in 2007 and 2011 (Case 2: Credit constraints exist, and the effect consists of binary and cumulative parts.

Note. The upper figure presents the predicted size of an effect of the pension on a labor market outcome for men in Feb–Mar 2007 and Oct–Dec 2011. The lower one presents the difference in the size between 2007 and 2011. The size is assumed to consist of a binary part, which depends on whether the pension is currently paid, and a cumulative part, which depends on how long it has been paid. In this figure, the cumulative part is assumed to be proportional to the length.

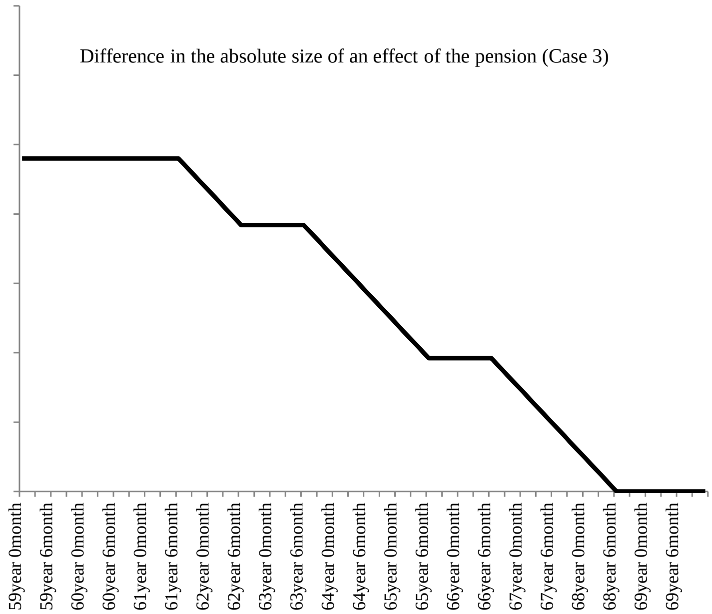


Figure 12: Difference in the predicted effect at each age between 2007 and 2011 (Case 3: No credit constraints are bounded.

Note. Shown is the difference in the predicted size of a pension effect on a labor market outcome for men between Feb–Mar 2007 and Oct–Dec 2011. The size is assumed to be proportional to the total lifetime benefit.

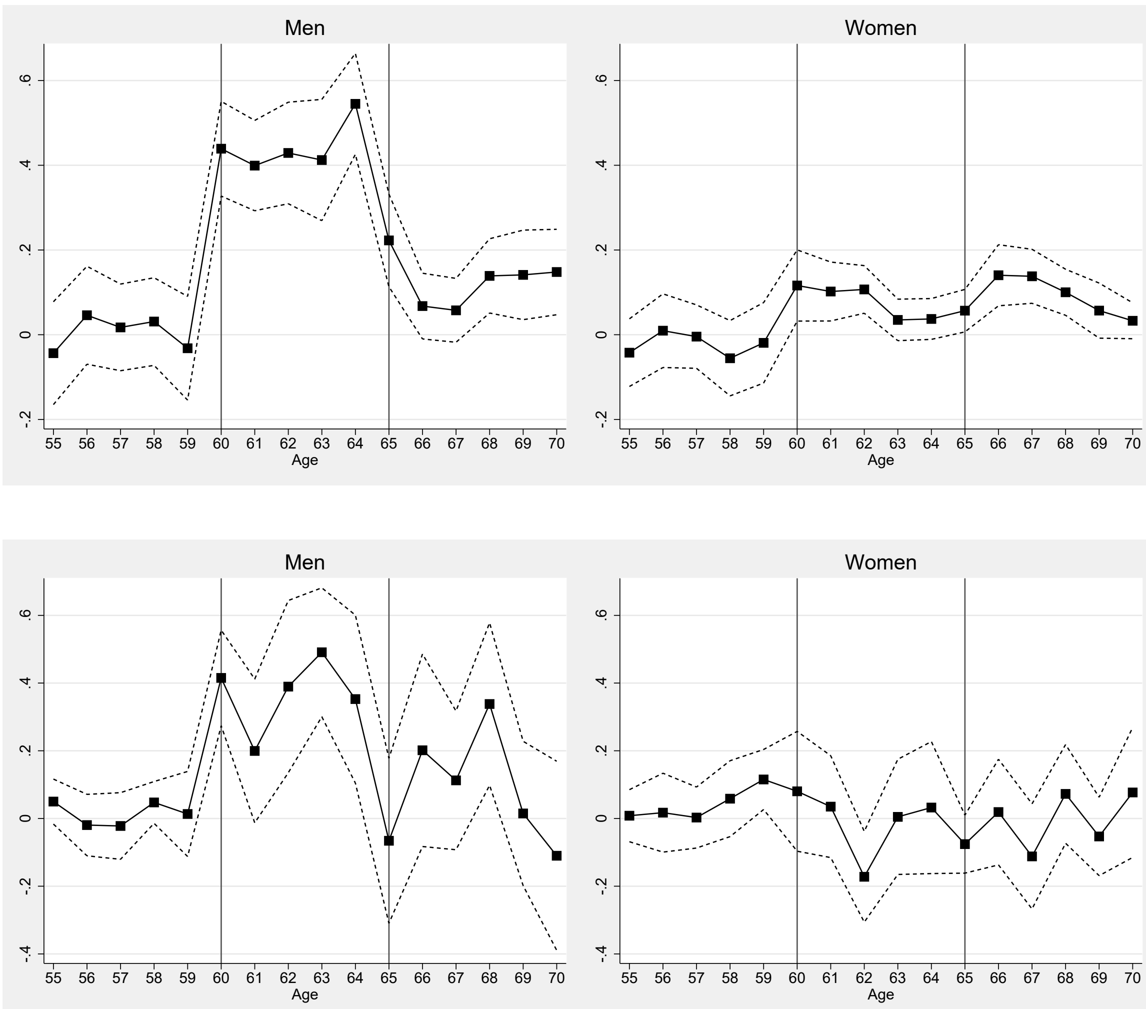


Figure 13: Differences in pension recipient rate between pre- and post-periods

Note. The upper figure uses the Quarterly Labor Force Surveys in the first quarters of 2008 and 2011–2015; and the lower uses the National Income Dynamics Study of wave 1 (2008), wave 2 (2010–2011) and wave 3 (2012). Shown are point estimates of differences in pension recipient rates between the prior and post periods of the reductions in the male eligible ages. The estimates are obtained as the coefficients $\delta_j (j = 55, \ldots, 70)$ in the equation 1, $y_{ijt} = \beta_j + \eta_t + \delta_j Post_t + \alpha' X_{ijt} + \epsilon_{ijt}$, which is run separately for men aged 55–70 and women aged 55-70. The controls, X_{ijt}, are a quartic polynomial of school years and province dummies. The upper figure uses the subsample of those *not employed* since the information of the pension status is available only for them, so the estimates represent the differences in the recipient rates *among not employed*. Robust standard errors are calculated. The dash lines show 95% confidence intervals. Note that the scales of y-axis differ between the panels.

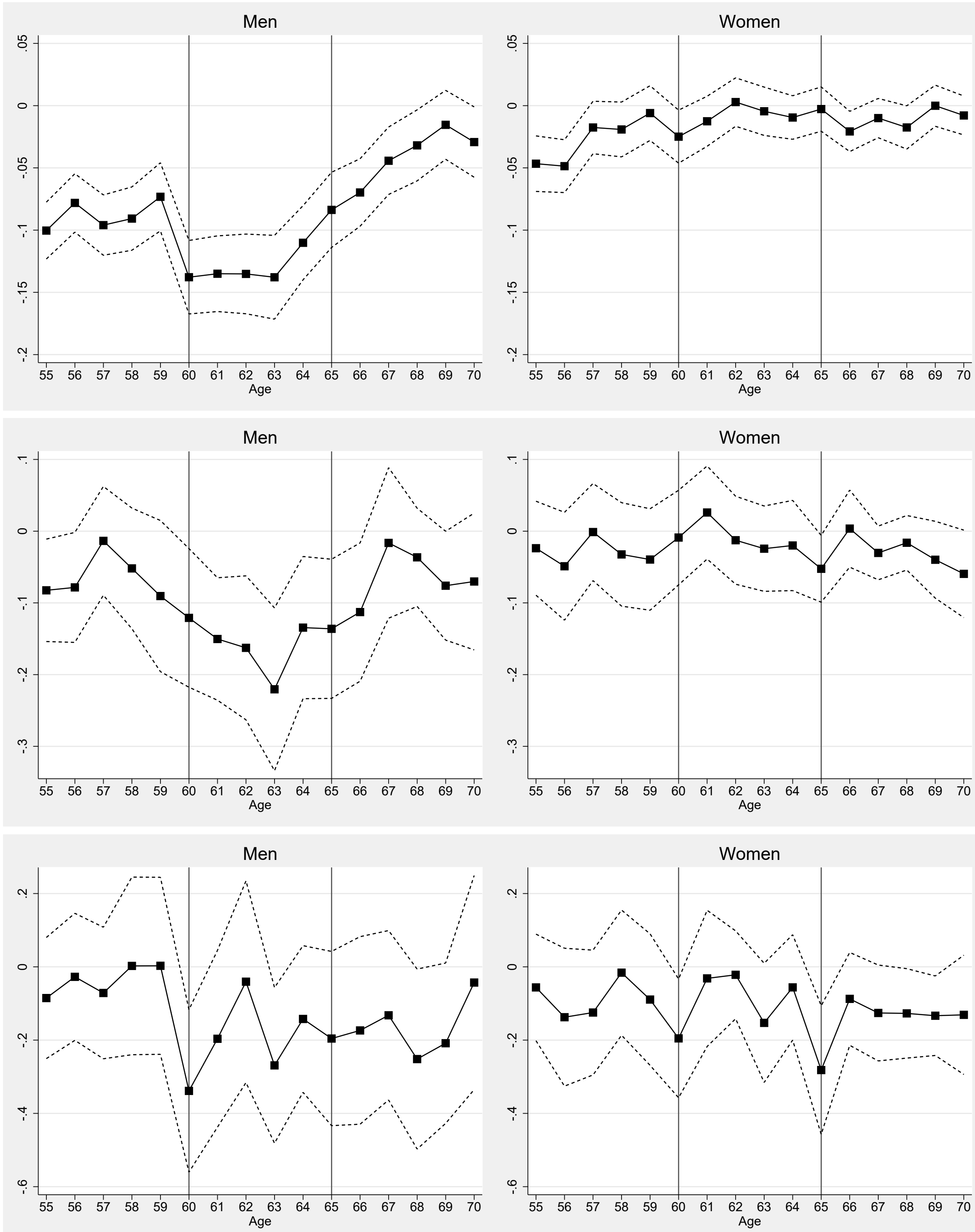


Figure 14: Differences in labor force participation rate between pre- and post-periods

Note. The upper figure uses the Community Survey 2007 and the Population Census 2011; the middle one uses the Quarterly Labor Force Surveys in the first quarters of 2008 and 2011–2015, and the bottom one uses the National Income Dynamics Study of wave 1 (2008), wave 2 (2010–2011) and wave 3 (2012). Point estimates of differences in labor force participation rates are shown. To obtain the estimates, the same estimation equation as in the note of figure 13 is run. Robust standard errors are calculated. The dash lines show 95% confidence intervals. Note that the scales of y-axis differ between the panels.

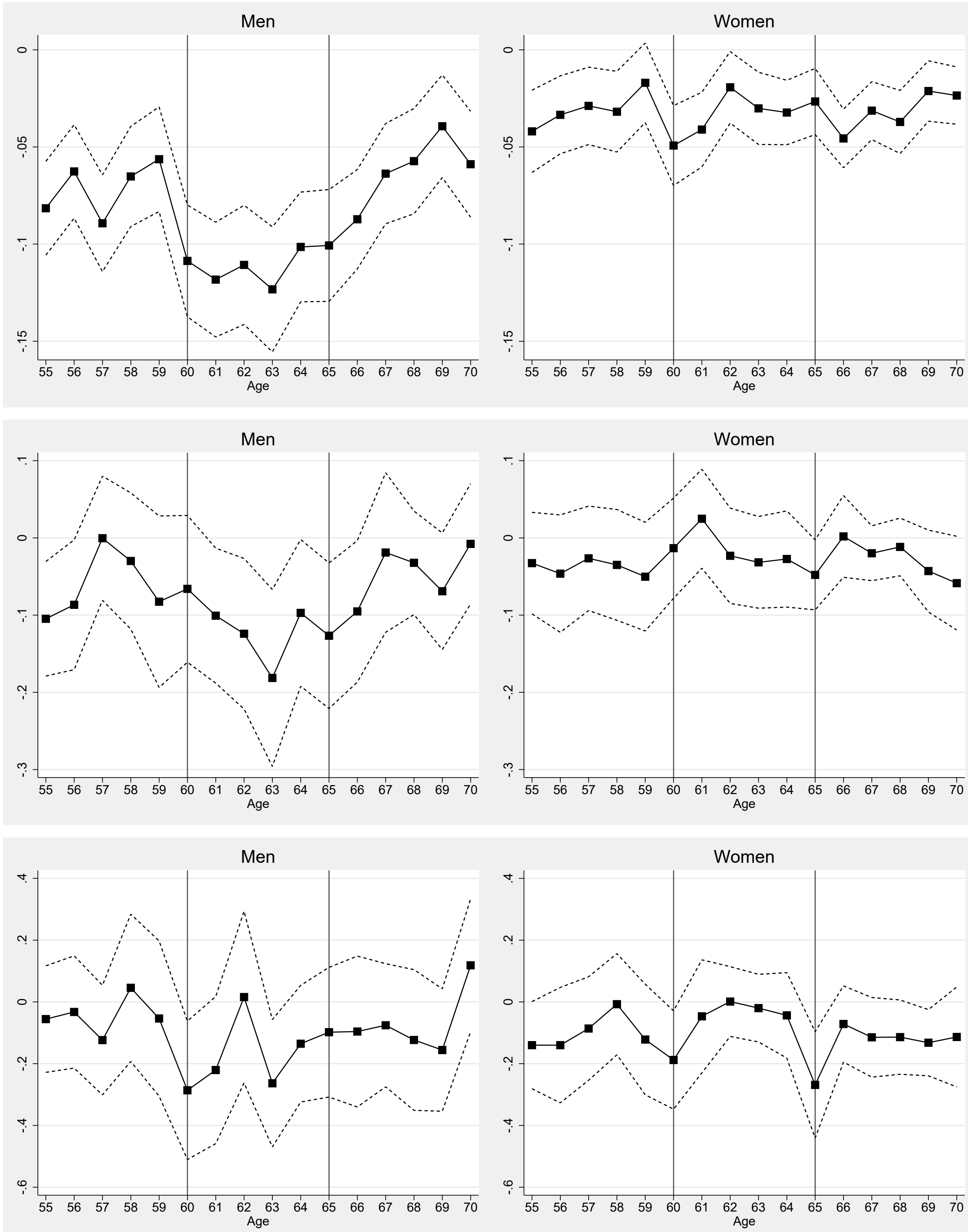


Figure 15: Differences in proportions of employed between pre- and post-periods

Note. The upper figure uses the Community Survey 2007 and the Population Census 2011; the middle one uses the Quarterly Labor Force Surveys in the first quarters of 2008 and 2011–2015, and the bottom one uses the National Income Dynamics Study of wave 1 (2008), wave 2 (2010–2011) and wave 3 (2012). Point estimates of differences in the proportions of the employed are shown. To obtain the estimates, the same equation as in the note of figure 13 is run. Robust standard errors are calculated. The dash lines show 95% confidence intervals. Note that the scales of y-axis differ between the panels.

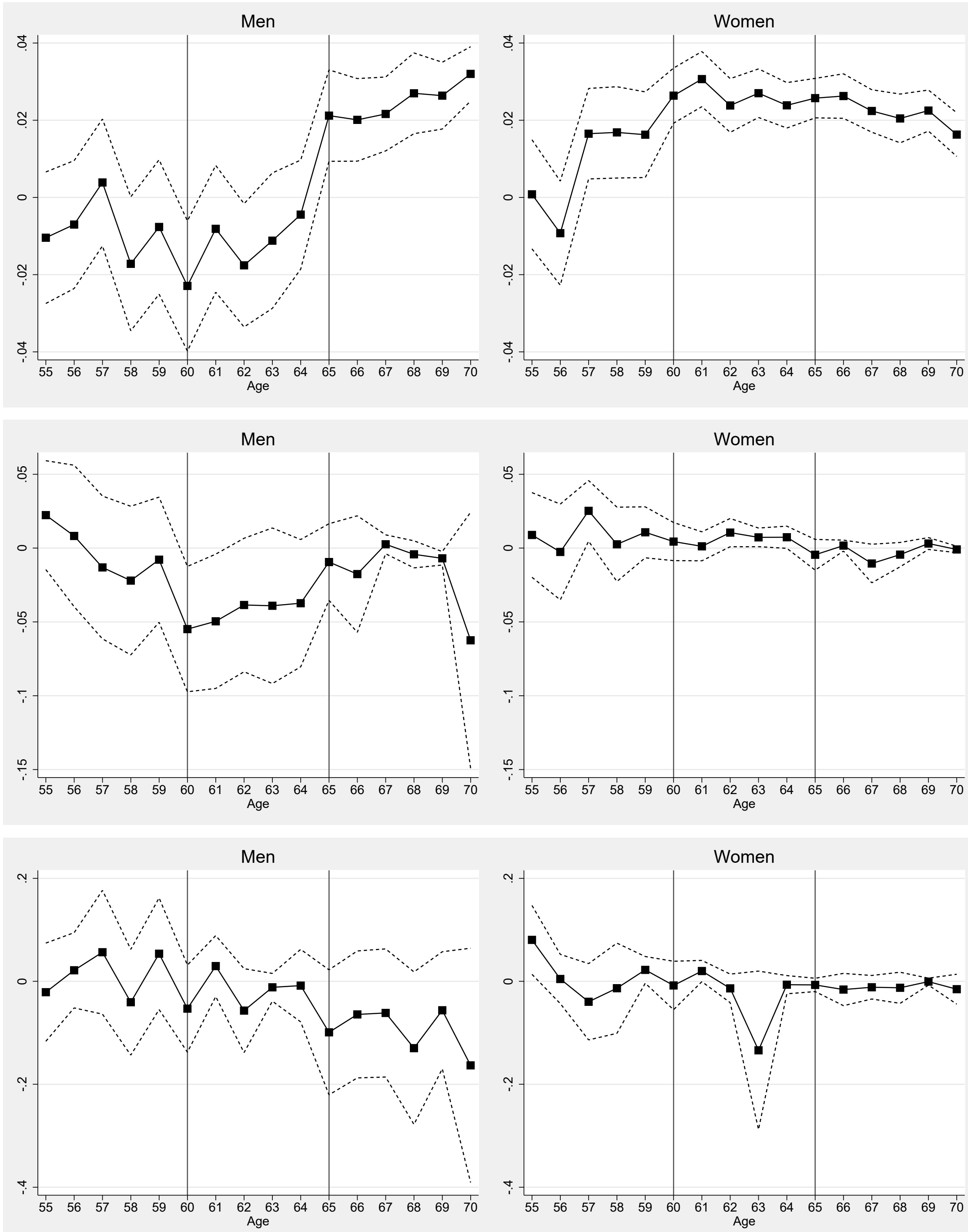


Figure 16: Differences in proportions of unemployed between pre- and post-periods

Note. The upper figure uses the Community Survey 2007 and the Population Census 2011; the middle one uses the Quarterly Labor Force Surveys in the first quarters of 2008 and 2011–2015, and the bottom one uses the National Income Dynamics Study of wave 1 (2008), wave 2 (2010–2011) and wave 3 (2012). Point estimates of differences in the proportions of the unemployed are shown. To obtain the point estimates, the same equation as in the note of figure 13 is run. Robust standard errors are calculated. The dash lines show 95% confidence intervals. Note that the scales of y-axis differ between the panels.

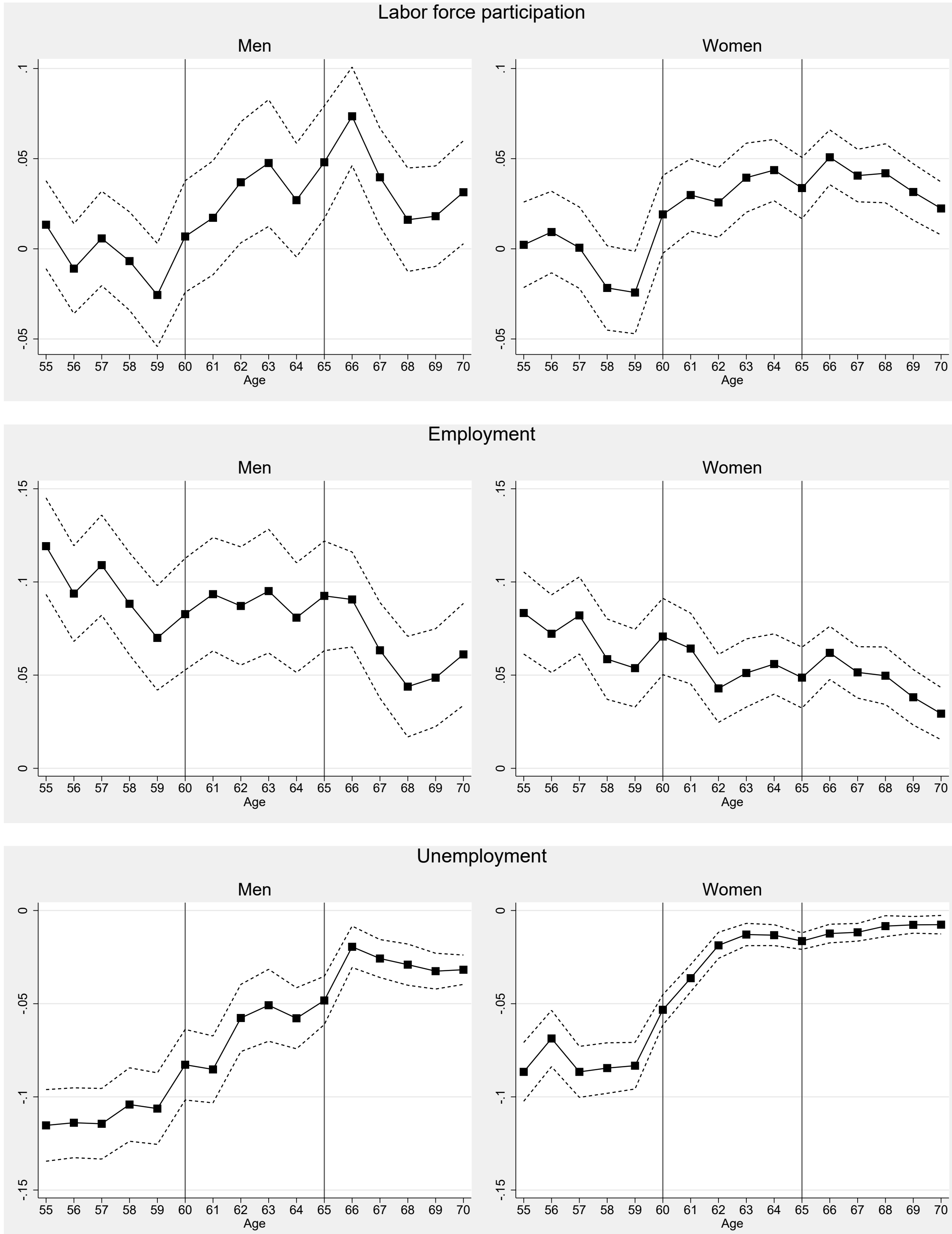


Figure 17: Prior trend from 2001 to 2007: Differences in labor force participation, employment and unemployment

Note. The datasets used are the Community Survey 2007 and the Population Census 2001. Point estimates of between-years differences are shown. Robust standard errors are calculated. The dash lines show 95% confidence intervals. Note that the scales of y-axis differ between the panels.

Table 1: The effect of the policy change on the pension take-up: Diff-in-diff estimates

Data: Type	National Income Dynamics Study			
Year	2008, 10–12	2008, 10–12	2008, 10–11	2008, 12
	(1)	(2)	(3)	(4)
Dummy for receiving the pension				
Age-eligible dummy	33.76 ***	33.94 ***	27.38 ***	38.36 ***
	(5.82)	(5.74)	(6.48)	(6.51)
R-sq	0.664	0.683	0.639	0.690
Pre-policy mean	14.84	14.84	14.84	14.84
N	1,788	1,781	1,193	1,210
Controls	No	Yes	Yes	Yes

Note. Shown are estimates by equation 3 with the dependent variable being a dummy for receiving the pension. The sample used consists of black males aged 55–70. Robust standard errors are in parentheses. Coefficients and standard errors are multiplied by 100 and should be interpreted as percentage points. All specifications includes age dummies and survey wave dummies) Additional controls in specifications (2), (3) and (4) are quartic polynomial of school years and province dummies. The pre-policy means are for men aged 60–64 in January–March 2008. Significance levels: *** = 1%, ** = 5%, * = 10%.

Table 2: ITT effects of the pension on labor market outcomes of elderly African men: Diff-in-diff estimates

Data: Type	Comm. Survey, Pop. Census		National Income Dynamics Study			
Year	2007, 11	2007, 11	2008, 10–12	2008, 10–12	2008, 10–11	2008, 12
	(1)	(2)	(3)	(4)	(5)	(6)
Labor force participation						
Age-eligible dummy	-6.56 ***	-5.77 ***	-9.91	-11.29 *	-9.40	-11.93
	(0.83)	(0.82)	(7.20)	(6.79)	(7.43)	(7.47)
R-sq	0.491	0.516	0.536	0.571	0.572	0.594
Pre-policy mean	42.88	42.88	51.13	51.13	51.13	51.13
Employment						
Age-eligible dummy	-4.89 ***	-4.09 ***	-10.66	-11.84 *	-12.43 *	-10.27
	(0.82)	(0.81)	(7.25)	(6.79)	(7.34)	(7.42)
R-sq	0.397	0.425	0.480	0.535	0.543	0.554
Pre-policy mean	35.98	35.98	47.81	47.81	47.81	47.81
Unemployment						
Age-eligible dummy	-1.67 ***	-1.68 ***	0.75	0.55	3.04	-1.66
	(0.45)	(0.45)	(2.39)	(2.44)	(2.64)	(2.51)
R-sq	0.096	0.099	0.067	0.083	0.080	0.100
Pre-policy mean	6.90	6.90	3.32	3.32	3.32	3.32
N	122,924	122,415	1,788	1,781	1,193	1,210
Controls	No	Yes	No	Yes	Yes	Yes

Data: Type	Quarterly Labor Force Survey						
Year	2008, 11–15	2008, 11–15	2008, 11	2008, 12	2008, 13	2008, 14	2008, 15
	(7)	(8)	(9)	(10)	(11)	(12)	(13)
Labor force participation							
Age-eligible dummy	-11.72 ***	-8.90 ***	-8.07 **	-11.81 ***	-10.40 ***	-5.83 *	-8.91 ***
	(2.69)	(2.64)	(3.39)	(3.29)	(3.48)	(3.37)	(3.38)
R-sq	0.520	0.546	0.549	0.554	0.554	0.557	0.561
Pre-policy mean	43.38	43.38	43.38	43.38	43.38	43.38	43.38
Employment							
Age-eligible dummy	-7.64 ***	-4.99 *	-4.19	-8.68 ***	-6.76 *	-1.36	-4.55
	(2.68)	(2.66)	(3.40)	(3.27)	(3.48)	(3.39)	(3.39)
R-sq	0.455	0.479	0.487	0.490	0.483	0.492	0.482
Pre-policy mean	37.35	37.35	37.35	37.35	37.35	37.35	37.35
Unemployment							
Age-eligible dummy	-4.09 ***	-3.91 ***	-3.88 ***	-3.14 **	-3.63 **	-4.46 ***	-4.36 ***
	(1.25)	(1.27)	(1.42)	(1.48)	(1.53)	(1.44)	(1.55)
R-sq	0.070	0.081	0.074	0.075	0.083	0.081	0.092
Pre-policy mean	6.02	6.02	6.02	6.02	6.02	6.02	6.02
N	13,831	13,567	4,617	4,663	4,738	4,727	4,218
Controls	No	Yes	Yes	Yes	Yes	Yes	Yes

Note. Shown are estimates by equation 3. The sample used consists of black males aged 55–70. Robust standard errors are in parentheses. Coefficients and standard errors are multiplied by 100 and should be interpreted as percentage points. All specifications includes age dummies and year dummies. (In specifications (3)–(5) using the NIDS, dummies for survey waves are included instead of year dummies.) Additional controls in specifications (1), (3) and (7) are quartic polynomial of school years and province dummies. In the case of the Quarterly Labor Force Survey, only the first quarter in each year is used. The pre-policy means are for men aged 60–64 in 2007 (the Community Survey) and 2008Q1 (the QLFS and the NIDS). Significance levels: *** = 1%, ** = 5%, * = 10%.

Table 3: LATE effects of the pension on labor market outcomes of elderly African men: 2SLS estimates

Data: Type	National Income Dynamics Study		
Year	2008, 10–12	2008, 10–12	2008, 10–11
2nd stage	(1)	(2)	(3)
Labor force participation			
Pension receipt dummy	-29.35	-33.26 *	-34.32
	(19.92)	(18.51)	(25.26)
Pre-policy mean	51.13	51.13	51.13
Employment			
Pension receipt dummy	-31.58	-34.88 *	-45.40 *
	(20.01)	(18.45)	(24.85)
Pre-policy mean	47.81	47.81	47.81
Unemployment			
Pension receipt dummy	2.23	1.63	11.09
	(7.04)	(7.13)	(9.80)
Pre-policy mean	3.32	3.32	3.32
1st stage			
Dep. var.= Pension receipt dummy			
Age-eligible dummy	33.76 ***	33.94 ***	27.38 ***
	(5.82)	(5.74)	(6.48)
R-sq	0.66	0.68	0.64
F	115.8	91.9	48.5
N	1,788	1,781	1,193
Controls	No	Yes	Yes

Note. Shown are estimates by equation 4. The sample used consists of black males aged 55–70. Robust standard errors are in parentheses. Coefficients and standard errors are multiplied by 100 and should be interpreted as percentage points. All specifications includes age dummies and year dummies. (To be precise, dummies for survey waves are included instead of year dummies.) Additional controls in specifications (2)–(4) are quartic polynomial of school years and province dummies. The pre-policy means are for men aged 60–64 in 2007 (the Community Survey) and 2008Q1 (the QLFS and the NIDS). Significance levels: *** = 1%, ** = 5%, * = 10%.

Table 4: Diff-in-diff estimates of placebo effects during 2001–2007

	Data: Type	Pop. Census 2001, Comm. Survey	
	Year	2001, 07	2001, 07
		(1)	(2)
Dep. Var. = Labor force participation			
	Pseudo-age-eligible dummy	1.72 **	1.15
		(0.86)	(0.85)
	R-sq	0.517	0.534
	Pre-policy mean	39.33	39.33
Dep. Var. = Employment			
	Pseudo-age-eligible dummy	0.49	-0.09
		(0.84)	(0.83)
	R-sq	0.386	0.408
	Pre-policy mean	25.59	25.59
Dep. Var. = Unemployment			
	Pseudo-age-eligible dummy	1.23 **	1.24 **
		(0.50)	(0.50)
	R-sq	0.153	0.156
	Pre-policy mean	13.74	13.74
	N	92,354	92,078
	Controls	No	Yes

Note. Shown are estimates by equation 3 using the Population Census 2001 and the Community Survey 2007. The sample used consists of black males aged 55–70. Robust standard errors are in parentheses. The pseudo-age-eligible dummy equals one if an age is equal to or greater than 65 in 2001 and 60 in 2007. If otherwise, it is zero. Coefficients and standard errors are multiplied by 100 and should be interpreted as percentage points. All specifications include age dummies and year dummies and province dummies. Additional controls in specification (2) are a quartic polynomial of school years. The pre-policy means are for men aged 60–64 in 2001. Significance levels: $^{***} = 1\%$, $^{**} = 5\%$, $^{*} = 10\%$.

Table 5: Diff-in-diff estimates of placebo effects for African women

Data: Type	Comm. Survey, Pop. Census		National Income Dynamics Study			
Year	2007, 11	2007, 11	2008, 10–12	2008, 10–12	2008, 10–11	2008, 12
	(1)	(2)	(3)	(4)	(5)	(6)
Labor force participation						
Pseudo-age-eligible	0.62	1.07 *	-1.76	1.19	1.75	0.64
dummy	(0.57)	(0.55)	(4.86)	(4.89)	(5.38)	(5.29)
R-sq	0.312	0.356	0.326	0.353	0.342	0.376
Pre-policy mean	16.83	16.83	25.08	25.08	25.08	25.08
Employment						
Pseudo-age-eligible	-0.62	-0.18	-1.03	1.92	0.67	2.18
dummy	(0.54)	(0.53)	(4.55)	(4.62)	(4.92)	(4.98)
R-sq	0.247	0.297	0.291	0.318	0.313	0.337
Pre-policy mean	15.62	15.62	20.99	20.99	20.99	20.99
Unemployment						
Pseudo-age-eligible	1.24 ***	1.25 ***	-0.74	-0.73	1.07	-1.54
dummy	(0.23)	(0.23)	(2.14)	(2.15)	(2.73)	(2.50)
R-sq	0.069	0.073	0.045	0.053	0.061	0.061
Pre-policy mean	1.20	1.20	4.08	4.08	4.08	4.08
N	169,700	169,029	3,161	3,156	2,104	2,131
Controls	No	Yes	No	Yes	Yes	Yes

Data: Type	Quarterly Labor Force Survey						
Year	2008, 11–15	2008, 11–15	2008, 11	2008, 12	2008, 13	2008, 14	2008, 15
	(7)	(8)	(9)	(10)	(11)	(12)	(13)
Labor force participation							
Pseudo-age-eligible	1.13	2.27	3.60	0.79	3.42	2.14	1.99
dummy	(1.84)	(1.78)	(2.25)	(2.32)	(2.29)	(2.39)	(2.35)
R-sq	0.361	0.406	0.378	0.407	0.399	0.420	0.420
Pre-policy mean	17.21	17.21	17.21	17.21	17.21	17.21	17.21
Employment							
Pseudo-age-eligible	1.07	2.08	3.06	0.76	3.13	3.01	1.03
dummy	(1.83)	(1.77)	(2.22)	(2.30)	(2.28)	(2.37)	(2.33)
R-sq	0.331	0.376	0.355	0.382	0.374	0.386	0.389
Pre-policy mean	17.01	17.01	17.01	17.01	17.01	17.01	17.01
Unemployment							
Pseudo-age-eligible	0.06	0.19	0.53	0.03	0.29	-0.87	0.96
dummy	(0.41)	(0.42)	(0.60)	(0.56)	(0.57)	(0.64)	(0.79)
R-sq	0.033	0.039	0.032	0.033	0.036	0.050	0.040
Pre-policy mean	0.20	0.20	0.20	0.20	0.20	0.20	0.20
N	20,733	20,502	6,885	6,942	7,119	7,162	6,174
Controls	No	Yes	Yes	Yes	Yes	Yes	Yes

Note. Shown are estimates by equation 3 using the sample of black females aged 55–70. Robust standard errors are in parentheses. The pseudo-age-eligible dummy was constructed as if the female minimum eligible age was reduced in the same manner as the male one was. Coefficients and standard errors are multiplied by 100 and should be interpreted as percentage points. All specifications includes age dummies and year dummies. (To be precise, in specifications (3)–(5), which use the NIDS, dummies for survey waves are included instead of year dummies.) Additional controls in specifications (1), (3) and (7) are quartic polynomial of school years and province dummies. In the case of the Quarterly Labor Force Survey, only the first quarter in each year is used. The pre-policy means are for women aged 60–64 in 2007 (the Community Survey) and 2008Q1 (the QLFS and the NIDS). Significance levels: *** = 1%, ** = 5%, * = 10%.

Table 6: Diff-in-diff estimates of placebo effects for white men

Data: Type	Comm. Survey, Pop. Census		National Income Dynamics Study			
Year	2007, 11	2007, 11	2008, 10–12	2008, 10–12	2008, 10–11	2008, 12
	(1)	(2)	(3)	(4)	(5)	(6)
Labor force participation						
Age-eligible dummy	-1.25	-1.67	-23.14	-23.11	-28.21 *	-16.93
	(1.53)	(1.51)	(18.91)	(15.62)	(17.04)	(18.07)
R-sq	0.669	0.682	0.691	0.727	0.743	0.739
Pre-policy mean	59.52	59.52	34.85	34.85	34.85	34.85
Employment						
Age-eligible dummy	-0.81	-1.21	-19.08	-20.96	-28.42	-13.93
	(1.54)	(1.52)	(19.32)	(15.88)	(17.26)	(18.70)
R-sq	0.647	0.662	0.676	0.723	0.744	0.728
Pre-policy mean	58.29	58.29	34.85	34.85	34.85	34.85
Unemployment						
Age-eligible dummy	-0.44	-0.46	-4.06	-2.15	0.22	-2.99
	(0.39)	(0.40)	(4.09)	(3.40)	(3.54)	(4.08)
R-sq	0.025	0.027	0.111	0.164	0.150	0.208
Pre-policy mean	1.22	1.22	0.00	0.00	0.00	0.00
N	36,642	36,163	264	263	187	187
Controls	No	Yes	No	Yes	Yes	Yes

Data: Type	Quarterly Labor Force Survey						
Year	2008, 11–15	2008, 11–15	2008, 11	2008, 12	2008, 13	2008, 14	2008, 15
	(7)	(8)	(9)	(10)	(11)	(12)	(13)
Labor force participation							
Age-eligible dummy	0.04	-0.73	4.42	1.32	-10.27	-1.64	1.67
	(5.74)	(5.80)	(7.52)	(7.06)	(7.04)	(7.55)	(7.52)
R-sq	0.663	0.679	0.694	0.695	0.702	0.700	0.666
Pre-policy mean	57.20	57.20	57.20	57.20	57.20	57.20	57.20
Employment							
Age-eligible dummy	-0.18	-0.85	4.50	0.67	-8.40	-3.50	1.73
	(5.81)	(5.84)	(7.60)	(7.09)	(7.16)	(7.58)	(7.58)
R-sq	0.639	0.658	0.666	0.677	0.674	0.681	0.641
Pre-policy mean	54.76	54.76	54.76	54.76	54.76	54.76	54.76
Unemployment							
Age-eligible dummy	0.22	0.12	-0.08	0.65	-1.87	1.86	-0.06
	(1.69)	(1.73)	(2.35)	(2.03)	(2.03)	(1.80)	(2.03)
R-sq	0.031	0.044	0.063	0.070	0.070	0.075	0.098
Pre-policy mean	2.44	2.44	2.44	2.44	2.44	2.44	2.44
N	3,242	3,200	1,047	1,111	1,138	1,129	955
Controls	No	Yes	Yes	Yes	Yes	Yes	Yes

Note. Shown are estimates by equation 3 using the sample of black females aged 55–70. Robust standard errors are in parentheses. The pseudo-age-eligible dummy was constructed as if the female minimum eligible age was reduced in the same manner as the male one was. Coefficients and standard errors are multiplied by 100 and should be interpreted as percentage points. All specifications includes age dummies and year dummies. (To be precise, in specifications (3)–(5), which use the NIDS, dummies for survey waves are included instead of year dummies.) Additional controls in specifications (1), (3) and (7) are quartic polynomial of school years and province dummies. In the case of the Quarterly Labor Force Survey, only the first quarter in each year is used. The pre-policy means are for men aged 60–64 in 2007 (the Community Survey) and 2008Q1. (the QLFS and the NIDS). Significance levels: *** = 1%, ** = 5%, * = 10%.