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Rearranging the Family? Household Composition Responses to Large Pension Receipts*

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Abstract: This paper considers how household composition responds to changes in household income. Using a semi-parametric regression discontinuity estimator, we exploit an age discontinuity in the benefit structure of South Africa's Old Age Pension Program. We find that household composition is an important and under-explored component of a household's response to changes in income. Overall, we observe modest but meaningful changes in the household structure surrounding elderly black South Africans. Our findings vary with the gender of the pension recipient. When a woman becomes pension eligible, we observe increases in the presence of young children and declines in the presence of women in their 30s. For men, we find increases in the number of school aged children and declines in the presence of men in their 30s. These results suggest that both researchers and policymakers should be wary of conditioning on household demographics in evaluating the effect of government programs or fluctuations in household resources.

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1. Introduction

Household structures in poor countries are substantially more complex than in rich economies. In poor countries, the co-residency of extended families is not unusual, and household membership is substantially more fluid. This paper considers how increases in income affect household composition. Much of the academic literature on income and household composition treats household composition as exogenous to the household's economic environment and considers its effect on household income. Household composition can influence a household's production capacity and labor supply (Singh, Squire, and Strauss 1986; Benjamin 1993), its members' preferences for consumption and investment (Chiappori 1988; Thomas 1994), and its ability to insure against risk (Kotlikoff and Spivak 1981; Rosenzweig 1988). However, in this study, we find that household composition responds to exogenous changes in income.

We use the benefit structure of the Old Age Pension Program in South Africa to examine the effect of income on the living arrangements of elderly black South Africans. The end of Apartheid brought parity in government pension benefits to blacks and whites in South Africa. However, blacks are substantially poorer than whites in South Africa, and with parity, the maximum pension benefit for a black South African became approximately twice the median per capita monthly household income of blacks (Case and Deaton 1998). Women become eligible for the pension at age 60, men at age 65. We study the impact of this large cash transfer on household composition by exploiting this age discontinuity in the pension program's benefit structure.

In our estimation strategy, we assume that in the absence of pension income, household composition would not naturally change discontinuously at the ages of eligibility – that is, for

women, household composition at age 60 would appear similar to household composition at ages 59 and 61. We then apply Porter's (1999) semi-parametric regression discontinuity estimator to data from the South African census of 1996 to explore changes in household composition at and around the age of pension receipt. We focus on age-eligibility rather than actual pension receipt, because pension take-up is a household choice whereas an individual's age presumably is not. This focus on changes in household composition at the age of pension eligibility does not capture changes in household composition that may occur in anticipation of the pension or with some delay. Our estimates are further attenuated if take-up of the pension is incomplete (as it certainly is) or does not coincide exactly with eligibility. Hence, our results likely represent a lower bound on the effect of the pension benefit on household composition.

We find modest but meaningful changes in household composition associated with pension eligibility. Overall, we observe a slight and insignificant increase in household size. Within the household, the dynamics are more substantive. First, contrary to most of the literature from higher income countries, we find no evidence of an increased propensity to live alone. Second, we observe declines in the presence of prime work age adults. The gender of the pension recipient is particularly important in our results. When a male becomes pension eligible, we find a decline in prime age men, but no effect on women. When a woman becomes pension eligible, we find declines in both prime age men and women, with the effect largest on women. Likewise, the dynamics for children vary with the gender of the pension recipient. The pension eligibility of men increases the presence of children (especially boys) ages 6-17. For pension eligible women, we see increases in young children (under 5) and in women ages 18-23. It is somewhat difficult to interpret these gender differences, but they are consistent with the widely held notion in South Africa that grandmothers help shelter and care for young children. It is

unclear whether the rise in boys associated with male pension eligibility is for schooling or for work, although much of the anecdotal evidence would point to the latter.

Our results are of interest for several reasons. First, they contribute to a better understanding of household composition choices in general. For reasons discussed in the next section, we cannot use our results to fully differentiate between competing models of household formation and decision-making. However, we do contribute new facts and patterns that models of household formation must be able to explain. Notably, and in contrast to previous literature on living arrangements of the elderly, we do not find an increasing propensity for the elderly to live alone associated with increased income. Second, we offer a cautionary note for researchers employing household-level empirical models. Our finding that household composition (if not size) changes with pension receipt implies that an analysis of the impacts of the pension program involving a naïve conditioning on household composition may produce misleading conclusions. More generally, our finding that household resources directly influence household composition extends this caution to household level models in which income effects are likely to be prevalent. Finally, changes in household composition are part of the household's optimal response to income changes. As such, government transfer policies that are conditioned on household composition may introduce inefficient distortions to this optimizing behavior.

The next section of this paper explores the theoretical and empirical evidence on how cash transfers influence living arrangements. The third section introduces the pension program that is the basis of this study's inference. Section 4 describes the data. Section 5 introduces the regression discontinuity estimator employed in this study and explains its limits. Section 6 presents results for the elderly, black population of South Africa. Section 7 discusses some remaining issues and concludes.

2. Income and Living Arrangements: Theory and Evidence

In this paper, a household is defined as a set of co-resident individuals. The composition of a household is an outcome of a joint decision between all actual and potential members of a household. Household composition itself may be a component of consumption (giving a potential member utility directly), an input to production, or both. In this section, we consider how changes in the income of an individual household member may influence living arrangements. We begin with an overview of the role played by household composition in the household. Within this framework, we consider how income may affect household composition and discuss how the nature of household decision-making influences the relationship between income and living arrangements.

In the literature on household composition in high income countries, most studies view household composition as consumption. The aged might prefer to live alone for privacy, independence, or greater freedom in their choices. Additional income or material wealth enables the aged to continue their independence.[1] In a series of papers, Costa (1997, 1999) finds evidence that either the elderly prefer to live alone or their children prefer them to live alone. Costa (1997) finds that Union Army pensions to veterans of the Civil War in the United States were associated with substantially increased demand for separate living arrangements for the elderly. Costa (1999) finds similar results using state-time variation in benefit levels of the Old Age Assistance program that preceded the Social Security Act in the United States. The literature from poorer countries echoes these results. DaVanzo and Chan (1994) analyzes cross-sectional variation in housing prices in Malaysia to argue that housing costs are a key factor in whether the elderly can afford to live independently. De Carvalho Filho (2000) considers the

[1] For examples, see Michael, Fuchs, and Scott (1980), Costa (1997), Costa (1999), and Boersch-Supan, McFadden, and Schnabel (1993) who emphasize wealth rather than income.

living arrangements of single, rural women in Brazil. He finds that an increase in pension income in these women increases their propensity to live alone.

Much of the literature on household composition in developing countries emphasizes the role household composition plays in production (or household resources). In the most basic model of household formation, households form to produce a non-tradable (home produced) good in which there are economies of scale (Becker 1973); the household's ability to produce this good varies with household composition. Even without a non-tradable good, incomplete markets or imperfect information can cause the household's production of tradable goods to depend on the household's on-hand labor. This idea probably owes to Chayanov (1926), and Singh, Squire, and Strauss (1986) survey formal models where household production in agriculture depends on the labor resident within the household. Whether or not agricultural production depends on household composition is still a debated question (Benjamin 1993). A household's composition also influences how well the household insures against risk. Kotlikoff and Spivak (1981) develop a model where individuals form families to self-insure against uncertain dates of death. Families overcome the moral hazard, adverse selection, and deception problems that plague public annuity markets. In addition, families can provide consumption insurance by diversifying sources of income within the household. Hence, a household with several different types of workers may be better insured than a more homogenous household (Rosenzweig 1988). Families may diversify income sources geographically by sending potential members away. This type of geographic diversification has been documented in both the migration and marriage literatures (Rosenzweig and Stark 1989). Lucas and Stark (1985) also report evidence on intra-family insurance by looking at remittance behavior.

In addition to household composition influencing a potential member's utility or the household's available resources, co-residency may influence a household's decision making. There are two general classes of models of household decision-making. First, the unitary household model treats a household as having a single decision maker (Becker 1981). Household composition may influence who this decision maker is. A subset of the unitary models allows altruism to enter the preferences of the unitary decision maker. Foster (1999) argues that co-resident family units care differently about the allocation of resources to each other than do otherwise equivalent family units living in separate households. In that case, household composition can influence the weight a unitary decision-maker puts on a household member's (or potential member's) preferences or well-being. Second, collective household models allow for multiple decision makers within the household (Bourguignon and Chiappori 1992). Whether the collective decision-making is modeled as a cooperative or a noncooperative process, the outcome of the model (hence the household's production and investment decisions) can depend on who is involved in the decision-making. If co-residents have a greater influence on household decisions than non-residents, then household composition will affect the allocation of resources.

There are a number of ways that the old age pension can influence household composition depending on the role one assigns to household composition in the household model. When household composition is viewed as consumption, the old age pension could increase consumption of the recipient's or the household's (depending on the decision framework within the household) preferred living arrangements. With additional cash income, migrants separated from the household for work may be able to return. Alternatively, adult children may be able to afford to take their parents into their home in order to care for them. If the pension

recipient (or her household) prefers to live with her children or grandchildren, compensation (cash or in kind) may be offered to keep these family members in the household.

On the production side, additional household income may enable the household to overcome credit constraints which prevent the household from making efficient changes to the household structure. Additional income may enable the household to hire additional labor for work inside the household or in the household's farm or business production. Children can be a source of labor that moves between households as well. In South Africa as in other African countries, many black children are fostered out – sent to live temporarily or permanently in a relative's household.[2] The pension income may enable an elderly household to take in children or grandchildren to provide non-tradable goods or services that improve the well being of their family. In data from the Cote d'Ivoire, Ainsworth (1996) finds that the receiving family's demand for child labor is an important determinant of the child fostering decision (even more so than the level of the sending family's resources).

Older grandchildren (or adult children) who had been living with grandparents (or parents) could now be free to migrate for work or education, either because they receive part of the pension or because the grandparent now has the resources to care for herself. Likewise, additional income may enable the household to send labor out to where it can get a greater return (moving to opportunity) if credit constraints prevented such a move before. Alternatively, the stability of the pension income may change the household's incentives to diversify against risk. Butcher (1993) emphasizes that child fostering may be a mechanism through which households smooth fluctuations in household resources. With the pension program, the elderly become an

[2] For examples, see Kayongo-Male and Onyango 1984, Beittel 1992, Spiegel, Watson, and Wilkinson 1996, Van der Waal 1996, and Sagner and Mtati 1999.

annuity and thereby may induce changes in the household's portfolio. This could lead to any number of changes in household living arrangements.

Finally, additional income to the elderly may change how decisions are made inside a household. Within a unitary household model, the pension income could make the elderly the primary decision-maker. Within a collective model, the influence that the pension recipient has on the allocation of resources may increase. This may lead to a shift towards the pensioner's preferred living arrangement even independent of the effect of additional income on consumption discussed above. A predatory variant on the collective model might posit that the elderly (perhaps because of age) have a weak position within the household. The large cash transfer makes the aged more valuable as an asset. Hence, predators move to the money or move pensioners to them because of the sudden increase in the value of the aged.

Thus, the theoretical effect of additional pension income on the living arrangements of the elderly is unclear. It will depend on the nature of household decision-making, the preferences of household decision-makers, the existence or non-existence of markets in goods consumed by decision-makers, and the presence of cash constraints. Most of the different mechanisms through which additional income can influence living arrangements are observationally identical. If we observe a 30 year old male move into a household with a pensioner, we cannot tell whether that reflects the preferences of the pensioner, is the resolution of a credit-constraint imposed exile, or is predatory behavior on the part of the 30 year old. Without a clear theoretical prediction, the effect of additional resources on household composition is ultimately an empirical question.

3. Research on the South African Pension Program

The old-age pension program in South Africa provides an ideal environment in which to explore how households respond to changes in economic resources. The political forces that led to the collapse of the apartheid system in the early 1990's brought many unexpected changes to the lives of South Africans, not the least of which was the vast increase in the pension receipts of almost all elderly blacks. Historically, blacks received only a fraction of the state pension that whites received, but in 1989 the government committed to making pensions equitable. By 1993 most blacks and whites eligible for the pension were receiving comparable amounts (Lund 1993, Van der Berg 1994). The maximum benefit in 1993 was 370 Rand per month (about $3 a day), about twice the median per capita monthly household income of blacks (Case and Deaton 1998). There has been a means test to the pension, but it is set as such a high level that it does not apply to most black South Africans. Further, the means test, if applied, does not take in to account income of other family members (Case and Deaton 1998), so there is no incentive to rearrange households in order to be eligible for the pension (as many have argued exists for social transfer systems such as Aid to Families with Dependent Children in the United States). Once eligible age is reached, however, pensioners appear to share their income with family members.

The earliest research on the South African Old Age Program emphasized its effect on household expenditures. Lund (1993) describes how markets develop around pensioners when they queue to receive their pensions. These markets tend to contain many goods that would either be shared by the pensioner with members of her household or would benefit individuals other than the pensioner. In addition, Lund finds that local stores tend to single out households with pensioners to make lines of credit available to the whole household. Case and Deaton (1998) find that pension income translates to increases in household expenditure across most

expenditure categories (alcohol and tobacco expenditures are a noticeable exception). They do not find that pension income is treated differently than other sources of income. This suggests, then, that pension income is shared with household members.

Other recent papers have examined the possible household behavioral responses to this large and arguably exogenous increase in income by looking at intrafamily transfers, health measures, and labor supply responses. Jensen (1998) investigates whether the pension receipts of rural black South Africans displace or 'crowd out' informal private transfers from migrant adult children; he finds that each Rand of the public pension income reduced such transfers by .2 to .4 Rand. He also finds that households receiving pensions were slightly less likely to have migrant members. Evidence in Duflo (2000) suggests that pensions received by women in black families had a large impact on the anthropometric status of girls, and a smaller, insignificant effect on that of boys; she found no effect of the pension on child nutrition when it is received by men. Bertrand, et al (2000) study household labor supply responses and find that the labor force participation of prime-age black men dropped measurably with the number of pensioners in the household. Notably, Bertrand et al find like Duflo a greater effect when the pensioner is a woman than when the pensioner is a man.

In the work cited above, as in much of the development literature focusing on household models, the analyses performed condition on household composition (although Bertrand et al fail to reject the hypothesis that there are no pension-related changes in the relevant demographic groups). There is a body of anecdotal evidence suggesting that part of the household labor supply response may be through drawing the least economically successful family members into the pensioner's household. Moller and Devey (1995) find in data from the 1993 Southern African Labour and Development Research Unit (SALDRU) household survey that 85% of

'elderly' households (which include a member aged 60 or more) have an unemployed person resident, while only 66% of 'young' households do. They interpret this as being consistent with economically less active individuals moving to rural elderly households where the cost of living is cheaper and there is a pension income. Sagner and Mtati (1999) present evidence that pensioners share pensions with their family in the hopes that the families will reciprocate later in times of need. They note, "Arguably, social pensioners act as magnets for economically weaker persons." (p. 399.)[3] And one pensioner noted, "If my granddaughter is around, she does all the housework for me and I make her happy when I get my pension money. I buy her nice things." (Moller and Sotshongaye p. 15).

The previous econometric work on the South African Old Age Pension program finds evidence that pensioners share their extra income, holding existing household structure constant. As the economic theory described in section 2 would predict, the descriptive evidence from South Africa suggests part of the household's response to additional income maybe to modify living arrangements. Because the pension program does not contain incentives to re-arrange households in order to receive treatment, it provides a uniquely appropriate environment in which to consider the effect of additional household resources on living arrangements.

4. Data

The empirical work in this study relies on the black population in the 1996 Population and Housing Census of South Africa. Statistics South Africa (Stats SA) administered the 1996 census to cover the night of October 9th to 10th; respondents either filled in the questionnaires themselves or were interviewed by enumerators. Questionnaires were translated into all 11 official languages of South Africa. Data were adjusted for undercount (with both household and person weights) on the basis of a nationwide post-enumeration survey. (Statistics South Africa.

[3] See also Moller and Sotshongaye 1996, Breslin, Delius, and Madrid 1997, and Sagner 1997, for example.

n.d. (a)) The data we use come from a 10% public use sample representing 40,578,900 individuals and 9,058,540 households. Seventy-seven percent of the population is black. Individuals not living in households but in hostels or institutions such as prisons and hospitals were counted in the census but are excluded from our analysis (Statistics South Africa. n.d. (b)).[4]

We focus on black households, because this group is most affected by the expansion of the pension program in 1993. We designate black households by the race of the household head. In a small percentage of cases where no head of household was indicated, we designated a black household by the race of the next most closely related family member (usually the spouse). The 1996 census allowed more than one household head to be reported; for the 0.4% of households with multiple heads we chose the oldest head for determining the race of the household (in the instance of a tie for oldest, we chose the oldest male). We are left with 2,587,285 (unweighted) individuals living in households with black heads.

Table 1 presents the mean and its standard deviation for various demographic characteristics of individuals living in black households with and without a person who is age-eligible for the pension. Table 1 previews many of the results that we find below. Pension eligible individuals live in larger households. Almost all of the pensioners live with some form of extended family. Hence, the average age of an individual in a pensioner's household is still under 30. There are more young children (0-5) and primary school age (6-11) children in households with a pension eligible individual. In addition, we observe slightly more persons of prime work age (30-39) in pension eligible households.

[4] 3% of the sample is dropped, because age information is missing for one or more household members. In less than 1% of the households we do not know the relationship between an individual and the household head; we continue to count these individuals in our demographic categories (75% of these individuals were aged 22 or younger). We dropped 62 individuals in households with at least one member indicating institutional residence (this being inconsistent with our sample definition).

This naive comparison of households with and without a pension eligible person is problematic. By definition, a pension eligible household contains at least one female over 60 or male over 65, and non-eligible households do not contain anyone that old. Older individuals will have older children, older grandchildren, and great-grandchildren. Hence, the differences in table 1 may owe entirely to age and have nothing to do with the pension. This problem of separating age differences from the effect of the pension benefit is the reason for using the regression discontinuity estimator developed in the next section.

5. Methodology

Our identification strategy in this paper is to test directly for changes in the living arrangements of the elderly at the ages of pension eligibility. The regression discontinuity design is used in situations where the probability of receiving treatment (in our case pension receipt) changes discontinuously at a known point. Unlike a pure experimental design, the "control" and "treatment" groups may vary in systematic ways – so it may be that women who are over 60 live in households which are different than those of women under 60, even if no pension program exists. However, it is still possible to estimate locally the impact of treatment under weak assumptions. Two conditions are needed for the regression discontinuity design. The first is that if there were no treatment, the outcome of interest would be continuous at the cut-off criterion. The second is that the probability of treatment varies discontinuously at this cut-off point. The case of the old age pension precisely fits this situation: while we expect that some household characteristics (number of small children in the household, for example) may vary with the age of the oldest woman, there is no a priori reason to think that at age 60 (and not 59, and not 61) household characteristics should change dramatically. However, we know from

the rules of the pension program that pension receipt should begin at age 60 for women (and at age 65 for men).

To implement the regression discontinuity approach, we use the partially linear estimator developed by Porter (1999). Porter's idea is to estimate the relationship between the outcome variable (household composition characteristics) and the treatment variable (age) nonparametrically, but to allow a parametric jump in the outcome variable at the point of discontinuity in treatment (age 60 for women, 65 for men). In our context, the household composition outcome variable of interest, y, for a person with a given gender and age *a* depends on resources and characteristics of the household and all of its potential members. We need the effects of these (unobserved) resources and characteristics on household composition to be smooth with respect to an individual's age. In that case, we write the expectation of the outcome variable *y* conditional on age and gender as:

$$(1) \qquad E\left[y \middle| a \right] = m(a) + d\alpha + \varepsilon$$

where *d* is an indicator that is 1 if a person is pension eligible, $E\left[\varepsilon \middle| A \right] = 0$ (where $a=A$ is the age of pension eligibility), and $m(\ )$ is continuous at the age of pension eligibility. Throughout this paper, we estimate both $m(\ )$ and α separately for each gender. Hence, for economy, we have suppressed gender subscripts in all elements of (1). Porter suggests rewriting this conditional expectation as:

$$(2) \qquad (y - d\alpha) = m(A) + \varepsilon \ .$$

We have a new dependent variable: $(y - d\alpha)$. α is estimated by finding a value of α that minimizes the averaged squared deviation between this new dependent variable and the nonparametric estimate of $m(A)$. Thus, Porter's estimator uses data from both above and below

the discontinuity in a kernel regression, with more weight given to data points closest to the discontinuity, to estimate the change in the conditional expectation of the household composition variable at the age when an individual becomes pension eligible.[5]

There are three issues that could cause our approach to give misleading results. First, households may respond to the pension either before or after actual eligibility. For example, if pension receipt tends to cause relatives to move in with an elder woman, they may do so when she turns 59, in anticipation of her becoming eligible. Alternatively, it may take a several years for household members to adjust their composition in response to pension eligibility. Both of these timing issues attenuate any measured program impact.

Second, there is a large amount of clustering in reported ages in our data, especially in ages rounded to the nearest decade such as 60. This poses several problems for our analysis. For example, many women younger than 60 (thus not pension eligible) may report their age as 60. We, then, miscode them as pension eligible. A second issue is that age clustering may induce a type of selection bias into our estimates if people who are likely to misreport their ages have different household characteristics than those who report accurately. Due to concern for these two biases, we omit individuals whose age is exactly that of the pension cutoff age. That is, we omit 60 year olds when estimating the pension impact for women, and 65 year olds when estimating the impact for men.

Third, retirement, absent the pension program, may complicate the application of the regression discontinuity methodology. If women stop working at 60 and men at 65 without the pension program, the effect of retirement is intertwined with our estimates of the effect of the

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[5] In our kernel estimation, we use an Epanechnikov kernel (specifically: $K(z) = (.75)\left(1 - .2z^2\right)/\sqrt{5}$ for $|z| < \sqrt{5}$) and a bandwidth of 2. Standard errors are calculated following the discussion in Porter (1999). For reference of future researchers interested in using the Porter regression-discontinuity methodology, we note that we compute the constant c_k in Theorem 1 (Porter 1999, pg. 12) to be 6.3582 for the indicated form of the Epanechnikov kernel.

pension program. This is impossible to test in our data, because the income effect of the pension may also lead to a labor supply response (retirement induced by the pension does not complicate our methodology). While we are concerned about this problem, we do not feel this substantively alters our conclusions below (especially for women). First, most black South Africans are not engaged in formal employment that would have a fixed retirement age. This is especially true for women, most of whom report no income prior to receiving the pension (below). Second, most black South Africans are sufficiently poor that they do not have the luxury of terminating employment at a specified age (in the absence of the pension). Typically, poor health and other problems associated with aging force the elderly out of employment, but there is no reason why we should expect health shocks for women at 60 versus 59 or 61.

To consider the magnitude of this retirement problem, we classify an individual as unlikely to have a formal retirement age if that individual is engaged in agriculture, fisheries, or what the South African Census classifies as "elementary occupations." We define this set of occupations as "informal" from the perspective of retirement. We believe this is a conservative estimate of the set of individuals who are unaffected by retirement ages. Individuals employed in many services, trades, or construction are also unlikely to have formal retirement programs at the age of pension eligibility, but we do not include them in this informal group. We hypothesize that formal retirement should only affect individuals engaged in formal work, while both formal and informal workers should experience the retirement induced by the pension program. Thus, we use the change in work that we observe in the informal sector as a baseline to compare to the change in work that we observe in the formal sector where formal retirement is more likely. Any extra change that we observe in the formal sector could be attributed to formal retirement. Of

course, this assumes that the pension affects formal and informal workers similarly and that the pension does not induce workers to switch from one category to another.

Formal retirement appears to be more likely to be a problem in the results for pension aged men than in the results for women. In the South African Census of 1996, 12.7% of black women aged 59 are working in what we have defined as informal work and are hence unlikely to face formal retirement (only 19% of black women are working at age 59, so 66% of these participate in informal work). We observe that the fraction of women in informal work declines to 6.8% of the population by age 61. This represents a relative decrease of 46%. For women who are working in occupations potentially subject to formal retirement, we observe a drop from 6.6% of the population aged 59 to 3.9% of the population aged 61, a relative decrease of 41%. Thus, formal retirement is unlikely a serious complication in our results when women become pension eligible. The numbers for men are less encouraging. 8.4% of black men age 64 are working in informal work at age 64 (26% are working; thus 32% of working black men age 64 engage in informal work). By age 66, the fraction of men working who are in our informal category, drops to 6.0%, a relative decline of 29%. For men working in possibly formal occupations, the drop is from 17.2% of the population aged 64 to 9.4% of the population aged 66, a relative drop of 41%. Thus, we regard our results for men with greater caution.

6. Results

6.1 Income

We expect pension eligibility to be associated with changes in household composition, because of the extra income associated with pension take-up. For this mechanism to be plausible, we must observe changes in income at the pension eligibility ages. Hence, we begin by examining the evidence that pension eligibility results in the receipt of pension income. The 1996 Census

does not ask about income by source of income. Instead, the questionnaire asks each member to select a category for her total income over the past year.[6] Thus, if a person was receiving the pension she should have at least as much income (on the order of R370/month) as was the pension payout. First, we consider the probability that a person reports not receiving any income in the previous year. Then, we examine the amount of income reported.

In figure 1, we focus on women. Using our semi-parametric regression discontinuity estimator, we regress an indicator that is 1 if a woman reports no earnings in the last year against age, allowing a change (as described in the previous section) at age 60 when a woman becomes pension eligible. Figure 1 plots the estimated relationship between zero earnings and age. In addition to the regression line, we have also plotted by age (unsmoothed) means of the indicator that a woman reports zero earnings.[7] Figure 1 shows that about 58% of women in their early 50's reported zero income during the previous year. This proportion declined slightly among women in their late 50's, and dropped dramatically - 12.5 percentage points – as women become age-eligible for the pension. The T-statistic for the null hypothesis that the jump at age 60 is equal to zero is over 15.

In fact, the extra earnings that individuals attain at pension age correspond (roughly) to the amount of the pension. In figure 2, we consider the probability of a woman reporting her income to be in a given income category. There are four income categories in figure 2: no income, R1-R200/month, R201-R500/month, and greater than R500/month. Category 3 (R201-

[6] Specifically: "Think of the past year (1 October 1995 to 30 September 1996) and the money each person received. **Please indicate each person's income category before tax.** Answer this question by indicating each person's weekly, monthly, or annual income. Include all sources of income, for example housing loan subsidies, bonuses, allowances such as car allowances and investment income. If the person receives a **pension or disability grant**, please include this amount." [bold and underline in original]

[7] In all pictures where we show the regression lines estimated by the Porter methodology, we also include the unsmoothed, by age means that are weighted to be nationally representative of the non-institutionalized population with a black head. Further, with the regression lines, we include the estimate of the discontinuity (alpha) and the T-statistics associated with the null hypothesis that the discontinuity (alpha) is zero.

R500) is the category that contains the pension level of monthly income. Panel A replicates Figure 1. Panel B shows that there are very few women reporting incomes of R1-R200/month, and there does not appear to be a strong change at the pension age. Panel C (where the maximum pension benefit falls) demonstrates a large jump of 13.3 percentage points at age 60, with a T-statistic of 16. We take this, as well as Panel A, as evidence that there are women receiving the pension at age 60. Finally, Panel D indicates a moderate (1.4 percentage point) and significant (T=2.33) drop in the fraction of women earning more than R500/month. This may be related to retirement, perhaps induced by pension receipt (given our discussion in the previous section). Results for men are presented in Figure 3. This shows patterns similar to those for women. There is a discontinuous drop (6.1 percentage points) in the fraction of men reporting zero income at the pension eligible age of 65 and an increase in men reporting the R201-R500/month category.

We note that for both men and women the measured effects of pension eligibility on income, while readily apparent, are not as strong as our knowledge of the pension rules would indicate they should be. For example, even after the discontinuity, about 33% of women still report zero income. We cannot identify whether this is due to misreporting or genuine pension non-receipt. This zero-income is not the result of smoothing in our regression estimator, because it is also evident in the unsmoothed means where zero-income rates persist at 20% even into the late 70's. This is further evidence that our estimation methodology, by focusing attention on the data just surrounding the eligibility point, is apt to produce coefficients of a muted magnitude.

6.2 Household Composition for the Population of Elderly Black South Africans

Most of the changes in household composition associated with pension eligibility appear to be from age and sex specific flows in and out of the household rather than through overall

increases or decreases in household size. Figure 4 applies our regression discontinuity approach to household size. Panel A considers household size against age for women, and panel B is for men. In panel B, there is clearly no change in household size when men become pension eligible. The increase in household size with pension eligibility is larger for the elderly women pictured in panel A. However, the increase of .068 people is not significant at conventional significance levels and corresponds to approximately a 1% increase in household size.[8] The point estimates of the size of the discontinuity are reproduced in table 2, row 1.

Most of the literature on elderly living arrangements emphasizes their preferences for independent living. While figure 4 suggests that household size on average does not change at pension eligibility ages, this could result from certain elderly households decreasing in size as the elderly begin living independently while other elderly households expand. However, we do not find any evidence of this. In table 2, rows 2 and 3, we consider whether an individual lives alone or with a spouse. Living alone or only with a spouse is extremely rare in South Africa (only 8% of women age 59 live alone or with only a spouse), and there is a general trend with away from living alone or only with a spouse. We do not observe any deviation from this trend associated with pension eligibility. One caution is that the elderly who live alone may continue to live independently prior to pension receipt in anticipation of the pension income; this would appear as no effect in our results. Curiously, however, in Lund (1993) there is no descriptive evidence of individuals being able to borrow in advance of pension eligibility even though pensioners have better access to credit once they have attained eligibility. In addition, since so

[8] In order to calculate percentage changes, we need to know what the mean household size (or any other dependent variable) would be at the age of pension eligibility in the absence of the pension program. To calculate this, we use the age trend before pension age to project the expected household size at the pension age. These projected means are the means reported in tables 2-5 and referenced henceforth in the text.

few elderly live alone in South Africa and, if anything, living alone is negatively associated with pension eligibility, we can rule out an increase in independence as a result of the pension.

Within the household, we find increases in young children and decreases in women 30-39 when women become pension eligible. Figure 5A plots the relationship between the number of children age 0-5 in the household and age of the elderly female. We see a large and statistically significant increase of .047 in these children at the age of pension eligibility, corresponding to a 6% increase in the number of young children in a household on average. This increase in the number of young children does not appear to be associated with an increase in births. In column 1, row 4 of table 2, we find that there is a small and statistically insignificant decrease in births in households where an elder woman becomes pension eligible. Thus, this increase in young children is attributable either to children being brought in or births occurring in anticipation of the pension. For the number of co-resident women of prime working age (30-39), we find large declines when an elder woman becomes pension eligible. This is evident in table 2, column 1, row 9 and in figure 6A. The number of co-resident women 30-39 declines by 13% when a woman reaches pension age. This pattern of a decline in women 30-39 and an increase in children 0-5 is consistent with prime age mothers sending their children to live in their grandmother's household so the mothers can work and with some prime age women leaving their elderly mother's households to work but leaving their children behind. Interestingly, the estimated departure of men age 30-39 associated with a woman achieving pension eligibility is much less (and statistically insignificant) than that observed for women (table 2, column 1, row 10, and Figure 6B). We find no evidence of an increase in primary school age children or children ages 12-17 when elder women become pension eligible (table 2, column 1, rows 6,7,and 8).

The patterns for men at pension eligibility appear different than for elder women. Figure 5B pictures the number of children 0-5 against the age of an elder man. We do not see any significant change in the number of children 0-5 at age 65. However, in table 2, column 2, row 6, we see large and statistically significant increases in the number of co-resident primary school age children. We also find increases in the number of boys and girls age 12-17 (rows 7 and 8), although the results are only significant at the 95% confidence level for boys. The results differ in both statistical significance and in the magnitude of the discontinuity estimate from what we observed in elder women. In figure 6C and table 2, column 2, row 9, we do not observe any change in the number of prime-age women when an elder man achieves eligibility. However, we observe a 19% drop in the number of men age 30-39 (figure 6D and table 2, column 2, row 10).

Figures 7 and 8 explore these age-gender differences in finer detail. We create a set of age group – gender counts and run a series of regression discontinuity treatment effect estimates for each of the different age group-gender possibilities. For example, we first estimate the impact of women's pension eligibility on the number of 0-5 year old girls in the household. We next estimate the impact of eligibility on the number of 6-11 year old girls and so on through age group 45-49. The by age group changes in the number of females associated with the pension eligibility of a woman are in figure 7. For each age group, we report the estimated change in the number of females in that group on top and the t-statistic associated with the hypothesis that the change is zero on the bottom. Each point in the figure represents the coefficient from one estimation. The point (X,Y) has the interpretation of "the number of X-year olds in the household increases by Y as a woman becomes pension eligible." With female pension eligibility, the largest changes appear to be an increase in women 18-23 and decreases in women 35-39. Of course the distribution of ages in the population is not perfectly uniform, and the

results in figure 7 would be easier to interpret if expressed as percentage changes. Hence, in figure 8, we present the semi-elasticities calculated by dividing the discontinuity estimate by the mean number of children of a given age and gender that are co-resident with elderly at the age of pension eligibility. Figures 8A and 8B correspond to the by age results when women become pension eligible, and Figures 8C and 8D are for pension eligible men. Figures 8A and 8C are age-by-age changes in the number of women household members, and 8B and 8D are changes in resident men. Thus, the semi-elasticities in figure 8A are based on the estimates of α in (2) pictured in figure 7.

In figure 8, we observe significant changes by age group for both male and female household members when women become pension eligible, but when men become pension eligible we observe significant changes by age group for males only. This finding is consistent with Bertrand et al (2000) who find evidence that women are more apt to share their income with both sexes. There are two main changes that appear to take place when women become pension eligible. First, when women become pension eligible, we observe increases in the number of 0-5 year olds (this is only significant for boys) and 18-23 year olds (figures 8A and 8B). This increase in 18-23 year olds and 0-5 year olds may be related, because it is not unusual for a black woman in the 18-23 year old group to have a child, and that child is likely to be below the age of 5. Second, we observe decreases in the number of women 35-39 and 45-49 as well as men 35-49.[9]

[9] In figure 8B we see a positive insignificant effect of a woman becoming pension eligible on number of male household members age 30-34 and a negative significant effect on men 35-39. Merging these two groups leads to the negative, insignificant effect of pension eligible women on number of men 30-39 in the household that we saw in table 2, column 1, row 10. Comparing figures 8A and 8B suggests that there are similar negative effects of pension eligible women on both men and women in these prime ages, rater than the contrasting effects that appeared using the broader age groups in Table 2.

We observe a different pattern of age group changes when men become pension eligible. First, when men become pension eligible, the increase in children 6-17 is for both boys and girls, but the magnitude of the effect is greater for boys. Second, the decline in the number of male household members in households where men become pension eligible appears to begin at age 18, and is statistically significant for men 24-39. The changes in women that we observe for male pension eligibility are generally smaller in magnitude and not statistically significant. The contrast in the results with male and female pension eligibility is striking. While for both male and female pension eligibility, we observe a departure in men and women who are at prime-working ages and have potentially completed fertility, we observe increases in young children and young (potential) mothers when women become pension eligible. However, we observe increases in ages of children who are potentially productive when men become eligible. Other studies of African populations find that these age groups are particularly affected by child fostering for work (e.g. Butcher 1993).

Our finding of the exit of prime age adults from pension eligible households is potentially important to the existing literature on the effect of income on consumption, remittance, and adult labor supply. In particular, we wish to know if the net outflow of prime age adults is related to adult earnings opportunities. To understand who these departing adults are, we divide the 30-39 age group into three education categories. In stratifying the data in this way, our assumption is that pension eligibility has no effect on the education of this group. We are missing education data for 2% of the census, so those individuals are necessarily dropped from this analysis. The three education categories are chosen by dividing the distribution of education among blacks in South Africa into 3 equally sized groupings. Our results, estimated separately for men and women, are in table 3.

As in table 2, the strongest effects of female pension eligibility are on women and of male pension eligibility are on men. In general, the net flows in household composition appear to be more or less equal across education categories, although results vary somewhat across demographic groups. When we consider female eligibility, the number of women in each education category declines. The magnitude of decline for the middle education group is somewhat smaller than that of the least educated or the most educated, but there is not a systematic tilting toward more or less educated individuals. For changes in the number of men 30-39 with female eligibility, we observe substantially smaller and uniformly insignificant changes for all education groups. With male pension eligibility, we observe modest but significant increases in the presence of least educated women, and small and insignificant declines in the other education groups. For men, we find declines of similar magnitude across all education categories. As such, we find no strong evidence that the net outflow of prime age adults associated with the pension comes from one particular education category, indicating that differential earnings opportunities (if associated with education) are not necessarily driving these flows.

6.3 For Non-Migrant Elderly

One obvious question is whether these results are driven by the mobility of the elderly upon receiving the pension or by changes in household composition conditional on the elderly not moving. The pension may be associated with migration in the elderly, so both explanations for changes in household composition are feasible. In fact, preliminary work indicates that receipt of the pension is associated with increased migration of the elderly, a topic we are examining further in work in progress (Edmonds and Miller 2001).

In order to try to understand whether our results hold true for the non-migratory elderly, we re-estimate selected models conditioning on those elderly who live in rural settings and who do not report moving within the past seven years. This selection drops approximately one half of the observations used in the previous section. The obvious problem with this conditioning is that part of the effect of the pension may be to induce the elderly to change their location. As such, the pool of non-migrating elderly who are pension eligible may be different than the pool of non-migrating elderly who are not yet pension eligible. Even absent an effect of the pension on elderly migration, the elderly who have not moved within the last seven years may be different that the more mobile elderly. Results that we find for this group therefore might not extend to the mobile elderly. Thus, this conditioning does not enable us to fully identify what part of our results above come purely from the migration of the elderly. Still, we believe it is informative to see whether our unconditional results also appear in the non-migratory elderly population. These results are in table 4.

When we consider the household composition results, we find that they are robust to conditioning on rural non-movers. Again, total household size does not seem to change. Also, for women pension eligibility is associated with large increases in the number of children 0-5 and decreases in the number of women 30-39. The magnitudes and statistical significance of these effects are comparable to the unconditional estimates. With men becoming pension eligible we find possible decreases in the number of young children (although in this sample they are not estimated to be statistically significant) and strong decreases in the number of men 30-39. Thus, the migration of the elderly does not appear to be the driving force behind our results in this paper.

7. Conclusion

The effect of additional income on the living arrangements of the elderly appear modest but substantive. Although changes in household size are small and statistically insignificant, we find interesting dynamics within the household. In contrast to much of the literature in other countries, we do not find evidence that the additional pension income leads to an increased propensity to live alone. Instead, our findings emphasize the role that the elderly play in supporting their extended families. One possible explanation for this might be the interconnectedness of extended families in South Africa (as is found in many very poor countries). While the additional pension income may be sufficient for the elderly to live independently, the aged feel responsible to the extended family, and hence are less apt to choose to live independently. Hoerger, Picone, and Sloan (1996) and Wolf and Soldo (1988) emphasize that in the United States, the marginal effect of income on the propensity to live alone is dwarfed by the importance of child attributes in the living arrangements of the elderly. It may be that the elderly and their extended family are altruistically linked in such a way that the needs of the extended family in the relatively destitute black South African population dwarf any desire of the elderly to live independently. Another explanation for the absence of any increase in elderly independent living could be that the structure of the black South African household is substantially more diverse than in most western, developed countries because of the coresidence of extended families. Thus, while most of the elderly begin living independently in the studies that find a positive relationship between income and independence, independence would be a radical change in household structure for elderly blacks in South Africa. Hence, instead of leaving their extended families, the elderly in South Africa are able to adjust the composition of their household in response to the additional income.

These adjustments take several forms, and they appear to depend on the gender of the pension recipient. First, individuals in their 30s seem to depart the household when the elderly become pension eligible. This age group is likely to have potentially the greatest market value as workers. We observe differences in the net outward movement of these prime age adults based on the gender of the pension recipient. When a male become pension eligible, we observe declines in men and an increase in uneducated women. When a female becomes pension eligible, we observe declines in both men and women, but the decline in the number of women is larger. This sex typing in the sharing of the pension benefits has been observed in other papers on the South African pension program (e.g. Duflo 2000, Bertrand et al 2000). Second, when elder women become pension eligible, we observe increases in young children (0 to 5) and in women that are of age to be these children's mothers (18 to 23). This is consistent with the caretaking role of grandmothers emphasized in some of the descriptive literature. We observe no changes in household composition for either of these groups when men become pension eligible. Third, when elder men become pension eligible, we see increases in older children (6-17). The increase is larger for boys than for girls. These children could be brought in for work or for schooling, but we do not observe any increase in the number of children of these ages when women become pension eligible.

These findings likely represent a lower bound on effect of the pension, because our approach can only consider changes in household composition local to the age of the onset of the pension. Nevertheless, the modest but meaningful changes in household composition have several implications. First, there is a large literature documenting that grandparents help their grandchildren (and hence that social policy directed at the elderly benefits other generations). This paper highlights two additional mechanisms through which elders assist other generations:

enabling individuals to move out of the household and others to move in. Second, changes in household composition appear to be an optimal household response to income changes. Hence, in addition to the usual distortions imposed by government transfer programs, policy that is conditioned on household composition may introduce an additional distortion to a household's optimal response to income changes. For example, targeting cash transfers to children but varying transfers with household size or limiting transfers to households with single parents may prevent individuals from adjusting their living arrangements in response to income fluctuations. Moreover, it is conceivable that policymakers may have preferences for the types of adjustments in household composition that appear to take place with cash transfers to the elderly. Third, the systematic changes in household composition in response to income changes highlight the fact that household members move in and out as a response to fluctuations in income. This ability to shift household composition has implications for both the consumption, insurance, and labor supply literatures that require further investigation.

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Table 1:
Household Characteristics

Variable	Individual lives in household without pension ager	Individual lives in household with pension ager
# Individuals - unweighted	1,881,267	706,018
Age	22.29	29.69
	(15.57)	(24.47)
Female	0.520	0.561
	(0.500)	(0.496)
Household size	5.64	7.59
	(2.94)	(3.71)
# kids 0-5 in HH	0.95	1.32
	(1.05)	(1.26)
# kids 6-11 in HH	0.98	1.25
	(1.05)	(1.26)
# women 30-39 in HH	0.40	0.41
	(0.55)	(0.64)
# men 30-39 in HH	0.28	0.30
	(0.51)	(0.58)

Standard deviations in parentheses. All means are weighted to be representative of the black headed population of households in South Africa. All data are from the 1996 population census, black headed households.

Table 2: Regression Discontinuity Estimates of the Effect of Pensions on Household Size

Variable	Women	[mean]	Men	[mean]
Household size	0.068	[5.83]	0.023	[5.79]
	(1.21)		(0.30)	
Lives Alone	-0.001	[0.05]	-0.008	[0.09]
	(0.17)		(1.28)	
Lives alone or with spouse only	0.006	[0.08]	-0.004	[0.15]
	(1.23)		(0.44)	
Births in HH in past year	-0.007	[0.09]	0.016	[0.08]
	(1.03)		(1.63)	
# kids 0-5	0.047	[0.78]	-0.006	[0.70]
	(2.54)		(0.24)	
# kids 6-11	-0.008	[0.86]	0.062	[0.76]
	(0.43)		(2.61)	
# girls 12-17	0.005	[0.40]	0.025	[0.37]
	(0.44)		(1.65)	
#boys 12-17	0.000	[0.38]	0.031	[0.37]
	(0.03)		(2.04)	
#women 30-39	-0.040	[0.31]	0.004	[0.25]
	(4.19)		(0.37)	
#men 30-39	-0.008	[0.26]	-0.048	[0.23]
	(0.93)		(4.30)	

T-statistics in parentheses. RD estimates of treatment effects estimate the jump in household size associated with turning 60 for women and turning 65 for men. RD methodology based on Porter (2000), as discussed in text. Conditional mean at treatment age of LHS variable in brackets. Data are from the 1996 census, black individuals living in black headed households. Estimates are based on samples of 1,377,502 females and 1,209,783 males, with 62,745 women and 31,588 men having nonzero weight in estimation of treatment effect.

Table 3: Regression Discontinuity Estimates of the Effect of Pensions on The Number of Co-resident Persons Age 30 to 39 by Gender and Education

	Woman at Pension Age	[mean]	Man at Pension Age	[mean]
Women 30-39				
No Education - Standard 3	-0.019	[0.12]	0.015	[0.09]
	(3.07)		(2.01)	
Standard 4 - Standard 6	-0.005	[0.11]	-0.007	[0.07]
	(1.06)		(1.18)	
Standard 7 and Above	-0.019	[0.11]	-0.004	[0.09]
	(3.29)		(0.56)	
Men 30-39				
No Education - Standard 3	-0.002	[0.10]	-0.020	[0.09]
	(0.27)		(2.96)	
Standard 4 - Standard 6	0.000	[0.06]	0.010	[0.06]
	(0.02)		(1.93)	
Standard 7 and Above	-0.005	[0.08]	0.016	[0.08]
	(0.96)		(2.40)	

T-statistics in parentheses. RD estimates of treatment effects estimate the jump in household size associated with turning 60 for women and turning 65 for men. RD methodology based on Porter (2000), as discussed in text. Conditional mean at treatment age of LHS variable in brackets. Data are from the 1996 census, black individuals living in black headed households. Estimates are based on samples of 1,377,502 females and 1,209,783 males, with 62,745 women and 31,588 men having nonzero weight in estimation of treatment effect.

Table 4: Regression Discontinuity Estimates for Non-migrant Pensioners
Conditioning on not moving within 7 years and living in Rural area

Variable	Women	[mean]	Men	[mean]
Household size	-0.100	[6.09]	-0.112	[6.17]
	(0.13)		(1.07)	
# kids 0-5	0.064	[0.89]	-0.063	[0.85]
	(2.42)		(1.89)	
# kids 6-11	-0.041	[1.00]	0.044	[0.89]
	(1.58)		(1.30)	
# women 30-39	-0.035	[0.26]	0.001	[0.22]
	(3.80)		(0.05)	
# men 30-39	-0.019	[0.22]	-0.048	[0.20]
	(1.69)		(3.42)	

T-statistics in parentheses. RD estimates of treatment effects estimate the jump in household size associated with turning 60 for women and turning 65 for men. RD methodology based on Porter (2000), as discussed in text. Conditional mean at treatment age of LHS variable in brackets. Data are from the 1996 census, black individuals living in black headed households. Estimates are based on samples of 691,908 females and 581,455 males, with 35,496 women and 17,599 men having nonzero weight in estimation of treatment effect.

Figure 1: Kernel regression estimates of the by age probability that a woman reports an income of 0 (with a discontinuity at age 60)

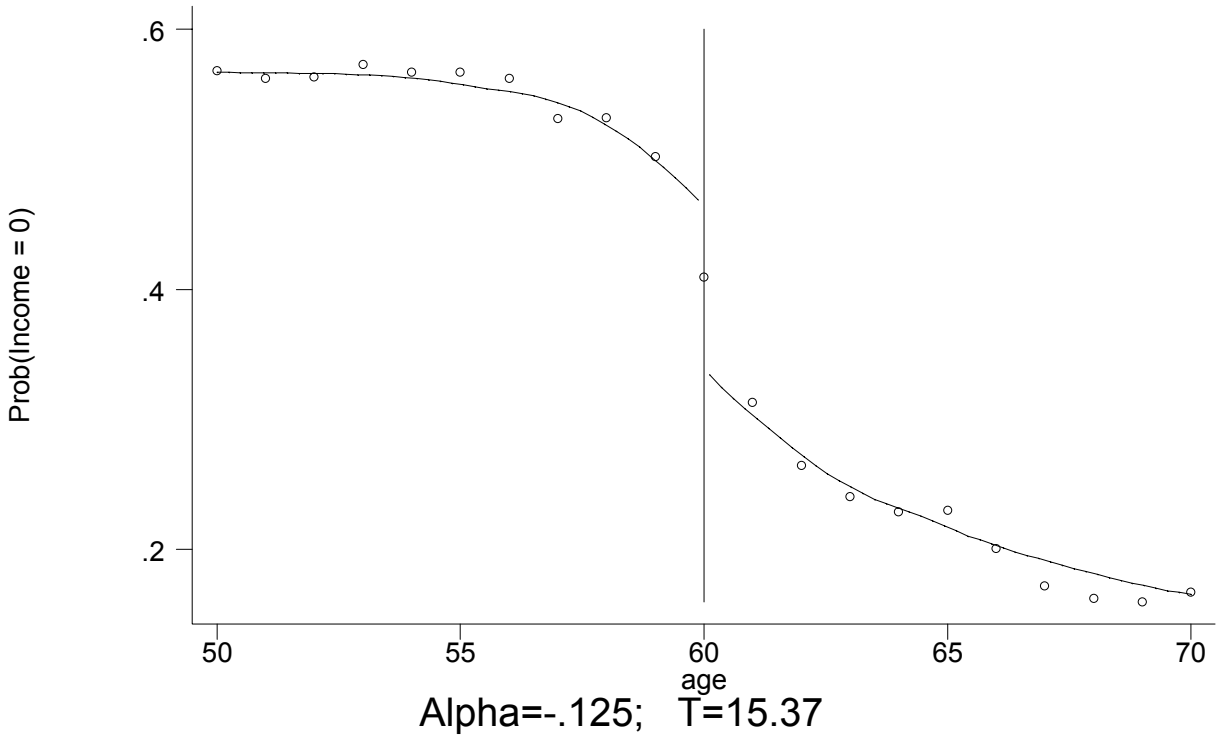
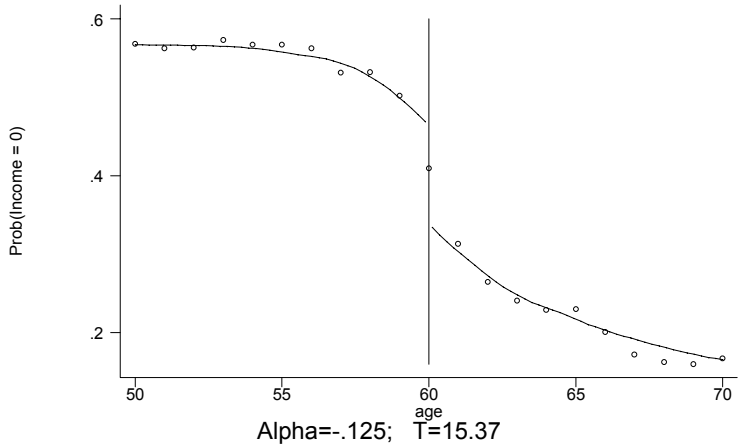
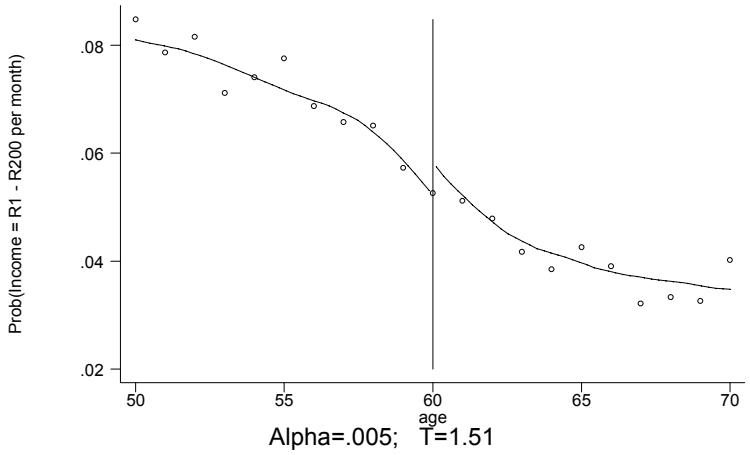


Figure 2: Kernel regression estimates of the by age probability that a woman reports being in a given income category (with a discontinuity at age 60)

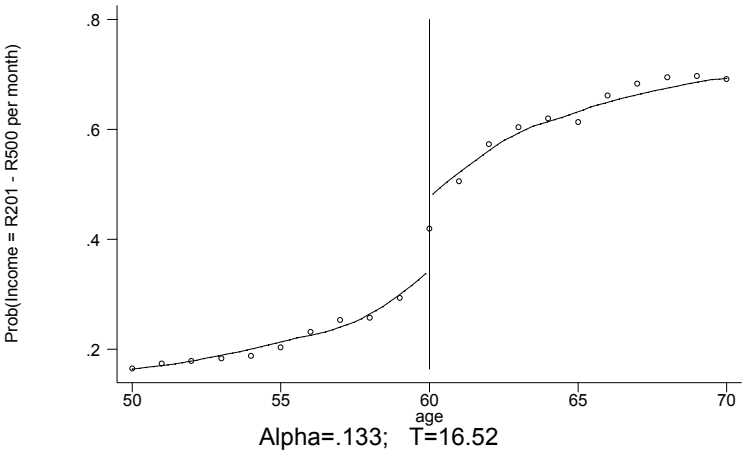
A. No Income



B. R1-R200 per month



C. R201-R500 per month



D. Greater than R500 per month

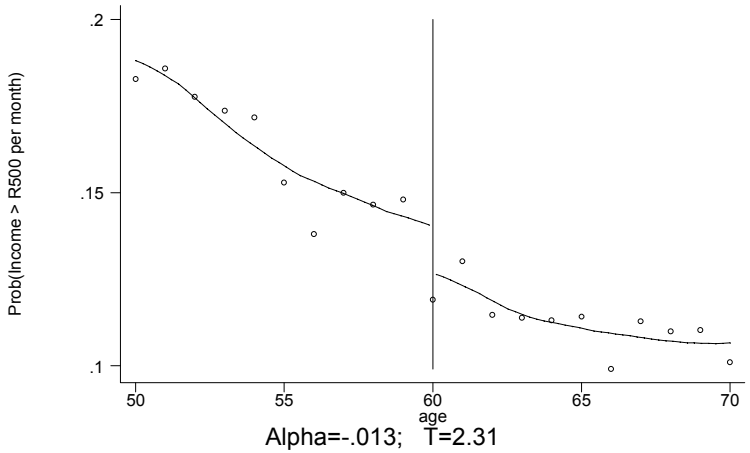
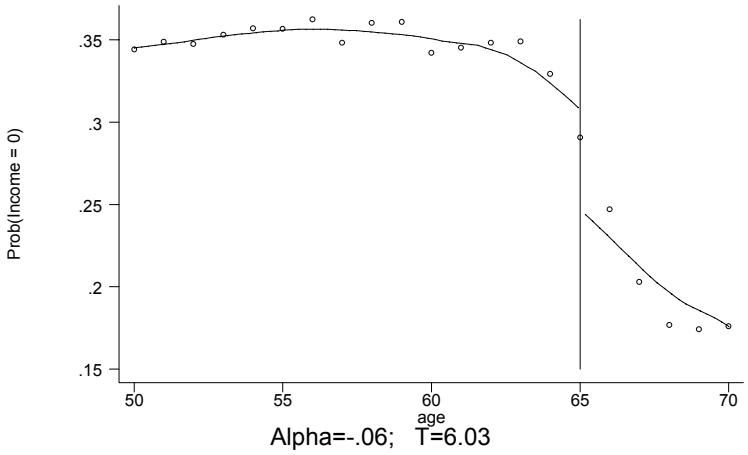
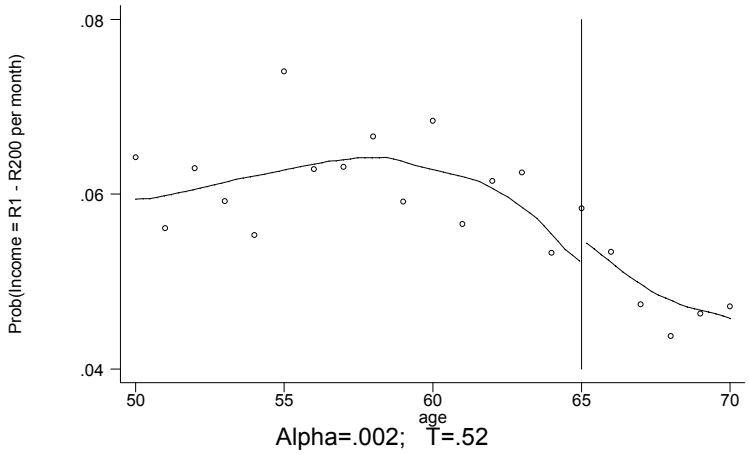


Figure 3: Kernel regression estimates of the by age probability of that a man reports being in a given income category (with a discontinuity at age 65)

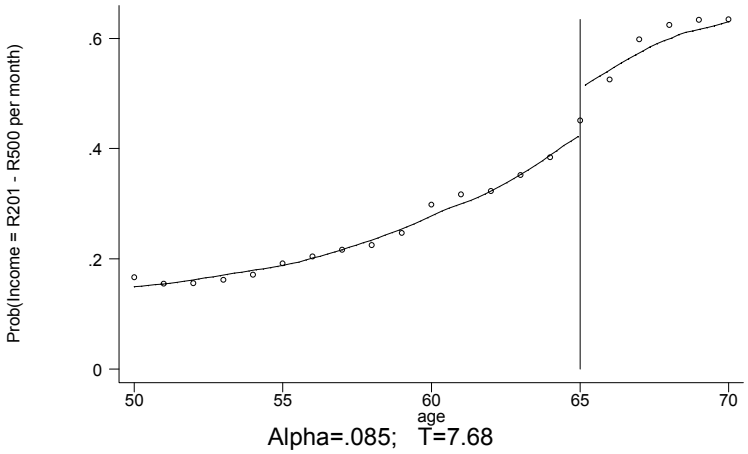
A. No Income



B. R1-R200 per month



C. R201-R500 per month



D. Greater than R500 per month

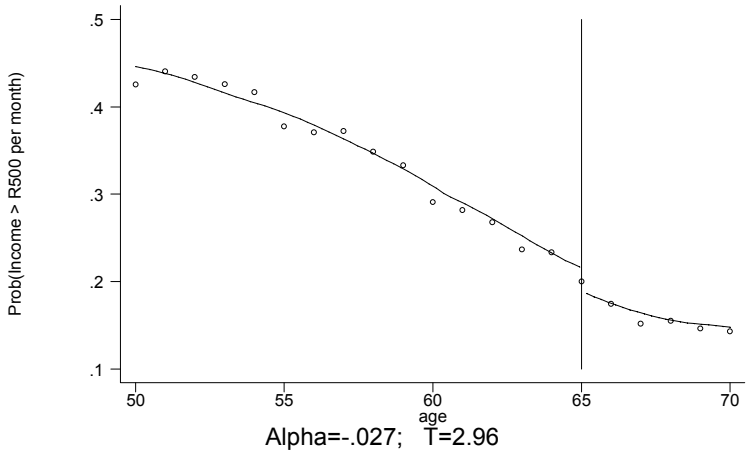


Figure 4: Kernel regression of household size on age by gender (with a discontinuity at the age of pension eligibility)

A. Women (pension eligibility at age 60) **B. Men (pension eligibility at age 65)**

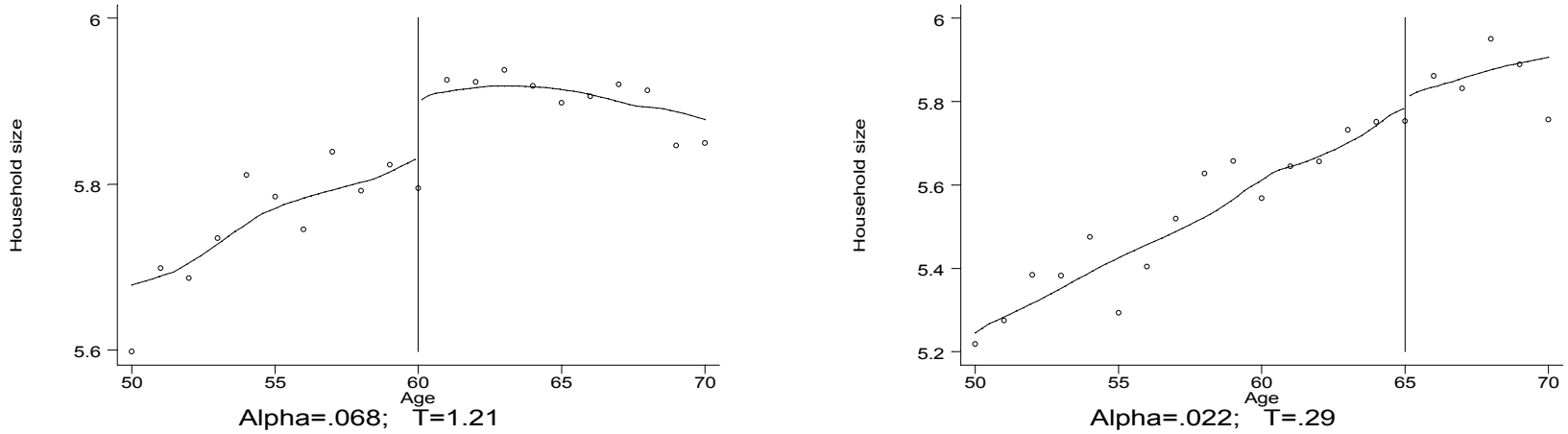


Figure 5: Kernel regression of the number of kids 0-5 in an individual's household on age by gender (with a discontinuity at the age of pension eligibility)

A. Women (pension eligibility at age 60) **B. Men (pension eligibility at age 65)**

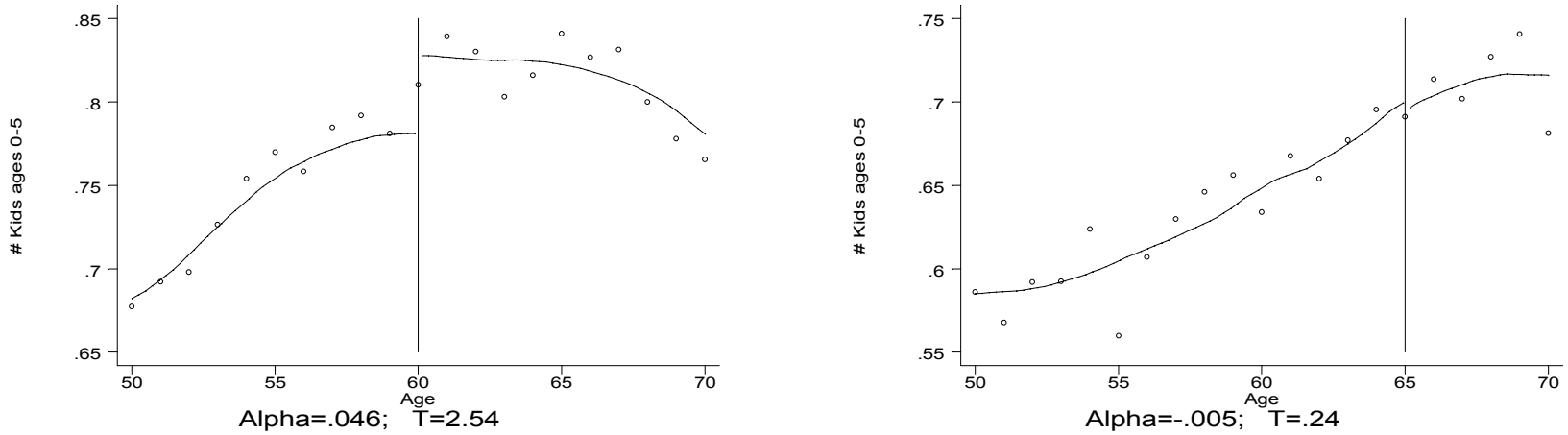
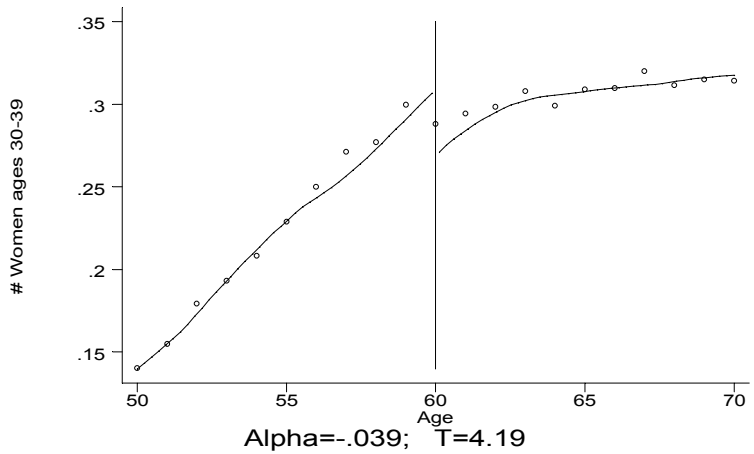
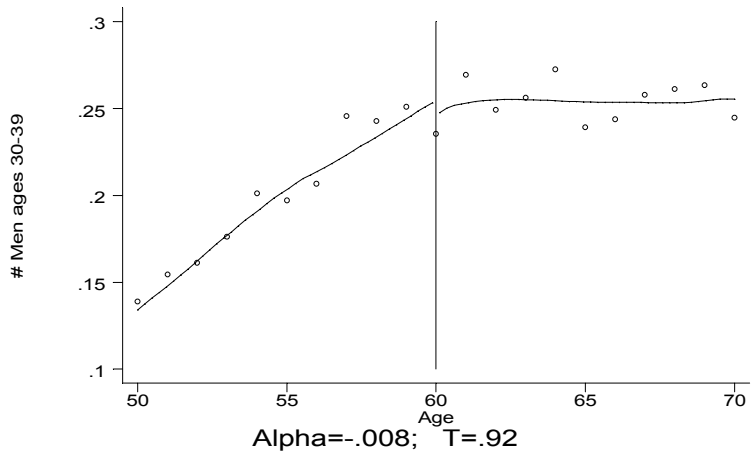


Figure 6: Kernel regression of the number of adults 30-39 in an individual's household on age by gender (with a discontinuity at the age of pension eligibility)

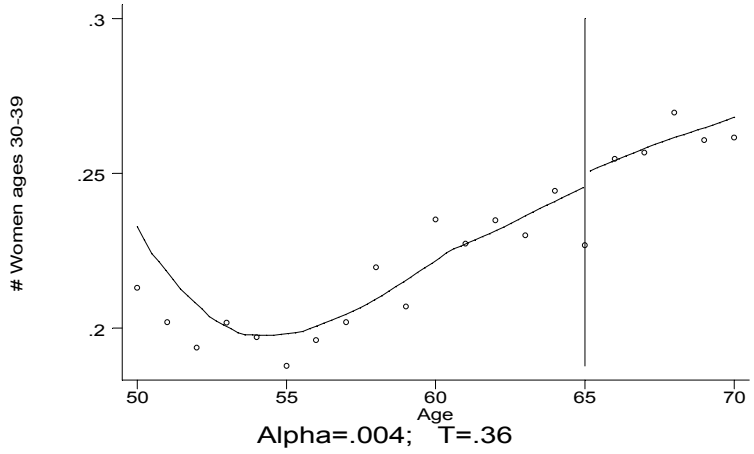
A. Number of females by age of woman



B. Number of males by age of woman



C. Number of females by age of man



D. Number of males by age of man

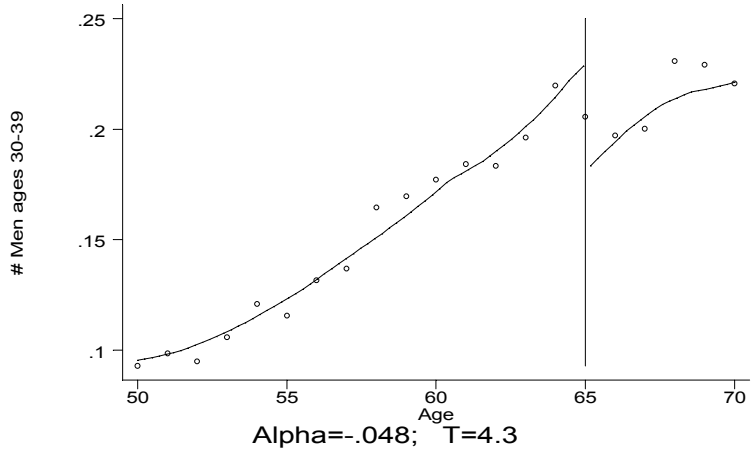
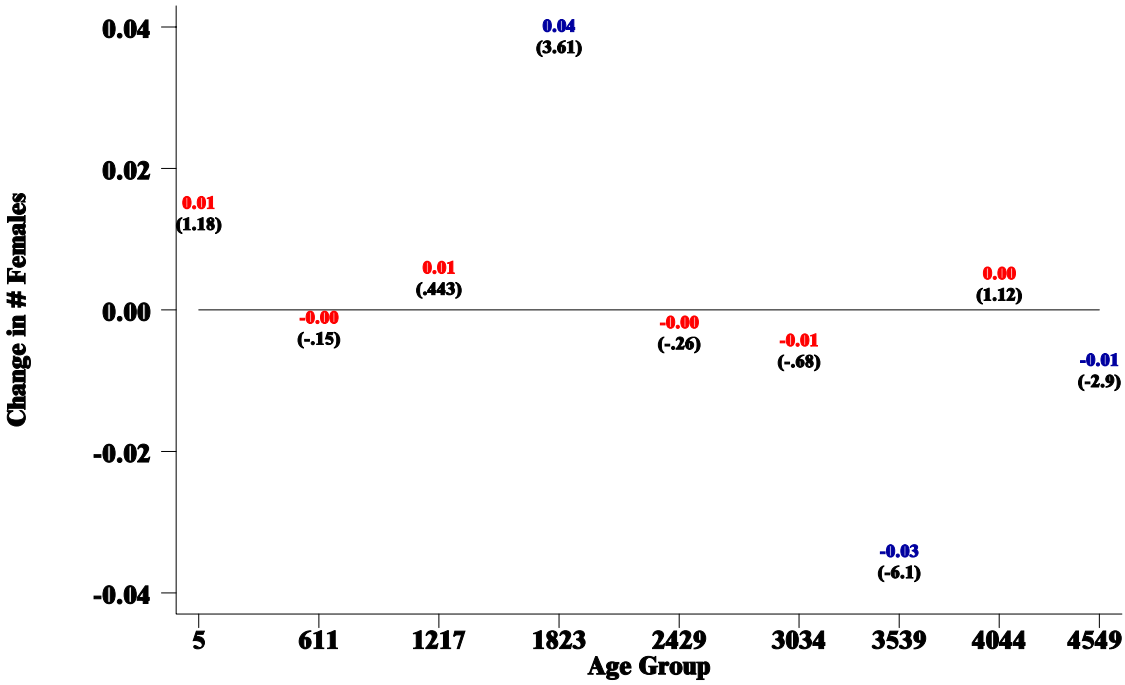


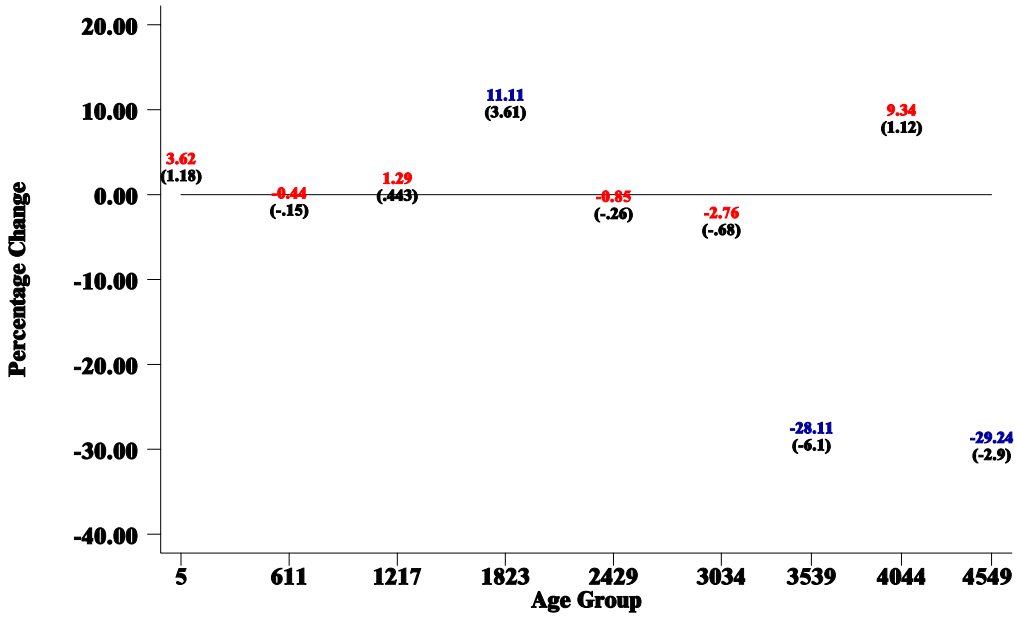
Figure 7: Age-by-age changes in the number of females when a woman becomes pension eligible at age 60



Estimates of treatment effects for different age categories. The first age category includes all females between the ages of 0 and 5. The second category includes ages 6-11, then 12-17, and so on. For each category, the top number is the estimate of the change in number of residents in that age category (alpha). T-statistics are in parenthesis.

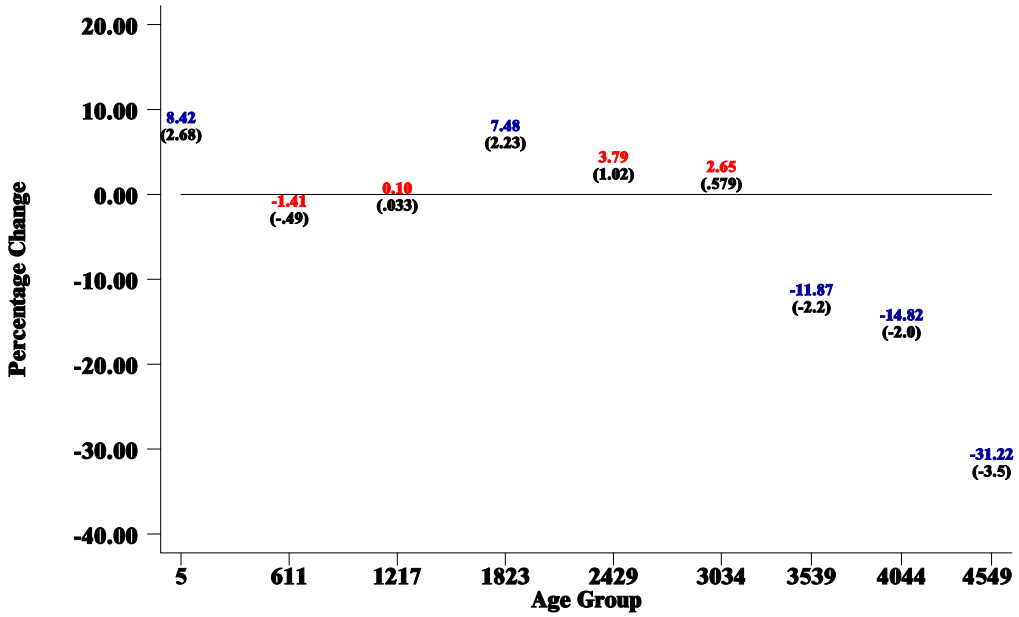
Figure 8: Age-by-age changes in residents when an individual becomes pension eligible

A. Percentage change in females by age for women age 60



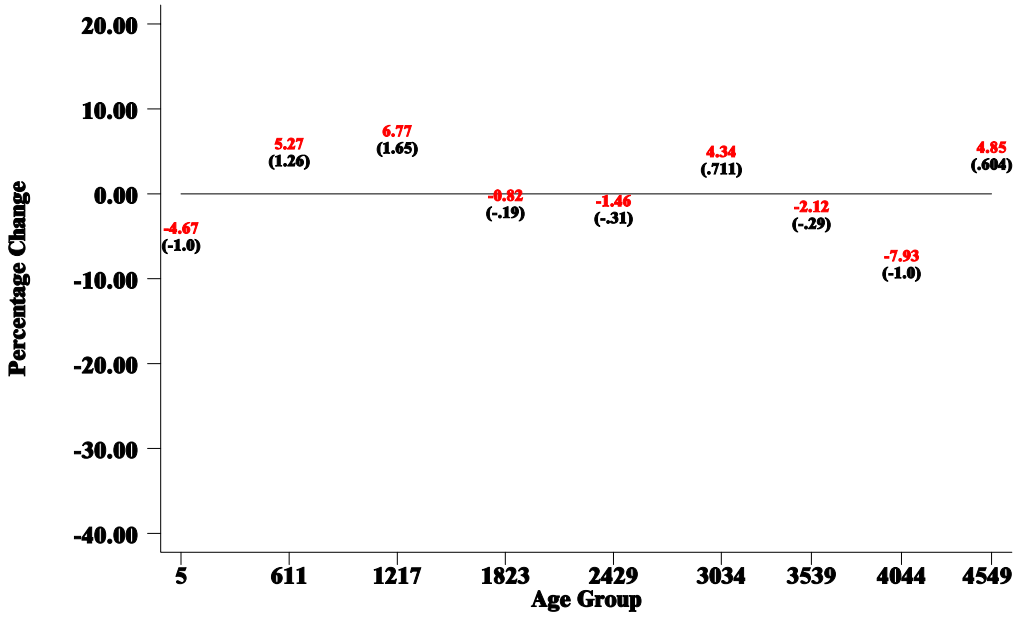
Estimates of the percentage change in the number of co-residents associated with pension eligibility for different age categories. The first age category includes all females between the ages of 0 and 5. The second category includes ages 6-11, then 12-17, and so on. T-statistics are in parenthesis.

B. Percentage change in males by age for women age 60



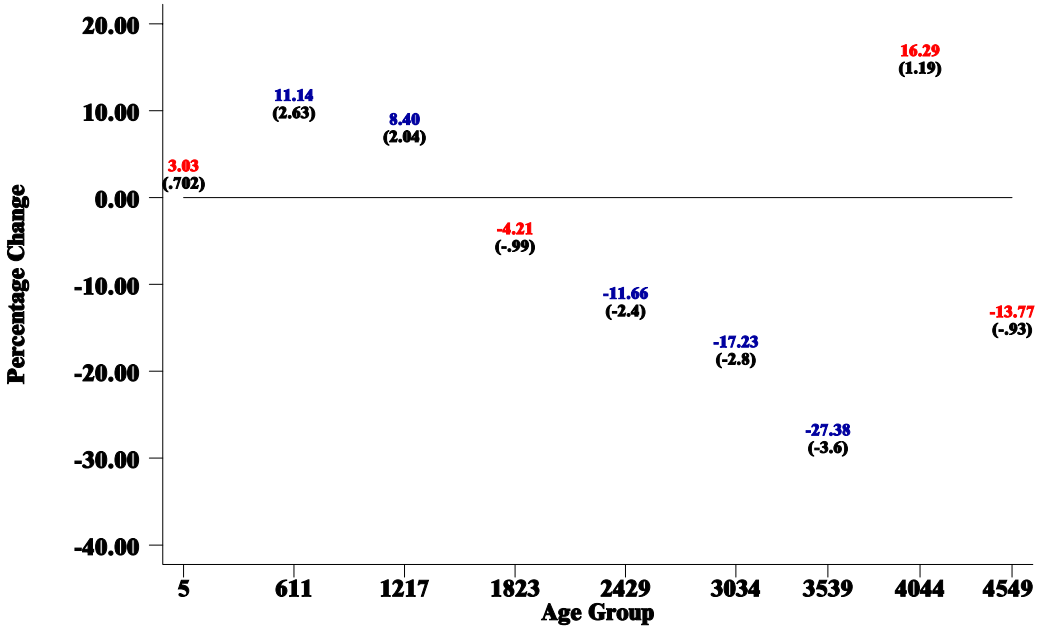
Estimates of the percentage change in the number of co-residents associated with pension eligibility for different age categories. The first age category includes all males between the ages of 0 and 5. The second category includes ages 6-11, then 12-17, and so on. T-statistics are in parenthesis.

C. Percentage change in females by age for men age 65



Estimates of the percentage change in the number of co-residents associated with pension eligibility for different age categories. The first age category includes all females between the ages of 0 and 5. The second category includes ages 6-11, then 12-17, and so on. T-statistics are in parenthesis.

D. Percentage change in males by age for men age 65



Estimates of the percentage change in the number of co-residents associated with pension eligibility for different age categories. The first age category includes all males between the ages of 0 and 5. The second category includes ages 6-11, then 12-17, and so on. T-statistics are in parenthesis.