Identifying Pure-Income Effects in an Empirical Model of Labour Supply: the case of the South African Social Pension

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Abstract

This paper investigates the income effects of the South African Social Pension. Using data from three waves of the the Labour Force Survey, we find that there appears to be a significant negative association between labour supply and pension receipt. However, we find little evidence to support the view that these results can be interpreted as pure income effects. Rather, the evidence suggests that the association is driven by age-cohort effects, which we argue reflects the burden of living with the elderly. We also report preliminary evidence which is suggestive of endogenous household formation in response to eligibility for the social pension.

Keywords: public transfers, pensions, labour supply

1 Introduction

A lively debate has emerged recently on the possible behavioural responses to the presence of pension income in the household, primarily in response to

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the findings made by Bertrand, Mullainathan, and Miller (2003), who showed that receipt of social pensions were associated with significantly negative hours and employment elasticities. Key contributions here include Alderman (1999), Posel, Fairburn and Lund (2005), and Ranchhod (2006). There is also a large literature which has looked at broader effects; notably, on savings behaviour, labour force status, nutritional status of dependents, household size, and household structure. Few of these contributions attempt to speak directly to the findings of Bertrand, Mullainathan, and Miller (2003), yet many contest the explicit interpretation provided by the authors that they have found uncontrovertible evidence showing disincentive effects of the social pension. In this paper, our goal is modest: we seek primarily to replicate the methods and results of Bertrand, Mullainathan, and Miller (2003) in order to more fully interrogate the interpretation they give to their findings. Thus, we examine whether pensions negatively affect labour supply within households which have age-eligible members. However, unlike their study, we use more recent data that come from the 2001 and 2002 Labour Force Survey. This is an important consideration if the effects observed in earlier data capture the shock of racial equalisation of the social pension rather than permanent effects. To examine this possibility the first part of our analysis recreates almost exactly their techniques. To identify the effect of the social pension we make use of the regression-discontinuity induced by the eligibility rules of the old-age pension programme.

Our estimates of the hours elasticity with respect to pension income are of the same sign as those found by Bertrand and colleagues, but are substantially lower in magnitude (about 20% lower) than their estimates . When subjected to a range of robustness tests, we find strong evidence to suggest that this effect is driven by age-cohort effects, which we interpret as evidence not of pure income effects, but of the burden of living with the elderly. Our interpretation appears to be further corroborated by changes in household composition that appear to be driven by access to the social pension.

The outline of this paper is as follows. Section 2 presents the standard theory surrounding the debate over the labour market effects of public policy. Section 3 provides an overview of the data and mean sample characteristics. Section 4 presents the main results of the paper, while sections sections 5 and 6 deal with causal inference and interpretation respectively. Concluding remarks are offered in section 7.

2 Neoclassical Theory of Labour Supply

2.1 Basic Framework

Consider the following canonical consumption-leisure framework. The decision over these goods are made by the agent as an outcome to the solution of the following constrained utility maximisation problem:

$$Max \quad U(G, L) \quad \text{such that} \quad P_G G = P_L H + V = P_L(T - L) + V \tag{1}$$

The agent's utility function depends on herlevel of consumption of goods G, and hours of leisure L. This utility function is well behaved. $U(G, L)$ is twice differentiable, where the first derivatives ($MU_L = \partial U/\partial L$ and $MU_G = \partial U/\partial G$) are assumed to be positive, and the utility function is concave ($\partial^2 U/\partial L^2 < 0$, $\partial^2 U/\partial L^2 < 0$ and $\partial^2 U/\partial L \partial G > 0$). We assume that the agent spends all available income. Thus the amount spent on consumption (price $P_G G$), must equal the sum of non-wage income (V) and wage income ($P_L H$) where H is the number of hours spent working, and is related to leisure as follows: $L + H = T$, where T is the maximum number of hours available to the agent. The slope of the budget line is the real wage (P_L/P_G). The optimal level of consumption and leisure is found where the marginal rate of substitution of leisure for consumption goods (MU_L/MU_L) is equal to the price ratio P_L/P_G, i.e. at the point of tangency between the agent's budget line and the highest attainable indifference curve. At this point, the marginal rate of substitution between consumption and leisure is equal to the slope of the budget line, i.e. the real wage. Equilibrium can occur either at an interior solution point, where $H > 0$, and $L < T$, or at a corner solution, where $H = 0$, and $L = T$. At the first type of solution point, the individual has made the decision to participate in the labour market, and has then chosen an optimal level of labour to supply. At the second point, the reservation wage $w*$ (the MRS_{LG}) is higher than the market wage (the slope of the budget line), and the agent decides not to participate in the labour market.

2.2 Effects of an Increase in Non-Labour Income

2.2.1 First-Order Effects: pure income effect

When an individual receives extra non-wage income V (e.g. pension income) the individual will tend to reduce labour supply, by substituting leisure for labour, under several strict conditions. Workers make their labour supply decisions based on information about wages, prices and non-wage income (Killingsworth, 1983). Assuming that wage rates and relative prices remain unaffected, an increase in non-wage income will allow the worker to afford the same bundle of goods, while working fewer hours. Under these conditions, the agent realises a higher level of utility, brought on by an outward shift of the budget line, resulting in a higher optimal choice of both leisure and consumption. Only under these conditions will a rise in V translate into a simple income effect, increasing consumption of both goods and leisure.

2.2.2 Second-Order Effects: reservation wages

Upon pension takeup, if the increase in non-wage income is not accompanied by a corresponding change in wage rates, the pension recipient's reservation wage will also unambiguously increase. This will tend to make the person accept fewer job offers, as a greater proportion of these offers will not exceed their reservation wage. The individual's bargaining power compared to that of their prospective employer will have increased, as will their bargaining power in the household.

2.2.3 Unobserved Heterogeneity in Preferences

Bertrand, Mullainathan, and Miller (2003) interpret their findings as indicative of pure income effects. It is intructive therefore to make clear what assumptions are required for this interpretation to be valid. We have made clear that a minimum requirement is that wage rates and prices must remain unchanged after the increase in non-wage income. However, even when these conditions hold, empirically identifying pure income effects requires very strong assumptions about preferences. An income effect might be non-monotonic or of the wrong sign for a given agent, depending on the shape of her indifference curve. If the individual has a preference for consumption (i.e. work) we may observe an increase in their labour supply after an increase in non-labour income.[1] Ruling out this possibility requires us to also rule out potential non-convexities in agents' indifference maps.

A further consideration is unobserved heterogeneity of preferences. If agents are assumed to differ according to their preferences, reservation wages, and relative bargaining power within households, an increase in non-wage income can have ambiguous aggregate effects. To illustrate, consider the following simple example: a male in the household reduces his supply of labour in response to an improvement in his reservation wage. If one assumes that the household operates according to the Unitary model, the implied comparative statics would see his spouse (implicitly) reducing her reservation wage in order to counter the loss of her partner's wage income, which in turn induces her to enter the labour force. If some households operate more or less according to this sort of "Arrow-Debreau" framework, while other households are better described say according to a "Separate Spheres" model, the overall effect of unearned income on labour supply cannot be separated from the effects of intra-household allocation. To be sure, there is no way to generalise a pure income effect (like that proposed in the simple model of individual labour supply outlined above), without providing an explicit account of intrahousehold bargaining over resources. Without such refinements, all that the model predicts (under those very strict conditions mentioned above) is an effect on one's own choice of work hours.

As we shall see in a moment, the data do allow a test of the externality effects of increases in non-labour income, but without a full account of the bargaining process underlying these observations, it is unclear whether any changes one might observe can be attributed to the effect of the income *per se*, or other unaccounted for factors.

3 Data and Summary Statistics

Our analysis is based on two waves of the Labour Force Survey – September 2001 and 2002. Two possible measures of work hours are permitted by the

[1] One way in which this might play out in South Africa is that upon reaching eligibility, pensioners may not wish to lose the prestige, and weight in household decision making that being the primary earner conveys. This preference in turn might be driven by the view that leisure is to be regarded as an inferior good when the individual has been unemployed for many years. Indeed, job satisfaction, pride in providing for family and other similar effects may all contribute to non-standard preferences regarding the choice of work over leisure (Killingsworth, 1983).

LFS: “the usual number of hours worked in a week”, and “the number of hours worked in the previous week”. We make use of the second measure, following Bertrand, Mullainathan, and Miller (2003) (hereafter BMM), in order to facilitate comparison between our results and theirs.

The sample used includes African individuals, who are of prime working age, and who live in three-generation households. Our definition of what constitutes a three generation household is somewhat different to that employed by BMM. Whereas the PSLSD data used by BMM permitted a direct assesment of whether households contains members of 3 distinct generations (i.e., grandparents, their children, and their grandchildren), this is not possible with the LFS. We therefore proxy three-generation households by using the yardstick of 30 years being equal to a generation. Thus in our analysis a household is deemed to be a three-generation household if it contained at least one member thirty years old or younger, plus one member in the 30-60 age cohort, and one member aged 60 and above. The sample is then truncated to only those individuals who are of prime working age (i.e., between the ages of 16 and 50 years).

Table 1 presents summary statistics relating to eligible and ineligible households, where eligibility is defined as whether a household contained a male aged 65 and above or a female aged 60 and above. Table 1 shows that there are statistically significant differences between eligible and non-eligible households on many of the observables. Of particular interest is the finding that average age in eligible households is about 2 years lower than average age in ineligible households. This is a key difference between our sample and the 1993 PSLSD data used by BMM. Later we will examine whether this information is instructive about how to interpret our results.

We find a higher rate of employment in pension ineligible households, and thus higher average hours worked and lower rates of unemployment. Eligible households have a 16.9% rate of employment, compared to 36.5% for ineligible households. This is a remarkable difference compared to that which is observed by BMM for Africans in three-generation households in the PSLSD. In part this is not all that surprising given what is now known about the rise in (involuntary) unemployment in the 1990s.

The strict unemployment rate for prime working-age individuals living in three generation households is nearly three times higher than the corresponding figure for for 1993, at 20.9%. In eligible households, strict unemployment is 25.3%, compared to 19.2% for ineligible households. This is partly a feature of the large proportion of eligible households who live in rural areas (58.5%

compared to 48.9% of ineligible households). Those in eligible households are more likely to act as safety nets for unemployed relatives (Klasen and Woolard (2000)), as seen by the much larger size of these households. The exclusion of migrant workers from the household in the Labour Force Survey also contributes to the larger unemployment rate in these households.

Weekly work hours (main job and other activities) are also higher than what was observed in the PSLSD data. Average hours worked in a week for individuals residing in both types of households is 27.5 hours, which is much higher than the levels found in the PSLSD data (an increase of between 3 and 6 times depending on the type of household). Eligible households have much lower labour supply than ineligible households (17.8 hours compared to 30.3 for ineligible households), and much greater differences in the number of employed people in the household. Pension eligible households have slightly higher levels of discouraged workers (7.1%) although the general level is low (6.4% in all households). The number of discouraged workers is about a third lower than that recorded in the PSLSD.

Individuals living in pension eligible households have slightly lower levels of education, which could support the hypothesis that pension income may induce migration, leaving households with less employable individuals. Posel and Casale (2002) find that for African female workers, an extra year of education raised the probability of migration by 3%. Only 22 percent of all prime working age Africans have completed matric, although this is roughly ten percentage points higher than the levels recorded in 1993. As expected, eligible households are much larger, and have lower income levels, than those households with no eligible elderly. Household size has fallen considerably since 1993. In all households, it fell by approximately 3 members. This could reflect increased migration since 1993. African three generation households have moved from living in predominantly rural areas (68% in 1993) to a more even split between rural and urban locations in 2001 (51% of all individuals live in rural areas).

The mean pension income in eligible households is close to the value of the state pension in 2001, or R640. Pension income plays a large role in eligible households. About 36% of household income in these households consists of pension income. Most pension income enters the household through female pensioners, as a much higher proportion of eligible women are present in these households than men. This pattern has not changed significantly since 1993. This stems from the eligibility rule, and the tendency of women to live longer than men.

4 Basic Results

We now present the key regression results of the effect of pension receipt on labour supply and employment. Equations (2) and (3) show the main equations of interest in this analysis.

$$h = \mathbf{x}\beta_{\mathbf{1}} + \gamma_1 y + u_1 \tag{2}$$

$$z = \mathbf{x}\beta_{\mathbf{2}} + \gamma_2 y + u_2 \tag{3}$$

h measures reported hours worked per week (including overtime and hours spent on other work for pay); z is an employment dummy that takes a value of one if the individual is employed, and is zero if strictly unemployed; and y is a pension variable. Following BMM, we use two pension measures – household pension income and household pension eligibility, where the latter is a dummy variable that is equal to one if the household has at least one person who is eligible for the state pension. Arguably, the second pension measure is the better of the two as it provides exogenous variation in pension income induced by the eligibility rules, whereas the income measure itself is counfounded potentially by endogenous take-up of the pension.

The vector $\mathbf{x}$ represents covariates. It contains the following individual level variables: a quartic in age (measured in years); binary variables for gender; a binary variable for whether matric has been completed or not; binary variables for location; the number of children in different age cohorts present in the household (ages 0-5, 6-15, 16-18, 19-21, 22-24); binary variables for provinces (the base category is the Western Cape); and a household size variable measured as the number of present household members.[2] These regressions are also run separately for men and women.

Preliminary estimates of equations (2) and (3) are presented in table 2 (see table 3 for a summary of these same regressions run separately by gender). The coefficient on household pension income has been multiplied by a factor of 1000, hence the coefficient in column one of -6.635 implies that an increase of household pension income of a thousand rand is associated with a fall in individual labour supply of 6.635 hours per week. Columns 1 to 3 report OLS estimates of the labour supply equation, while columns 4 to 6 report OLS

[2]Changes in household structure due to pension receipt may imply that household size should be modeled endogenously. As the effects on household composition are our secondary focus, we make this simplifying assumption in the first part of the analysis. However, we revisit this issue below when discussing causal inference.

estimates for employment status. Maximum likelihood estimates of these identical regressions are presented in tables 4 and 5. While there are some interesting differences between the linear and non-linear model specifications, we will not focus on these as we cannot make a comparison of such findings against those of BMM. For most of what follows therefore, we will restrict our attention to the results reported in tables 2 and 3.

One of our objectives is to investigate the implied elasticities associated with the estimated effects of pension income. However, our binary measure of the pension variable (which we use in the regressions reported in columns 2 and 5) is not useful for such a purpose. For this reason, we also provide instrumental variable (IV) estimates of pension income. The advantage of this approach is that it allows us to use a pension measure that is continuous (thus permitting the calculation of an elasticity) while exploiting the exogenous variation provided by the eligibility rules (as in the measure adopted in columns 2 and 5). Our IV is measured by summing all men and women in the household who are eligible to receive a social pension.[3] Since the number of eligible men and women in the household is strongly correlated with the level of household pension income, these variables pass a neccesary condition to be considered valid IVs.[4] To the extent that household size (and composition) is orthogonal to individual labour supply, which we assume to be true for the moment, then the IV also meets the sufficient condition of redundancy in the labour supply and employment equations. These IV estimates are reported in columns 3 and 6 of tables 2 to 5.

Our results are similar to those of BMM for both the hours worked and employment status models, though the effect on hours is stronger than on employment. Living in a pension eligible household reduces hours worked by 7.7 hours, or employment probability by 15.6%. An increase of household pension income of a thousand rand reduces hours worked by 8.5 hours and employment probability by 17.3%. As in BMM, the non-instrumented results underestimate the effect of pension income on hours worked and employment probability. The absolute values of these coefficients are much lower than the estimates for 1993 (12.3 and 17.0 respectively). In contrast however, the employment probability coefficients have increased since 1993.

If we accept this IV strategy as valid, what sort of economic magnitudes

[3]Using estimates based on income measures of the pension (whether instrumented or not) are easier to work with in terms of converting to elasticities.

[4]These variables are significant at the one percent level in the reduced form equation for pension income (not reported here).

does pension income translate into, in terms of forgone earnings due to the reduction in work hours? To get a sense of this, we calculate the (pension) income elasticities of labour supply and employment probability. Table 6 shows a comparison of the 1993 and 2001 hours worked and employment probability income elasticities. On average there are 1.9 prime working age people in a three-generation African household.[5] Thus if individual income rises by R1000, hours worked drops by -8.512 * 1.9 which is a drop of -16.17 hours per week. Average hours in a pension eligible household are 44.69 (conditional on working), and average individual income is 974 (per prime working age individual in the household). Thus an elasticity of hours to income is -8.512 * 1.9 *(0.974/44.69) = -0.35. The elasticity for employment is similarly calculated as -0.173 * 1.9 *(0.974/0.311) = -1.03. If we are to believe these results, it would suggest that the labour supply elasticity has fallen since the 1993 figure of -0.53, while the employment elasticity has almost doubled in size from -0.55 in 1993. Although we do not attempt to explore the reasons for this change, if these magnitudes are correct, one hypothesis for such a dramatic increase in the negative employment elasticity is the rise in labour market participation witnessed recently.

Table 3 presents a gendered breakdown of these results. Of note is the finding that men are affected much more than women by living in a pension eligible household. In column 3 we notice that an extra thousand rand of pension income reduces male hours worked by 12.525 hours per week, while having a matric has only a small positive effect on male work hours (1.847). By contrast, education seems to play almost as strong a role in explaining labour supply for women. These coefficients translate in pension income elasticities of -0.52 and -0.17 for men and women respectively and employment elasticities of -1.37 and -0.65 for men and women respectively.[6] It appears that male labour supply and employment elasticities have not changed remarkably since 1993, with a slight drop in labour supply elasticity, and a small increase in the employment probability elasticity. For women however, labour supply elasticity has reduced to a third of the size, and the employ-

[5] According to BMM, the PSLSD contains on average 4.7 prime-age African workers per three-generation household. Household size has also fallen significantly since 1993. The percentage of prime age workers in households in 1993 and 2001 was 51% and 25% respectively.

[6] Hours worked elasticities: -12.525 * 1.9 *(0.974/44.69) and -4.208 * 1.9 *(0.974/44.69). Employment elasticities: Men -0.230 * 1.9 *(0.974/0.311) and Women -0.109 * 1.9 *(0.974/0.311).

ment probability elasticity has increased more than four times. [7]

5 Is the Pure Income Effect Identified?

The validity of the above identification strategy rests on two key assumptions: (i) that the IV is (sufficiently) correlated with pension income; and (ii) that it is redundant in the labour supply and employment probability equations. The first of these assumptions is testable and we saw from the results discussed in section 4 that it was supported by the data. The second assumption on the other hand is untestable – if there are other unobservable variables that it turns out to be correlated with, the assumption will not hold and the endogeneity problem will remain unresolved. This possibility is effectively equivalent to a failure of the redundancy assumption. Therefore a quick test of identification is to look for systematic differences between the members of pension-eligible and pension-ineligible households that might conceivably be correlated with our IV. Stated differently, if we can conceive of reasons for why the variable *number of age-eligible men and women in a household* should be included in either equation 1 or 2, then the identifcation strategy is no longer valid.

A key result of Posel, Fairburn, and Lund (2006) is that household pension receipt is correlated with household in-migration of younger women with fewer skills, and household out-migration of older, prime age working women with more experience. From table 1, we see that the average age in eligible households is significantly lower, and the strict rate of unemployment is higher. Eligible households are also comparatively bigger. While we cannot ascribe a causal relation between these variables and take-up, if (the anticipation) of take-up of the pension induces shifts in household composition (i.e., if household composition and size is endogenous to pension take-up) the BMM identification strategy is weakened, if not invalidated. At best we will not be able to separate the pension effect or other correlated or contextual effects.

In this section we repeat the confoundedness experiment presented in BMM by exploiting the eligibility rules more directly to construct regression discontinuity estimates of the pension effect. Although the local-average treatment effect here is not of direct interest (as it only tells us about the

[7] The elasticities reported in BMM for hours worked are: -0.66 men and -0.43 women, and employment probability: -0.98 men and -0.14 women. See their table 10.

effects of the pension in the neighbourhood of the discontinuity) the one advantage of doing this exercise is that it allows an indirect assessment of the plausibility of the IV estimates, The basic idea is that if the effects of the pension reported in table 2 are true effects, then we would expect to observe a large drop in average hours worked for individuals living with elderly people that are close to or at the eligibility threshold, but not for those individuals who live with elderly people just below the threshold. To implement this, we contruct a disaggregated version of the IV used above. We do this for men and women 50 and above in 5-year cohort intervals. These variables will be used together with binary variables that capture whether there are any ineligible elderly in the household as well as whether there are any eligible elderly in the household.

Before proceeding to the results, it is important to point out that we need to make sure that we account for deviations from the eligibility rule, since identification is driven entirely by these rules. Following BMM, the deviation from the eligibility rule in the region measures the percentage of people who live in households which receive pension income where men between the ages of 60 and 65 are present, and no other eligible elderly in the household are present. This is a provincial level variable which reflects the extent to which the eligibility rules are strictly enforced in a given province. We would expect delivery to be standardised in the provinces by 2001. In 1993 a not insignificant fraction of people reported receiving the pension despite not having reached the correct age for their gender. Case and Deaton (1998) find that approximately a quarter of men between 60 and 65, and a tenth of women between 55 and 60 report receiving a pension. An important question that arises out of this finding is the extent to which this deviation is due to the mis-measurement of age as opposed to active deviation from the implementation of the eligibility rules by public officials administering the pension at a local level. We use the method reported by Case and Deaton (1998) to fine-tune our estimate of the proportion of deviant recipients for the 2001 Labour Force Survey. This method lowers the estimate considerably to approximately 0.5% of men, and 0.42% of women within 5 years of 65 and 60 respectively, who are receiving a pension illegally. We then construct a deviation variable from these estimates and include it as an additional control in the final specification (column 6) we report in table 7.

We find a negative effect of the presence of eligible elderly on hours worked in columns 1a and 1b in table 7, where the presence of an eligible elderly person in the household reduces individual hours worked by between 7.7 and

9.5 hours. These effects are similar in size to the coefficients found in table 2. As in 1993, the presence of a non-eligible elderly person in the household has no significant negative effect on hours worked. On the face of it, this would seem to suggest that the IV estimates are valid.

However a more disaggregated analysis of this variable reveals otherwise. The most interesting finding, and a key point of departure between our results and those of BMM, is the fact that non-eligible elderly women appear to reduce their labour supply well in advance of becoming eligible for the pension (witness the statistically significant coefficient of -0.70 on *Women in the household 50-55*). We take this to mean that the key results reported in BMM tables 2-4, and our replications thereof (tables 2, 3 and 7) do not identify pure income effects of the social pension. To be clear, such effects may very well be present, but to our knowledge, there have not been any successful attempts at cleanly seperating the negative income elasticity of the pension, from other effects having to do with living with elderly people. Our results seem to suggest that individuals living with elderly african women who are not old enough (in some instances, ten years too young) to receive a state pension also work fewer hours than individuals without such persons in their household. This suggests that at least some part of the negative effect on hours worked (and by extension, the large negative elasticities attributed to the social pension) have to do with the unobserved burden of living with the elderly.

6 Discussion and Future Directions

The preceeding section suggested that the negative elasticities reported should not be interpresetd as pure income effects of the pension, since an apparently significant negative effect of the social pension on work hours is detected even before the age of eligibility, which seems to increase monotonically after the age of eligibility. We intrepreted this evidence as the pension variable reflecting at least partly the burden of living with the elderly. In this section, we explore whether part of the explanation might relate to out-migration once someone in the household reaches the age of eligibility. Our intuition here is motivated by the findings of Posel, Fairburn, and Lund (2006) which we find intuitively compelling.

To investigate this, we exploit the panel structure of the LFS, albeit in a limited way, by using the two waves of the LFS as a pooled cross-section.

Our ultimate objective is to seperate temporal shifts in household size, from spatial shifts (where space is here defined as eligible versus ineligible households). What follows is a first attempt at investigating whether such an endeavor is worth undertaking: i.e., are the differences in mean household sizes we observe between the two cross-sections, a) statistically significant, and b) statistically significant when conditioned on household eligibility?

To accomplish this, we employ the following double-difference estimator of household size:

$$size = \beta_0 + \delta_0 Time + \beta_1 z + \delta_1 Time \times z + u \tag{4}$$

$Time$ is equal to unity in 2002; $size$ measures household size; z is equal to one if there is a positive number of pension eligible people in the household. δ_0 captures the effect of time on changes in household size over the two periods, for both groups eligible and ineligible; β_1 captures the effect of eligibility on average household size; and δ_1 is the main coefficient of interest. If this coefficient is significant, then there are significant differences in household size, between eligible and ineligible households, having accounted for temporal changes in $size$.

The results of these preliminary investigations indicate that household size is bigger for eligible households, and the change in household size for eligible households is bigger than the corresponding change in ineligible households.[8] Also, both changes are negative – both types of households are seeing members exit from the household, although more exit from eligible households. This is consistent with the findings of Posel, Fairburn, and Lund (2006) although we would caution against a causal interpretation at this point: i.e., that becoming eligible for the pension *causes* out-migration of prime-age working women. Such a conclusion, of course, is only possible where we can observe a change in eligibility status for both eligible and ineligible status.[9]

Having noted this caveat, we find $\delta_1 = -0.13$. This implies that controlling for the effect of time, eligible households tend to be on average 13% smaller and this difference is statistically significant. While these findings need to be verified for the dynamic case, taken together with our findings discussed in section 5, it would seem that the assumption of exogenous household

[8]The tables on which these differences are based are not included here but are available on request from the authors.

[9]Stated differently, we cannot rule out selection bias is when both counterfactuals (mean household size for eligible and ineligible households in alternative states of the world) cannot be computed.

size is deserving of further study. If these patterns turn out to be true for the dynamic case, it would suggest that the BMM identification strategy is invalid.

7 Conclusion

We find that although there appears to be a significant negative association between labour supply and pension receipt, our results are much smaller than those reported in previous work. We also find little evidence to support the view that these results can be interpreted as pure income effects. The discrepency between our findings and those of Bertrand, Mullainathan, and Miller (2003) suggests that their work on the subject might have been capturing short term effects of the pension (such as the shock induced through racial equalisation). While it is clear that further research is required in order to pin down the causal mechanisms at work, we have argued that what might appear to be pure income effects of the social pension are potentially conflated by the effect of living with the elderly as well as changes in household composition driven by access to the social pension.

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Table 1: Descriptive Statistics

	All		Eligible		Ineligible	
Variable	Mean	S.D	Mean	S.D	Mean	S.D
Age	29.5	9.9	27.9	9.2	30.1	10.1
Employed	0.311	0.463	0.169	0.374	0.365	0.481
Hours Worked	27.53	26.79	17.84	25.05	30.34	26.62
Unemployed	0.209	0.407	0.253	0.435	0.192	0.394
Discouraged	0.065	0.247	0.072	0.258	0.063	0.242
4th grade or more	0.885	0.319	0.890	0.312	0.883	0.321
8th grade or more	0.646	0.478	0.655	0.475	0.643	0.479
Matric or more	0.221	0.415	0.214	0.410	0.223	0.416
Household Size	5.81	3.38	7.64	3.83	5.14	2.92
Rural	0.508	0.500	0.580	0.494	0.482	0.500
Total Income	1930	2663	1849	2276	1949	2744
Pension Income	233	409	662	460	53	197
No of Eligible Women	0.230	0.442	0.866	0.429	0	0
No of Eligible Men	0.088	0.286	0.330	0.478	0	0

Note: Eligible households refer to those households that contain at least one pension eligible person, i.e. one man over the age of 64 or one woman over the age of 59). Ineligible households refer to households with no pension eligible members. The sample is further restricted to African individuals that are of prime working age (between 16 and 50 years of age), and who live in three-generation households.

Table 2: Effect of Pension Income on Hours Worked and Employment Probability

	Hours Worked			Employment Status		
	OLS	OLS	IV	OLS	OLS	IV
	Pension Uptake	Pension Eligibility		Pension Uptake	Pension Eligibility	
Variables	(1)	(2)	(3)	(4)	(5)	(6)
Pension Income x 1000	-6.635	–	-8.512	-0.136	–	-0.173
	(0.652)	–	(0.780)	(0.013)	–	(0.015)
Household Eligibility	–	-7.702	–	–	-0.156	–
	–	(0.560)	–	–	(0.011)	–
Female	-5.610	-4.393	-5.648	-0.079	-0.053	-0.079
	(0.446)	(0.404)	(0.412)	(0.008)	(0.007)	(0.008)
Age	-20.052	-22.356	-20.316	-0.523	-0.679	-0.528
	(4.952)	(4.511)	(5.065)	(0.096)	(0.090)	(0.094)
Age^2	0.975	1.126	0.989	0.025	0.033	0.025
	(0.235)	(0.212)	(0.240)	(0.005)	(0.004)	(0.004)
Age^3	-0.019	-0.023	-0.019	-0.000	-0.001	-0.000
	(0.005)	(0.004)	(0.005)	(0.000)	(0.000)	(0.000)
Age^4 x 1000	0.132	0.164	0.134	0.003	0.004	0.003
	(0.036)	(0.032)	(0.036)	(0.001)	(0.001)	(0.001)
Matric or more	2.475	1.586	2.475	0.070	0.052	0.070
	(0.474)	(0.445)	(0.465)	(0.009)	(0.008)	(0.009)
R^2	0.12	0.14	0.12	0.15	0.17	0.15
n	14484	18307	14484	14491	18318	14491

Note: *Hours worked* is measured per week, and *Employment Status* is equal to 1 if the person is employed, and 0 if unemployed. In columns 3 and 6 we instrument for endogenous household pension income using the number of eligible men and women present in the household as instruments. The coefficient on household pension income has been multiplied by a factor of 1000. Hence the coefficient in column one of -6.635 implies that an increase of household pension income of a thousand rand is associated with a fall in individual labour supply of 6.635 hours per week. An increase in household pension income of a R1000 reflects the effect of the addition to the household of between one and two pensioners. Standard errors have been corrected for clustering within households and are reported below the regression coefficients in parentheses. The list of explanatory variables includes binary variables for the number of children in the different age groups present in the household (ages 0-5, 6-15, 16-18, 19-21, 22-24), and indicator variables for province and rural/urban location. Household size is also included. The sample is restricted to African individuals who are of prime working age (between 16 and 50 years of age), and who live in three-generation households.

Table 3: Gender Effect of Pension Income on Hours Worked and Employment Probability

	Hours Worked			Employment Status		
	OLS	OLS	IV	OLS	OLS	IV
	Pension Actual	Pension Eligibile		Pension Actual	Pension Eligibile	
Variables	(1)	(2)	(3)	(4)	(5)	(6)
Men						
Pension Income x 1000	-9.730	–	-12.525	-0.186	–	-0.230
	(0.916)	–	(1.131)	(0.017)	–	(0.021)
Household Eligibility	–	-10.198	–	–	-0.186	–
	–	(0.798)	–	–	(0.015)	–
Matric or more	1.856	1.273	1.847	0.058	0.046	0.057
	(0.705)	(0.636)	(0.696)	(0.013)	(0.011)	(0.013)
R^2	0.11	0.14	0.11	0.13	0.16	0.12
No Obs	7157	9096	7157	7164	9108	7164
Women						
Pension Income x 1000	-3.542	–	-4.208	-0.084	–	-0.109
	(0.788)	–	(1.067)	(0.017)	–	(0.020)
Household Eligibility	–	-5.141	–	–	-0.123	–
	–	(0.686)	–	–	(0.014)	–
Matric or more	3.043	1.767	3.045	0.083	0.058	0.083
	(0.600)	(0.578)	(0.617)	(0.012)	(0.011)	(0.012)
R^2	0.13	0.13	0.13	0.19	0.18	0.19
n	7327	9211	7327	7327	9210	7327

Table 4: Maximum Likelihood Estimates of the Effect of Pension Income on Hours Worked and Employment Probability

	Hours Worked			Employment Status		
	Tobit	Tobit	IV	Probit	Probit	IV
	Pension Uptake	Pension Eligibility		Pension Uptake	Pension Eligibility	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Pension Income x 1000	-8.087	–	-19.072	-0.162	–	-0.200
	(0.813)	–	(1.644)	(0.016)	–	(0.017)
Household Eligibility	–	-9.548	–	–	-0.172	–
	–	(0.713)	–	–	(0.012)	–
Female	-5.687	-4.455	-10.928	-0.091	-0.062	-0.091
	(0.478)	(0.444)	(0.836)	(0.010)	(0.008)	(0.009)
Age	-27.872	-31.715	-53.744	-0.570	-0.722	-0.576
	(5.912)	(5.472)	(10.582)	(0.116)	(0.103)	(0.110)
Age^2	1.377	1.590	2.657	0.027	0.035	0.028
	(0.273)	(0.252)	(0.497)	(0.005)	(0.005)	(0.005)
Age^3	-0.027	-0.032	-0.053	-0.001	-0.001	-0.001
	(0.005)	(0.005)	(0.010)	(0.000)	(0.000)	(0.000)
Age^4 x 1000	0.194	0.234	0.375	0.004	0.005	0.003
	(0.039)	(0.037)	(0.074)	(0.001)	(0.001)	(0.001)
Matric or more	3.049	2.141	5.824	0.080	0.061	0.080
	(0.518)	(0.494)	(0.949)	(0.010)	(0.009)	(0.010)
Pseudo R^2	0.04	0.04	–	0.12	0.13	0.12
% correctly predicted	–	–	–	67.4%	70.1%	73%
n	14484	18307	14484	14491	18318	14491

Dependent variables are same as in tables 2 and 3. *Hours Worked* equation estimated with Tobit model instead of OLS. *Employment Status* equation estimate with Probit model rather than the OLS. IV is measured in the same way as in tables 2-3. Standard errors (in parentheses) have been corrected for within-household clustering. Pseudo R^2 measures refers to McFadden's pseudo R^2 measure. The probit coefficients in column 4 and 5 are marginal effects, calculated at the means for continuous variables, and differential changes for binary explanatory variables. Tobit results in columns 1 through 3 are also marginal effects: partial derivatives of the expected value of hours worked with respect to the vector of characteristics, computed at the means of the explanatory variables. The Probit IV and Tobit IV estimation follows the method of Newey (1987). The coefficients in column 3 and 6 are not corrected for cluster effects.

Table 5: Maximum Likelihood Estimates of the Gender Effect of Pension Income on Hours Worked and Employment Probability

	Hours Worked			Employment Status		
	Tobit	Tobit	IV	Probit	Probit	IV
	Pension Uptake	Pension Eligibility		Pension Uptake	Pension Eligibility	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Men						
Pension Income x 1000	-11.945	–	-25.632	-0.214	–	-0.257
	(1.211)	–	(2.266)	(0.021)	–	(0.023)
Household Eligibility	–	-12.310	–	–	-0.201	–
	–	(1.027)	–	–	(0.016)	–
Matric or more	2.407	1.763	4.198	0.064	0.053	0.064
	(0.766)	(0.708)	(1.329)	(0.014)	(0.012)	(0.014)
Pseudo R^2	0.03	0.04	–	0.10	0.13	0.09
% correctly predicted	–	–	–	65.7%	70.2%	72%
No Obs	7157	9096	7157	7164	9108	7164
Women						
Pension Income x 1000	-4.505	–	-11.219	-0.104	–	-0.130
	(0.939)	–	(2.371)	(0.020)	–	(0.024)
Household Eligibility	–	-6.806	–	–	-0.139	–
	–	(0.852)	–	–	(0.016)	–
Matric or more	3.663	2.437	7.686	0.099	0.070	0.099
	(0.647)	(0.638)	(1.349)	(0.014)	(0.013)	(0.014)
Pseudo R^2	0.08	0.03	–	0.15	0.14	0.15
% correctly predicted	–	–	–	69.5%	70.0%	74%
No Obs	7327	9211	7327	7327	9210	7327

Table 6: Hours Worked and Employment Probability Elasticities

	Hours	Employment
	<u>All</u>	
1993 OLS	-0.53	-0.55
2001 OLS	-0.35	-1.03
2001 MLE	-1.98	-1.19
	<u>Men</u>	
1993 OLS	-0.66	-0.98
2001 OLS	-0.52	-1.37
2001 MLE	-1.06	-1.53
	<u>Women</u>	
1993 OLS	-0.43	-0.14
2001 OLS	-0.17	-0.65
2001 MLE	-0.46	-0.77

Estimates for 1993 are from Bertrand, Mullainathan, and Miller (2003) whereas estimates for 2001 are based on own calculations from table 2 which uses the Labour Force Survey.

Table 7: Effect of Living with the Elderly on Labour Supply (Regression Discontinuity Design)

Variable	OLS (1a)	MLE (1b)	OLS (2a)	MLE (2b)	OLS (4a)	MLE (4b)	OLS (6a)	MLE (6b)
Eligible Elderly in HH	-7.702	-9.548	–	–	–	–	–	–
	(0.560)	(0.713)	–	–	–	–	–	–
NonEligible Elderly in HH	-0.556	-0.731	–	–	–	–	–	–
	(0.910)	(0.948)	–	–	–	–	–	–
Persons in HH 50-55	–	–	-0.347	-0.423	-0.332	-0.408	–	–
	–	–	(0.181)	(0.202)	(0.182)	(0.203)	–	–
Women in HH 50-55	–	–	–	–	–	–	-0.701	-0.836
	–	–	–	–	–	–	(0.259)	(0.284)
Men in HH 50-55	–	–	–	–	–	–	0.037	0.005
	–	–	–	–	–	–	(0.284)	(0.317)
Persons in HH 55-60	–	–	-0.144	-0.163	-0.123	-0.141	–	–
	–	–	(0.204)	(0.233)	(0.206)	(0.234)	–	–
Women in HH 55-60	–	–	–	–	–	–	-0.407	-0.483
	–	–	–	–	–	–	(0.307)	(0.346)
Men in HH 55-60	–	–	–	–	–	–	0.142	0.191
	–	–	–	–	–	–	(0.336)	(0.375)
Persons in HH 60-65	–	–	-0.456	-0.506	-0.537	-0.606	–	–
	–	–	(0.214)	(0.238)	(0.225)	(0.250)	–	–
Women in HH 60-65	–	–	–	–	–	–	-0.962	-1.091
	–	–	–	–	–	–	(0.287)	(0.325)
Men in HH 60-65	–	–	–	–	–	–	-0.432	-0.604
	–	–	–	–	–	–	(1.282)	(1.386)
n6065m X deviation from	–	–	–	–	–	–	16.422	21.421
eligibility rule in region	–	–	–	–	–	–	(37.300)	(40.339)
Persons in HH over 65	–	–	-0.840	-0.955	–	–	–	–
	–	–	(0.144)	(0.162)	–	–	–	–
Persons in HH 65-70	–	–	–	–	-0.463	-0.530	–	–
	–	–	–	–	(0.254)	(0.280)	–	–
Women in HH 65-70	–	–	–	–	–	–	-0.682	-0.817
	–	–	–	–	–	–	(0.335)	(0.368)
Men in HH 65-70	–	–	–	–	–	–	-0.087	-0.064
	–	–	–	–	–	–	(0.450)	(0.512)
Persons in HH over 70	–	–	–	–	-0.866	-0.982	–	–
	–	–	–	–	(0.185)	(0.207)	–	–
Women in HH over 70	–	–	–	–	–	–	-1.185	-1.435
	–	–	–	–	–	–	(0.254)	(0.285)
Men in HH over 70	–	–	–	–	–	–	-0.192	-0.072
	–	–	–	–	–	–	(0.337)	(0.375)
R^2	0.14	0.04	0.13	0.04	0.13	0.04	0.13	0.04

Note: All reported MLE estimates are marginal effects for the tobit model. The R^2 value for the tobit models is an Anova Based Fitness Measure. Tobit results have been corrected for heteroscedasticity. Coefficients in parentheses refer to robust standard errors corrected for within-household correlation in hours worked. The regressions contain the same additional regressors as in table 2.