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Are poor households credit-constrained or myopic? Evidence from a South African panel*

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Abstract

Credit constraints are an almost ubiquitous assumption in development economics. Yet direct evidence for credit constraints is limited, and many observations consistent with credit constraints are equally compatible with precautionary saving or myopic (non-forward-looking) consumption. Using household panel data and a source of widely anticipated income in South Africa, this paper first tests and rejects the standard consumption model with perfect capital markets. The standard model enriched with credit constraints is then contrasted with precautionary saving and myopic consumption as alternative explanations for the observed expenditure pattern. The standard model with credit constraints cannot be rejected in favour of precautionary saving or myopic consumption.

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1 Introduction

Credit constraints are a standard assumption in development economics, but there is surprisingly little robust evidence that such constraints are present and binding on household behaviour. Furthermore, many observations that are consistent with credit constraints are equally consistent with precautionary saving or a myopic model in which consumption depends only on current income. Recent work on credit constraints in rich countries has exploited the idea that under standard assumptions and perfect capital markets, there should be no jump in consumption associated with an anticipated increase in income. This paper applies that methodology to a large panel data set of South African households. Having established that expenditure jumps are present, further evidence is presented indicating that the jumps are more likely to be due to credit constraints than to precautionary saving behaviour or myopic consumption.

Credit market imperfections are central to current development theory. And in models of household behaviour in developing countries, such imperfections are

frequently operationalised as a constraint on the amount of credit available to the household (often taken to be zero). But given how often credit constraints are assumed, there is little clear evidence of their presence. In his entry on development economics in the *New Palgrave Dictionary of Economics*, Ray (2008) discusses credit market failure and constrained credit as a key theme of the field, but writes that "[t]he direct empirical evidence on the existence of credit constraints is surprisingly sparse".

In fact, many observations that are consistent with credit constraints are explained equally well by myopia or precautionary saving. In particular, this applies to observations that consumption responds to anticipated increases in income, as well as to observations of 'excess sensitivity' of consumption with respect to temporary income shocks that have only a small impact on permanent income. Apart from the direct interest in establishing whether credit constraints are present, this relates to the long-standing debate on whether poor households optimise within their constraints ("poor but rational") or not (Duflo 2006). Finding that the consumption behaviour of poor households is not guided by forward planning could have large consequences for the study of development.

This paper uses data on a large panel of black South African households to investigate the response of expenditure to a large anticipated increase in income. The public, non-contributory Old Age Pension scheme is widely known in the population, and take-up amongst Africans is high and well predicted by age. An approximate Euler equation is derived, from which it is clear that a jump in expenditure associated with an anticipated increase in income is as consistent with precautionary saving as it is with a binding credit constraint. The consumption jump can also be explained by myopia. The prediction that consumption should not respond to the anticipated income increase is empirically rejected, even when allowing for household-specific trends. Furthermore, evidence is presented that households are more likely to save when a member reaches the age of eligibility, making the precautionary saving story less likely. Finally, a change in the age at which the Child Support Grant lapses is used to identify an anticipated decrease in household income. This allows for a test of credit constraints versus myopic consumption behaviour. Credit constraints cannot be rejected in favour of myopic consumption.

The main contributions of this paper are as follows. It is one of a small number of studies that provide explicit tests of the standard consumption model with perfect capital markets using expenditure data from a developing country. Uniquely, the tests exploit a large panel data set and a large and highly anticipated source of income. The use of fixed-effects estimation and clean identification provides for a robust rejection of the standard model with perfect capital markets amongst black South Africans. This is also the first paper to attempt to differentiate between credit constraints, precautionary saving and myopic consumption as explanations of expenditure jumps in a developing country. By showing that credit constraints is a more likely explanation of household behaviour than non-forward-looking consumption, the paper also contributes to the debate on whether people in developing countries are 'poor but rational'.

The remainder of the paper is organised as follows. The next section provides an overview of related literature. In the theory section, an approximate Euler equation that incorporates credit constraints and a precautionary saving motive is derived from a standard consumption model, and the relevant implications for empirical work are drawn out. The evidence section provides background information about the Old Age Pension and the Child Support Grant in South Africa and describes the data before presenting the empirical specifications and results. Finally, there is a brief conclusion.

2 Related literature

Hall (1978) extended the standard additive consumption model[1] to the case with stochastic earnings. The paper's influential central finding is that, conditional on today's consumption, no other information available today is predictive of consumption tomorrow.[2] In other words, consumption should respond to income changes only to the extent that these reflect changes in permanent income. This

[1]These models assume intertemporally additive utility, constant discount factors and consumers who maximise (expected) present discounted utility over the remainder of the lifetime. I follow Browning & Lusardi (1996) in referring to these as standard consumption models, though their origins lie in Modigliani's life-cycle and Friedman's permanent income hypothesis models.

[2]Parker (1997) presents a consumption model that combines Hall's (1978) framework with the durable-goods model of Mankiw (1982), and derives similar predictions.

prediction has subsequently been exploited by many economists to construct tests of the perfect capital markets assumption. One line of research looks at the effect on consumption of anticipated or unanticipated changes in income (see Browning & Lusardi (1996) for a review). Recently, two papers have argued that households in the United States and Spain adjust their consumption in response to significant and regular changes in annual income, but not in response to irregular and small payments because the computational cost of doing so outweighs the potential utility gain. Hsieh (2003) exploits the annual payout of petroleum dividends to residents of Alaska as an anticipated change in permanent income. He does not find excess sensitivity of consumption related to the petroleum payout, but on the other hand he confirms earlier results when he finds that the same households do over-react to tax refunds, which are smaller and harder to predict on average. Similarly, Browning & Collado (2001) find that the consumption patterns of Spanish households working in sectors with regular bonus payments do not differ significantly from those of households in other sectors. The findings in this paper go the other way, indicating that at least in one developing country context, the standard model with perfect capital markets is rejected even with respect to a very large and highly predictable increase in income.

Flavin (1985) tests credit constraints versus myopic consumption using data from the United States, but relies on an instrumental variable technique and aggregate data rather than a source of anticipated income in a panel of households. Altonji & Siow (1987) wrote one of the first papers to exploit the prediction that forward-looking but credit-constrained agents should react differently to positive and negative changes in income. This paper follows theirs in attempting to test between myopic consumption and the standard model with and without credit constraints, but they rely on an instrumental variables approach rather than a well-known and specific source of anticipated income. They find no strong evidence of either credit constraints or myopic consumption in the American panel they look at (PSID), whereas this paper provides a strong rejection of the standard model with perfect capital markets in a panel of black South African households. Deaton (1992) investigates how closely simple 'rules of thumb' can approximate the optimal solution to the standard consumption model. These rules of thumb are partly a function of current income and are thus related to myopic consumption

as defined here.

Deaton (1991) studies savings behaviour in the presence of credit constraints, and uses simulations to study the effectiveness of precautionary saving in smoothing household consumption.

Banerjee (2003) surveys the evidence for capital market imperfections in developing countries. One of the most striking pieces of evidence he presents is a wide dispersal between interest rates on saving and borrowing in many parts of the developing world. Banerjee & Duflo (2005) call this 'suggestive, albeit indirect, evidence of credit constraints'.

One of the few articles to provide a direct test of a prediction from the standard consumption model in a developing country is by Paxson (1992). The starting point for her influential paper is the prediction that the propensity to consume unexpected income should be much smaller than that for expected income. In other words, a large proportion of windfall gains should be saved. In line with the prediction, she finds that saving amongst agricultural households in Thailand is responsive to positive income shocks resulting from rainfall variation. Rosenzweig & Wolpin (1993) find evidence of credit constraints among farmers in India, and Banerjee & Duflo (2008) do the same for Indian firms. Morduch (1990) tests for credit constraints in rural India.

One strand of the literature looks at self-reported credit constraints. Typically, respondents are flagged as credit-constrained either if they report having tried to borrow and being refused, or if they report being discouraged from applying for credit because they thought they would be refused. For instance, Barham, Boucher & Carter (1996) uses this type of survey information to look at whether the presence of credit co-operatives in regions of Guatemala relaxes credit constraints. Rasmussen (2002) takes a similar approach using data from South Africa. She finds that a high proportion of the households in her sample are credit-constrained.

There is a growing body of literature exploiting the South African Old Age Pension as a source of exogenous income variation. Case & Deaton (1998) look at how various expenditure categories respond to the pension income, but they do not analyse total expenditure, nor do they focus on credit constraints. Edmonds (2006) finds that receiving the pension leads to a decrease in child labour and an increase in school attendance. Duflo (2003) finds that young girls living with a

pension-eligible grandmother are in better health than those who do not. Both Edmonds and Duflo argue that their findings are indicative of credit constraints. But their use of cross-section data rules out household fixed effects, and they do not attempt to rule out myopic consumption as an alternative explanation of their findings.

3 Consumption with credit constraints and income uncertainty

An agent lives for T periods. In each period t she receives an income

$$y_i = \overline{y}_i + \varepsilon_i,$$

where $\overline{y}_i$ is deterministic (but may vary over time) and ε_i is a series of shocks with expectation 0. Income is the only source of uncertainty in the model. The agent sets consumption c_t so as to maximise

$$\mathrm{E}_t \sum_{i=t}^{T} \delta^{i-t} u\left(c_i, Z_i\right)$$

subject to

$$\begin{aligned} w_i &= R w_{i-1} + y_i - c_i \\ w_i &\geq 0 \end{aligned}$$

for $t \leq i \leq T$. Here, $\delta \in (0,1)$ is the discount factor, R is the gross interest rate and w_t denotes the agent's wealth at the end of period t. Instant utility u is a function of consumption c_t and household characteristics Z_t. For simplicity, household characteristics are assumed to evolve deterministically. The instant utility function is three times differentiable with respect to consumption and satisfies $u' > 0$, $u'' < 0$ and $u''' \geq 0$. (Here and later, primes indicate differentiation with respect to consumption.) Concavity implies risk aversion, while a positive third derivative gives the agent a precautionary saving motive.

The first constraint is just the evolution of wealth: wealth at the end of period i, w_i, is the wealth remaining at the end of the previous period, w_{i-1}, augmented by the gross interest rate R, and adjusted for current-period labour income y_i and consumption. The second is the credit constraint: the requirement that wealth is non-negative at the end of every period implies that the agent cannot borrow. Note, however, that the following analysis is robust to allowing some borrowing, i.e. a negative rather than a zero lower bound on wealth.

The first-order condition with respect to consumption at time t yields the Euler equation

$$u'(c_t, Z_t) = \delta R \, \mathrm{E}_t \, u'(c_{t+1}, Z_{t+1}) + \lambda_t,$$

where λ_t is the Kuhn-Tucker multiplier for the credit constraint. λ_t is always non-negative, and it is positive when the credit constraint binds in period t.

Because of the uncertainty in future income, there will be uncertainty in future consumption. Consumption uncertainty cannot be made precise at this level of generality, but it is always possible to write

$$c_i = \overline{c_i} + e_i$$

for future periods i with $\overline{c_i} = \mathrm{E}_t \, c_i$ and $\mathrm{E}_t \, e_i = 0$. Denote $\sigma_i^2 = \mathrm{Var}(e_i)$. Then a Taylor expansion and taking expectations yield

$$u'(c_t, Z_t) \approx \delta R \left[u'(\overline{c_{t+1}}, Z_{t+1}) + \frac{\sigma_{t+1}^2}{2} u'''(\overline{c_{t+1}}, Z_{t+1}) \right] + \lambda_t.$$

Rearranging gives the approximate Euler equation

$$u'(c_t, Z_t) - \delta R u'(\overline{c_{t+1}}, Z_{t+1}) \approx \delta R \frac{\sigma_{t+1}^2}{2} u'''(\overline{c_{t+1}}, Z_{t+1}) + \lambda_t. \tag{1}$$

This equation is informative of the expected evolution of consumption, and its interpretation motivates the empirical specification used in this paper.

3.1 Implications

Since $u' > 0$ and $u'' < 0$, the left-hand side of (1) is positively related to the expected increase in consumption between periods t and $t+1$. If there were no credit constraints, no uncertainty and no changes in the household characteristics, the right-hand side would be zero and consumption would evolve smoothly according to the shape of the utility function and the magnitude of δR. A higher discount factor δ ('more patience') or a higher interest rate R tends to decrease today's consumption relative to that of tomorrow.

A binding credit constraint, $\lambda_t > 0$, introduces a positive term on the right-hand side and therefore tends to reduce current consumption relative to future consumption. Uncertainty in future consumption will also reduce current consumption relative to the future (precautionary saving). Changes in household characteristics also influence the consumption path.

Several previous empirical tests of the standard consumption model with perfect credit markets have relied on the prediction that only lifetime wealth should matter, and therefore an anticipated change in income should not have any effect on consumption. Rejections of this prediction have been taken as evidence of credit constraints. In the light of (1) it is clear that such constraints are indeed a possible explanation for a change in consumption levels associated with an anticipated change in income. A household that expects an increase in income would, in the absence of credit constraints, smooth its consumption by borrowing before the higher income stream begins. But a credit constraint may prevent the consumption increase from happening until the higher income is realised. In terms of the equation, recall that a binding credit constraint is associated with $\lambda_t > 0$, and that $\lambda_t = 0$ corresponds to the absence of a binding constraint in period t. Clearly, a binding constraint is associated with a steeper increase in consumption.

However, (1) makes it clear that a break in the household consumption path at the time of an anticipated income increase can also be due to a change in consumption uncertainty. The above framework is too general to study this effect precisely, but it seems plausible that the increase in income can reduce the household's need for buffer savings and thereby allow consumption to increase.

Though a jump in consumption can be explained by either credit constraints or

a reduction in buffer savings, looking at the stock of savings should help distinguish between the two mechanisms. A credit-constrained household will necessarily reduce its savings to zero in the period before the income increase (otherwise the constraint would not be binding), so that savings will either stay constant at zero or increase when income increases. By contrast, if the jump in expenditure is due to a reduced need for buffer savings, then clearly savings could decrease at the time of the increase in income.

A break in the household consumption path may also be observed if the above model is misspecified. The core assumptions of the standard consumption model are that utility is temporally additive and that consumers maximise expected utility. This can be contrasted with a myopic consumption model, a basic version of which stipulates that households consume a fixed proportion of their disposable income. Considered as a theory of household behaviour, this may be simplistic. But it does capture a sense of myopia or lack of forward planning, and can therefore serve as a benchmark against which to measure the meticulously planning households of the standard models. Specifically, the consumption of a myopic household will track that of its income, regardless of whether income increases or decreases, and regardless of whether the change is anticipated or not.

The above framework rules out borrowing, but not saving (otherwise the household would simply consume its income in each period). This is probably realistic, though the return on savings may be low and uncertain. This gives rise to an asymmetry of predictions: in the presence of constraints on credit (but not on savings), an anticipated *increase* in income may be associated with a simultaneous increase in consumption. But an anticipated *decrease* in income, other things equal, should not be accompanied by a steep drop in consumption in the standard model, because the household is able to smooth its consumption profile by first accumulating and then drawing down savings. In contrast, the consumption of a myopic household will track income across both increases and decreases.

The rest of this paper documents a jump in expenditure associated with an anticipated increase in income, and then attempts to explain this observation. Panel data allow the use of household fixed effects to take account of time-invariant, household-specific consumption determinants. This makes it possible to focus on credit constraints, precautionary saving and myopic consumption as possible

explanations for the observed behaviour.

4 Evidence

4.1 The Old Age Pension and the Child Support Grant

The South African Old Age Pension is a non-contributory, means-tested public benefit scheme, originally set up in the 1920s to provide a retirement income for the minority of white workers who did not have pension arrangements through their employer. In 1989, the government committed to removing racial inequalities in the system, and from 1993 onwards the new scheme was fully operational throughout the country. In rural areas, the money is distributed from a sophisticated network of mobile paypoints reaching each beneficiary once a month. Today, the pension is a major source of income[3] amongst the poor all over the country, and in many households the main breadwinner is an Old Age Pensioner. See Case & Deaton (1998) for more detail.

To qualify for the pension, the applicant has to be old enough and also pass a means test. Only people whose wealth and income are below certain levels qualify. The means test is set at a level where most *white* elderly fail to qualify, but a large majority of *black* elderly easily pass the hurdle. Hence for black people the pension can largely be regarded as a universally available and secure monthly income stream, beginning at the lower qualifying age (60 for women, 65 for men) and continuing until death. It is independent of the economic situation of family members, except the spouse for married pensioners. From an econometric perspective, it provides a convenient source of nearly exogenous income variation amongst people around the qualifying age.

Given its popularity, reach and financial importance, the pension is a highly anticipated source of income. The qualifying age limits are also widely known (Duflo 2003, Case & Deaton 1998).

The Child Support Grant was introduced in 1998 to replace the earlier state

[3]In 2000, the pension paid out ZAR 540 per month. This is approximately 1.7 times the median per-capita total black household income of ZAR 320 per month computed from the Income and Expenditure Survey 2000.

maintenance grant. It is paid to the primary care-giver (usually the mother) of young children. It is in principle means-tested, and a large proportion of black mothers with age-eligible children qualify. Initially it only covered children up to and including the age of six, but from 2002 there was a gradual expansion to older children. For the 2001 and 2002 surveys used in this study, the upper age limit was six years. When the 2003 survey was conducted, the upper age limit was eight years. Grant take-up increased rapidly during the period studied, in part due to considerable effort by the government. In 2002, the grant paid out ZAR 110 per month per child, which is more than one-third of the median per capita household income amongst blacks. Rosa & Mpokotho (2004) provide more detail on the Child Support Grant.

4.2 Data

Since 2000, Statistics South Africa has conducted a Labour Force Survey twice a year. It is designed as a rotating panel survey, but the longest series of longitudinal household observations is substantially shorter than it could have been, due to several 'fresh starts' in which completely new samples were drawn. This study uses the September 2001, September 2002 and September 2003 waves of the survey. The surveys are nationally representative. Only households headed by a black ('African', in South African terminology) person are considered here. After dropping 175 households due to missing data on the gender, age or population group of one or more household members, there remains 36,192 households in the panel, of which 11,964 are observed twice and 6,651 are observed three times.

An expenditure variable was constructed as follows. In each of the survey rounds the households were asked, "What was the total household expenditure in the last month? Include everything that the household and its members spent money on, including food, clothing, transport, rent and rates, alcohol and tobacco, school fees, entertainment and any other expenses." The answer options were: R 0–399, R 400–799, R 800–1,199, R 1,200–1,799, R 1,800–2,499, R 2,500–4,999, R 5,000–9,999, R 10,000 or more, "Don't know" or "Refuse". The expenditure variable used in this paper was constructed by taking the midpoint of each interval. For the "R 10,000 or more" option, expenditure was set to R 15,000. The last two

response options were coded as missing values. It is the logarithm of the resulting number that is used here as the measure of monthly household expenditure.

A household is coded as age-eligible for the Old Age Pension if there is at least one woman aged 60+ or one man aged 65+ in the household, and age-eligible for the Child Support Grant if at least one child below the upper age limit is present in the household at the time of observation. The validity of age eligibility as a predictor of take-up is confirmed in Figure 1, where take-up of the Old Age Pension is plotted against the age of the oldest male and female household members. Age eligibility is clearly a strong predictor of pension take-up, though perhaps more cleanly for women than for men.

Descriptive statistics are presented in Table 1. Most of the variables appear to be stable over time, but the increasing take-up of the Child Support Grant is discernible.

For the analysis reported in Figure 4 only, data from the Income and Expenditure Survey 2000 were used instead.

4.3 Testing the unconstrained standard model

This section looks at how household expenditure responds to the anticipated income stream from the Old Age Pension. The results can be previewed in Figure 2, in which the logarithm of monthly household expenditure is plotted against the age of the oldest female and male household members. For both sexes, expenditure increases from the age of 20 until it reaches a maximum around the age of 40 for women and a few years later for men, after which it decreases. However, at 60, the age of female pension eligibility, there is a clear break and expenditure as a function of the age of the oldest female seems to rise rapidly before levelling off. There is a similar trend break for men around the age of 65, though again the onset is perhaps not as cleanly defined.

The main empirical specification is given by

$$Y_{it} = \alpha_i + \mu_t + \gamma L_{it} + X_{it}\beta + \varepsilon_{it}.$$

Here, Y_{it} is the expenditure of household i at time (survey) t. On the right-hand side, α_i is a household fixed effect, μ_t is a time fixed effect, and L_{it} is an indicator

for whether household i receives, or is age-eligible for, the Old Age Pension in year t. X_{it} is a vector of control variables, and ε_{it} is the error term.

The coefficient of interest is γ. Given the specification, it is only identified within households that move from a non-pension to a pension state or vice versa. Disregarding deaths and other changes in household composition, this will only happen when a female household member turns 60, or a male household member turns 65. In the eligibility specification, the identifying assumption is that conditional on household demographics and other controls there is nothing special about turning 60/65 other than that one becomes eligible for the Old Age Pension. This assumes that all other changes to do with 'entering retirement' are either a direct consequence of qualifying for the pension, or are controlled for in the regression.

Throughout the analyses, age eligibility rather than actual take-up will be the preferred right-hand-side variable. This is due to possible endogeneity in the take-up decision: it may be that a person's decision about when to apply for the Old Age Pension is co-determined with other unobserved characteristics that may also impinge on expenditure.

The main results are presented in Table 2. All regressions, here and later, include household and time fixed effects, and standard errors are robust and clustered at the level of the primary sampling unit.

In column 1, household expenditure is regressed on a dummy variable indicating whether the household receives the Old Age Pension (1) or not (0). The coefficient of interest is positive at .195 and highly significant. This suggests that receipt of the pension increases a household's expenditure by approximately 20%. Column 2 regresses expenditure on a binary indicator for whether at least one household member is age-eligible for the pension. The coefficient drops to .159 but remains highly significant. In column 3, variables indicating the presence of male and female age-eligible household members are included separately. The male and female eligibility coefficients are .147 and .159, respectively. Both are significantly different from zero, but they are not significantly different from each other.

All these coefficients are estimated within households. The fixed effects control for unobservables that are invariant over time, but previous work has suggested that household composition can change in response to pension receipt (Edmonds, Mammen & Miller 2004). It has been suggested that mothers with young children

are more likely to move in with the grandmother when the latter starts receiving the Old Age Pension, while mothers with older children may leave their children in the care of the grandmother and migrate for work.

This is not a problem for the present purposes, since the main right-hand-side variable, age eligibility, is exogenous to such changes. Changes in household composition is a consumer choice, and if they occur at the time of pension eligibility then that is arguably another indication of credit constraints: if it were possible to smooth consumption, it would not be necessary to wait for pension eligibility in order to effectuate changes in household composition. Nevertheless, in column 4 a full set of household demographics are included, namely separate variables for the number of male and female family members in the age groups 0–4, 5–14, 15–24, 25–34, 35–44, 45–54 and 55+. The coefficient drops to 0.074 but remains significant at the 1% level. The drop in magnitude suggests that about half the observed increase in expenditure at the age of pension eligibility is related to changes in household composition.

This paper uses expenditure as the main dependent variable, whereas the theory's predictions are in terms of consumption. The two are often the same. However, given the imprecise wording of the survey question on which the expenditure variable is based, there is a possibility that the expenditure data used may include debt repayment. There may therefore be a concern that the observed increase in expenditure at the age of pension eligibility may reflect the repayment of loans rather than an increase in consumption. However, there are no data on debt principal or repayment with which to investigate this. But households were asked whether they, in the 12 months prior to the interview, incurred any debt, and also whether there is a mortgage on their dwelling. One may therefore attempt to study the expenditure of households without debt by focusing on the subsample of observations where no debt had been incurred in the past 24 months, and where the dwelling is not mortgaged at any observed point. This rules out short-term debt and the most important type of long-term debt, though non-mortgage debts incurred more than 24 months ago and still outstanding cannot be ruled out.

The restriction reduces the sample size dramatically, but for the remaining households consumption should be closely aligned with expenditure. The results of this regression are reported in column 5. The coefficient remains significant,

and is not significantly different from the full-sample result.

Though the regressions reported so far allow for household fixed effects, the Euler equation suggests that households may be on different consumption *trends* depending on their discount factor and/or household characteristics. Because some households are observed three times, it is possible to put first-differenced expenditure on the left-hand side. The household fixed effects would then capture idiosyncratic household trends, and the regression would test for whether pension eligibility leads to a break in expenditure trend. The result of this regression is reported in column 6, and the highly significant coefficient of .258 indicates that becoming eligible for the pension increases the year-on-year growth in consumption by as much as 29%.

The expenditure variable was also regressed on annual demographic variables for household members within five years of the qualifying age. For each integer n from -5 to 5, a variable was constructed that counts the number of household members no more than n years younger than the qualifying age. Thus the variable indexed by $n = -5$ equals the number of women aged 55 or more plus the number of men aged 60 or more, and the variable indexed by $n = 0$ counts the number of household members who qualify for the pension. If qualifying for the pension is linked to increased expenditure but general ageing in this window is not, then one would expect the coefficient on the variable indexed by $n = 0$ to be positive, but none of the others.

The results of this regression are presented graphically in Figure 3. The estimated coefficients are marked by dots, and the bars show the 95% confidence intervals. It is clear that only the coefficient at the qualifying age ($n = 0$) is significantly different from zero, and it is positive. The regression also included household and year fixed effects. Running separate regressions by sex (not reported) yield equivalent results for women, whereas in the male-only specification none of the coefficients are significant.

There may remain a concern that the distinction between consumption and expenditure is not adequately addressed. Therefore, an analysis similar to the one in Figure 3 was undertaken using a different data set. The Income and Expenditure Survey 2000 breaks down financial transactions in more detail, so that it is possible to create a consumption variable that explicitly excludes loan payments or deposits

into saving/investment accounts. However, this variable is only available in a cross-section, implying that computing within-household estimates are not possible. The logarithm of annual consumption was regressed on the same set of variables as before, and the results are presented in Figure 4. Standard errors were again robust and clustered at the level of the primary sampling unit.

It is clear that the only variable that is significant is at +1, and it is positive. This implies that having a female aged 61+ or a male aged 66+ significantly increases consumption relative to younger age groups. Though this indicates the presence of a positive discontinuity in consumption as expected, it does seem to occur one year after the qualifying age as opposed to simultaneously with it. A possible explanation for this relates to the way age information was collected for this survey. As well as age, the year of birth was recorded. But it appears that rather than collecting age information separately, it has been calculated by subtracting the year of birth from the year of the survey (2000). However, since the survey took place in October, this will overstate the age for a significant proportion of the sample. It is therefore likely that some people recorded as being 61 years of age were in fact only 60 at the time of the survey.[4]

Because the cross-sectional data set rules out household fixed effects, a similar regression was also run in which household demographics (the same set of variables as before) were controlled for, along with two binary variables indicating the level of education of the household head (indicators for having any education at all and having completed high school). The qualitative results did not change. As above, when the regression was run separately for men and women, the qualitative result was retained for women, but for men there were no significant coefficients.

In order to interpret these results, refer back to the Euler equation (1). Basic household demographics are controlled for in column 5, so these cannot account for the expenditure response. Though it is difficult to rule out unobserved characteristics that change over time, the household fixed effects capture anything that is time-invariant at the household level. The Euler terms involving the discount

[4]If one assumes that the survey was done in mid-October, and birthdays are uniformly distributed across the year, then 12.5% of the population will not yet have had their birthday in the current year. However, in the survey less than 2% of household members appear to not yet have had their birthdays in the survey year, suggesting that age has been understated by one year for a significant proportion of the sample. This could explain the regression result.

factor and interest rate may be responsible for an increasing or decreasing trend in expenditure, though the year fixed effects will capture any changes that affect the sample as a whole and the first-differenced results allow for household-specific trends.

The results presented here indicate that expenditure responds forcefully to pension age eligibility. The standard explanation for this observation would be that the households faced binding credit constraints the previous year. But if qualifying for the pension is associated with a decrease in consumption uncertainty (which seems likely) as well as an increase in income, then these results cannot rule out the possibility that the observed change in consumption is due to a reduced need for precautionary saving: households that previously saved for a rainy day may dissave once they receive the pension, and this may boost expenditure. Finally, myopic consumption that simply tracks income can also explain these findings.

Some readers may be concerned that a major life change ('retirement') might systematically coincide with the qualifying age for the pension and bias the results. However, for the majority of the black population who are either unemployed or employed in the informal sector, there is no well-defined concept of 'retirement' in the sense of reaching the end of an employment contract. This is illustrated by the fact that only 2.6% of black women in the 60–64 age group receive employment-related pensions, and the equivalent number for black men in the 65–69 age group is 7.2%. If anything, withdrawal from the labour market is much more likely to be *caused* by the Old Age Pension income (Ranchhod 2006), or to have already happened (Bertrand, Mullainathan & Miller 2003), than to coincide with it in a way that might bias the identified effect on consumption.

4.4 Credit constraints versus precautionary saving

Within the general consumption theory outlined earlier in this paper, the results so far are compatible with either credit constraints or precautionary saving, or both. If a household is fairly certain of an increase in future income, it would be expected to smooth consumption by borrowing before the change. But income from the Old Age Pension is not just sizeable, it is also guaranteed for life. Employment income is probably uncertain in comparison, especially in the informal labour market.

Therefore, becoming eligible for the pension is likely to be associated with increased consumption security and a reduced need for precautionary saving. A reduction in the need for buffer savings may increase consumption around the time of pension eligibility even in the absence of credit constraints.

However, a consideration of household savings may allow a distinction to be made. If the jump in expenditure is solely due to credit constraints, then one would expect the household to draw down savings in anticipation of becoming eligible for the pension. In the absence of uncertainty, the household will as far as possible try to smooth consumption across the income discontinuity, and this is inconsistent with maintaining positive savings. Put differently, a credit constraint cannot truly bind as long as savings are positive. With a binding credit constraint, pension take-up should therefore be associated with an increase, or at least not a decrease, in household savings.

On the other hand, assume that the jump in consumption is entirely due to precautionary saving. Entering a phase of greater consumption certainty reduces the need to keep a buffer of savings. Therefore, the prediction is a fall in savings. This suggests a way to distinguish between the two mechanisms: if behaviour is primarily guided by a credit constraint, savings would be either constant (at zero) or increase when income increases. By contrast, if the expenditure response to increased income is due to precautionary saving, then household savings could decrease when income increases.

Only binary information on savings is available, in the form of the household's response to the question "Does this household, or a household member, own any of the following financial assets?" Yes/no responses are provided to each of the following options: bank savings accounts, stokvels (ROSCAs), pension plans or retirement annuities, unit trusts, stocks or shares, cash loans expected to be repaid, life insurance, or any other savings. It is therefore possible to construct an overall savings indicator as follows: the household is coded as having savings (of *some* form) if it answered 'yes' to at least one of these options, and otherwise not.

Clearly, the analysis is limited to identifying households that move from being savers to non-savers or vice versa. Households that adjust their amount of savings from one positive amount to another cannot be discerned from those that do not adjust their (positive) savings at all.

Table 3 presents the results of regressions where the dependent variable is the binary savings indicator. The precautionary saving story predicts a reduction in savings, while the credit constraints hypothesis predicts either no change or an increase in household savings. It is therefore natural to take the credit constraints story as the null hypothesis in these regressions, to be rejected in favour of precautionary saving if the coefficient on pension take-up or eligibility is negative and significant.

Regressing the savings indicator on pension take-up in column 1 yields a highly significant coefficient of .07. Regressing the savings indicator on pension eligibility in column 2 results in a smaller coefficient of .037, but it remains significant at the 1% level. In column 3, male and female age eligibility are considered separately. The female coefficient is .039 and highly significant, whereas the male coefficient drops to .023 and is not significantly different from zero. Column 4 presents the results of a regression in which the sample is restricted to observations of households that did not incur any debt in the last 12 months prior to the survey. The coefficient on pension eligibility is .033 and marginally significant. Including a full set of household controls and using the full sample in column 5 results in a positive coefficient of .018, but it is not significantly different from zero.

One potential issue with these findings is related to how the respondents understand the question about savings, in particular the last option, 'Any other savings'. There may be a tendency not to count modest amounts of cash kept in a purse or wallet as savings, whereas the same amount held in a bank account may be thought of as such. The resulting measurement error may lead to under-reporting of savings for those households that do not use a bank account or other non-cash savings technologies. If households are less likely to hold their savings in cash after the increase in income, then an estimate of the effect of the income on savings may be biased upwards. However, using 'Any other savings' as the dependent variable does not change the results substantially (not reported).

In summary, the savings indicator seems to increase or stay constant in response to the increased income. There is therefore no evidence of a decline in savings at the age of pension eligibility, and the null hypothesis of credit constraints cannot be rejected in favour of precautionary saving.

4.5 Credit constraints versus myopic consumption

Above, it was found that household expenditure, even when de-trended, responds positively to an anticipated increase in income. Moreover, the response of household savings to the anticipated increase in income seems to point in the direction of credit constraints rather than precautionary saving as an explanation. However, this conclusion still rests on the assumptions of the standard model from which the Euler equation was derived. In this section, the aim is to contrast the general framework of forward-looking but credit-constrained households with the stark myopic alternative in which households simply consume a fixed fraction of their income.

More advanced versions of myopic consumption models are considered in the literature, but they share the critical feature that current consumption is a positive function of current-period income. Deaton (1992) considers several rules of thumb of this type, of which his benchmark rule is 'consume all cash on hand up to mean income, and 30 per cent of any excess'.

A myopic model of consumption compares well to the standard model as far as the results so far are concerned: for myopic households, an increase in income is accompanied by an increase in consumption, even if the change is anticipated. Furthermore, if consumption is a fixed proportion of income, and savings are the difference between income and consumption, then savings must also be a fixed proportion of income. Therefore, the basic myopic consumption model predicts that household savings will increase when income does. This prediction is supported by the findings in the previous section.

But an anticipated *reduction* in income may help distinguish between the models: if consumers are credit-constrained but forward-looking, then an anticipated decrease in income should not have any effect on expenditure because the technology required to smooth consumption is savings, not credit. In contrast, the myopic consumption model predicts that expenditure should decrease in line with income.

It is important to be clear that this strategy of differentiating between credit constraints and myopic consumption relies on the presence of an effective savings technology. Besley (1995) provides a number of reasons why savings may not be an

available or attractive way to achieve smooth consumption in a typical developing country. Anyone can save cash at home or on their person, but the expected return on savings may be zero or negative if there is a considerable risk of loss, appropriation by other household members or theft. Even during apartheid, the South African post office offered a savings account that was available to the whole population. But it is possible that distance or transport cost to the nearest branch, mistrust in the system or illiteracy may have made this product unattractive to many potential savers. Since the end of apartheid, most if not all banks have offered savings accounts available to anyone, though high fees and fear of intimidation may still act as barriers. Informal savings devices, for example in the form of stokvels (ROSCAs), remain popular amongst black South Africans. Taken together, it seems reasonable to assume that some savings technology was available to the black population in the period studied, even if returns may have been low. This is supported by the observation that in the sample, 48% of households report having savings in some form.

The South African Child Support Grant is available to 'primary carers' (predominantly mothers) of young children. There is a means test that will be ignored here because most black households easily pass it. Instead, the focus is once again on the age cut-offs. Originally, the grant covered any number of biological children up to and including the age of six. So as a first pass, one could look for a drop in household expenditure when children turn seven. If there is such a drop, it could be taken as evidence for myopic behaviour rather than credit constraints.

But this is not entirely satisfactory. In the previous section, identification relied partly on the assumption that households where a 59-year-old woman (or 64-year-old man) is present is not substantially different from a household with a 60-year-old woman (or 65-year-old man), except that the latter household would be age-eligible for the Old Age Pension. It is less clear that a child with a six-year-old is comparable to a seven-year-old. Young children develop fast and progress through the educational system, so arguably an age difference of a single year may be associated with considerable changes in household behaviour even if other circumstances were fixed.

For this reason, a change in the age cut-off for the Child Support Grant will be exploited. From 1 April 2003, the upper age limit went up from six to eight

years. So in the first two rounds of the survey data, collected in September 2001 and September 2002, only households with children up to and including six years old were age-eligible for a Child Support Grant. But in the final round used here, conducted in September 2003, households with children up to and including eight years old were age-eligible.

The situation is illustrated in Figure 5, which shows the proportion of households reporting receipt of the Child Support Grant as a function of the age of the youngest household member. Perhaps the first thing one notices is the rapid expansion in take-up of the grant in the period. For households in which the youngest child is aged 0–6 years, take-up in 2003 was greater than it was in 2002, which in turn was greater than it was in 2001. But the change in the upper age limit for grant eligibility is also clearly visible: in 2001 and 2002, when children up to and including the age of six were eligible, take-up was very low for households in which the youngest child was seven or older. In 2003, when children aged up to and including eight were eligible, similarly low take-up rates are not reached until age nine. The identification of the results presented in this section relies on this change in the eligibility of seven-year-olds between 2002 and 2003.

The validity of the identification is confirmed in Table 4. A binary variable indicating whether a household had taken up the Child Support Grant is regressed on the number of children aged seven in the household and an interaction of this variable with an indicator for the year 2003. The identification strategy rests on the assumption that households with seven-year-olds in 2003 are more likely to be in receipt of the Child Support Grant than households with seven-year-olds in 2002, because the former are eligible for the grant whereas the latter are not. This is borne out by the results presented in columns 1–3. Column 1 presents the basic result. Annual and household fixed effects are controlled for throughout, as before. Column 2 adds yearly count variables for children aged 0–9, and column 3 adds general household demographics. The coefficient is positive and highly significant throughout.

If the household consumption pattern is determined by myopia, then expenditure for a household with a child aged seven in 2003 should be higher than the expenditure for a household with a child aged seven in 2002, because only the former household is eligible for the grant. For forward-looking households that

may face constraints on credit but not on savings, the presence of a seven-year-old in 2003 should have the same impact on consumption as the presence of a seven-year-old in 2002. It is true that the former may expect higher lifetime earnings, but this should be captured by a household fixed effect, and any macro-trend is captured by year fixed effects.

To test the null hypothesis of forward-looking households, a regression of the form

$$Y_{it} = \alpha_i + \mu_t + \gamma C_{it} + \delta C_{it} * I_{t=2003} + X_{it}\beta + \varepsilon_{it}$$

was run. Here, C_{it} denotes the presence of a child age seven in household i and year t. The indicator $I_{t=2003}$ is 1 if $t = 2003$ and 0 otherwise. If households are subject to constraints on credit but not on savings, the coefficient δ should be zero: there is no difference between the consumption behaviour of a household with a child who turns seven in 2002, and that of a household with a child who turns seven in 2003, apart from what can be explained by the household fixed effect, the macro-trend or changes in other characteristics. But myopic households will reduce consumption in line with the lapse of the Child Support Grant. Therefore, a household with a seven-year-old in 2002 should decrease its consumption compared to the previous year, whereas a household with a seven-year-old in 2003 should not. In other terms, the coefficient δ is zero under the standard model with credit constraints, and positive under the myopic model.

Table 4 presents the results of this regression. In column 4, the regression is run without control variables apart from household and survey fixed effects. The coefficient of interest is negative but not significantly different from zero. In column, 5 further controls are added in the form of variables counting the number of children in age categories from zero up to nine years of age. The coefficient of interest is still not significant. In column 6, household demographics are added as control variables, but the coefficient of interest changes little and remains insignificant.

Thus the coefficient is negative and insignificant for all three regressions. So the null hypothesis of a standard model with forward-looking households and credit constraints cannot be rejected in favour of the myopic consumption model.

Though it is likely that the relevant population was reasonably well informed of

the basic qualifying requirements for the Child Support Grant, the above analysis implicitly assumes that the *change* in the qualifying age was also known. The analysis does not permit a distinction between the behaviour of forward-looking but 'surprised' households that learnt that their seven-year-olds were still receiving the grant in 2003, from that of myopic (non-forward-looking) households. Both types of households would be expected to consume more than their counterparts in 2002. The results presented here, a negative coefficient not significantly different from zero, are therefore indicative of forward-looking households that were aware of the changes in the rules. This is probably realistic and reflects the importance of the grant in South African society.

5 Conclusion

Many economists are quick to point to credit constraints when they suspect failures of the standard consumption model. Indeed, the assumption of constrained credit is a staple of economic development theory. Yet direct and rigorous tests of this assumption, of the type that has recently gained popularity in analyses of rich-country data, are scarce. Moreover, many findings consistent with credit constraints are equally compatible with precautionary savings or myopic consumption. This paper tests and rejects the standard model with perfect capital markets using data on a panel of black South African households. It is one of the first papers to present evidence that credit constraints, rather than precautionary saving or myopic consumption, drive the observed excess sensitivity of consumption to anticipated income changes in a development context.

At the micro-level, inefficient credit markets hamper consumption smoothing and restrain production through inefficient allocation of resources and investments. At the macro-level, these investment misallocations are widely believed to impede growth. Credit markets matter for development, and effective development policy must be sensitive to the precise mechanisms that determine consumption behaviour.

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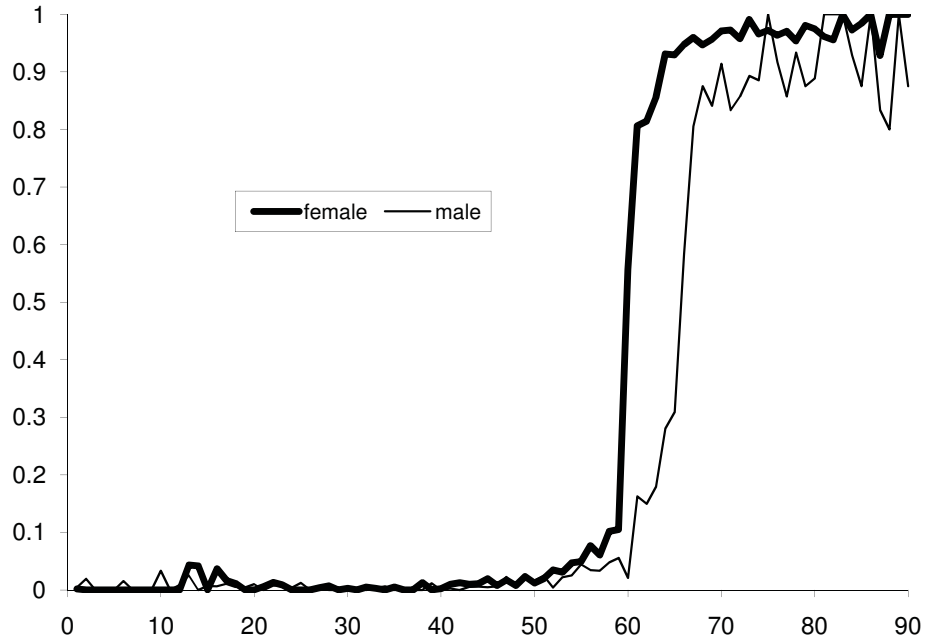
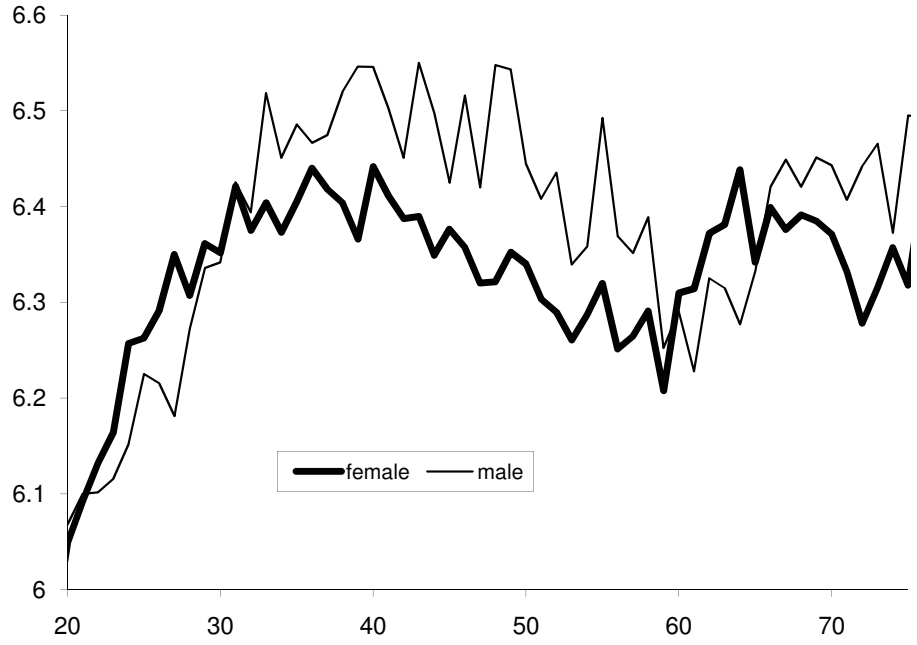
Figure 1. Household take-up of the Old Age Pension by age of oldest member.



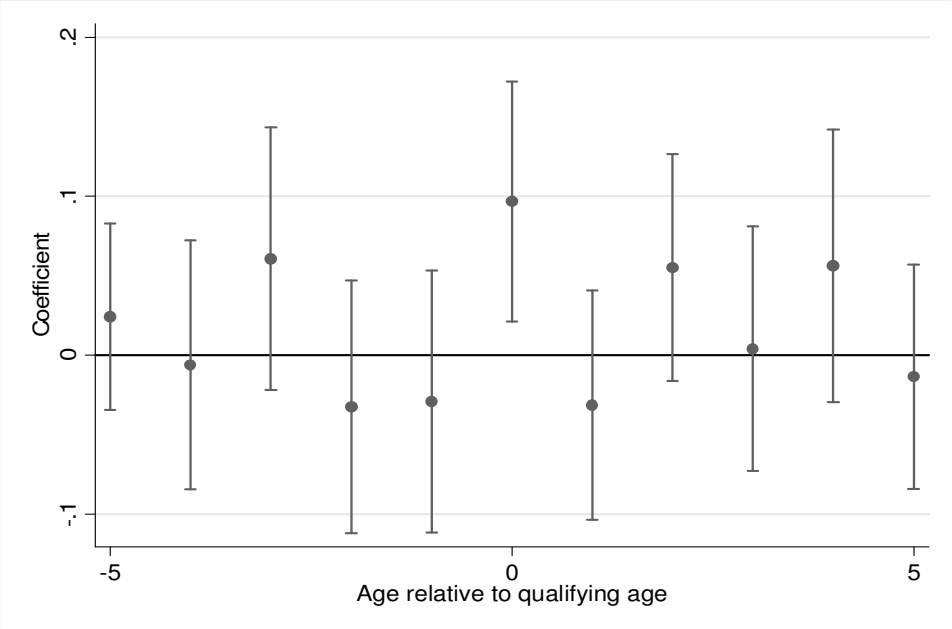
The thick line shows the fraction of households that receive the Old Age Pension, as a function of the age of the oldest female household member. In order to focus on take-up as a function of female eligibility, households with men aged 55+ are excluded from this graph. Similarly, the thin line shows the household take-up rate of the pension as a function of the age of the oldest male household member. Households with women aged 55+ are excluded from the graph. Recall that women are age-eligible from the age of 60, and men from the age of 65. Age eligibility is clearly a strong predictor of pension take-up. The data is from the September 2003 survey.

Figure 2. Log expenditure by age of oldest woman/man.



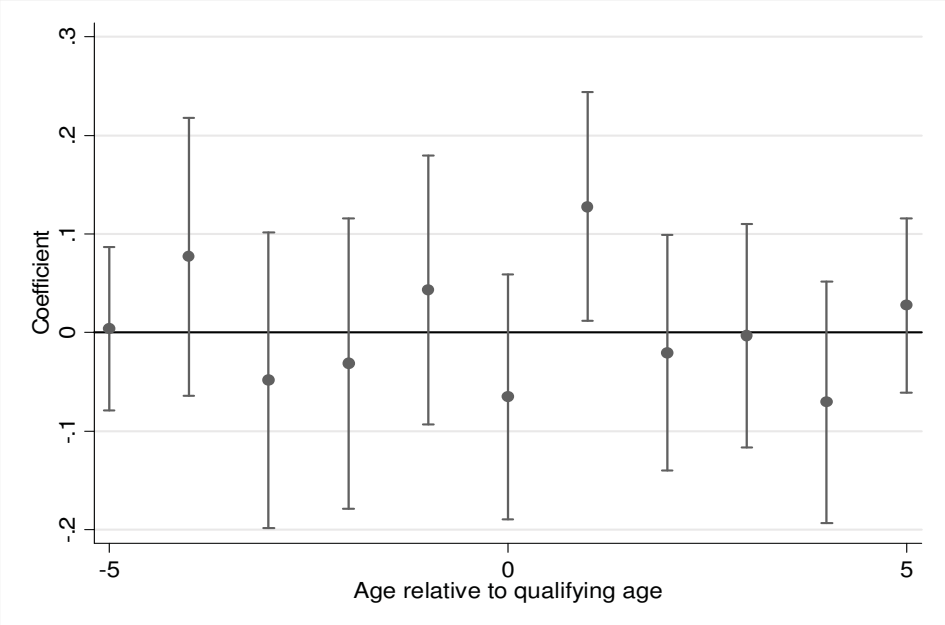
The thick (thin) line shows the average of the logarithm of monthly expenditure as a function of the age of the oldest female (male) household member. The chart pools data for all three survey waves.

Figure 3. Expenditure by annual age coefficients around the Old Age Pension qualifying age.



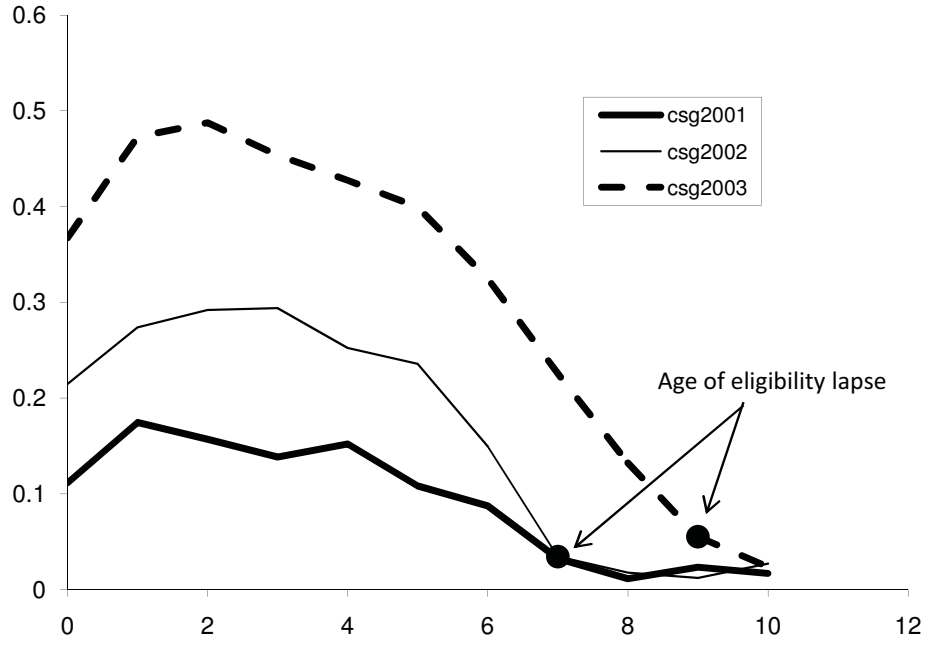
The plot is similar to Figure 3, but based on data from the 2000 Income and Expenditure Survey. This makes it easier to isolate consumption as opposed to expenditure, but the data are only cross-sectional. The above chart is based on a regression of log annual household expenditure on household members a certain number of years away from pension qualification. Including household demographics and education of the household head does not change the qualitative result that the variable at +1 is positive and uniquely significant. The significance at +1 rather than at 0 may be due to age being systematically understated in the survey. The regression pools data on men and women.

Figure 4. Expenditure by annual age coefficients around the Old Age Pension qualifying age.



The plot is similar to Figure 3, but based on data from the 2000 Income and Expenditure Survey. This makes it easier to isolate consumption as opposed to expenditure, but the data are only cross-sectional. The above chart is based on a regression of log annual household expenditure on household members a certain number of years away from pension qualification. Including household demographics and education of the household head does not change the qualitative result that the variable at +1 is positive and uniquely significant. The significance at +1 rather than at 0 may be due to age being systematically understated in the survey. The regression pools data on men and women.

Figure 5. Take-up of Child Support Grant by age of youngest household member.



The take-up rate of the Child Support Grant, by the age of the youngest child. The rapidly increasing popularity of the programme between 2001 and 2003 is clear, as is the expansion in age eligibility from 0-6 years in 2001 and 2002 to 0-8 years in 2003.

Table 1. Descriptive statistics.

	2001	2002	2003
Log monthly expenditure	6.22	6.28	6.40
	[.90]	*[.93]*	*[.92]*
The household receives the Old Age Pension	0.20	0.21	0.21
	[.40]	*[.41]*	*[.40]*
A household member is age-eligible for the pension	0.21	0.22	0.21
	[.41]	*[.41]*	*[.41]*
A female household member is age-eligible for the pension	0.18	0.18	0.18
	[.38]	*[.39]*	*[.38]*
A male household member is age-eligible for the pension	0.07	0.07	0.07
	[.26]	*[.25]*	*[.25]*
The household receives the Child Support Grant	0.06	0.10	0.17
	[.23]	*[.30]*	*[.37]*
The household has savings in any form (binary)	0.46	0.48	0.51
	[.50]	*[.50]*	*[.50]*
Household size	3.98	3.98	3.76
	[2.79]	*[2.78]*	*[2.65]*
Observations	21,096	20,026	20,336

Number of observations	61,458
Linear regressions. The dependent variable is 1 if the househol	36,192
Number of households observed twice	11,964
Number of households observed three times	6,651

Table 2. Expenditure response to the Old Age Pension.

	1	2	3	4	5	6
Dependent variable	**Expend-iture**	**Expend-iture**	**Expend-iture**	**Expend-iture**	**Expend-iture**	**Δ Expend-iture**
Household receives the pension	.195*** *[.0199]*					
Household is age-eligible for the pension		.159*** *[.0203]*		.074*** *[.0254]*	.161*** *[.0740]*	.258*** *[.0634]*
A female household member is age-eligible for the pension			.159*** *[.021]*			
A male household member is age-eligible for the pension			.147*** *[.0265]*			
Household fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Household demographics	No	No	No	Yes	No	No
Sample	Full	Full	Full	Full	No debt	Full
Observations	60506	60532	60532	60532	8564	24123
Households	35877	35889	35889	35889	4162	17742

Linear regressions. The dependent variable is 1 if the household has any savings of any form, and 0 otherwise. All regressions include household fixed effects and year fixed effects. In column 4, the sample is restricted to those households who reported not taking out any debt in the previous 12 months. In column 5, controls are included for the number of female and male household members in each of the age groups 0-4, 5-14, 15-24, 25-34, 35-44, 45-54 and 55+. Robust standard errors, shown in brackets, are clustered at the level of the survey primary sampling unit. *** significant at 1%, ** significant at 5%, * significant at 10%.

Table 3. Response in household savings to the Old Age Pension.

	1	2	3	4	5
Dependent variable: Household has savings in any form (binary indicator)					
Household receives the pension	.07*** *[.0132]*				
Household is age-eligible for then pension		.0371*** *[.0133]*		.0325* *[.0177]*	0.0178 *[.0156]*
A female household member is age-eligible for the pension			.0389*** *[.0141]*		
A male household member is age-eligible for the pension			0.0231 *[.0185]*		
Household fixed effects	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Household demographics	No	No	No	No	Yes
Sample	Full	Full	Full	No debt	Full
Observations	61409	61458	61458	41339	61458
Households	36168	36192	36192	28653	36192

Linear regressions. The dependent variable is 1 if the household has any savings of any form, and 0 otherwise. All regressions include household fixed effects and year fixed effects. In column 4, the sample is restricted to those households who reported not taking out any debt in the previous 12 months. In column 5, controls are included for the number of female and male household members in each of the age groups 0-4, 5-14, 15-24, 25-34, 35-44, 45-54 and 55+. Robust standard errors, shown in brackets, are clustered at the level of the survey primary sampling unit. *** significant at 1%, ** significant at 5%, * significant at 10%.

Table 4. Response in expenditure to expansion in Child Support Grant eligibility.

	1	2	3	4	5	6
Dependent variable	**Grant take-up (binary)**			**Log monthly expenditure**		
Presence of child aged 7	-0.00714 *[.011]*	0.0249 *[.0154]*	0.0223 *[.0154]*	0.0316 *[.0213]*	0.000327 *[.0332]*	-0.00725 *[.0329]*
Presence of child aged 7 in 2003	.142*** *[.0183]*	.169*** *[.0178]*	.168*** *[.0178]*	-0.0352 *[.0328]*	-0.0233 *[.0331]*	-0.0236 *[.0329]*
Household and year fixed effects	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*
Controls for number of children aged 0, 1, 2, ... , 9	*No*	*Yes*	*Yes*	*No*	*Yes*	*Yes*
General household demographics	*No*	*No*	*Yes*	*No*	*No*	*Yes*
Observations	61403	61403	61403	60532	60532	60532
Households	36166	36166	36166	35889	35889	35889

OLS. Columns 1-3 verify that the presence of a child aged seven in 2003 is associated with a greater probability of receiving the Child Support Grant than the presence of a child aged seven in 2001 or 2002. In columns 4-6, the logarithm of household expenditure is regressed on a binary variable indicating the presence in the household of a child aged seven, and an indicator for the presence of a child aged seven in 2003. The coefficient on the latter should be near zero under the null hypothesis of forward-looking households and liquidity constraints. The null hypothesis is not rejected. In column 2, controls are included for the number of children aged 0, 1, ..., 9 in the household. In column 3, controls are included for the number of female and male household members in each of the age groups 0-4, 5-14, 15-24, 25-34, 35-44 and 45-54 and 55+. Year and household fixed effects are used throughout. Robust standard errors, shown in brackets, are clustered at the level of the survey primary sampling unit. *** significant at 1%, ** significant at 5%, * significant at 10%.