Child Labor and Schooling Responses to Anticipated Income in South Africa[*]

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Abstract: Forward looking, unconstrained households make child labor and schooling decisions considering their permanent income and weighing the relative returns to child time in various potential activities. The timing of anticipated changes in income should have no effect on child labor and schooling in a setting where households can borrow against permanent income. However, this study documents large increases in schooling attendance and declines in total hours worked when black South African families become eligible for fully anticipatable social pension income. As an explanation, the data are most consistent with liquidity constraints for black elder males forcing rural families into less schooling for boys than they would choose absent the constraint, perhaps because of schooling costs.

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1. Introduction

This study considers to what extent child labor and schooling decisions are made in a rational, forward looking environment where families are unconstrained in their ability to weigh the return to child labor against alternative uses of the child's time. This standard human capital investment (HC) model of child labor and schooling decisions is in Shultz (1960) and has been formalized in an intertemporal setting by Ben-Porath (1967). However, Baland and Robinson (2000) emphasize that in the presence of liquidity constraints, children may work even if the return to working is below that of alternative uses of the child's time. In their terminology, child labor is then 'inefficient' from the perspective of the child's family, and policies which affect the relative return to work or schooling may have little overall influence on child labor supply or schooling.

There is a literature on the applicability of the HC model to schooling decisions that debates the significance of liquidity constraints on education. In the U.S. context, Card (2001) has argued that the general finding that estimates of the returns to schooling increase in an instrumental variable setting is consistent with liquidity constraints influencing college matriculation decisions although Carneiro and Heckman (2002) are skeptical of this interpretation. Kane (1994) and Ellwood and Kane (2000) find more direct empirical evidence of credit constraints in higher education decisions in the U.S., but Cameron and Heckman (1998, 2001) argue that their findings are more indicative of the effects of long-term family background factors. This argument may be relevant as well in the developing country evidence in Jacoby (1994) who examines liquidity constraints by comparing progress through schooling across Peruvian households that differ in their asset holdings and Beegle, Dehejia, and Gatti (2003) who argue for liquidity constraints by comparing how child labor responses to unanticipated income shocks vary across Tanzanian households that differ in their asset holdings.

Several recent studies have questioned the applicability of the unconstrained HC model to schooling and child labor decisions by documenting a correlation between unanticipated income shocks to the household and child labor supply (Jacoby and Skoufias 1997, Beegle, Dehejia, and Gatti 2003, Duryea, Lam, and Levison 2003, Guarcello, Mealli, and Rosati 2003, and Yang 2003). Household responses to unanticipated changes in income are of considerable interest in their own right, but four issues complicate their interpretation as evidence of a credit constraint. First, economic shocks will be associated with changes in the relative return to child time in various activities. For example, recovering from a flood may raise local wage rates as the community recovers; it may lower them as a consequence of a lost harvest; or it may cause the school to close. Moreover, since most children work within their own homes, a within community association between child labor or schooling's response to the shock and household assets may reflect how assets and child time enter into the household's production function. Second, for unanticipated changes in household income to affect child labor and schooling, there need to be both insurance failures and credit failures. Thus, studies of household responses to unanticipated changes in income cannot answer questions about the relevance of credit constraints under normal conditions. Third, the extent to which a change in family income is unexpected is difficult for the econometrician to identify. A perennial question in studies of crop shocks for example is the extent to which the disaster is unanticipated. Fourth, even if a researcher can separate the predictable and unpredictable parts of household income, it is never clear if this separation corresponds to what the family decision-maker perceives. For these latter two reasons, recent tests of the consumption smoothing life-cycle/permanent income hypothesis models have tended to examine household responses to well-defined, anticipated income changes (e.g. Parker 1999, Souleles 1999, 2002).

This idea of the present study is that variation in permanent income should influence child labor and schooling decisions, but absent factors such as credit constraints, child labor and schooling should not be affected by the timing of that income when the timing and the amount of income are fully anticipated. Child labor and schooling responses to anticipated income are considered in the context of a large social pension program in South Africa. The black population of South Africa is substantially poorer than the white population. Hence, when the relatively meager white Old Age Pension program (OAP) was extended to other South Africans at the end of apartheid, the social pension became a large cash transfer. There is a means test in the OAP that binds for most white households but affects few black households (Case and Deaton 1998). The primary determinant of the cash transfer in black households is the age of the beneficiary (and thereby does not depend on family background or unobservables), and there is little uncertainty about the benefit level (Alderman 1999). Moreover, the South African social pensions are so large (for black households, the 1999 benefit of 520 Rand per month is 125 percent of black median per capita income) that they are well-known and highly anticipated by recipients (Lund 1993). Further, black South African households are typically multi-generational, so there is ample scope for this pension income to be shared with children. In fact, other studies that use the OAP to identify income effects have documented sharing of the pension income with co-resident adults (Lund 1993, Case 2001, Bertrand, Mullainathan, and Miller 2003), across households (Jensen 2004, Edmonds, Mammen, and Miller 2005), and with children (Duflo 2000, 2003). Consequently, the South African social pension seems an unusually clear environment in which to consider the degree to which child labor and schooling are influenced by factors such as credit constraints in addition to HC type behavior.

This study examines the response of schooling and child labor to the timing of income by comparing child activities in households that are eligible for the OAP to households that are nearly eligible. The idea is that these two types of households face similar permanent incomes but differ in the timing of the income. A third order polynomial expansion in the age of the oldest man and woman is used to control for differences between eligible and nearly eligible households and thereby to construct estimates of how child labor and schooling change at the age of pension eligibility. In data from the 1999 Survey of the Activities of Youth in South Africa, there appear to be large changes in schooling when a male elder transitions from nearly eligible to eligible for the social pension. In rural areas, the schooling attendance of children 13-17 rises with male eligibility to nearly 100 percent. Hours worked drop substantially, and completed schooling appears to be increasing in the time a child has lived with a male eligible for the pension. Boys generally have lower schooling attendance rates, are more active in market work (defined as wage work, self-employment, or work in the family farm or business), and lower school completion rates absent the pension. Schooling and child labor appears to be more sensitive to the timing of income for boys than girls, especially with male pension eligibility. One exception is in total hours worked which includes domestic chores - girls tend to work more hours than boys to begin with, and girls experience larger declines in hours. These changes in work and schooling appear only with male eligibility, and the changes in child labor and schooling that occur with male eligibility approach levels of child labor and schooling observed when children live with a nearly eligible elder female. Hence, the patterns in the data are consistent with gender differences in access to credit. Two other pieces of evidence are consistent with the idea that liquidity constraints drive this association between child labor, schooling, and the timing of income. First, the association between child labor, schooling, and the timing of income are largest in households where the pensioner is uneducated. Second, child labor declines and schooling increases only when the first pensioner in the household becomes eligible for the pension benefit. Moreover, the data suggest a decline in the perception that children cannot afford school with male pension eligibility. Thus, direct and indirect schooling costs may be the expense that keeps children out of school when families are liquidity constrained.

The next section provides background on the OAP and reviews the existing research on household responses to the program. Section 3 describes the data, and section 4 presents the basic empirical strategy. Section 5 contains the findings. Section 6 discusses several factors that may influence the interpretation of the results in this study including measurement error in age, endogenous household composition, and retirement absent the pension, and considers various explanations for timing of income effects on child labor and schooling. Section 7 concludes.

2. Background on the Old Age Pension Program (OAP)

The OAP has three important attributes that make it useful for examining child labor and schooling responses to anticipated income. First, pension eligibility is largely determined by the age of resident elders. A woman is pension eligible if she is age 60 or older. A man is pension eligible at age 65 or older. There is a means test in the pension formula that is important in the white population, but in practice its impact on benefit determination for black South Africans is minimal because of the relative depravation of the black population.[1] Second, neither the activities of children nor their presence influences pension eligibility. Hence, the OAP does not create any incentive to change household composition or alter the activities of household members in order to receive the pension (though the behavioral response to the income may alter composition as well as the activities of members). Third, grandparents often reside with a grandchild in black South African households. In the data used in this study, 42 percent of all black children 5-17 in South Africa co-reside with an individual between the ages of 50 and 75 and 24 percent live with a person who is age eligible for the pension. Hence, there is ample scope for the sharing of pension income with co-resident children and their parents.

The academic literature on the OAP largely focuses on using the age eligibility of the pension benefit formula to identify income effects, and it finds substantial evidence that elderly

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[1] A number of authors have observed that the means-test does not bind for most black African households (Case and Deaton (1998), Alderman (1999), Case (2001), Jensen (2004), Duflo (2000, 2003), and Bertrand, Mullainathan, and Miller (2003)). While these studies work with data from 1993, the means-test in the pension benefit formula has not changed substantively between 1993 and 1999, and the income at which the means-test begins has increased since 1993. The means-test is based on the *individual* wage income of the recipient, but most elder blacks do not pay income taxes and thereby have no incentive to declare income for the calculation of the means-test. In the South African tax code, *individuals* age 65 and older do not pay income tax so long as their personal income is below R47,222 per year. Less than 9 percent of the nearly eligible population in the data in this study report total *household* income at or above R47, 222. Even if reported, relatively few elder blacks report incomes near the pension age which are high enough to be affected by the means-test, and in the data used in this study all but 4 percent of pension recipients report total *household* income in a category at or above the maximum pension benefit of 520 Rand per month. The means-test only begins when official *individual* incomes exceed 30 percent of the maximum grant. It does not include the income of other household members other than the spouse, and it explicitly does not include spouse's pension income. Consequently, if the means test were implemented regardless of an individual's tax status, it would only affect pensioners with a formal sector income above 156 Rand per month. When the dataset used in this study is restricted to households in the nearly eligible population, only 42 percent of *households* have per capita incomes above 156 Rand per month when all sources of income are considered (most of which would not be reported to tax authorities).

pension income is shared with family members. Lund (1993) provides narrative evidence of pensioners sharing pension income through the types of expenditures they make with the income, and Case and Deaton (1998) formally show that the marginal propensity to spend pension income appears similar to that of other income for a host of expenditure categories. Duflo (2000, 2003) compares the nutritional status of children who co-reside with a pension eligible individual to that of children who do not. She finds evidence that girls who live with a pension eligible female have substantially better weight for height and height for age than girls who do not. Case (2001) notes that additional pension income is associated with improved health benefits in older cohorts as well. The sharing of the pension income appears to extend beyond direct expenditures. Bertrand, Miller, and Mullainathan (2003) note declines in labor supply for prime age adults with increases in pension income, and. Edmonds, Mammen, and Miller (2005) observe changes in household composition associated with pension eligibility. Even outside of a given residence, Jensen (2004) presents evidence that pension income is shared in part with family members outside of the household as it does not fully crowd out private, inter-household transfers. With the exception of Edmonds, Mammen, and Miller (2005), all of the studies focus on identifying the effects of increases in household income as a result of the pension, rather than the timing of pension income as considered in the present study.

A persistent theme in the academic literature on the OAP is that the effect of income given to men is different from that of income given to women. Case and Deaton (1998) for example note that female headed households spend less on alcohol and tobacco. Duflo (2000, 2003) observes improvements in child nutrition with the pension entirely in girls and only with female pension eligibility. The declines in labor supply with income among prime age adults documented in Bertrand, Miller, and Mullainathan (2003) are largest among males when women are eligible for the pension. Edmonds, Mammen, and Miller (2005) only consider changes in household composition with female pension eligibility, but they report that in supplemental work, they did not observe significant changes in household composition with male pension eligibility. Because of these findings, this study will allow the changes in child labor and schooling examined herein to vary with the gender of the pension recipient.

3. Data and Preliminary Findings

This study compares the child labor and schooling status of children who co-reside with a pension eligible individual to children living with an elder who is nearly eligible. Data for this study comes from the June 1999 Survey of the Activities of Youth in South Africa (SAYP). The SAYP was conducted by Statistics South Africa with the help of the International Labor Organization and is the first nationally representative survey in South Africa to collect detailed data about the allocation of child time. The survey also collects basic information about all of the members of a child's household including age eligibility and participation in the OAP.

As is typical for ILO child labor surveys, the SAYP was conducted in two phases (Statistics South Africa 2000). In the first stage, randomly selected households from within randomly selected census enumeration areas (PSU) are interviewed to assess basic household characteristics and whether any children are working in the household. In the case of the SAYP, 25 urban and 50 rural households are interviewed from each PSU. In the second stage, a random sample of households where at least 1 child reports working are interviewed with detailed questions about the activities of children including detail on hours worked in 9 different types of activities. Of the 10,480 sampled black households with children in phase 1, 74 percent had at least one child participating in an economic activity. In this second stage, 5 urban and 10 rural households are interviewed in each PSU. This study combines phase 1 and phase 2 data. Detail on the activities of children come from phase 2 of the SAYP, but participation status comes from phase 1. In combining phase 1 and 2 data, households interviewed in phase 2 are weighted to reflect that they are a simple random sample of households with working children in the PSU in order to retain the randomness of selection into the dataset.

Table 1 presents summary statistics for the outcome measures used in this study. The focus of this study is on black children in black headed households, because they are the group least likely to be affected by the means-test in the pension benefit formula and the group where schooling is lowest, child labor highest. The first column of table 1 presents summary statistics for the entire population of black children in black headed households. The top panel of table 1 includes all children for which data is available (ages 5-17). Of the 15,485 black children age 5-17 in black headed households, half are female, 88 percent attend school, 85 percent work in either domestic chores or market work, and 29 percent are active in market work which includes wage employment, self employment, and work in the family farm or business. 2 percent of children age 5-17 are active in market work without attending school. On average children work 1.4 hours per day in domestic and market work, and 1 percent work 40 or more hours per week in market work.[2] 21 percent of children 5-17 have completed primary school, and they average 4.5 years of completed schooling.

The sample in this study is restricted to rural children who live with an elder that is between 50 and 75. Summary statistics for children who live with an elder between 50 and 75 are in column 2 of Table 1, and they appear similar to the full population. Summary statistics for rural children who live with an elder between 50 and 75 are in column 3 of table 1. Access to credit is more apt to be an issue in rural South Africa, and work is more prevalent and schooling attainment is lower in rural South Africa. The results below focus only on rural children, because the data reveal little evidence (in magnitudes or statistical significance) of changes in schooling or child labor with pension eligibility in urban areas. Hence, the data summarized in column 3 of Table 1 are the data used in most of the rest of the study.

[2] Phase 1 of the SAYP provides information on hours worked for individuals who work 0 hours, and phase 2 provides detail on hours worked when at least 1 child in the household is working. Thus, combined, they provide detail on total hours worked for a nationally representative sample. However, eligibility for phase 2 of the SAYP depends on at least 1 child in the household working in market work or at least 1 child working typically an hour or more per day in domestic work. Thus, it is not possible to compute total hours for children that work in domestic work for less than an hour per day who are in households where no child works in market work. This corresponds to just under 5 percent of the sample. This study assumes that these children work 0 hours in domestic work per day. The hours worked results are not sensitive (in statistical significance or in coefficient magnitudes) to making this assumption, to imputing hours work to these children, or to omitting them. It is also important to emphasis that none of the results that use schooling attainment or indicators about participation in types of work or school are affected by this problem; only hours worked is affected.

The bottom panel of table 1 focuses on children age 13 and above, and the main changes in child activities observed in the data are in this age range. Age 13 is chosen as a cut, because it is the age a student would normally have completed primary and entered lower secondary. It also seems a logical choice given the age trends observed in the data.[3] Age 17 is the upper limit in the dataset. 92 percent of rural children age 13-17 who co-reside with an elder 50-75 attend school (column 3, bottom panel). School attendance is higher in this subsample than in the top panel because of the low schooling rates of children under 7. Of rural children 13-17 living with an elder, 96 percent work. 47 percent are engaged in market work, 4 percent work full time in market work, and 3 percent work in market work without attending school. On average, these children work 2.5 hours per day, almost 18 hours per week. Nearly half of rural children 13-17 have completed primary school.

The test of child labor and schooling responses to anticipated income in this paper relies on comparing households that have similar permanent incomes who vary in the timing of that income. Columns 4 and 5 of Table 1 present summary statistics that mimic this idea. Column 4 contains means for children living with an elder within 2 years below pension eligibility (an elder woman age 58 or 59, an elder man age 63 or 64). Column 5 describes the data for children living with an elder within 2 years above pension eligibility (an elder woman age 61 or 62, an elder man age 66 or 67). All of the main findings of this study are previewed by comparing columns 4 and 5 for rural children age 13-17.

Rural children age 13-17 living with a person who is pension eligible (column 5) are 8 percentage points or nearly 10 percent more likely to attend school than a child who is nearly eligible (column 4). Their attendance is higher despite the fact that school attendance is decline in the age of the oldest person in the household. This finding and age trend are evident in the raw data plotted in figure 1. It pictures the school attendance rate of rural children 13-17 against the age of the oldest man in the household. Pension eligibility begins at age 65 and is

[3] Schooling attendance is increasing between ages 5 and 8, flat between ages 9 and 12, then decreasing in age. The probability the child works in either market or domestic work is increasing through age 12, then relatively flat.

represented in the figure by a vertical line. There is a clear, downward trend in schooling attendance with the age of the oldest man in the household until age 65, when attendance rates increase dramatically. The means in Figure 1 differ from those in columns 3 and 4 of Table 1, because Figure 1 only plots schooling against age of the oldest man, rather than including both sexes as in Table 1. In the results below, the findings will be driven largely by male pension eligibility.

These higher school attendance rates are not mirrored in a decline in the probability a child works (which essentially does not change with pension eligibility in the dataset). However, the probability that a child is engaged in market work full time is 5 percentage points or 83 percent lower in households with a pension eligible individual (column 5) relative to a household with a person who is nearly eligible (column 4). The probability a child engages in market work without school is 4 percentage points or 67 percent lower in pension eligible households. Again, these observed differences in the data are against the general trend of increasing market work without schooling in the age of the oldest person in the household. Figure 2 plots the probability a child engages in market work without attending school against the age of the oldest man in the household. Because it only focuses on male elders, the means in Figure 2 differ from Table 1. Note that prior to male pension eligibility, the probability a child engages in market work without attending school is increasing, but very few children who live with a pension eligible male work in market work without at least attending school. This additional school attendance and declines in market work without school for children living with a pension eligible individual are also associated in Table 1 with 31 percent fewer hours worked per day and a 24 percent increase in primary school completion rates. Of course, the difficulty with all of these comparisons in table 1 is that the children in column 5 differ systematically from the children in column 4 in that the column 5 children live with older elders. The methodology in the next section aims to deal with this problem.

4. Empirical methodology

Since pensioners are not randomly distributed among households, the comparison of children in households with pensioners to children in households where the elder is nearly pension eligible raises two concerns. First, take-up of the pension may be an endogenous household decision. This endogenous pension take-up problem is addressed by focusing on age eligibility rather than actual take-up. Second, households with pensioners may differ systematically from non-pension households. They are older on average than are households without a pensioner and are more apt to contain multiple generations. Moreover, the presence of an elder may influence the time allocation of children in any number of ways.

Systematic differences between households with and without pension eligible individuals influence the empirical work in two ways. First, the sample is limited to children that co-reside with an elder between the ages of 50 and 75. Restricting the sample in this manner means that the effect of a pensioner of a given gender is identified by comparing the effect of a person who is near but below pension age to a person who is of pension age. An obvious concern in this approach is that the pension indicator may capture age trends in addition to the effect of pension income on child labor. Second, this study allows for differences in child labor with the age of the elder by including a polynomial expansion in the ages of the oldest male and female in the household in each regression. Changes in child labor and schooling with pension eligibility are captured by indicators that the elder male, female, or both is pension eligible. The basic regression approach is thus:

(1) $$H_{ij} = \alpha_0 + \alpha_1 EM_i + \alpha_2 EMF_i + \beta_1 PE_i + \beta_2 PEF_i + \beta_3 PEMF_i + \pi\left(AM_i, AF_i\right) + \varepsilon_{ij}$$

where *H* is one of the outcome variables summarized in table 1 for child *j* in household *i*, *PE* indicates that a person in the household is pension eligible, *PEF* indicates there is a pension eligible female in the households, and *PEMF* indicates that there is both a pension eligible male and female in the household. $\pi\left(AM_i, AF_i\right)$ is a third order polynomial expansion in the age of the oldest man and age of the oldest woman and all of their interactions.[4] Because the sample is restricted to households with at least 1 person 50 to 75, EM is an indicator that there is a male 50 to 75 in the household and EMF indicates that there is both a male and a female age 50 to 75 in the household. A Female 50 to 75 is the omitted category. Standard errors are clustered at the age of oldest man - age of the oldest woman cell level throughout.[5]

The polynomial in the age of the oldest man and woman controls for changes in permanent income (perhaps because of mortality risk), labor demand, and other factors that vary with age of elders. Thus, the coefficients on pension eligibility compare pension eligible households to nearly eligible households with approximately the same expected permanent income (ignoring differences owing to discounting) but that differ in the timing of that income. Absent credit constraints, any permanent income effects that differ owing to the presence of a male or female elder in the household will be captured by α_1 and similarly for α_2 and two elder households. Hence, β_1 is a timing of income effect. In particular, the coefficient on the indicator of a pension eligible individual in the household, β_1, is interpreted as the change in schooling or child labor associated with moving an elder from nearly eligible to a pension eligible male. β_2 indicates how the change in schooling or child labor varies from β_1 if the pensioner is female rather than male. 4 percent of children live in households with multiple pension eligible elders, and β_3 is the additional incremental change in schooling or child labor if

[4] Precise coefficient estimates are sensitive to the order of the polynomial, but the statistical significance and sign of the coefficients for all of the main results below (columns 3-5 of table 3, table 5) do not vary with whether a second, third, or fourth order expansion is used. A third order is chosen, because with some of the more fine data cuts below such as columns 4 and 5 of table 3, higher order terms drop out.

[5] It would be preferable to cluster standard errors on either the age of the eldest male elder or female elder (separately). Either approach yields smaller standard errors than the approach in the text. Card and Lee (2004) suggest an additional inflation factor for the standard errors that relies on estimating the cell variance of the raw data and comparing that to the variation generated by the model. The number of age of the oldest man - age of the oldest woman cells in the data varies depending on what households are included in each regression, but the ratio of observations to cells can be very small. For example, in the first regression in table 3, there are 3708 observations and 1074 clusters (the density for each cluster is obviously not uniform given the selection rule of having at least one member between 50 and 75 - in this region it is not unusual to observe cells with more than 30 observations). Hence, the Card-Lee adjustment is not applied in this paper. Only the schooling attendance results for boys are statistically significant at 5 percent with the Card-Lee adjustment. The small ratio of observations to cells also makes clear that identification relies on the functional form of the polynomial expansion and is not semi-parametric.

the household contains both a male and a female pension eligible person. In sum, β_1 is the change in H associated with moving an elder male from ineligible to eligible, $\beta_1 + \beta_2$ is the change in *H* associated with moving an elder female from ineligible to eligible, and $\beta_1 + \beta_2 + \beta_3$ is the change in *H* associated with moving from no eligible pensioners to a male and a female pension eligible individual.

The focus of the empirical work in the next section is on whether the declines in hours worked and increases in schooling evident in the raw data are robust to controlling for underlying age trends as in (1). If the coefficient on pension eligibility is to be interpreted as a timing of income effect, it is important to show that pension eligibility is associated with an increase in pension take-up and household income. These results are in Table 2. Column 1 reports sample means and standard errors for each variable in the table. Column 2 contains estimates of β_1 from (1) and its standard errors, column 3 contains β_2, and column 4 contains β_3. In the first row, the dependent variable in (1) is an indicator that a child's family reports receiving pension income. Having a pension eligible individual in the household increases the probability that the family reports pension income by 46 percentage points. In the data, the probability that a household without a pension eligible individual reports receiving the pension is 5 percent, so this is a large increase in take-up. The probability of reporting take-up is slightly higher when women become eligible than when men become eligible and with multiple eligible individuals although neither difference is statistically significant. The observation that β_1 is not 1 is consistent with the take-up patterns found in other papers where authors have attributed this to leakage in pension participation, misreporting, measurement error in age, and the influence of factors other than age in the eligibility rules. To the extent that take-up is not perfect, estimates of the true effect of pension take-up are attenuated.

This rise in take-up is also associated with increased income. The income data in the SAYP are poor, but reported incomes increase substantially with pension eligibility.[6] Reported income is nearly 110 percent higher with the first pension eligible person in the household and 115 percent higher if there are two pension eligible individuals (relative to none) although the 95 percent confidence interval on the increase in income when 2 persons become eligible ranges from 25 percent to 205 percent. The reported rise in family income is higher when the pensioner is female than male, albeit not in a statistically significant way.

For the approach in (1) to be valid, predetermined child and family attributes should not change discretely with pension eligibility. The remaining rows of Table 2 contain tests for the comparability of covariates around pension eligibility for available covariates. Though changes in household composition may be important mechanisms for how household react to the realization of anticipated income, the present data do not indicate any statistically significant changes in household size, number of children age 17 and under, or fraction female with pension eligibility. Location changes also do not appear to be associated with pension eligibility as neither urban status (urban households are added for this row only) nor provincial location appears to change with pension eligibility. Consistent with this, there do not appear to be a large number of reported moves as the probability a child has moved with pension eligibility does not change in a statistically significant way. Likewise, neither the child's gender nor age appear to change with pension eligibility, but all subsequent empirical work will also include a third order polynomial in child's age and a control for gender in estimating (1).

5. Main Findings

The realization of anticipated income may change net savings, but schooling and child labor are determined by returns to each and permanent income if households make decisions about child labor and schooling in an unconstrained, forward looking environment. As was clear

[6] The questionnaire collects income by asking for total household income in the last year from one individual. Rather than recording an income amount, the survey asks the respondent to select one of several broad categories and non-response is an issue. The income variable is coded by taking the log of the midpoint of each income category and non-responders are omitted.

in the raw data discussed in section 3, there appear to be substantive changes in schooling attendance and time spent working associated with the realization of fully anticipatable income. The patterns in the raw data in Table 1 are robust to the specification of (1) and a wide range of controls, and are contrary to the predictions of the basic unconstrained human capital investment model.

5.1 Schooling

Results of estimating (1) with school attendance as the dependent variable are in Table 3. $(\alpha_1, \alpha_2, \beta_1, \beta_2, \beta_3)$ are included in the table. The bottom panel of the table reports linear combinations (and the corresponding standard errors) of some of the coefficients. For rural children 5-17, having an elder, nearly eligible male present is associated with schooling attendance that is 5 percentage points lower than having a nearly eligible elder female present. Male pension eligibility is associated with a nearly 5 percent increase in school attendance. Combined then, schooling attendance for a children 5-17 living with a pension eligible male is about the same as schooling attendance for a child living with a nearly eligible female (this is the hypothesis that α_1 (Elder Male) + β_1 (PE in HH) =0). Columns 2 and 3 split the full sample by age. As noted in the raw data, the variation in schooling is much greater amongst older children, and these patterns of increasing schooling with male pension eligibility are larger in magnitude and statistically significant for children age 13-17 (column 3).

Moreover, the data do not present evidence of similar changes in school attendance at other ages of elder men. Figure 3 contains estimates of β_1 from equation (1) and 95 percent confidence bounds for the estimate, allowing for changes in school attendance at each age between age 55 and 75. That is, (1) is re-estimated 21 times, pretending that male pension eligibility is at a different age each time. Prior to age 65 when men become pension eligible, there are no observed statistically significant changes in schooling attendance for rural children 13-17. The largest change in school attendance occurs with male pension eligibility at age 65. Estimates of the change in schooling are also considerably more precise at age 65 than at younger ages. It is not surprising, given the smoothing inherent in the underlying polynomial expansion, that we observe smaller, statistically significant changes in school enrollment at ages after 65. The estimate of a significant discontinuity at age 70 is obvious in the raw data of figure 1 as well because of an unusually low attendance rate associated with age 69. However, the raw data in figure 1 clearly suggest that the break in the age trend occurs at age 65.

The increase in schooling in rural areas with male pension eligibility appears to be largely among males. Columns 4 and 5 of Table 3 split the rural sample by the child's gender. While none of the results for females are statistically significant, the column 4 findings for male children are both larger in magnitude and statistically significant. Male pension eligibility is associated with an 18 percentage point increase in schooling attendance for boys. Female eligibility is associated with a small, insignificant increase in attendance, and joint male and female eligibility is similar in magnitude to male eligibility alone. In interpreting the differences in the effects of male eligibility for boys versus girls, it is important to bear in mind that girls have higher attendance rates than boys do in the nearly eligible population. Hence, there is more scope for increases in schooling among boys, although we observe statistically significant increases in female attendance when both elder males and females are present and eligible (last row, column 5). A similar issue complicates the interpretation of the differences observed with the gender of the pension eligible individual. Schooling attendance rates of boys living with a nearly eligible female are nearly 7 percentage points higher than that of boys living with a nearly eligible male (top row, column 4). When combined, boys living with pension eligible males have higher school attendance than boys living with nearly eligible females (and no males), but the difference is not statistically significant (third row from bottom). Hence, the data suggest rising school attendance for boys with male eligibility although the data are consistent with the assertion that the rise in attendance for boys with male eligibility puts their school attendance on par with boys who live with a nearly eligible female (and no elder male).

5.2 Child labor

Increases in schooling do not necessarily translate into declines in child labor, and in the present case, the data suggest the substantive changes are in hours worked rather than in participation rates. The bottom row of table 2 shows that pension eligibility is not associated with a significant change in the probability that there is at least one working in the household works (and the households is thereby eligible for phase 2 of the SAYP). At the individual level, Table 4 looks at the association between work status and pension eligibility in the sample of rural children age 13-17. The set-up of Table 4 mirrors table 3. In the first three columns, results are reported from estimating equation (1) with an indicator that the child works in either market or domestic work (any work) as the dependent variable. In the last 3 columns, the dependent variable is an indicator that the child works in market work. Boys and girls are pooled in columns 1 and 4. Columns 2 and 5 focus only on boys (3 and 6 refer to girls). There are no statistically significant changes in work status with either male or female eligibility. Only girls who move from no eligible individuals to two pension eligible individuals experience a significant change in the probability the child works. That said, estimates of the changes in market work status with pension eligibility are large, especially for girls. However, these large coefficients have large standard errors. The data do not suggest that they are statistically different than zero.

While statistically significant changes in work status are not observed, the data suggest substantial declines in hours worked. The first three columns of Table 5 contains the results of estimating (1) for rural children 13-17 with average hours worked per day in the last week as the dependent variable. Again, male pension eligibility is associated with declining hours worked, and the declines observed with female eligibility are smaller. Also, as before, the changes observed in boys with male pension eligibility are not statistically different from the changes observed in girls. However, the hours worked results contain two differences. First, unlike the participation and attendance results, the change in total hours worked with female eligibility is not statistically different than the change observed with male eligibility. Second, the magnitude of the decline in total hours worked is larger for girls than boys. In nearly eligible rural households, the average girl 13-17 works 2.5 more hours per day than the average boy who works 1.25 hours per day. Thus, the scope for a decline in hours worked is much greater for girls than boys.

In fact, the additional hours worked by girls are largely in domestic chores. The declines in hours in market work are larger for boys, who work more hours in wage work, self-employment, and in the family farm or business. Columns 4-6 report estimates from (1) with hours worked in market work as the dependent variable. Boys work nearly 0.7 fewer hours per day in market work with male pension eligibility. This accounts for 75 percent of the decline in total hours worked for boys associated with male pension eligibility. In contrast, the declines in average hours in market work are not statistically significant for girls, and changes in hours in market work are only 45 percent of the change in total hours worked for girls. The last three columns of Table 5 contain estimates of (1) with an indicator that the child works at least 40 hours per week in market work (full time) as the dependent variable. Boys who live with an elder male are more likely to work full time in market work, and there is a significant decline in full time work for boys with male pension eligibility. As with schooling, the data are consistent in general with the claim that with male pension eligibility, hours worked in boys living with a pension eligible male decline to levels similar to that observed among girls living with a nearly eligible female.

5.3 Schooling Attainment

If these declines in child labor and increases in schooling are real, schooling attainment should be affected. In particular, the longer a child has been exposed to a pensioner during its schooling years, the greater the child's attainment should be (after controlling for age). Table 6 considers the relationship between years of completed schooling (columns 1-5), primary school completion (columns 6-8) and length of schooling exposure to a pension eligible individual for children in rural households. The length of a child's exposure to a pension eligible individual is calculated by first identifying the number of years that the child has been eligible for school (age minus 6). The child's exposure to a pension eligible individual is the minimum of the number of years the child has been eligible for school and the number of years a resident elder has been pension eligible.[7] To mimic equation (1), controls are also included for the number of school years a child has been exposed to an elder age 50 or more. This is computed in the same way as school years with a pensioner except using the number of years a resident elder has been age 50 or more in place of pension eligible. Table 6, then, replicates the idea of equation (1) by regressing schooling attainment on a third order polynomial in the child's age, an indicator for the child's gender (in columns 1 and 6 only), a third order polynomial in the age of the oldest man and woman, the number of schooling years the child has been exposed to an elder (row 1), and elder female (row 2), and two elders (row 3), and the number of schooling years the child has been exposed to a pensioner in the household (row 4), a female pensioner (row 5), and the number of schooling years the child has been exposed to both a male and a female pensioner (row 6).

Each additional year of exposure to a male pension eligible individual increases schooling attainment and primary school completion rates for boys, but no such effects are evident for girls or female pension eligibility. Surprisingly, while the changes in schooling *attendance* with male pension eligibility are generally larger for boys than girls, they were not statistically different. The larger improvements in *attainment* for boys with male pension eligibility are statistically significant. Column 1 contains the results for children 5-17 in rural areas. Each additional schooling year of exposure to a male pensioner is associated with a tenth of a year of a grade completed. Columns 2 and 3 split the sample by the child's gender, and it is clear that the patterns observed in the pooled sample are for boys. It is worth nothing that even as early as age 8, girls average more than a fifth of a year of additional completed schooling relative to boys, and the additional gain in completed schooling for boys for each year of male

[7] Universal extension of the white old age pension program to the black population was not achieved until 1993. Hence, the number of years a resident elder has been pension eligible has a maximum of 6 years.

pension eligibility is approximately one fifth of a school year among both younger and older boys (columns 4 and 5). These increases in schooling attainment for boys may reflect their relative lagging in completing schooling. The completed schooling of girls in the nearly eligible population of male elders is more than a year and a half higher than the completed schooling for boys. These increases in schooling attainment also show up in primary school completion rates. For primary school completion, the sample is restricted to children age 13 and up, because it would be very unusual for a child younger than that to have completed primary school. Each additional schooling year of exposure to a male pensioner increases the probability that a male completes primary school by 3 percentage points (column 7).

5.4 Gender of the Pensioner

In sum, all of the findings suggest that there are relatively large changes in schooling status, hours worked, and schooling attainment with male pension eligibility. With the exception of total hours worked, the observed changes are largest for boys. The differences in results by gender are consistent with patterns of time allocation such that individuals who work more (attend school less) experience larger declines in hours (increases in schooling). If the schooling and child labor supply decisions were efficient in the Baland and Robinson (2000) sense that they reflect the balancing of returns to work against returns to school and permanent income, whether or not the family has yet to receive the anticipated pension income should not affect the allocation of child time. Thus, the schooling and child labor results in this section are inconsistent with the unconstrained HC model of schooling and child labor supply.

The observation that the response of child labor and schooling to the timing of income is greater for male pension eligibility than female pension eligibility at first appears a surprise given the existing literature on the OAP which generally finds larger changes with pension income to women. Duflo (2003) for example observes a connection between income to women and investments in girls. This study is similar in that it finds a connection between the timing of income to men and investments in boys. However, typically the results for boys and girls with male pension eligibility are within a 95 percent confidence interval. Hence, the more substantive observation is that schooling appears to increase and time working decrease more with male pension eligibility than with female eligibility. There are several plausible explanations for the findings associated with male eligibility in this study. Edmonds (2004) discusses several other behavioral models that might generate different child labor and schooling responses for men. One obvious possibility is that men are more apt to be liquidity constrained. There are several reasons men might be more liquidity constrained including mortality risk, gender differences in access to credit or credit programs, and gender differences in behavior that make men a greater credit risk. Consistent with this idea of gender differences in illiquidity is that the timing of income effects observed with male eligibility largely move children to the hours worked and school attendance rates that exist in households with nearly eligible women.

6. Identification and Interpretation

6.1 Identification of Timing of Income Effects

The above results are based on comparing households that currently receive the pension to households that will receive the pension in the near future after controlling for general patterns in child labor and schooling associated with the age of the oldest man and woman in the household. This section highlights issues of cohort effects, lead and lags in program response, measurement error in age, endogenous household composition, retirement absent the pension, and price effects of the means-test that should be discussed in the context of this study.

First, pensioners are older and thereby will have older children living with them. If the elder's age mapped one to one with the child's age, it would be impossible to separate child cohort effects from pension eligibility effects. However, child and pensioner ages are not perfectly correlated, and the results for male pension eligibility in tables 3-5 are not sensitive to whether child age is included with the polynomial in the age of the oldest man and woman.

Second, leads and lags in program effects and measurement error in age may attenuate estimates of the change in child labor and schooling with pension eligibility. To see how leads or lags may be relevant, consider a liquidity constrained parent deciding whether to educate a child in a setting with important diploma effects in the return to schooling. If a parent knows that in two years time, the girl will be needed to work and hence she will not complete primary school, a parent may choose to keep the child at home now. However, if the parent anticipates receipt of the pension income in two years, then the parent may elect to educate the child today even though the parent cannot access the income itself. Classical measurement error in age can also attenuate any observed effects of the timing of income, because there are likely to be several households defined as pension eligible that are not eligible and some households classified as ineligible that receive the pension.[8]

Third, some households respond to the pension by re-arranging. Thus, the observed effects of the pension on child labor may reflect changes in household composition rather than a change in child labor supply. To be clear: the concern is *not* that the timing of income induces changes in household composition which produces an indirect relationship between the timing of income and changes in time allocation. This sort of variation is what this study hopes to capture as it reflects a timing of income effect. Instead, the concern is that maybe the time allocation of existing children are not affected but new children enter the household and make it appear that schooling status and child labor changes even though it does not in reality. Three pieces of evidence suggest that this is not a first order concern. First, in table 2 (next to last row), there is no evidence of an association between pension eligibility and whether a child reports having moved within the last two years. Second, in rural households, after controlling for the age of the oldest man and woman in the household as in equation (1), male eligibility is associated with a 0.02, statistically insignificant increase in children 13-17 (of just over 1 percent). Hence, the magnitudes of the changes in the number of children are too small to generate the results observed in this study. Third, neither the magnitude nor statistical significance of the basic

[8] A related problem without a clearly defined bias is age-heaping. One option would be to omit pensioners at the ages of pension eligibility from the analysis. While this has some effect on the coefficient estimates, the basic sign and statistical significance of the main results of are not altered by omitting individuals at pension age.

results differs substantively if basic household composition controls are included. This is evident in Table 7. The two main findings of this study (increases in schooling and declines in hours worked for rural children 13-17) are reproduced, adding controls for household composition to the basic specification used in table 3-5. Specifically, in addition to the specification of earlier tables, province effects and controls for the residency status of the child's mother, father, grandmother, grandfather, household size, number of children, and fraction of children female are included in Table 7, yet the results hardly vary from earlier specifications that did not control for living arrangements.

Fourth, retirement, absent the pension program, may complicate the interpretation of this study's findings. If women stop working at 60 and men at 65 without the pension program, the effect of retirement is intertwined with the effect of the pension program. Note: the problem is not that the pension may induce individuals to retire and thereby influence child labor and schooling. If, for example, liquidity constraints prevent individuals from retiring and this thereby influences child labor supply, the results of this paper capture an indirect effect of the timing of income on child labor. Rather, an identification problem arises if retirement *absent* the pension occurs at ages 60 for women and 65 for men. There are several reasons why this identification problem may not be substantive in the present discussion. First, most black South Africans are not engaged in formal employment that would have a fixed retirement age (Edmonds, Mammen, and Miller 2005). Second, most black South Africans are sufficiently poor that they do not have the luxury of terminating employment at a specified age (in the absence of the pension). Typically, poor health and other problems associated with aging force the elderly out of employment. Third, one would expect that the changes in the activities of children owing to the retirement of elders would be more important in domestic chores than in work outside of the household or market work more generally. However, much of the observed declines in hours worked are in market work rather than domestic work. Hence, nothing in the data or the present context highlights retirement absent the pension as an important omitted variable.

Fifth, the means-test in the pension benefit formula could create incentives to retire at age 60 for women and age 65 for men. Child labor could then be affected by this price induced retirement as with retirement absent the pension. As discussed in footnote 1 of section 2, the data do not suggest much scope for the means-test to be applied to black South Africans (9 percent of control households report *household* incomes at the level where *individual* incomes could be taxed), nor is there any evidence that the means test binds (96 percent of pension recipients reports incomes at or above the maximum benefit). Moreover, the discussion in the next section suggests that the pension eligibility effects are largest amongst the least educated households who are least apt to be affected by the means-test. Thus, the data do not suggest that means-test induced retirement is likely to be a significant source of bias.

6.2 Interpretation of Timing of Income Effects

There are several reasons why child labor and schooling might respond to the realization of anticipated income. First, the assumption that parents are rational and forward looking in their time allocation decisions might be incorrect. If parents are myopic, then the results of this study would reveal an income effect, not a liquidity constraint. Given the strong evidence that households in developing countries can smooth consumption over predictable seasonal variation in income in Paxson (1993) and Jacoby and Skoufias (1998), it seems reasonable to assume that households are not myopic in their decisions with respect to a doubling of income like the Old Age Pension. Moreover, if households were myopic, it becomes difficult to understand the evidence that suggests schooling decisions are sensitive to anticipated returns to education such as Foster and Rosenzweig (1996).

Second, the unconstrained HC model has a single household decision-maker. It is possible that in a more sophisticated model of household decision-making, peer pressure to share income among family members may increase when pension cash is on-hand. However, if individuals are not liquidity constrained, their personal standard of living and bargaining power in the household should not depend on cash on hand. It seems plausible to assume that peer pressure to share income and wealth is driven by an individual's relative living standards rather than her cash on hand. Thus, the peer pressure story works best with other market imperfections that are inconsistent with the unconstrained HC model.

Third, in the context of consumption responses to anticipatable social security tax changes in the U.S., Parker (1999) argues that a boundedly rational household may allow spending to track income provided that this strategy does not make the household too much worse off than the fully rational strategy. This behavior occurs, because it is costly for the individual to re-optimize her consumption plan with each change in income. Given the magnitude of the cash transfer stemming from the old-age pension, it is hard to imagine that this nearly rational rule of tracking child labor and schooling to income would not make the household much worse off.

Fourth, liquidity constraints on the parent's ability to transfer resources between time t and time $t+1$ can cause the timing of income to affect child labor and schooling decisions because a constrained household's investment decisions that yield a return in time $t+1$ depend on the marginal utility of consumption in time t. Three pieces of evidence suggest a scope for liquidity constraints to be the correct interpretation of the findings in this study. First, the changes in child labor and schooling appear to occur largely when one person (the male in most specifications) and the extra incremental change in child labor and schooling associated with a second pension eligible person typically has no additional incremental effect on child labor and schooling beyond having one man eligible. Second, the magnitudes of the results for male pension eligibility (absent a female pensioner) are larger than found for female pension eligibility (absent a male pensioner). Men may be more likely to be liquidity constrained overall for several reasons discussed in section 5.4, not the least of which is the elevated mortality risk faced by men. This mortality risk should be priced into the interest rates individuals face in borrowing against future income and could lead to prohibitively high interest rates for men. Third, less educated elders may be more likely to be liquidity constrained. If liquidity constraints are in play, the data should suggest larger pension effects on child labor supply and schooling in households where the elders are relatively less educated (assuming that myopia, peer influence, and the propensity to exhibit boundedly rational behaviors do not vary with education). In the context of black South Africa, even the wealthiest 10 percent of the population are still very poor and maybe liquidity constrained. Thus, a failure to find variation with elder education does not necessarily reject liquidity constraints as an explanation.

To explore this, the specification of equation (1) is estimated separately for households where elders have some schooling and do not have some schooling. Just under half of rural children 13-17 live with elders who have never completed any schooling. The basic results of this study are reported for children living with uneducated elders in the first three columns of Table 8. For this group, schooling attendance increases slightly more and hours worked declines more than observed in the general population. In contrast, for the group living with elders with some education, the improvement in schooling is smaller and the decline in hours worked is smaller and statistically different. Thus, the declines in child labor and increases in schooling are largest in households that are most apt to be liquidity constrained.

There are many reasons why schooling and child labor may respond to the realization of anticipated income if households are liquidity constrained. One possibility is that the receipt of anticipated income lowers the marginal utility gained from the child's economic contribution. A second option is that the pension enables families to spend more on education. Case and Deaton (1998) have documented an increase in education spending associated with pension income, and Case and Ardington (2005) hypothesizes that an inability to pay school fees and other expenses is critical for understanding why South African orphans have lower schooling enrollment. A third option is that income affects improvements in child nutrition as in Duflo (2004). Improvements in nutrition might raise the child's relative productivity in schooling if schooling rewards nutrition more than work. Two other possible explanations have already been discussed. Changes in household composition may follow the realization of income (Edmonds, Mammen, and Miller 2005) and influence child time allocation as in Edmonds (2005), but based on tables 2 and 7, the data do not support an important role for shifts in household composition in explaining the findings. Spillovers in child time allocation from changes in adult or elder labor supply might also be important. For example, if adult labor supply declines as in Bertrand, Mullainathan, and Miller (2003), the presence of the adult in the household might lead to a relative decline in child productivity in domestic chores, freeing time for school. Since the changes in hours worked for boys are largely in market work, spillovers from changes in adult labor supply that affect the value of child time in domestic chores are unlikely to be behind the changes in schooling and hours worked observed herein (although a similar story for market work cannot be excluded).

It is not possible to separately identify changes in schooling and hours worked that owe to declining marginal utility of the child's contribution, direct schooling costs, or changes in child productivity. However, the data can provide some insight. Phase 2 of the SAYP asks children why they are not attending school. Thus, in the selected sample of children in households with at least one working child (who are eligible for the phase 2 questionnaire), the data can provide a sense of what children perceive as the reason for improving schooling with male pension eligibility. A natural concern in using this sample is that selection into phase 2 might be correlated with pension eligibility although this does not appear to be the case (bottom row of Table 2). The relevant question asks why children did not attend school. Answers include schooling affordability, illness, work, and "other" reason. Table 9 contains estimates of (1) with a dependent variables that is an indictor that is 1 if the child reports not attending school for a given (column heading) reason and 0 if the child did not attend for a different reason or if the child attended school. Each column label in table 9 indicates the reason for not attending school. Descriptive statistics for each dependent variable are at the bottom of the table.

With male pension eligibility, affordability and "other" reasons decline significantly for boys as explanations for not attending school. In the entire sample, only 4 boys and 2 girls report not attending school because of a need to work. Hence, it is unsurprising that the data do not suggest an effect of male pension eligibility on this reason for not attending school. Illness is also relatively rare, so there is no conclusive evidence about the role in schooling decisions played by the value of the child's economic contribution or changes in health status / productivity. That said, these findings are consistent with Case and Ardington's (2005) findings among orphans that schooling expenses influence enrollment. Caution is merited in interpreting these results. When a child replies that he does not attend school because his family cannot afford it, it is tempting to assume that this is a statement about the direct costs of schooling, but it could just as easily reflect opportunity costs (e.g. the marginal utility of the child's economic contribution). However, there is at least suggestive evidence that the timing of income effects on schooling and hours in work stem from liquidity constraints that keep families from affording direct and indirect schooling costs.

7. Conclusion

This study finds that anticipated large cash transfer to the elderly in South Africa appear to be associated with increases in schooling and declines in hours worked. The average rural South African child living with an elder that is not yet pension eligible spends almost 3 hours per day working. In the data, pension income to an elder male is associated with over an hour less work per day. These declines in hours worked occur simultaneously with increases in school attendance (to nearly 100 percent for rural boys). In turn, these declines in time spent working and increases in school attendance are also associated with increasing schooling attainment and primary school completion rates, especially for boys, in the length of time that the child has lived with a pension eligible male. These changes in hours worked and schooling with male pension eligibility lead to levels of work and schooling that are similar to what the data report for nearly eligible elder women. Hence, the results herein would follow from a model where men are credit constrained to a greater degree than are women. There is some suggestive evidence that these credit constraints influence schooling because of an inability to afford schooling.

The main focus of this study has been to test whether schooling and child labor supply decisions concord with the predictions of an unconstrained human capital investment model. The data in this study seem consistent with an important role for liquidity constraints in these child time allocation decisions. Baland and Robinson (2000) show that if a family faces liquidity constraints, then child labor is inefficiently high (from the family's perspective), because child labor supply is determined by the marginal utility of consumption rather than the relative return to educational investments. The results of this paper are consistent with a story where an inability to borrow against future income forces households to under-invest in education. Receiving large cash transfers weaken this cash constraint, and hence children work less and attend school more.

The policy implications of the finding that the activities of children are affected by the timing of income are substantial. A great deal of the literature on the determinants of child labor, going back as far as Marx, argues that parents send their children to work, because the return to working is greater than the return to not-working. This view then suggests that, if ending child labor is a policy priority, countries should focus on interventions directly related to child labor. However, the results of this study are consistent with the view that, for some households, work stems more from market imperfections than from high market returns to child labor or low returns to education. Hence, in these households, influencing the activities of children may best be accomplished by attacking poverty, making schooling more affordable, developing markets, and building financial intermediaries. With substantive liquidity constraints, actions design to prohibit child labor directly (such as trade sanctions) may impose costs on the poorest parents rather than the most callous and could, in the end, serve to exacerbate child labor and worsen schooling (Ranjan 2001).

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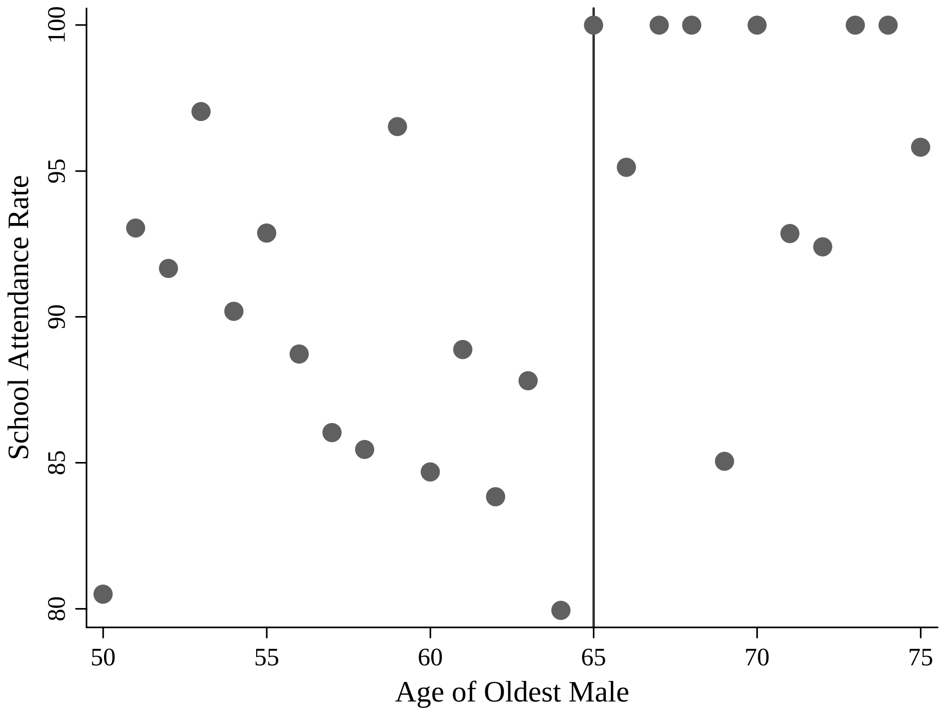
Figure 1: Schooling Attendance Rates and Age of the Oldest Man in the Household

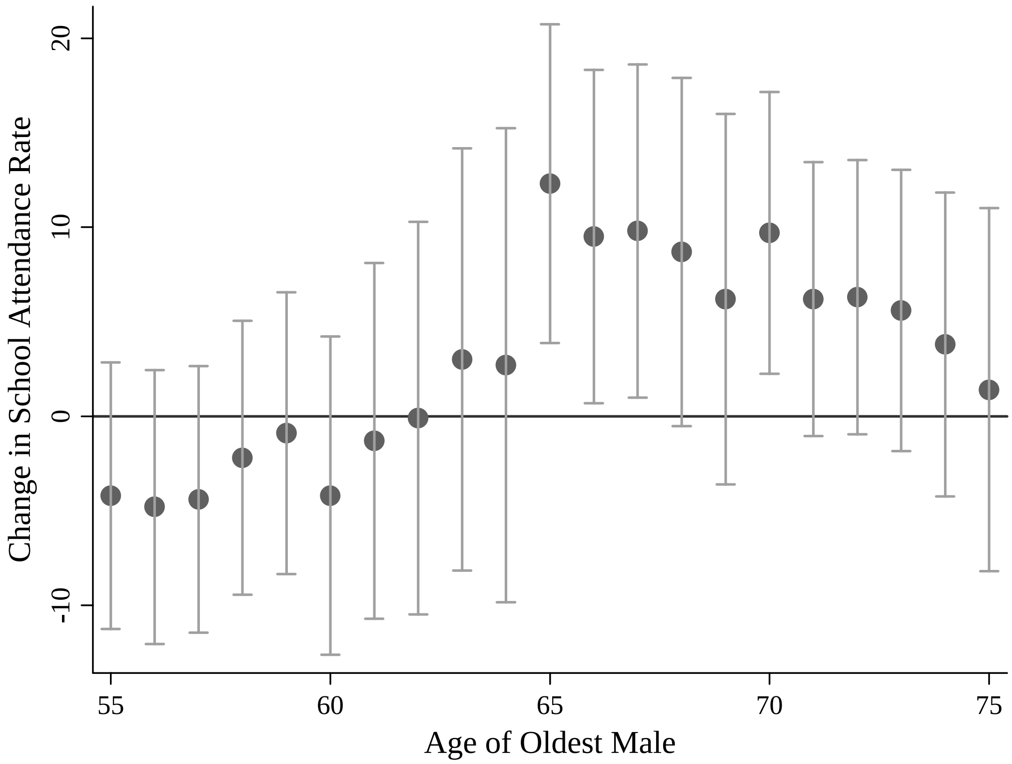


Sample means by age of the oldest male. Limited to rural children 13-17

Figure 2: Market Work while not also Attending School Participation Rates and Age of the Oldest Man in the Household

Sample means by age of the oldest male. Limited to rural children 13-17.

Figure 3: Changes in Schooling Attendance with Alternative Age Discontinuities



Equation (1) with an indicator of school attendance as the dependent variable is replicated pretending that male pension eligibility occurs at indicated ages. Point estimates and 95 percent confidence intervals are pictured for the coefficient on the indicator that a person in the household is pension eligible. Limited to rural children 13-17.

Table 1: Summary Statistics, African Children in African Headed Households

	Full Sample		Person of Age 50-75		Rural Households		Rural Households Eldest w/i - 2 yrs of PE		Rural Households Eldest w/i + 2 yrs of PE	
	1		2		3		4		5	
	mean	s.e.	mean	s.e.	mean	s.e.	mean	s.e.	mean	s.e.
Full Sample, 5-17										
N	15,485		6,363		3,708		404		275	
Age	10.99	(0.04)	11.09	(0.06)	11.02	(0.08)	10.81	(0.25)	11.10	(0.35)
Female	0.50	(0.00)	0.49	(0.01)	0.49	(0.01)	0.47	(0.03)	0.49	(0.03)
Attends School	0.88	(0.00)	0.89	(0.01)	0.89	(0.01)	0.85	(0.02)	0.90	(0.02)
Works	0.85	(0.01)	0.85	(0.01)	0.88	(0.01)	0.91	(0.02)	0.91	(0.02)
Market Work	0.29	(0.01)	0.31	(0.01)	0.37	(0.02)	0.41	(0.05)	0.39	(0.05)
Market Work without School	0.02	(0.00)	0.02	(0.00)	0.02	(0.00)	0.05	(0.02)	0.02	(0.01)
Hours Worked per Day	1.44	(0.06)	1.55	(0.08)	1.88	(0.11)	2.30	(0.39)	1.86	(0.31)
Hours in Market Work	0.38	(0.02)	0.42	(0.03)	0.54	(0.05)	0.77	(0.17)	0.70	(0.18)
Full Time Market Work	0.01	(0.00)	0.01	(0.00)	0.02	(0.00)	0.04	(0.02)	0.01	(0.01)
Educational Attainment	4.52	(0.04)	4.52	(0.06)	4.35	(0.08)	4.08	(0.21)	4.53	(0.31)
Completed Primary School	0.21	(0.01)	0.20	(0.01)	0.18	(0.01)	0.16	(0.03)	0.22	(0.04)
Ages 13-17										
N	5,705		2,427		1,387		148		112	
Age	14.96	(0.02)	14.96	(0.03)	14.93	(0.04)	14.82	(0.12)	14.94	(0.17)
Female	0.50	(0.01)	0.49	(0.01)	0.49	(0.01)	0.50	(0.05)	0.52	(0.05)
Attends School	0.93	(0.00)	0.93	(0.01)	0.92	(0.01)	0.85	(0.04)	0.93	(0.02)
Works	0.95	(0.00)	0.95	(0.01)	0.96	(0.01)	0.98	(0.01)	0.97	(0.01)
Market Work	0.37	(0.01)	0.39	(0.02)	0.47	(0.02)	0.51	(0.06)	0.47	(0.06)
Market Work without School	0.03	(0.00)	0.03	(0.00)	0.03	(0.01)	0.06	(0.03)	0.02	(0.01)
Hours Worked per Day	1.96	(0.08)	2.06	(0.12)	2.49	(0.16)	2.95	(0.49)	2.05	(0.30)
Hours in Market Work	0.59	(0.04)	0.64	(0.05)	0.82	(0.07)	0.99	(0.25)	0.81	(0.19)
Full Time Market Work	0.03	(0.00)	0.03	(0.01)	0.04	(0.01)	0.06	(0.03)	0.01	(0.01)
Educational Attainment	7.44	(0.05)	7.29	(0.07)	7.03	(0.10)	6.82	(0.28)	7.39	(0.25)
Completed Primary School	0.52	(0.01)	0.49	(0.02)	0.45	(0.02)	0.41	(0.05)	0.51	(0.06)

Weighted to be nationally representative for indicated population. Standard errors corrected for clustered and stratified sample design. Columns 3-5 are rural households only.

Table 2: Validity Tests

	Descriptive Statistics		Regression Results					
			Pension Eligible (PE) in Household		PE is Female		Male and Female PE in HH	
	Mean	S.E.	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.
	1		2		3		4	
Family Attributes								
Reports Pension Income	0.485	(0.020)	0.461	(0.080)**	0.006	(0.092)	0.110	(0.076)
Log Income^	7.736	(0.120)	1.070	(0.429)**	-0.439	(0.465)	0.519	(0.392)
Household Size	7.971	(0.191)	-0.496	(0.855)	2.031	(1.593)	-1.868	(1.748)
# children age <=17	4.401	(0.137)	-0.335	(0.609)	1.564	(1.104)	-1.330	(1.218)
Fraction Children Female	0.493	(0.009)	0.013	(0.050)	0.053	(0.059)	0.066	(0.060)
Elders with no Education	2.905	(0.162)	-0.709	(0.714)	-0.855	(0.833)	0.226	(0.919)
Location								
Urban^^	0.335	(0.024)	0.054	(0.064)	-0.067	(0.075)	0.115	(0.081)
Province:								
Western Cape	0.003	(0.002)	-0.002	(0.005)	0.008	(0.007)	-0.004	(0.003)
Eastern Cape	0.196	(0.031)	-0.011	(0.080)	-0.097	(0.091)	-0.099	(0.074)
Northern Cape	0.011	(0.005)	-0.006	(0.009)	0.017	(0.012)	0.000	(0.015)
Free State	0.076	(0.017)	-0.015	(0.044)	0.042	(0.052)	-0.064	(0.052)
KwaZulu-Natal	0.228	(0.037)	-0.047	(0.075)	0.125	(0.095)	0.121	(0.117)
North West	0.082	(0.018)	-0.010	(0.034)	-0.032	(0.039)	0.074	(0.054)
Gauteng	0.020	(0.007)	-0.022	(0.018)	0.011	(0.016)	-0.006	(0.019)
Mpumalanga	0.161	(0.030)	0.001	(0.080)	0.046	(0.084)	-0.053	(0.073)
Northern Province	0.223	(0.035)	0.112	(0.090)	-0.120	(0.110)	0.032	(0.094)
Child Attributes								
Gender	0.492	(0.009)	0.047	(0.054)	0.004	(0.062)	0.068	(0.067)
Age	11.019	(0.075)	-0.139	(0.407)	-0.289	(0.508)	0.500	(0.556)
Not Moved within Last 2 Yrs	0.963	(0.006)	0.031	(0.032)	-0.008	(0.032)	0.025	(0.027)
Selected into Phase 2	0.826	(0.014)	0.012	(0.055)	0.027	(0.062)	-0.003	(0.065)

Sample Size: 3,708 except for ^ where income information is missing for 28 households and ^^ which also includes urban households (6,363). All regressions also include third order polynomial in age of oldest man and age of oldest woman, a dummy that the household includes an elder male, and a dummy that the household includes an elder male and female. Standard errors corrected for clustering on age of oldest man / woman cell. OLS results. ** Significant at 5 percent. * significant at 10 percent.

Table 3: School Attendance and Pension Eligibility

Rural Children Only

Gender	Both	Both	Both	Male	Female
Age Range	5-17	5-12	13-17	13-17	13-17
	1	2	3	4	5
Elder Male	-0.051	-0.035	-0.070	-0.068	-0.051
	(0.038)	(0.042)	(0.056)	(0.062)	(0.070)
Elder Male and Female	-0.001	-0.013	0.008	0.046	-0.011
	(0.029)	(0.030)	(0.044)	(0.048)	(0.069)
Pension Eligible (PE) in HH	0.045	-0.028	0.123	0.175	0.078
	(0.039)	(0.053)	(0.043)**	(0.047)**	(0.059)
PE is Female (PEF)	-0.055	-0.006	-0.100	-0.149	-0.051
	(0.042)	(0.057)	(0.045)**	(0.057)**	(0.069)
Male and Female PE in HH (PEMF)	0.060	0.019	0.095	0.116	0.097
	(0.039)	(0.051)	(0.045)**	(0.058)**	(0.067)
Observations	3708	2321	1387	710	677
Adj. R2	0.210	0.290	0.030	0.010	0.050
Elder Male + PE in HH	-0.006	-0.063	0.053	0.107	0.027
	(0.057)	(0.068)	(0.035)	(0.057)	(0.073)
PE in HH + PEF	-0.010	-0.034	0.023	0.026	0.027
	(0.026)	(0.028)	(0.051)	(0.087)	(0.102)
PE in HH + PEF+PEMF	0.049	-0.015	0.118	0.141	0.124
	(0.042)	(0.052)	(0.078)	(0.045)**	(0.052)**

* significant at 10%; ** significant at 5%. All regressions also include female indicator, age, age squared, age cubed, and a third order polynomial in the age of the oldest man and age of oldest woman in the household. Standard errors in parenthesis are corrected for age of the oldest man / age of the oldest woman cell clustering. All regressions weighted to correct for sample design. "Elder Male + PE in HH" adds the coefficient on the elder male dummy to the pension eligible individual in household dummy in order to test whether schooling with male pension eligibility differs from schooling with a nearly eligible female. "PE in HH + PEF" adds the coefficient from pension eligible individual to that from pension eligible female to compute how schooling changes with female eligibility. "PE in HH + PEF + PEMF" adds all three pension eligibility coefficients to compute how schooling changes in moving from no pension eligible elder to both a male and female eligible elder.

Table 4: Work Status and Pension Eligibility
Sample limited to rural children 13-17

Dependent Variable: Participation in . . .	Any Work			Any Market Work		
Gender	Both	Male	Female	Both	Male	Female
	1	2	3	4	5	6
Elder Male	0.011	0.040	-0.016	0.113	0.055	0.145
	(0.037)	(0.068)	(0.025)	(0.099)	(0.134)	(0.131)
Elder Male and Female	-0.026	-0.055	-0.002	-0.098	-0.071	-0.122
	(0.031)	(0.055)	(0.021)	(0.076)	(0.098)	(0.120)
Pension Eligible (PE) in HH	-0.035	-0.056	-0.034	-0.108	-0.003	-0.191
	(0.033)	(0.072)	(0.023)	(0.097)	(0.136)	(0.116)
PE is Female (PEF)	0.017	0.032	0.022	0.000	-0.098	0.138
	(0.033)	(0.067)	(0.029)	(0.116)	(0.154)	(0.149)
Male and Female PE in HH (PEMF)	-0.030	-0.023	-0.059	0.023	0.086	-0.023
	(0.037)	(0.051)	(0.043)	(0.118)	(0.143)	(0.157)
Observations	1387	710	677	1387	710	677
Adj. R2	0.020	0.040	-0.010	0.020	0.000	0.020
Elder Male + PE in HH	-0.024	-0.016	-0.051	0.005	0.052	-0.046
	(0.047)	(0.053)	(0.032)	(0.152)	(0.094)	(0.191)
PE in HH + PEF	-0.018	-0.024	-0.013	-0.108	-0.101	-0.053
	(0.027)	(0.096)	(0.046)	(0.078)	(0.139)	(0.154)
PE in HH + PEF+PEMF	-0.048	-0.048	-0.072	-0.085	-0.014	-0.076
	(0.035)	(0.047)	(0.023)**	(0.114)	(0.205)	(0.109)

* significant at 10%; ** significant at 5%. See notes to table 3.

Table 5: Daily Hours Worked and Pension Eligibility

Limited to rural children 13-17

	Avg. Total Hours per Day			Avg. Market Hours per Day			Full Time in Market Work		
Gender	Both	Male	Female	Both	Male	Female	Both	Male	Female
	1	2	3	4	5	6	7	8	9
Elder Male	0.565	0.204	0.796	0.508	0.338	0.530	0.106	0.122	0.074
	(0.567)	(0.725)	(0.844)	(0.379)	(0.444)	(0.575)	(0.045)**	(0.044)**	(0.082)
Elder Male and Female	-0.569	-0.748	-0.300	-0.484	-0.554	-0.282	-0.033	-0.031	-0.022
	(0.507)	(0.556)	(0.797)	(0.284)*	(0.265)**	(0.594)	(0.042)	(0.021)	(0.101)
Pension Eligible (PE) in HH	-1.025	-0.920	-1.249	-0.630	-0.689	-0.562	-0.074	-0.089	-0.059
	(0.462)**	(0.531)*	(0.721)*	(0.313)**	(0.361)*	(0.449)	(0.035)**	(0.030)**	(0.057)
PE is Female (PEF)	0.444	-0.024	1.196	0.136	-0.057	0.543	0.030	0.065	-0.002
	(0.648)	(0.743)	(0.993)	(0.381)	(0.459)	(0.525)	(0.039)	(0.034)*	(0.065)
Male and Female PE in HH (PEMF)	0.104	0.475	-0.414	0.261	0.733	-0.160	-0.006	0.010	-0.006
	(0.696)	(0.771)	(1.016)	(0.421)	(0.508)	(0.578)	(0.038)	(0.045)	(0.051)
Observations	1387	710	677	1387	710	677	1387	710	677
Adj. R2	0.130	0.230	0.000	0.070	0.110	0.030	0.130	0.230	0.040
Elder Male + PE in HH	-0.460	-0.716	-0.453	-0.122	-0.351	-0.032	0.032	0.033	0.015
	(0.771)	(0.942)	(0.910)	(0.394)	(0.300)	(0.742)	(0.036)	(0.02)*	(0.105)
PE in HH + PEF	-0.582	-0.944	-0.053	-0.494	-0.747	-0.018	-0.044	-0.024	-0.061
	(0.476)	(0.573)	(1.228)	(0.247)**	(0.505)	(0.300)	(0.051)	(0.046)	(0.032)*
PE in HH + PEF+PEMF	-0.477	-0.469	-0.467	-0.233	-0.014	-0.179	-0.049	-0.015	-0.067
	(0.616)	(0.764)	(0.690)	(0.454)	(0.588)	(0.521)	(0.019)**	(0.051)	(0.045)

* significant at 10%; ** significant at 5%. See notes to table 3. "Full Time in Market Work" is an indicator that the child works at least 40 hours per week in market work.

Table 6: Length of Pension Exposure and Schooling Attainment
Rural children only

Dep. Variable	Years of Completed Schooling					Primary School Completed		
Age Range	5-17	5-17	5-17	5-12	13-17	13-17	13-17	13-17
	Both	Male	Female	Male	Male	Both	Male	Female
	1	2	3	4	5	6	7	8
School Years with an Elder	-0.046	-0.060	-0.051	-0.039	-0.036	-0.005	-0.016	-0.006
	(0.033)	(0.043)	(0.044)	(0.063)	(0.085)	(0.010)	(0.015)	(0.012)
Sch. Years with a Female Elder	0.030	0.030	0.040	0.099	-0.040	-0.004	0.003	-0.005
	(0.031)	(0.043)	(0.043)	(0.057)*	(0.089)	(0.012)	(0.018)	(0.017)
Sch. Years with Two Elder	0.015	0.017	0.008	-0.027	0.116	0.014	0.011	0.007
	(0.031)	(0.037)	(0.042)	(0.045)	(0.083)	(0.011)	(0.016)	(0.015)
Sch. Yrs with Pension Eligible (PE) Person	0.095	0.176	0.003	0.221	0.188	0.029	0.033	0.009
	(0.056)*	(0.064)**	(0.089)	(0.064)**	(0.109)*	(0.016)*	(0.024)	(0.021)
Sch. Years PE Female	-0.089	-0.176	0.011	-0.203	-0.212	-0.013	-0.020	0.006
	(0.062)	(0.073)**	(0.097)	(0.073)**	(0.135)	(0.020)	(0.028)	(0.028)
Sch. Years PE Male & Female	-0.038	-0.076	-0.041	-0.145	0.104	0.006	0.009	-0.023
	(0.066)	(0.090)	(0.078)	(0.081)*	(0.159)	(0.022)	(0.030)	(0.031)
Observations	3688	1862	1826	1153	709	1383	709	674
Adj. R2	0.690	0.680	0.710	0.620	0.120	0.140	0.120	0.130

* significant at 10%; ** significant at 5%. All regressions also include female indicator, age, age squared, age cubed, and a third order polynomial in the age of the oldest man and age of oldest woman in the household. Standard errors in parenthesis are corrected for age of the oldest man / age of the oldest woman cell clustering. All regressions weighted to correct for sample design.

Table 7: Basic Results with Household Composition Controls

Sample restricted to rural children age 13-17

Dep Var.	Currently Attends School 1	Average Hours Worked Per Day 2
Pension Eligible (PE) in HH	0.114	-1.041
	(0.042)**	(0.480)**
PE is Female	-0.106	0.575
	(0.046)**	(0.677)
Male and Female PE in HH	0.116	0.442
	(0.044)**	(0.710)
Observations	1387	1387
Adj. R2	0.060	0.170

* significant at 10%; ** significant at 5%. All regressions also include dummy variables indicating province, household size, residency status of child's mother, father, grandmother, grandfather, number of children 17 and under, and fraction of children female. As before, all regressions include an indicator for presence of elder male, indicator for presence of elders of both sexes, female indicator, age, age squared, age cubed, and a third order polynomial in the age of the oldest man and age of oldest woman in the household. Standard errors in parenthesis are corrected for age of the oldest man / age of the oldest woman cell clustering. All regressions weighted to correct for sample design.

Table 8: Basic Results and the Education of Elders
Sample restricted to rural children aged 13-17

	Elders without any education		Elders with some education	
	Currently Attends School	Average Hours Worked Per Day	Currently Attends School	Average Hours Worked Per Day
Dep Var.	1	2	3	4
Pension Eligible (PE) in HH	0.147	-1.993	0.132	-0.251
	(0.056)**	(0.822)**	0.069*	(0.526)
PE is Female	-0.153	1.843	-0.041	-1.029
	(0.059)**	(1.008)*	(0.066)	(0.769)
Male and Female PE in HH	0.165	-1.149	0.062	1.560
	(0.075)**	(1.006)	(0.057)	(0.930)*
Observations	692	692	695	695
Adj. R2	0.040	0.090	0.040	0.190

* significant at 10%; ** significant at 5%. All regressions also include male elder indicator, male and female elder indicator, female indicator, age, age squared, age cubed, and a third order polynomial in the age of the oldest man and age of oldest woman in the household. Standard errors in parenthesis are corrected for age of the oldest man / age of the oldest woman cell clustering. All regressions weighted to correct for sample design.

Table 9: Reasons for Not Attending School

Limited to children in households with at least 1 working child (interviewed in Phase 2 of the SAYP)

Limited to rural children 13-17

Reason for not attending school	Affordability		Illness		Work		Other	
Gender:	Male	Female	Male	Female	Male	Female	Male	Female
	1	2	3	4	5	6	7	8
Elder Male	0.008	-0.060	-0.001	-0.038	-0.026	0.006	0.066	0.129
	(0.026)	(0.030)**	(0.016)	(0.031)	(0.018)	(0.009)	(0.060)	(0.065)**
Elder Male and Female	-0.014	0.055	0.017	0.079	0.013	-0.010	-0.046	-0.102
	(0.021)	(0.027)**	(0.012)	(0.038)**	(0.011)	(0.007)	(0.044)	(0.058)*
Pension Eligible (PE) in HH	-0.044	-0.007	-0.037	-0.054	-0.027	0.000	-0.054	-0.011
	(0.021)**	(0.049)	(0.026)	(0.038)	(0.021)	(0.005)	(0.027)**	(0.025)
PE is Female (PEF)	0.037	-0.056	0.005	0.075	0.015	0.003	0.082	0.036
	(0.032)	(0.057)	(0.021)	(0.035)**	(0.015)	(0.004)	(0.046)*	(0.038)
Male and Female PE in HH (PEMF)	-0.045	0.009	0.009	-0.040	-0.018	0.004	-0.067	-0.058
	(0.030)	(0.031)	(0.031)	(0.051)	(0.016)	(0.004)	(0.038)*	(0.040)
Observations	498	482	498	482	498	482	498	482
Adj. R2	-0.010	0.000	-0.010	0.020	-0.010	-0.010	0.000	0.040
Elder Male + PE in HH	-0.035	-0.067	-0.039	-0.092	-0.053	0.005	0.012	0.118
	(0.032)	(0.043)	(0.037)	(0.056)	(0.023)**	(0.003)	(0.034)	(0.071)
PE in HH + PEF	-0.007	-0.064	-0.032	0.021	-0.011	0.002	0.028	0.026
	(0.032)	(0.034)*	(0.018)*	(0.024)	(0.009)	(0.006)	(0.074)	(0.041)
PE in HH + PEF+PEMF	-(0.052)	-(0.055)	-(0.024)	-(0.019)	-(0.029)	(0.006)	-(0.038)	-(0.033)
	(0.023)**	(0.065)	(0.039)	(0.066)	(0.036)	(0.013)	(0.023)	(0.041)
Descriptive Statistics								
Phase 2 Sample	0.016	0.022	0.010	0.015	0.007	0.003	0.025	0.046
	(0.005)	(0.009)	(0.005)	(0.006)	(0.004)	(0.002)	(0.007)	(0.010)
Eldest w/i - 2 yrs of PE	0.028	0.032	0.016	0.022	0.006	0.000	0.042	0.042
	(0.014)	(0.015)	(0.010)	(0.012)	(0.006)	(0.000)	(0.017)	(0.017)

* significant at 10%; ** significant at 5%. See notes to table 3. Descriptive statistics are means (standard errors in parenthesis) for the dependent variable indicated by the column heading and the population subgroup indicated by the row.