Cash grants and HIV risk: evidence from South Africa's Old Age Pension[1]

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Abstract

Recent experimental evidence has raised hopes for unconditional cash transfers as a technology for HIV prevention for young women. We assess the impact of a large unconditional cash grant - South Africa's Old Age Pension - on pregnancy and HIV incidence, exploiting the sharp change in household income occurring when an elder attains pension eligibility age. We utilize longitudinal data (2000-2011) on members of over 11,000 households in a rural area of KwaZulu-Natal, with 29% adult HIV prevalence. We find no effect of Old Age Pensions on HIV risk among women in the household. Point estimates for pregnancy are precisely estimated. Our results suggest that the impact of cash grants on HIV risk may be dependent on context or program design.

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Introduction

In spite of the rapid scale up of antiretroviral treatment, HIV incidence remains high in many areas of southern Africa (Bärnighausen, et al. 2009; Zaidi et al. 2013), as policy makers continue to grapple with the challenge of preventing new infections. One promising avenue is the use of unconditional cash transfers to reduce the economic dependence of young women on risky sexual relationships. In a recent World Bank experiment in Malawi, small monthly cash transfers to young women and their households lowered HIV incidence rates by 60%, by reducing the prevalence of risky sexual partnerships, particularly with older men (Baird, 2010; Baird, 2012). Qualitative evidence links poverty and gender inequality as key determinants of HIV infection among young women (Hunter 2010). However, despite the increasing popularity of unconditional cash transfers as a poverty alleviation strategy (Hulme et al. 2010), to our knowledge, no other study has evaluated the effect of unconditional grants on HIV risk. Further evidence is needed before scaling up cash grants as an HIV prevention strategy.

This paper assesses the impact of a large unconditional cash transfer program – South Africa's Old Age Pension – on HIV risk, measured directly through HIV incidence, as well as sexual behaviors and pregnancy as proxies. South Africa's Old Age Pensions provide generous financial support to all women and men over sixty years (the eligibility age for men was recently lowered from 65 years), with assets below a threshold. In an era of declining formal sector employment, these grants have been an important stop-gap measure to maintain livelihoods and to prevent immiseration of entire households (Case & Deaton 1998; Leibbrandt & Levinsohn 2011). There is a robust literature on the effect

of the Old Age Pension on labor supply and household formation, and the grants have been shown to improve nutrition, health, and education outcomes among children in households of pensioners (Ardington, et al. 2007, Duflo 2000, Woolard & Leibbrandt 2010). The Old Age Pension is delivered monthly to 2.5 million elder South Africans, and is equivalent to nearly twice median income. Previous research suggests that grants to female pensioners are shared with other household members (Duflo 2000). No study to date has analyzed the effect of these pensions on HIV incidence or pregnancy rates among household members of pensioners.

The relationship between income and HIV risk has long been debated. Several studies have found positive correlations between socioeconomic status and HIV status (Mishra 2008, Fortson 2011), suggesting that poverty might be protective against HIV, rather than a risk factor. On the other hand, qualitative evidence suggests that poverty may lead to increased transactional sex and to greater risk of infection (Bassett and Mhloyi 1991; Farmer 1993; Hunter 2010). A recent empirical study found that regional income shocks resulting from droughts led to increased rates of HIV infection, with transactional sex the likely pathway (Burke et al. 2011).

One difficulty in estimating the causal effect of income on HIV risk lies in the possibility that unobserved factors (e.g. discount rates, risk aversion, self-esteem) might drive both income and HIV infection risk. The sharp changes in household income occurring when an elder reaches pension age provide a source of identification to estimate causal effects. Previous studies have found substantial recombination of households when an elder receives a pension, with an increase in young mothers with small children, and a decline

in the number of working age women (Edmonds et al. 2005). As a result, women in households with an elder aged 60 cannot be directly compared to women in households with an elder aged 59 in cross-sectional data. To eliminate bias from changing household composition, we use the age of a woman's biological mother and grandmother as an instrument for the number of pension-eligible elders in the woman's household. Using this approach, we allow household recombination to mediate the relationship between a mother's (or grandmother's) pension eligibility and her daughter's (or granddaughter's) HIV risk; but we avoid possible selection bias.

A second difficulty in testing this relationship is that data on HIV incidence are scarce. Cross-sectional data on HIV prevalence have been collected in household surveys, such as the Demographic and Health Surveys. However, it is difficult to assess the effect of income as a time-varying exposure on an outcome that is assessed only at one point in time.[2] We use longitudinal data from a large population-based surveillance system in rural South Africa, with repeated HIV biomarker collection. These data provide a rare opportunity to measure HIV incidence in a general population with very high HIV infection rates – 29% of adults (15-49 years) are HIV-positive (Bärnighausen et al. 2009; Zaidi et al. 2012). Our hypothesis was that pension-income in the household would reduce HIV infection risk for women.

To complement our analysis of HIV incidence, we also assessed the effect of pensions on pregnancy. Unprotected sex with an HIV-positive partner is the most common mode of

[2] One approach would be to define the exposure as the average exposure for an individual across different exposure periods; but this approach runs the risk of an inferential problem akin to *ecological fallacy*: individuals with higher average incomes could be at greatest risk in relatively low-income periods.

HIV acquisition for women in sub-Saharan Africa. Pregnancy is correlated with HIV risk and has been used as a biomarker outcome in other HIV prevention studies (Dupas 2011). Pregnancy has some advantages as an outcome: it is less prone to non-response; it can be approximately dated rather than being defined as occurring within an interval; and it is more common than HIV infection, leading to greater statistical power. Of course, many pregnancies occur in relationships in which both partners are HIV-negative. Additionally, many pregnancies are desired, and the incidence of pregnancy may rise with income as people can better afford to provide for their children (Becker 1960). Since demand for children may vary by age, we estimate our pregnancy models separately for different age groups – women 15-19, 20-24, 25-34, and 35-44. Most pregnancies to adolescent women in KwaZulu-Natal are reported to be unplanned (Manzini 2001), and our hypothesis was that pension income in the household would reduce the incidence of pregnancy among younger women, although we hypothesized that pregnancy rates might increase with household income, among older women of childbearing age.

Data

We utilized twelve years of longitudinal data (2000-2011) on all adult female members of all households in a 432 km^2 surveillance area in rural KwaZulu-Natal. We limited the analysis to person-time of women of reproductive age, 15 – 44 years. The surveillance area is largely rural and is located in one of the poorest districts in South Africa; in 2011, 29% of adults were HIV positive (Zaidi et al. 2012). Data were collected by the Africa Centre for Health and Population Studies (Africa Centre), a research center affiliated with

the University of KwaZulu-Natal, and funded by the Wellcome Trust (www.africacentre.ac.za). Demographic data were collected for all resident *and non-resident* household members, via semiannual household visits. In annual surveys, interview teams also collected data on HIV biomarkers (2003-2011). To rule out selective entry into the cohort due to pension eligibility, we limited the sample to individuals who came under surveillance prior to 1 January 2003. Data were analyzed as a quarterly panel from 1 January 2003 to 31 December 2011 for all women aged 15-44 years at the start of that quarter.

Data on pregnancies were collected prospectively on all resident and non-resident cohort members via routine demographic surveillance, with pregnancy dates and outcomes reported either by the woman herself or by a household proxy respondent. Data are intended to include abortions and miscarriages, though underreporting is likely, as fewer than 3% of reported pregnancies ended in miscarriage or abortion. Approximate pregnancy start dates were identified based on the date when the pregnancy ended and the estimated duration of the pregnancy at that point. We constructed an indicator variable, *PregStart*, which took the value 0 if a person did not become pregnant in that quarter, and the value 1 if a person did become pregnant in that quarter. Observations after the pregnancy start date and up to and including the quarter in which the pregnancy ended were coded as missing because the woman was not eligible to become pregnant in those periods. We also censored *PregStart* one year prior to the end of follow-up (the last date a woman was observed in the demographic surveillance), because most pregnancies that started during this time would not have been observed.

HIV biomarkers were obtained by collecting blood from finger pricks and preparing dry blood spots. HIV status was determined by antibody testing using a broad-based HIV-1/HIV-2 ELISA and confirmatory ELISA (Tanser et al. 2007). Only resident cohort members and a 10% sample of non-resident cohort members were eligible for HIV biomarker collection. Consent rates in the HIV biomarker collection were low, with only about half of respondents providing biomarkers. HIV data are snapshot data, i.e. at any given HIV test date, we learn whether a respondent is HIV-positive, but not when she seroconverted. We imputed quarter of seroconversion by drawing from a uniform distribution between date last observed to be HIV-negative and date first observed to be HIV-positive. We constructed an indicator variable, *HIVinfected*, that took the value 0 if a person did not contract HIV in that quarter and 1 if a person did contract HIV in that quarter. Observations prior to first negative HIV test and observations after date of serocoversion (for converters) or last negative HIV test (for non-converters) were censored, because people were not eligible to be observed as a seroconverter if they were not under observation or if they had already seroconverted. Although there is certainly measurement error in pregnancy start dates and imputed dates of HIV seroconversion, we expect this measurement error to be random, leading to reduced precision, but no bias in our point estimates.

Our exposure of interest was the number of older female household members (aged 45-74) who were eligible for an Old Age Pension in a given quarter. For sake of clarity, we refer to these older female household members who were within +/- 15 years of pension eligibility as “household elders”, to distinguish them from the “index women” (aged 15-44) whose reproductive health outcomes we are interested in. Quarterly data on index

woman were merged with information on their household memberships; these data were then merged with data on all women 45-74 years who were members of those households in that quarter. We defined *NumPenEligHH* as the number of household elders (ages 45-74) who were of pension eligibility age in that quarter.

To assess take-up of the grant conditional on eligibility, we also merged data on eligibility of household elders with snapshot data on reported pension receipt, collected through annual household socioeconomic (HSE) surveys (2001-2011). Pension receipt was only observed in about 12% of quarterly observations; in contrast age-eligibility for the pension was always observed. As we describe below, take-up is very high in this community – nearly 80% of age-eligible female elders do receive a pension – and so we use eligibility in place of actual receipt to increase precision in our analysis.

We used age of biological mother and age of biological maternal grandmother as instruments for household membership with an elder of pension age. Data on biological parent-child relationships were collected in the Africa Centre's population surveillance. By merging twice on the identity of a respondent's mother and retaining dates of birth, we constructed exposure variables for age of an index woman's mother and maternal grandmother at the start of each quarter during which the index person was under observation. (We excluded fathers and grandfathers because many fewer paternal linkages existed in the data). Not all women could be linked to mothers and/or grandmothers, as some mothers and grandmothers had already died and others were not members of a household under surveillance between 1 January 2000 and 1 January 2003. Further, we were only able to link index women to biological grandmothers if the mother

was alive and under surveillance in 2000-2002. These restrictions imply that our estimated treatment effects may not be generalizable to women without maternal matches in the surveillance. We constructed indicator variables for whether data on mother's age or maternal grandmother's age were missing, and controlled for these in our regressions.

The Africa Centre obtained ethics approval for data collection, linkage, and analysis, from the Biomedical Research and Ethic Committee of the University of KwaZulu-Natal, in agreement with the Research Office of the KwaZulu-Natal Department of Health.

Descriptive statistics for the sample are presented in Table 1. The final sample included data on 21,670 women, contributing 523,745 person-quarter observations. We observed 451,829 non-missing person-quarter observations of *PregStart* and 60,682 non-missing person-quarter observations of *HIVinfect*. There were 14,139 incident pregnancies and 433 incident HIV infections observed during follow-up.

First Stage: Age of Biological Mother and Pension-Eligible Elders in the Household

There is substantial recombination of households when South African elders reach pension eligibility age (Edmonds et al. 2005). If household composition is endogenous to pension eligibility, then we cannot simply compare reproductive health outcomes of women in households with elders who have attained pension eligibility with women in households with elders who have not yet attained eligibility age, since these women may not be comparable on important but unobserved characteristics (such as desired fertility).

To solve this identification problem, we used data on age of biological mother and maternal grandmother as instruments for the number of pension-eligible elders in the household. These variables were defined prior to any elder attaining pension eligibility – date of birth cannot be manipulated – and they allowed us to assess outcomes for women regardless of household membership choice.

In the first stage of our instrumental variables framework, we estimated how the probability of being in a household with a pension-eligible elder changed with biological mother's (and grandmother's) age. We estimated regression-discontinuity models, of the form:

(1)
$$g\left(NumPenEligHH_{iq}\right) =$$

$$\beta_1 \, f\left(MAge_{iq}\right) + \beta_2 \, f\left(MAge_{iq}\right) 1\left[MAge_{iq} > 60y\right]$$

$$+\gamma_{FS} 1\left[MAge_{iq} > 60y\right] + \delta \, 1\left[MAge_{iq} = .\right]$$

where *i* and *q* denote subscripts for woman and quarter respectively. $MAge_{iq}$ denotes mother's age, *f(.)* is a continuous function and *1[.]* are indicator functions. For clarity, we have suppressed grandmother's age in equation (1). The outcome, $NumPenEligHH_{iq}$, is the number of elders in the household who are age-eligible for the pension.

Figure 1 displays the distributions of biological mother's and maternal grandmother's age in our data. There was no observable discontinuity in the density in either of these

variables at age 60. This is to be expected, since it would be very hard to manipulate dates of birth decades earlier.

First stage regression results describing the relationship between age of mother and/or grandmother and number of pension-eligible household members are presented in Table 2. In estimating regression-discontinuity models, we specified $f(MAge_{iq})$ as a linear function, and we varied the bandwidth – limiting the range in the instrument to those women with mothers and/or grandmothers within 10 years or 5 years of the threshold. As shown in Table 2, pension-eligibility for a biological mother increased the number of pension-eligible elders in the household of the index person by about 0.85 persons. When a maternal grandmother reached pension age, this increased the number of pension-eligible elders in the household by about 0.69 persons. These estimates are highly statistically significant.

Although we have some data on actual pension receipt, it is snapshot data and so we only have information for a small fraction of the person time that we are interested in. However, there is a very strong relationship between eligibility and pension receipt. Figure 2 shows take-up of the pension as a function of age. (This figure was plotted at the level of the household elder, in contrast to the rest of the analysis, which was conducted at the level of the index woman.) Nearly 80% of pension-eligible elders report that they actually received a pension. Coverage increases about 50% at the 60 year old eligibility threshold, and further increases thereafter, for the next five years.

HIV and pregnancy: reduced form

In reduced form models, we estimated the impact of mother's (and grandmother's) age directly on the incidence of HIV and pregnancy. We estimated models identical to equation 1, but replaced the outcome with $HIVinfect_{iq}$ and $PregStart_{iq}$ – indicators for whether the index woman seroconverted (acquired HIV) or became pregnant during that quarter. Because both HIV infection and pregnancy are relatively rare events, a linear probability model could yield biased results. In most of our specifications, we estimated Poisson regression models, specifying *g(.)* as a log link function; we present linear probability estimates, for comparison. The covariates in these reduced form models were identical to equation (1). That is, each model included separate linear terms on either side of the 60-year threshold, and the covariate of interest was the indicator for whether the index person's mother (or grandmother) was greater than 60 years old.

Figure 3 plots pregnancy rates against age of the index woman's biological mother. There was no evidence of a discontinuity in the incidence of pregnancy when a biological mother turned 60 years. Tables 3 and 4 regression results for pregnancy and HIV incidence. In pregnancy models estimated pooling all age groups (Table 3), we found a precisely estimated null result, with risk ratios all very close to unity. Disaggregating by age revealed some heterogeneous effects (Table 4). For women over 20 years, there was no evidence of any effect. However, among young women, ages 15-19, the risk of pregnancy increased by about 50% ($p < .05$) when their mother reached pension age. No change in pregnancy rates was observed when a grandmother attained pension age.

Figure 4 plots HIV incidence against mother's age. The relationship appears to be non-monotonic, however, there is again no evidence of a discontinuity at age 60. In regression models (Table 5), mother's pension eligibility was associated with increased risk of HIV infection, and grandmother's pension eligibility was associated with decreased risk of HIV infection; however, these results were not statistically significant.

Conclusion

While there has long been interest in the socioeconomic determinants of HIV risk, very few studies have used causal designs to analyze the effect of income on HIV risk. This study used the exogenous assignment of Old Age Pensions to assess the effect of household income on the incidence of HIV infection and pregnancy among younger women. We used the availability of linked data on biological mothers and grandmothers to avoid bias from household recombination. The number of pension eligible elders in the household increased sharply when a mother or grandmother becomes pension eligible; and the probability that an elder actually received a pension rises sharply with eligibility. In contrast to our initial hypotheses, we found no evidence for a decline in the risk of HIV infection or pregnancy when an elder became pension-eligible; in fact, we observed a rise in teen pregnancy when a mother became pension-eligible. These results suggest that the effectiveness of unconditional cash transfer programs to reduce HIV risk may depend substantially on context and program design. For years, public health scholars have called for "structural approaches" to HIV prevention. This study should inform the debate about whether, and if so how unconditional cash transfers should be implemented as a "technology" for HIV prevention.

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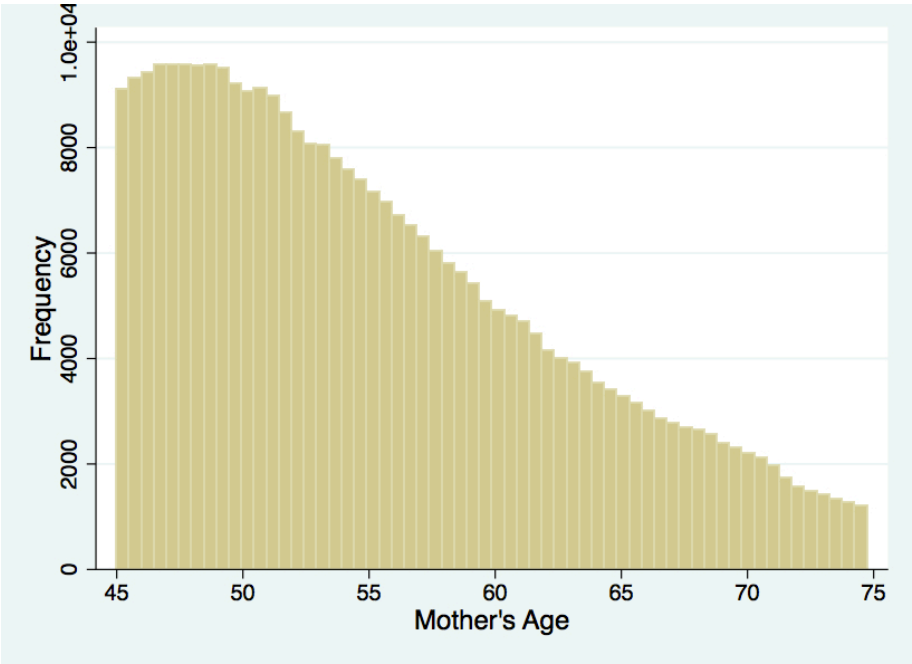
Figures

Figure 1, Distribution of ages of biological mother and grandmother



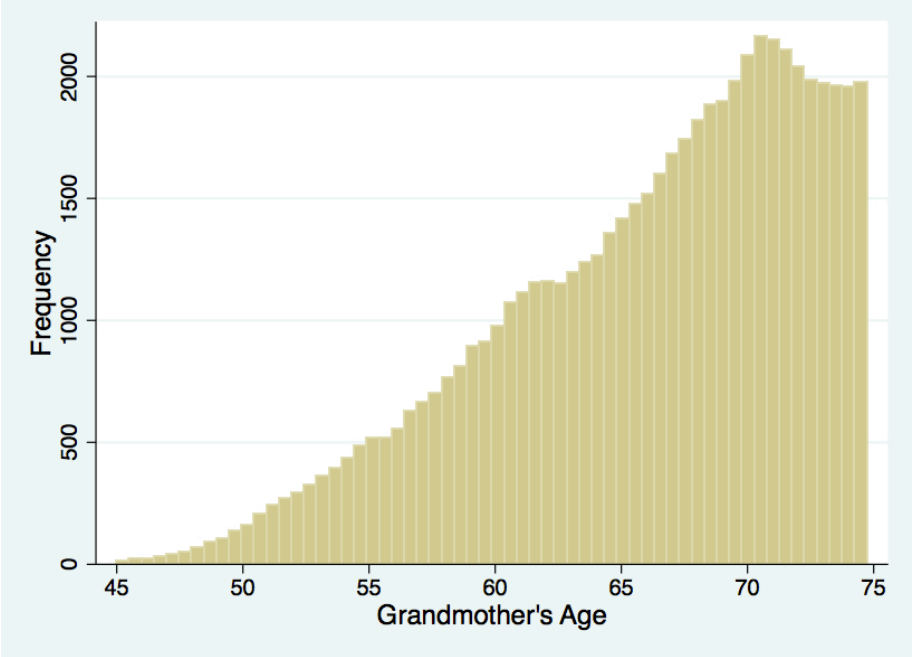
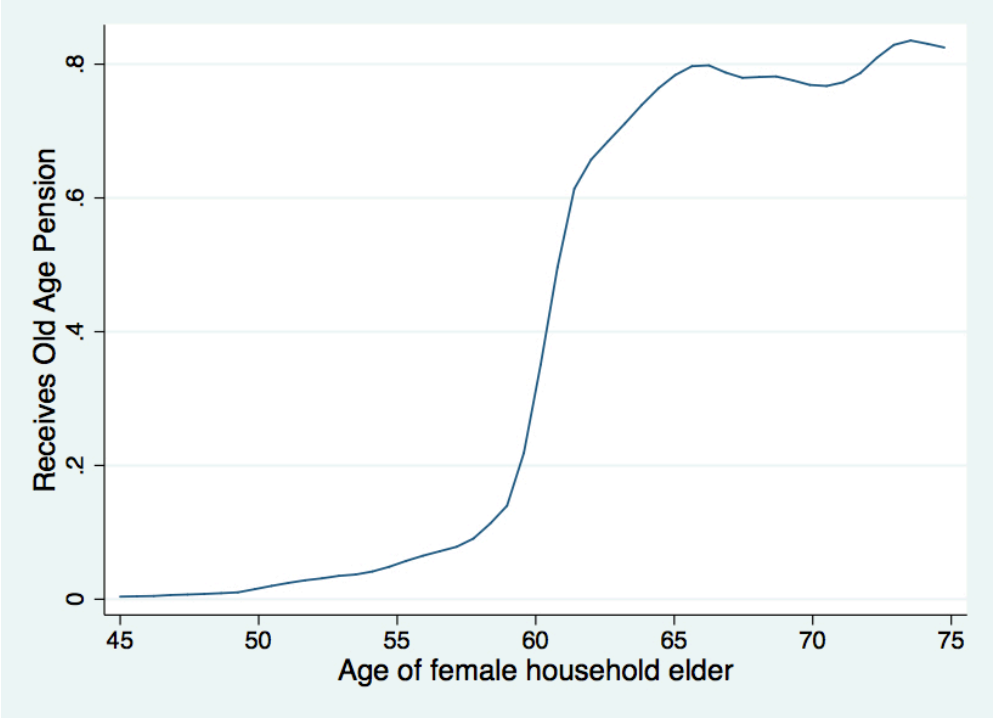


Figure 2, Take-up of the Old Age Pension by age of household elder.



Note: Official age of eligibility is 60 years.

Figure 3, Reduced form: pregnancy incidence and age of biological mother

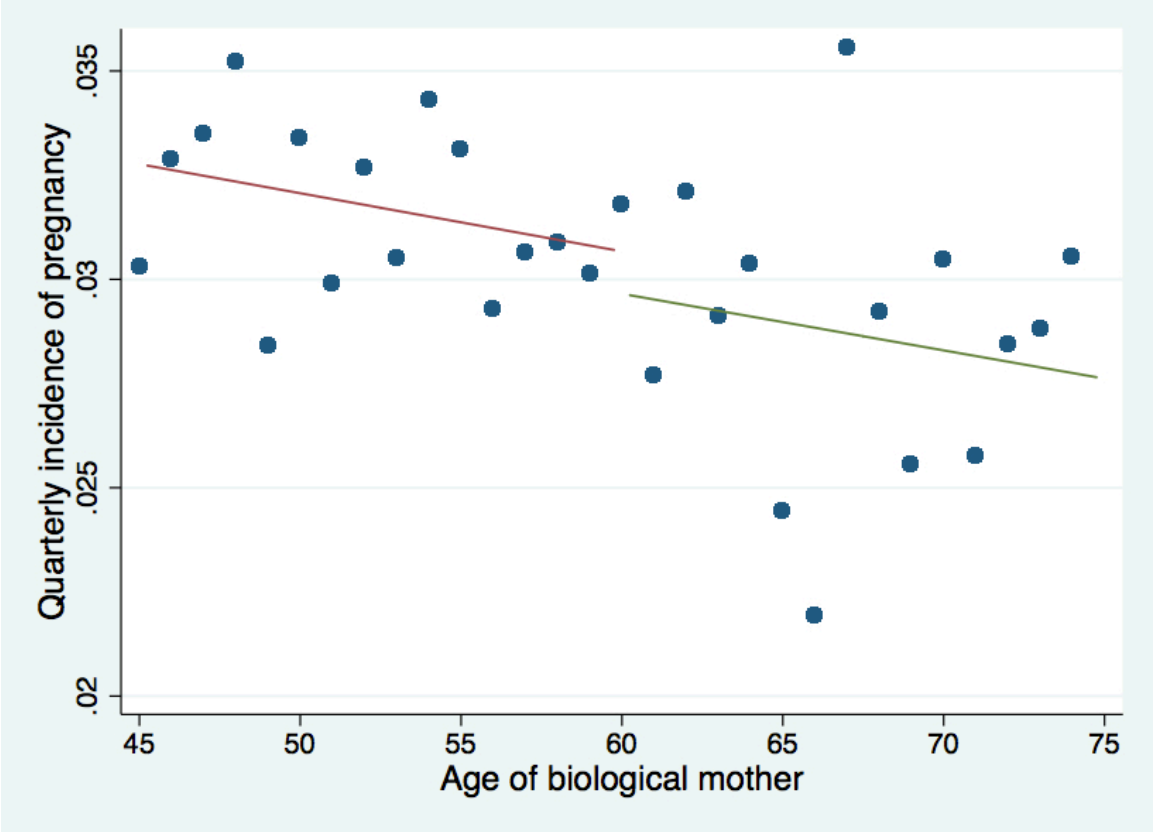
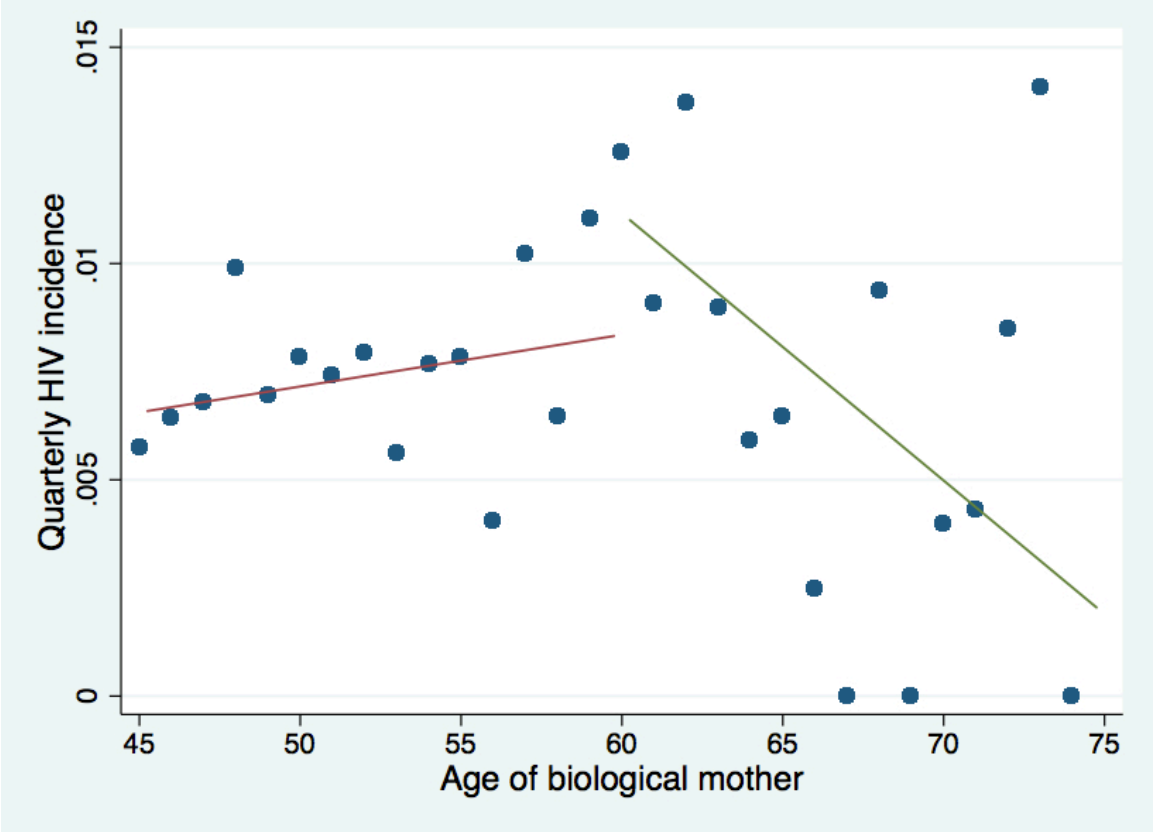


Figure 4, Reduced form: HIV incidence and age of biological mother



Tables

Table 1, Descriptive Statistics

Women, 15-44 years	21,670
Women-quarter observations	523,745
Pregnancy observations (person-qtrs)	451,829
Incident pregnancies	14,139
Pregnancy incidence rate	12 per 100 person-years
HIV observations (person-qtrs)	60,682
Incident HIV infections	433
HIV incidence rate	2.8 per 100 person-years

Table 2, First Stage Regressions

	Age of biological mother and number of OAP-eligible household members			**Age of maternal grandmother and number of OAP-eligible household members**		
Bandwidth	**+/- 15**	**+/- 10**	**+/- 5**	**+/- 15**	**+/- 10**	**+/- 5**
RD Coef.	0.86	0.84	0.83	0.69	0.69	0.69
Std. Error	(0.01)	(0.01)	(0.01)	(0.02)	(0.03)	(0.03)
Individuals	12,428	8980	5319	3111	2281	1339
Observations	279,288	183,128	87,966	52,888	34,342	16,441

Note: All models control for linear terms on either side of the threshold. All coefficients are statistically significant at $p < .01$. Rather than using actual receipt of the OAP, we use eligibility for the OAP as our "treatment". The reason is that OAP data is snapshot data and thus only available for a small fraction of the time that we would like to observe pregnancy outcomes. Figure 1 shows that take-up of the OAP among eligible elders approaches 80%. ** $p < .01$, * $p < .05$

Table 3, Reduced Form Results – Pregnancy, All Ages

	Age of biological mother and incidence of pregnancy			Age of maternal grandmother and incidence of pregnancy		
Bandwidth	**+/- 15**	**+/- 10**	**+/- 5**	**+/- 15**	**+/- 10**	**+/- 5**
Linear probability models (coefs and s.e. multiplied by 100)						
RD Coef.	-0.01	0.01	0.03	-0.27	-0.40	-0.82
Std. Error	(0.14)	(0.18)	(0.25)	(0.34)	(0.40)	(0.57)
Individuals	12,151	8778	5161	3062	2247	1308
Observations	248,461	162,774	78,190	47,997	31,245	15,039
Poisson regression models						
Risk Ratio	0.97	1.01	1.01	0.91	0.86	0.74
z-stat	(0.66)	(0.09)	(0.68)	(0.78)	(1.01)	(1.51)
Individuals	12,151	8778	5161	3062	2247	1308
Observations	248,461	162,774	78,190	47,997	31,245	15,039

Note: All models control for linear terms on either side of the threshold. ** $p < .01$, * $p < .05$

Table 4, Reduced Form Results – Pregnancy, By Age

[Poisson Regression Models]

	Age of biological mother and incidence of pregnancy			**Age of grandmother**
Bandwidth	**+/- 15**	**+/- 10**	**+/- 5**	**+/- 15**
15-19 years				
Risk Ratio	1.54	1.62	1.29	0.85
z-stat	(3.03)**	(2.98)**	(1.10)	(1.02)
Individuals	5819	3268	1377	2743
Observations	60,353	31,309	11,979	30,141
20-24 years				
Risk Ratio	0.94	0.90	0.89	0.94
z-stat	(0.66)	(1.05)	(0.79)	(0.22)
Individuals	6820	4436	2302	1321
Observations	72,422	44,476	20,728	12,974
25-34 years				
Risk Ratio	0.92	0.96	0.91	0.38
z-stat	(1.15)	(0.45)	(0.75)	(1.12)
Individuals	6438	5217	3192	467
Observations	93,060	69,495	35,445	4714
35-44 years				
Risk Ratio	1.00	1.16	1.71	Insufficient
z-stat	(0.01)	(0.68)	(1.91)	Data
Individuals	1779	1438	967	
Observations	22,581	17,494	10,038	

Note: All models control for linear terms on either side of the threshold. ** p < .01, * p < .05

Table 5, Reduced Form Results – HIV incidence, all Ages

[Poisson Regression Models]

	Age of biological mother and HIV incidence			Age of grandmother and HIV incidence
Bandwidth	**+/- 15**	**+/- 10**	**+/- 5**	**+/- 15**
Risk Ratio	1.41	1.73	1.37	0.52
z-stat	(1.21)	(1.74)	(0.77)	(0.96)
Individuals	2378	1547	784	587
Observations	31,666	19,670	9131	5936

Note: All models control for linear terms on either side of the threshold. ** $p < .01$, * $p < .05$