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Abstract

The rapid increase in adult mortality due to the AIDS epidemic in sub-Saharan Africa raises great concern about potential intergenerational effects on children. This article estimates the impact of AIDS-related adult mortality on primary school attendance in rural Kenya using a panel of 1,266 households surveyed in 1997, 2000, and 2002. The paper distinguishes between effects on boys' and girls' education to understand potential gender differences resulting from adult mortality. We also estimate how adult mortality affects child schooling *before* as well as after the death occurs. The paper also estimates the importance of households' initial asset levels in influencing the relationship between adult mortality and child school attendance. We find that all of these distinctions are important when estimating the magnitude of the effects of adult mortality on child school attendance. The probability that girls in initially poor households will remain in school *prior* to the death of a working age adult in the household drops from roughly 88% to 55%. Boys in relatively poor households are less likely than girls to be in school *after* an adult death. By contrast, we find no clear effects on girls' or boys' education among relatively non-poor households, either before or after the timing of adult mortality in the household. We find a strong correlation between working-age adult mortality in our data and lagged HIV-prevalence rates at nearby sentinel survey sites. The evidence indicates that rising AIDS-related adult mortality in rural Kenya is adversely affecting primary school attendance among the poor. However, these results measure only short-term impacts. Over the longer run, whether school attendance in afflicted household rebounds or deteriorates further is unknown.

JEL classification: O12, O15, J10, Q12

Keywords: HIV/AIDS, Education, Kenya

1. Introduction

The onset of the AIDS epidemic in the past two decades raises serious questions about long term human capital development, including education, and the future quality of the workforce in areas where the disease is particularly acute, such as in many countries of Africa. There is growing concern over the effect of high AIDS-related adult mortality and illness on child welfare and the disease's effects on their potential over the long run to support themselves as adults and to contribute to their countries' development (UNICEF 1999; Bell, Devarajan, and Gersbach 2003).

Available empirical evidence indicates that school enrollment rates are lower among AIDS orphans compared with non-orphans (World Bank 1999; Ainsworth, Beegle, and Koda 2002; Case, Paxson, and Ableidinger 2002). A long list of anecdotal case studies also supports this evidence (Gachuhi 1999; Nyambedha, Wandibba, and Aagard-Hansen. 2001; Guest 2001; USAID 2002). However, using 39 nationally representative data sets collected in the 1990s from 28 countries, mostly in Sub-Saharan Africa, Ainsworth and Filmer (2002) show that the difference in enrollment rates between orphaned and non-orphaned children varies greatly across countries and wealth levels within a country. These studies focus on the education status of orphans, regardless of the timing of parental deaths. The impact of parental mortality, however, could depend greatly on the timing of such mortality. More importantly, most children faced with the death of adults in their households do not become orphans, and so the broader question of how children's schooling is affected by the death of adults in their households remains largely unknown.

This paper, therefore, focuses on the impact of working-age adult mortality on child primary school attendance. A major difficulty in measuring the impact of adult mortality, especially mortality attributable to AIDS, is that it is caused by behavioral choices rather than by random events. Individuals and households incurring adult mortality are more likely to display certain characteristics. For example, especially in the early years of the epidemic in sub-Saharan Africa, evidence suggests that men and women with higher education and income were more likely to contract HIV than others because they were more likely to have numerous sexual partners (Ainsworth and Semali 1998; Gregson, Waddell, and Chandiwana 2001).[1] If prime-age mortality remains correlated with individual and household characteristics such as social status, wealth, and mobility – which are also important determinants of school enrollment – failure to control for these characteristics may generate biased estimates of the impact of adult mortality on school attendance. We overcome this problem, to some extent, by estimating household- and child-fixed effects models using panel data. Evans and Miguel (2004) took a similar approach in a recent paper, using a five-year panel of 17,000 children in Busia district of western Kenya.

The paper is organized as follows. We first identify potential pathways by which adult mortality may affect child school attendance and the timing of each pathway. Second, we estimate re-interview models to assess the degree to which sample attrition is a problem and use the inverse probabilities of being re-interviewed as weights to control for attrition in the subsequent analyses. Then, using household fixed-effects to control for time-invariant unobservables, we estimate the

[1] As information about HIV transmission spreads, however, it is believed that educated people are more likely to change their behavior in ways that reduce their vulnerability to the disease compared to less educated people.

impacts of adult mortality on child school attendance in rural Kenya after stratifying sampled households and their children by wealth and gender. Kenya is one of the most heavily HIV-infected countries in the world: 13.5 percent of adults aged 15 to 49 are estimated to be living with HIV in June 2000 (NASCOP 2001).

Two sources of data are used: a three-year nationwide panel data set of 1,422 rural households, collected in 1997, 2000, and 2002, and secondary data on HIV-prevalence rates from 13 sentinel surveillance sites, collected annually between 1990 and 1999 by the National AIDS and STDs Control Programme (NASCOP). The data are used to estimate the effect of working-age adult mortality since the first survey on school attendance of children aged 7 to 14 in the second and third surveys.

The study highlights three major findings: First, children's school attendance is adversely affected by the death of working-age adults among the bottom half of sample households ranked by initial asset levels in 1997, but no significant effects are detected among households in the top half of the asset distribution. Second, working-age adult mortality negatively affects school attendance even before the death in poor households. The negative impact is larger among girls than boys. This suggests that children, especially girls, are sharing the burden of caring for sick working-age adults and/or that school fees tend to be among the first expenditures curtailed in relatively poor households after one of their prime-age members becomes chronically ill. By contrast, school attendance among boys in relatively poor households drops sharply after the adult member dies but not for girls. These results indicate the importance of differentiating between gender, impacts before as well as after the occurrence of mortality, and initial wealth conditions of the household in empirical assessments of the effects of adult mortality. Third, school attendance among the poor is negatively correlated with lagged provincial HIV-prevalence rates, even after controlling for child fixed-effects. This result suggests that AIDS is indirectly affecting child school attendance in ways other than through the death of household members, or that unobserved time-varying factors correlated with changes in HIV-prevalence rates are impeding school attendance.

The next section of the paper describes the sample and panel data used in the analysis. Section 3 presents the conceptual framework and some descriptive analyses. Estimation strategies and variables are discussed in Section 4. Estimation results are in Section 5, followed by conclusions in Section 6.

2. Data

2.1. Sampling

The study uses a three-year panel of rural household surveys implemented by Egerton University's Tegemeo Institute in 1997, 2000, and 2002.[2] The sampling frame for the surveys was prepared in consultation with the Central Bureau of Statistics, although its agricultural sample frame was

[2] Please see Yamano and Jayne (2003) for a detailed description of the sample and analyses on the impacts of adult mortality on household composition, farm production, and assets.

unavailable. First, 24 districts were purposively chosen to represent the broad range of agro-ecological zones and agricultural production systems in Kenya. Next, all non-urban divisions were assigned to one or more AEZs based on secondary data. Third, proportionally to population across AEZs, divisions were chosen purposively from each AEZ. Fourth, within each division, villages and households within selected villages were randomly selected. As a result, a total of 1,578 households were chosen from 24 districts. Thus, the sample selection is not a choice based survey which has been used by some AIDS studies, such as the World Bank's Kagera survey (Ainsworth, Beegle, and Koda 2002; Ainsworth and Dayton 2003), which oversampled households afflicted by HIV/AIDS.

From the original sample, we exclude two pastoral districts (40 households) that differ substantially from other zones and had high rates of attrition. Thirty-eight households having farm sizes greater than 20 acres were also excluded, to maintain the study's focus on small-scale households. Of the 1,500 remaining households, 1,422 households (94.8 percent) were able to be re-interviewed in 2000, and of those, 1,266 households (84.4 percent) were re-interviewed in 2002 (Table 1).[3] After two rounds of interviews with two to three year intervals, the attrition rate remains

Table 1. Sample Households and Working-Age[a] Adult Mortality in Rural Kenya

Province	Original sample households in 1997	Percentage of households re-interviewed among the original (1997) households		Percentage of households incurring working-age mortality		HIV prevalence at urban sentinel sites in 1990-94[b]
		in 2000	in 2002	between 1997-2000	between 2000-2002	
	(A)	(B)	(C)	(D)	(E)	(F)
	- # -	- (%) -	- (%) -	- % -	- % -	- % -
Coastal	91	(96.7)	(78.0)	6.8	5.6	13.2
Eastern	242	(96.3)	(88.8)	4.3	1.9	6.3
Nyanza	280	(93.6)	(87.5)	12.2	3.3	12.6
Western	303	(95.7)	(89.8)	4.8	2.6	15.8
Central	181	(96.1)	(90.6)	2.9	2.4	9.6
Rift Valley	403	(93.1)	(74.2)	0.4	1.7	12.3
Total	1,500	1,422 (94.8)	1,266 (84.4)	5.8	2.5	
Districts in Nyanza Province						
Kisumu/Siaya	188	(94.1)	(93.6)	16.9	4.0	21.6
Kisii	92	(92.4)	(75.0)	2.4	1.4	3.6

Source: Tegemeo Institute/Egerton University Agricultural Monitoring and Policy Analysis Household Surveys in 1997 and 2000. National AIDS and STDs Control Programme (NASCOP 2001)
Note: (*a*) Working-age is defined as 15-49 for women and 15-54 for men. (*b*) The average percentage of pregnant women who visited the urban sentinel-surveillance sites and tested HIV positive in 1990-1994. Data are taken from 11 urban sentinel-surveillance sites (NASCOP 2001).

[3] Out of the 78 households in the 1997 survey that were not interviewed in the 2000 survey, 48 of them were found and re-interviewed in the 2002 survey, indicating that non-contact rather than household dissolution was the main cause of attrition. Only households that were interviewed in each of the three surveys were included in this analysis.

reasonably low. Attrition rates of similar surveys in developing countries range from 5 to 30 percent after two rounds (Alderman et al. 2001; Thomas, Frankenberg, and Smith 2001).

In Table 2, we present initial characteristics of households (from the 1997 survey) categorized by attrition status. Mean per capita incomes from the initial survey in 1997 are higher among households that subsequently attrited than among households remaining in the sample. On the other hand, the proportion of poor households in 1997 is higher among attrited households than non-attrited households. There appears to be no clear difference in household wealth between households that exited the sample and those remaining in the sample. We investigate potential attrition bias in more detail later in the paper.

Table 2. Household Characteristics Stratified by Attrition Status

	All	No-Attrition Households	Attrition Households	
		Interviewed in all three surveys	Attrited in 2000, contained in 1997 sample	Attrited in 2002, contained in 1997 and 2000 sample
	- mean (sd.) -	- mean (sd.) -	- mean (sd.) -	- mean (sd.) -
Household Wealth in 1997				
Per Capita Income in 1997 (USD)	419.5	407.4	463.8	495.8
	(548.2)	(490.2)	(1023.5)	(645.7)
Percentage of Poor Households	50.0 %	49.7 %	50.0 %	52.6 %
	(50.0)	(50.0)	(50.3)	(50.1)
Working-Age Adult Mortality				
At least one in 1997-2000	5.47	5.84	n.a.	5.13
	(22.7)	(23.5)		(22.1)
At least one in 2000-2002	2.13	2.53	n.a.	n.a.
	(14.5)	(15.7)		
Primary School Attendance				
Children aged 7-14 in 2000	89.7	89.8	n.a.	89.2
	(30.3)	(30.3)		(31.1)
Children aged 7-14 in 2002	88.4	88.4	n.a.	n.a.
	(32.1)	(32.1)		
Reasons for Attrition			- number -	- number -
Temporary not available			45	n.a. [a]
Moved away			19	n.a. [a]
Dissolved			9	n.a. [a]
Refused			5	n.a. [a]
Number of Households	1,500	1,266	78	156

Source: Tegemeo Institute/Egerton University Agricultural Monitoring and Policy Analysis Household Surveys in 1997 and 2000.
Note: (*a*) The 2002 survey did not recode the reasons for attrition.

2.2. Working-age Adult Mortality

When enumerators re-visited the 1997 sampled households in 2000 and 2002, they asked about each individual in the demographic roster of the 1997 survey, including their whereabouts and basic socioeconomic information, including school attendance. Thus, we have information on all of the original household members from the 1997 survey even if some of them were no longer residents of the households in 2000 and 2002. If the household contained new members not listed in the 1997 survey, the enumerators added the new members' names on the demographic roster, obtained socio-demographic information about these individuals, and included them in the roster for subsequent surveys. Enumerators also asked about new members who may have joined the households since the previous survey but left before the next revisit survey. Thus, we should be able to capture deaths of household members who had returned to sampled households to receive terminal care. Through this process, information was obtained on individuals who had passed away since the preceding survey. After collecting information about how many months in the last 12 months each individual spends at the household, we define household members as individuals who spent more than six months at the household. If the husband resided at the household less than six of the past 12 months, then for purposes of this study, the household is considered female headed.

Further questions were asked about deceased members on the cause and year of deaths. Because it is very difficult to identify the exact cause of deaths, we have asked respondents to identify the cause of deaths as "disease," "accident," "old age," "other (specify)." Among working-age adults, defined below, about 93 percent of deaths were caused by "disease" and the remaining 7 percent (9 out of 126 working-age deaths) were caused by accidents. In this paper, we include all deaths in our analysis. However, to assess the robustness of our results to the inclusion of accidental deaths, we ran analogous regression models after excluding the nine cases of accidental death. The results in this paper are virtually identical when excluding these nine cases of accidental deaths among the working-age adults. Moreover, as discussed below, evidence clearly indicates that a high proportion of working-age adult deaths in this sample were due to AIDS.

Our analysis focuses on adults between the years of 15 to 54 for men and 15 to 49 for women because these correspond to the age ranges reported by the World Health Organization for HIV prevalence, and are the age ranges hardest hit by the disease (NASCOP 2001). Between the 1997 and 2000 surveys, 84 adults in these age ranges (which we hereafter refer to as "working-age" adults for shorthand) passed away among the 1,422 sampled households. In the 2002 survey, we found 42 additional cases of mortality in these age ranges among the 1,266 households since the 2000 survey. Thus, the total number of deceased working-age adult members is 126. Of those, about 36 percent were found in Nyanza province, where only 12 percent of the sampled households reside.

In Nyanza province, the survey covered three districts: Kisumu, Siaya, and Kisii districts. In Kisumu and Siaya districts, 16.0 percent of the sampled households experienced at least one working-age adult death in between the 1997 and 2000 surveys. In Kisii district, which is farther away from the main highways, more rural, and culturally different from the Luo dominated areas of Kisumu and Siaya, only 2.2 percent of sampled households had experienced at least one working-age death since the 1997 survey (Table 1, column D). These differences in death rates across districts are broadly consistent with the differences in HIV-prevalence among pregnant women testing HIV-positive at urban sentinel surveillance sites in these districts (Table 1, column F).

In Kisumu City, the third largest city in Kenya and the capital of Nyanza province, about 22 percent of pregnant women tested HIV positive at a sentinel-surveillance site from 1990-94. Kisumu and Siaya districts are well connected to Kisumu city with busy highways that also connect Kisumu with Nairobi to the east and with Kampala, Uganda, to the west. Along highways in Kenya, especially the Trans-Africa Highway that connects Mombasa with Kampala via Nairobi, high HIV-seroprevalence rates have been found among high-risk people, e.g., truck drivers and commercial sex workers (Carswell, Lloyd, and Howells 1985; Mbugua et al. 1995). HIV appears to spread into rural communities from truck stops via interactions between high-risk people and local people, including adolescents who interact with high-risk people (Nzyuko, et al. 1997).

Additional factors are believed to have contributed to a high HIV prevalence rate in Kisumu and Siaya districts. Such factors include the relative importance of polygamous households (which increases the risk of HIV transmission between a husband and wives), Luo peoples' traditional practices, such as widow inheritance (which increases HIV transmission from a widow to a new husband), sexual relations embedded in economic transactions (such as those associated with fish trading in which sexual favors are common among fishermen and women traders), and male non-circumcision (which increases men's risk of HIV infection per sexual contact). By contrast, the district town of Kisii, which is further from the Trans-Africa Highway and home to the Kisii ethnic group, only 3.6 percent of pregnant women tested HIV positive at a sentinel surveillance site from 1990 to 1994 (NASCOP 2001).

Therefore, the difference in the numbers of working-age adult deaths between Kisumu/Siaya districts and Kisii district is consistent with the difference in HIV prevalence rates between these districts. To examine this point further, we estimate the association between lagged HIV-prevalence rate at the nearest urban-sentinel-surveillance site and the probability of experiencing working-age adult death in Section 5.

2.3. Primary School-Age Children Included in the Analyses

In the 2002 and 2002 surveys, respondents were asked about each household member's schooling, including the number of years spent in school prior to the survey and whether each child was regularly attending school at the time of survey. Unfortunately, the 1997 survey did not ask about school attendance. The results on school attendance in columns D and E of Table 1 are taken from 2,565 and 2,107 children aged 7 to 14, who had yet to complete their primary school education by the time of the 2000 and 2002 surveys, respectively.

Before proceeding, we need to explain the criteria for determining which children are included in the analysis. First, this sample includes children who had left the sampled households since the 1997 survey. Excluding them from the final sample may cause an under-estimation of the negative impact of adult mortality on school attendance because it is more likely that children who left (or were sent away from) their households after experiencing adult mortality were less likely to be attending school than those who remained in the sampled households. In the 2000 and 2002 survey, enumerators asked about the schooling of children listed on the 1997 survey even when the children had left the sampled households. Among the 2,565 and 2,107 children in the 2000 and 2002 sample, 8.3 and 11.8 percent of children left their households (either permanently or temporarily) since the 1997 survey, respectively.

Second, despite our efforts to keep track of all of the sampled households and individuals, we lost some children in the final samples due to attrition. For instance, between the 2000 and 2002 surveys, we lost track of 345 children aged 7 to 14 years old. Out of the 345 children, 259 are in the 156 households that could not be re-interviewed in 2002. If these households suffered a higher incidence of working-age mortality between 1997 and 2002, we would have sample selection bias when estimating the impact of working-age adult mortality on child schooling. To control for attrition bias at the household level, we use the inverse probability method as discussed in Section 4.

Third, to maintain our focus on primary school attendance, we exclude from the final sample 28 and 96 children who were between 7 and 14 years old in 2000 and 2002, respectively, because respondents indicated that they had already completed primary school prior to the surveys. Finally, we are unable to determine which sampled children might be orphans because the surveys did not ask about the whereabouts of their biological parents.

Despite these shortcomings, the panel data provide us a rare opportunity to measure the impacts of working-age adult mortality on children's school attendance before and after the death of an adult household member. It is important to understand the impact of working-age adult death *before* as well as after the occurrence of a death in light of the fact that care giving is a labor intensive activity that often falls on children and may therefore affect their ability to attend school.

3. Primary Education and Working-age Adult Mortality

3.1. Primary Education in Kenya

Since independence, Kenya has successfully raised the proportion of children enrolled in primary school. The gross primary school enrollment rate (GER) increased from 62 percent in 1970 to 115 percent by 1980.[4] The gender gap in enrollment rates between male and female students also declined in this period from about 20 percent in 1970 to 10 percent in 1980. These improvements are largely due to newly built schools and free education for grades 1 to 4 introduced in 1974. However, a formal cost-sharing system was introduced in 1988 to ease the financial burden on the education system. Since then the GER has declined to about 90 percent (World Bank 2004). At the time of the surveys used in this paper the GER was 87.6 percent in 2000 and 92.1 percent in 2002.[5]

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[4] Enrollment rates can exceed 100% when some proportion of students repeats the same grade. As a result, the total number of students in primary school can exceed the number of students who would be in primary school if they all advanced normally.

[5] In January 2003, the newly elected Kenyan president Mwai Kibaki introduced Free Primary Education, which has initiated a massive increase in primary enrollment. However, at the time of the last survey used in this paper, this was not expected and hence does not affect results in this paper.

3.2. Impact of Working-age adult mortality on Schooling: Conceptual Framework

How does working-age mortality reduce child schooling? There are three main economic factors influencing child school enrollment: (a) the financial costs of schooling, such as school fees and books, relative to households' resources; (b) the opportunity costs of children' time; and (c) the expected returns from school. The potential effects of working-age mortality on child schooling depends on how working-age mortality affects these factors (World Bank 1999).

First, medical expenditure associated with prolonged illness and eventual funeral costs reduce the financial resources of the household (Barnett and Blaikie 1992; Lundberg, Over, and Mujinga 2000). Depending on the household's initial income and wealth and impact of death on these variables, the household may withdraw children from school. Because afflicted households' financial situation may not recover quickly after the death of the sick member, children's school attendance may be affected long after the death occurs.

Second, AIDS-related adult mortality is likely to negatively affect child schooling by increasing the opportunity costs of children's time. As one family member becomes chronically ill, another household member (usually a female member and often an older girl) must devote more time to care giving.[6] As a result, a care-giving female member may need to reduce time devoted to her usual activities. If a girl is expected to help care for the sick, she may stay home instead of going to school. The increased demand for care-giving labor, however, will disappear if the sick adult passes away.

The household also experiences an increased demand for labor previously provided by the sick member. The way in which households adjust to internal labor supply shocks varies according to the resources of the households. For example, households with sufficient income may hire additional workers to meet residual labor needs. Some afflicted households are able to attract additional members to at least partially offset the loss of another member (Beegle 2003; Ainsworth, Ghosh, and Semali 1995). Households that are poor and/or face high opportunity costs on their scarce resources may be most likely to call upon their children to discontinue school to provide labor to the household after an adult passes away. An increased demand for labor among non-sick members will not disappear even after the death of the sick member.

Third, because of its effect on life expectancy, HIV/AIDS in high-prevalence countries may alter individuals' time preference for money (McPherson 2001). Life expectancy in Kenya in 1999, for instance, is 48 years, compared to 58 years in 1993 (World Bank 1995; 2002). Part of the decline in life expectancy is due to increased infant mortality caused by mother-to-child HIV infection, yet there is still a considerable decline in the expected years of work among young adults. The decline in the expected years of work is expected to change parents' expectations about the lifetime return

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[6] For instance, Ainsworth and Dayton (2003) found a decline in Body Mass Index among elderly in non-poor households after an adult death. They suggested that the decline could be explained by a diversion of household resources to the medical care and/or increased demands on the time of elderly to care for sick AIDS patients in non-poor households.

from their children's education both to themselves (e.g., intergenerational transfers from children to parents in their old age) and to their children.

We lack suitable instruments to empirically distinguish the impacts of one factor from the others. However, this conceptual framework provides some insight into the potential pathways by which adult mortality may affect child schooling. We consider two hypotheses in particular. The first hypothesis is that children in poor households are more vulnerable to working-age adult mortality because of the first and second factors. To test this hypothesis we stratify the sample into two groups based on households' initial asset levels and compare model results for both groups. According to this hypothesis, we expect to find larger impacts of working-age adult mortality among the relatively poor half of the sample than the less-poor half.

The second hypothesis is that adult mortality affects boys and girls differently. In Kenya, as in most parts of eastern and Southern Africa, girls tend to assume more care giving responsibilities than boys (Opiyo 2001). Thus, if the second pathway is important, we expect to find a larger negative impact on girls' schooling than that of boys prior to adult mortality. Thus, we also stratify the sample by gender.

AIDS may affect child schooling not only through direct impacts on afflicted households but also through community effects. For instance, areas with high HIV prevalence rates suffer reduced profitability of businesses such as commercial farms, outgrower schemes, and non-farm businesses relying on wage labor (e.g., Fox et al. 2003; Rugalema 1999). Reduced economic profitability may translate into wage job contraction, fewer economic opportunities, and lower household incomes in such areas, which may indirectly affect children's school attendance. For these reasons, studies measuring the effects of AIDS on child schooling only through the pathway of how afflicted households themselves respond may underestimate the actual impact of the disease.

3.3. Descriptive Analysis

Based on data collected from the 2000 and 2002 surveys, we categorize children in the sample according to whether they incurred the death of a working age adult in their household in either of the two survey intervals: September 1997 – May 2000 and May 2000 – May 2002. In the following discussion, we call the interval between the 1997 and 2000 surveys as the first period and the interval between the 2000 and 2002 surveys as the second period. This distinction allows us to measure the differential effects on child schooling *before* as well as after the occurrence of adult mortality.

In Table 3, we group children by whether they suffered the death of a working-age adult in their household in either the first or second period. Among these different groups of children, Table 3 shows differences in school attendance rates, years of schooling, age, the school advancement rate (measured as a ratio of the actual grade attained over the grade that would have been attained had the child advanced one grade each year), and numbers of sample children in each category. In columns G and H, we present difference-in-differences (DID) estimators comparing children residing in households not experiencing any working-age adult mortality over the 1997-2002 period (the control group) and children of households that experienced it in 1997-2000 and in 2000-2002.

Table 3. Primary School Attendance and Working-age Adult Mortality

	Did household incur Working-age Mortality?						Dif-in-Dif [b]	
	NO		**YES I** mortality occurred between 1997-2000		**YES II** Mortality occurred between 2000-2002		$\Delta Y_{Yes\ I} - \Delta Y_{No}$	$\Delta Y_{Yes\ II} - \Delta Y_{No}$
	2000 (A)	2002 (B)	2000 (C)	2002 (D)	2000 (E)	2002 (F)	(G)	(H)
All Children								
Attendance (%)	90.1	88.8	91.1	81.6	70.0	89.6	-8.2*	20.9**
Grade (years)	3.50	2.88	3.58	2.86	2.78	2.38	-0.10	0.21
Age (years)	10.7	10.8	10.7	10.9	10.5	10.7	0.09	0.13
Grade Progress [a]	0.79	0.58	0.83	0.61	0.60	0.44	-0.01	0.04
Number	2,358	1,934	157	125	50	48		
Poor								
Attendance (%)	88.4	86.9	86.4	72.1	61.3	83.3	-12.8*	23.6**
Grade (years)	3.25	2.66	3.62	2.44	2.03	1.90	-0.59	-0.22
Age (years)	10.6	10.7	10.8	10.7	10.1	10.5	-0.27	-0.01
Grade Progress [a]	0.75	0.53	0.80	0.57	0.47	0.33	-0.01	0.04
Number	1,170	951	66	43	31	30		
Non-Poor								
Attendance (%)	91.8	90.6	94.5	86.6	84.2	100	-6.8	16.9
Grade (years)	3.75	3.09	3.55	3.09	4.00	3.17	0.19	-0.01
Age (years)	10.8	10.8	10.6	10.9	11.1	10.9	0.33	0.01
Grade Progress [a]	0.82	0.63	0.85	0.63	0.81	0.62	-0.02	0.11
Number	1,188	983	91	82	19	18		
Children who were aged 7-14 in both 2000 and 2002								
Attendance (%)	90.9	91.5	90.0	85.6	69.7	93.9	-5.1	23.6**
Grade (years)	2.83	3.40	2.66	3.20	2.00	2.73	-0.03	0.15
Age (years)	9.75	11.5	9.88	11.6	9.42	11.4	-0.09	0.24
Grade Progress [a]	0.80	0.60	0.78	0.58	0.57	0.46	0.00	0.09
Number	1,422		90		33			

Source: Tegemeo Institute (Nairobi)/Michigan State University Agricultural Monitoring and Policy Analysis Household Surveys in 1997, 2000, and 2002.

Note: The schooling information is not available from the 1997 survey. ** indicates 1 percent significance level; * indicates 5 percent significance level.

(a) Grade Progress is defined as a ratio of the *achieved* highest grade over the *projected* highest grade under the normal school progress, measured as Grade / (Age - 6).

(b) ΔY_{No} is calculated by subtracting the number in column A from the number in column B.
$\Delta Y_{Yes\ I}$ is calculated by subtracting the number in column D from the number in column C.
$\Delta Y_{Yes\ II}$ is calculated by subtracting the number in column F from the number in column E.

Average attendance rates[7] are 90.1 and 88.8 percent in 2000 and 2002, respectively, among the control group children. But among children who experienced working-age adult mortality in the first period (Table 3, column C and D), the mean attendance rate declines by 9.5 percentage points from 91.1 percent in 2000 to 81.6 percent in 2002. According to the DID estimator (column G), this decline is statistically different by 8.5 percentage points from the small decline in attendance rates among the control group children. This may suggest that children are dropping out of school, after experiencing adult mortality in their households. When we stratify the sample into two groups ranked by initial value of productive assets – hereafter, poor and non-poor – we find a larger decline in enrollment among the poor (from 86.4 to 72.1) than non-poor (from 94.5 to 86.6). The DID estimators indicate that the decline among the poor is statistically different from the change among the control group children.

On the other hand, children who experienced working-age adult mortality in the second period (between 2000 and 2002) were less likely to be in school in 2000 (i.e., before the death occurred) than in 2002 (Table 3, column E and F). The average attendance rate jumps from 70.0 percent in 2000 to 89.6 percent in 2002, after experiencing adult mortality. This increase in attendance is statistically larger than the change in attendance among the control group children by 20.9 percentage points. To investigate whether the comparison of different children across surveys artificially creates this large increase in attendance, we restrict the sample to children who were aged 7 to 14 in both 2000 and 2002 at the bottom of Table 3. The large increase in attendance remain significant, indicating that the attendance rate increased among the same children from 2000 to 2002 after the children experienced the working age adult mortality in 2000-2002. This may indicate that children were not in school prior to the timing of adult mortality due to the added burden of caring for the ill member and/or to temporarily provide other labor activities resulting from the ill member's inability to work and the household's inability in the short run to attract other members to compensate for the lost labor.

The findings in Table 3 suggest the following possible process: When an adult starts to become ill from AIDS, the probability drops that a school-age child will attend school. Because medical expenses tend to rise at the same time that earnings from the sick member declines, we expect that adjustments in school attendance will be greater for relatively poor households. After the death of the sick member, the child's likelihood of being in school rises because care giving needs have subsided. However, depending on whether the household has been able to attract additional labor to take over tasks formerly handled by the deceased, the opportunity cost of children's labor may still be perceived to be too high to warrant an immediate return to school. Relatively poor afflicted households may also be less able to continue incurring school fees. By contrast, children in better-off households are less subject to a dramatic change, because of their households' greater ability to draw upon savings and assets to maintain desired expenditures in the face of economic shocks.

The findings in Table 3 are only bivariate associations, which may be spuriously driven by regional differences or household characteristics. Thus, to measure the differences in attendance rates due to adult mortality, holding other factors constant, we need to employ multivariate

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[7] Throughout this paper, we use the "attendance" to signify whether the child was actually in school, as opposed to "enrollment," which does not necessarily imply regular presence in school.

techniques. In addition, because HIV infection is influenced by behavioral choices of household members, the possible correlation between working-age adult mortality and unobserved factors in schooling also needs to be controlled for.

4. Estimation Strategies and Variables

Our estimation procedure followed four steps described below. We first estimate a re-interview model to obtain probability of being interviewed in the 2000 and 2002 surveys. We use the inverse of probability weighing (IWP) method to correct for attrition in the following models. Second, we develop a working-age mortality model to identify factors associated with children experiencing the death of a working-age member in their household over the entire survey period. Third, we estimate primary school attendance models to examine the effects of mortality on schooling both before and after the timing of the death. Lastly, to examine the robustness of the findings, we estimate household- and child-fixed effects models (conditional logit and linear probability models) of child school attendance.

4.1. Estimation Strategies

The Re-Interview Model

As we discussed in the data section, the panel data used in this study suffer from some attrition (5.2 and 15.6 percent in the 2000 and 2002 surveys, respectively). If sample attrition occurs randomly, then we do not need to worry about selection biases caused by attrition, although efficiency will be lost because of a reduced sample size. But if sample attrition occurs systematically, then attrition may create selection biases. Because it is possible that the AIDS epidemic contributes to sample attrition in our data, there is a possibility that the sample attrition leads to selection biases.

To overcome this problem, we use the inverse probability weighting (IPW) method (Wooldridge 2002). The IPW method assumes that the probability of being re-interviewed (non-attrition) as a function of observable information is the same as the probability of being re-interviewed as a function of observables plus unobservables that are only observable for non-attrited observations. Thus, the IPW method works well if two conditions hold: if the observations on observed variables are strong predictors of non-attrition and if the observations on unobserved variables are not strong predictors of non-attrition.[8]

In our study, we use enumerator team categorical variables to predict re-interview. Three to four enumeration teams conducted the surveys of sampled households. Each enumeration team was headed by a supervisor who was authorized to decide whether enumerators give up in trying to contact designated households.

[8] Alternatively, we have included an attrition dummy for households that were interviewed in 2000 but could not be reinterviewed in 2002. The coefficient on the attrition dummy indicates whether there are unobserved differences between the attrited and non-attrited households after controlling for the observed characteristics based on the 2000 data, following the method proposed by Fitzgerald, Gottschalk, and Moffitt (1998). The attrition dummy was never significant in any models.

In short, we can write our re-interview model as:

$$Prob\ (R_{h\,t} = 1) = f\,(HIV_{t\text{-}j},\ X_{h\ 1997},\ T_{h\ t}, P) \qquad (1)$$

where $R_{h\,t}$ is one if a household h is re-interviewed at time t, conditional on being interviewed in the previous survey, and zero otherwise; $HIV_{t\text{-}j}$ is the average lagged-HIV-prevalence rate at the nearest surveillance site in 1990-94 for the observations in the 2000 survey and in 1992-96 for the observations in the 2002 survey; $X_{h\ 1997}$ is a set of household characteristics in the 1997 survey; $T_{h\,t}$ is a set of enumeration team dummies; and P is a set of five provincial dummies. Note that all of the variables are observable even for households that were not re-interviewed after the 1997 survey. We match the lagged HIV-prevalence information from 1990-1994 (1992-1996) with the observations of the 2000 (2002) survey because of the 6-10 year average survival time after HIV sero-conversion (Whitworth et al. 2003; Todd 2003).

We estimate Equation (1) with Probit for attrition between the 1997 and 2000 surveys, obtaining predicted probability, $\cdot_{2000}$. The observations include 1,500 sample households in the 1997 survey. Then, we estimate the same model for 1,422 sample households that were interviewed in 2000 and obtain the predicted probability, $\cdot_{2002}$. For observations in the 2000 survey, the inverse probability weight is $1/\cdot_{2000}$. But for the observations in the 2002 survey, the inverse probability weight is a product of $(1/\cdot_{2000})$ and $(1/\cdot_{2002})$ because the 2002 observations survived attrition twice. We apply these weights to models discussed below.

The Working-Age Adult Mortality Model

Next, we estimate the determinants of working-age adult mortality at the household level. The working-age adult mortality model is

$$Prob\ (D^{IP}_{h\,t} = 1) = f\,(HIV_{t\text{-}j},\ X_{h\ 1997}, P) \qquad (2)$$

where $D^{IP}_{h\,t}$ is a dummy variable that equals one if household h experienced working-age adult morality in between the previous survey and the survey at time t. We call $D^{IP}_{h\,t}$ the "immediate-past" mortality variable. Other variables are as defined as before. We estimate this model with Probit by using the inverse probabilities from the equation (1) as weights. We estimate this model for pooled data as well as the 2000 and 2002 survey separately. By estimating the coefficient of the lagged-HIV prevalence, we can establish the relationship between working-age adult mortality and the lagged HIV-prevalence rate, to indicate the extent which working-age mortality found in the panel data is associated with AIDS.

The Primary-School Enrollment Model

The economic theory and estimation models on schooling have been discussed in numerous papers (Strauss and Thomas 1995; Glewwe 2002). As we discussed in Section 2, the impact of the working-age adult mortality may start a long time before the death because of taking care of sick members and medical costs or a long time after the death because of reduced financial resources and labor. To measure the total impact of the adult mortality, we start with an adult mortality dummy variable, $D^{All}_{h\,t}$, which equals one for all of the children who experienced working-age adult mortality in our analysis (i.e., in between the 1997 and 2002 surveys). Thus, a base model for our analysis is

$$Prob\ (S_{it} = 1) = f\,(D^{All}_{h\,t}, HIV_{t-j}, X_c, X_{h\,1997}, P) \qquad (4)$$

where S_{it} is a dummy which takes one if child i attends school at time t, X_c is a set of child characteristics, and the other variables are as defined as before.

Several points should be clarified about this equation. First, we continue to use initial household characteristics in 1997, $X_{h\,1997}$, instead of contemporary household characteristics at year t, $X_{h\,t}$. Contemporary values of $X_{h\,t}$ for the years 2000 and 2002 are quite likely to be affected by $D^{All}_{h\,t}$. For instance, when a male household head dies, it is possible that his household transitions from being relatively wealthy to relatively poor. To measure the full impact of his death on his children's schooling, we need to compare his children with children like his, who resided in relatively wealthy households, not in poor households. Therefore, we hold the household characteristics at the initial level.

Second, in this estimation model, the lagged-HIV prevalence rate at the nearest surveillance site is expected to pick up broader community effects of the AIDS epidemic on child school attendance (separate from the direct effect via afflicted households). However, the lagged HIV-prevalence rate could be correlated with many regional characteristics. For instance, we know from previous studies that HIV prevalence rates tend to be high in areas with major trunk roads where there is a steady influx of outsiders. Thus, we should be cautious when we interpret the results on this variable.

Next, we disaggregate the adult mortality dummy variable into three to ascertain whether children's school attendance is affected prior to and/or after the occurrence of a death. Therefore, in addition to $D^{IP}_{h\,t}$, which measures the impacts of the adult mortality immediately after the mortality, we include two categorical variables in the models to account for time of death relative to the observations on child schooling. The "less immediate-past mortality" variable ($D^{LIP}_{h\,t}$), measures the impact of mortality in the household occurring between 1997 and 2000 on child school attendance in 2002, a three- to five-year lag. The "future-mortality" dummy variable ($D^{F}_{h\,t}$), on the other hand, indicates whether child i experiences a working-age adult death in the time interval before the next survey. For example, if a child experiences a working-age death between the 2000 and 2002 surveys, then the "future-mortality" dummy would take a value of one in 2000 and zero in 2002, while the "immediate-past" mortality dummy would take a value of zero in 2000 and one in 2002, and the "less-immediate-past" mortality variable would be zero in both 2000 and 2002. Thus, the estimation model includes three categorical variables on the adult mortality:

$$Prob\ (S_{it} = 1) = f\,(D^{LIP}_{h\,t}, D^{IP}_{h\,t}, D^{F}_{h\,t}, HIV_{t-j}, X_c, X_{h\,1997}, P) \qquad (5)$$

We estimate models in the equation (4) and (5) with Probit using the inverse probabilities as weights. After estimating these models by using the pooled data, we stratify the data into two groups ranked by initial wealth in 1997, and by gender, to obtain better understandings on the pathways that the adult mortality affects the child schooling.

Before proceeding, we clarify the notation for survey periods. In the first period (1997 to 2000), which we call $t=1$, the schooling dummy (S_{i1}) indicates whether child i is in school at the time of the 2000 survey. Analogously, S_{i2} indicates whether child i is in school at the time of the 2002 survey. The "immediate-past" working-age adult mortality dummy, $D^{IP}_{h\,1}$ ($D^{IP}_{h\,2}$) indicates whether child i resides in a household experiencing at least one working-age death in the period between the 1997 and 2000 (2000 and 2002) surveys.

Lastly, we re-estimate the equation (5) using fixed effects. Household fixed effects models purge the potential correlation between working-age adult mortality variables and unobserved household-fixed effects from the estimates while still being able to estimate the effects of child-specific characteristics, such as relationship to the household head. (We also estimated child fixed effects models but do not report these results because they are similar to the ones from the household fixed effects models.) We estimate these fixed effects models with conditional logit models as well as linear probability models with heteroskedasticity robust standard errors. Since the dependent variable is a binary variable, conditional logit models are appropriate. But because conditional logit models cannot include observations unless there are variations in the dependent variable within group, our sample size declines substantially, and so we also estimate linear probability model with household fixed effects to investigate the robustness of estimation results.

When estimating the equation (5) with the household fixed effects model, however, we cannot include time-invariant variables. In addition, we need to exclude one of the three working-age adult mortality variables ($D^{LIP}_{h\,t}$, $D^{IP}_{h\,t}$, $D^{F}_{h\,t}$) because we have only two observations for each child. Therefore, we exclude $D^{LIP}_{h\,t}$ from the equation (5) when we estimate the household fixed effects model.

The School Advancement Model

We undertake an additional set of analyses focusing on school advancement, i.e., the successful progression of children from one grade to the next. This analysis provides an alternative indication of the effects of working-age adult mortality on child education. It is possible that the effects of adult mortality could be better captured in the school advancement than in school attendance. The grade repetition rate is quite high in Kenya: it is estimated that only a third of children who have been in the education system for eight years are actually in Grade Eight (World Bank 2004). One measurement of school advancement is the ratio of the grade actually achieved over the grade that would be achieved under normal school advancement without repetition. In this paper, this is measured as G_{it} = (the highest grade attained) / (age - 6). We estimate this school advance model with household- and child-fixed effects models:

$$G_{it} = f\,(D^{LIP}_{h\,t}, D^{IP}_{h\,t}, D^{F}_{h\,t}, HIV_{t-j}, X_c) \qquad (6)$$

If we find any negative impacts of the working-age adult morality in school attendance, then we should find similar results in the school advance model because the negative impacts on school attendance should delay the school advancement. The reverse does not necessarily hold: even if we do not find any negative impacts on school attendance, we may still find negative impacts in school advancement.

4.2. Variables

As discussed in Section 2, we have excluded children who had already completed primary school. Thus the dependent variable, S_{it}, is an indicator of whether a child is attending a primary school who is eligible for enrollment. The "immediate-past mortality" dummy, $D^{IP}{}_{h\,t}$, equals one if one or more working-age adults in the household passed away since the previous survey. The HIV-prevalence variable, $HIV_{t-j,}$ is the average ratio of pregnant women who visited the nearest urban sentinel surveillance site and tested HIV positive in 1990-94 for the 2000 samples and in 1992-96 for the 2002 samples. Data from 11 such sites, maintained by the National AIDS and STDs Control Programme (NASCOP), were matched at the district-level to all households in the sample (Appendix Table A2). While unobserved time invariant effects are controlled for using child fixed-effects models, it is still possible that the lagged HIV-prevalence term may be picking up the effects of time variant unobservables. With this caveat, we include lagged HIV-prevalence rates in the models to account for broader community level effects of AIDS on child schooling.

Child characteristics include the age of the child, its squared term, and a gender dummy variable for girls. We also include two dummy variables indicating whether the child's relationship to the household head is as a grandchild, nephew, or niece, or as a distant relative or a non-relative who lives in the household.

Because the effects of working-age mortality are potentially devastating, household characteristics that are usually considered exogenous or fixed in most household models could be rapidly and severely affected by adult mortality. Because of this, we feel it is inappropriate to include in the models current values of variables such as asset levels, family size, or landholding size. For this reason, we use household characteristics based on values from the initial 1997 survey. These household characteristics are two education dummies (one for having some primary education but not completing it and one for completing primary education) for the most educated man in the household in 1997; two education dummies for the most educated woman in the household in 1997; a binary variable for female-headed households in 1997; the acres of land owned in 1997 in logs; and the value of asset holdings in 1997 in logs.

We also include two community variables designed to control (albeit imperfectly) for access to markets and services, and interactions outside the village community: distance of the village to the nearest bus/taxi stop, and distance to the nearest piped water outlet. Finally, five provincial dummies are included, with Eastern Province as the reference province.

5. Results

5.1. Determinants of Re-Interview

We first discuss results of the re-interview model (Table 4). The results indicate that the attrited households differ from non-attrited households in terms of their1997 values of observed variables.

The 1997 household characteristics are jointly significant as the determinants of re-interview. Especially, polygamous households are more likely to be re-interviewed. This may be because it is easy to find large households for enumerators. The number of men in the household is also positively associated with re-interview in the 2002 survey. The enumeration team dummies are also jointly significant; suggesting that differences in enumeration team effort could be strong predictors of re-interview.

The lagged HIV prevalence variable is negatively associated with re-interview. This may suggest that AIDS exacerbates the attrition. Households who have suffered from adult mortality due to AIDS may have moved away or dissolved, although the lagged HIV prevalence rate is also correlated with other factors that could cause sample attrition. For instance, access to large cities could be correlated with the lagged HIV prevalence but also with migration. In any case, the results in Table 4 suggest the importance of controlling for attrition, as is done in the remainder of the analysis.

5.2. Determinants of Households Afflicted by Prime-age Adult Mortality

Table 5 presents results on the determinants of working-age adult mortality. We present both models with and without controlling for attrition. We find that coefficients are stable across models with and without controlling for attrition. In general, standard errors are smaller when we control for attrition as expected from the theory (Wooldridge 2002).

The results indicate a positive correlation between the lagged HIV prevalence rate and the working-age adult mortality. A simple simulation based on the results of the pooled model (column B) indicates that the probability that a household experiences working-age adult mortality increases by 2.0 percentage point, from 2.1 to 4.1 percent, when the HIV prevalence rate increases from 5.0 percent to 13.5 percent (from the 1990 level to the 2000 level).[9] When we estimate the model for the 2000 and 2002 survey separately, we find a strong correlation between the lagged HIV prevalence rate and the working-age adult mortality in 2000 but not in 2002. This could be because of a small number of observations of working-age adult mortality between the 2000 and 2002 surveys: only 2.5 percent of sample households experienced at least one working-age adult death during this two year period.

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[9] These probabilities are calculated after changing the value of the HIV prevalence variable from one value to another, using mean sample values of all other variables except the HIV prevalence.

Table 4. Household-level Re-interview Model (Probit[a])

	Pooled Model	Attrited in 2000, contained in 1997 sample	Attrited in 2002 survey contained in 1997 and 2000 sample
	(A)	(B)	(C)
Lagged HIV Prevalence Rate			
Ratio of *HIV+* Pregnant Women in 1990-94 / 1992-96	-0.071	0.002	-0.529
	(0.65)	(0.02)	(2.75)**
Household Characteristics in 1997			
Male: Some Primary School (=1)	0.008	-0.020	0.028
	(0.35)	(0.62)	(0.84)
Male: Primary Finished (=1)	-0.019	-0.024	-0.016
	(0.93)	(1.03)	(0.48)
Female: Some Primary School (=1)	0.007	-0.009	0.026
	(0.51)	(0.52)	(1.16)
Female: Primary Finished (=1)	0.002	-0.012	0.025
	(0.14)	(0.79)	(1.03)
Female Headed in 1997 (=1)	-0.035	-0.058	-0.032
	(1.15)	(1.32)	(0.68)
Polygamous Household in 1997 (=1)	0.060	0.021	N.A.[b]
	(3.73)**	(1.42)	
Number of Male Adults in 1997	0.007	0.002	0.013
	(1.58)	(0.39)	(1.84)
Number of Female Adults in 1997	-0.005	-0.002	-0.005
	(1.18)	(0.39)	(0.68)
Land Tenure in 1997 (=1)	-0.001	-0.005	0.007
	(0.13)	(0.80)	(0.75)
ln (Landholding Size in Acres in 1997)	0.005	-0.007	0.021
	(0.86)	(1.15)	(2.26)*
ln (Asset Value in 1997)	0.001	0.002	0.001
	(1.15)	(1.63)	(0.36)
Distance to Bus Stop in 1997 (km)	0.001	0.003	-0.000
	(0.78)	(1.33)	(0.07)
Distance to Piped Water in 1997 (km)	0.000	-0.000	0.001
	(0.25)	(0.77)	(1.30)
Year 2002 (=1)	-0.012	N.A.	N.A.
	(0.54)		
Enumeration Team Dummies			
Team 2 in 2000	0.028	0.233	N.A.
	(0.96)	(10.09)**	
Team 3 in 2000	-0.043	-0.036	N.A.
	(1.76)	(1.86)	
Team 4 in 2000	-0.062	-0.039	N.A.
	(2.20)*	(1.38)	
Team 2 in 2002	-0.133	N.A.	0.164
	(4.47)**		(3.34)**
Team 3 in 2002	-0.028	N.A.	-0.117
	(0.83)		(2.79)**
Province Dummies Included[c]	YES	YES	YES
Joint Test for Team Effects (χ^2)	39.1 [0.00]**	111.5 [0.00]**	24.6 [0.00]**
Joint Test for HH Characteristics (χ^2)	17.2 [0.14]	14.5 [0.34]	21.2 [0.04]*
E[y]	0.920	0.948	0.890
Number of Households	2,922	1,500	1,422

Note: Numbers in parentheses are absolute z-scores, calculated using heteroskedasticity robust standard errors clustered for households. ** indicates 1 percent significance level; * indicates 5 percent significance level. Estimated coefficients are marginal changes in probability. (*b*) Because all the polygamous households were re-interviewed in 2002, the dummy is excluded in column C. (*c*) Five province dummies are included but not reported in this table.

Table 5. Factors Associated with Households Afflicted by Working-age Adult Mortality (Probit[a])

	Pooled Model		Adult mortality in 1997-2000		Adult mortality in 2000-2002	
		Corrected for attrition		Corrected for attrition		Corrected for attrition
	(A)	(B)	(C)	(D)	(E)	(F)
Lagged HIV Prevalence rate						
Ratio of *HIV*+ Pregnant Women	0.261	0.269	0.516	0.520	0.053	0.053
	(3.61)**	(3.74)**	(4.07)**	(4.57)**	(0.71)	(0.70)
Household Level Variables						
Male: Some Primary (=1)	-0.010	-0.010	-0.005	-0.006	-0.014	-0.014
	(0.78)	(0.71)	(0.23)	(0.23)	(1.13)	(1.26)
Male: Primary Finished (=1)	-0.011	-0.011	-0.021	-0.020	-0.004	-0.005
	(0.77)	(0.74)	(0.81)	(0.81)	(0.28)	(0.36)
Female: Some Primary (=1)	0.004	0.005	0.010	0.010	0.001	-0.000
	(0.37)	(0.44)	(0.53)	(0.55)	(0.11)	(0.00)
Female: Primary Finished (=1)	-0.005	-0.004	0.006	0.006	-0.015	-0.015
	(0.46)	(0.41)	(0.35)	(0.36)	(1.29)	(1.39)
Female Headed in 1997 (=1)	0.002	0.001	-0.016	-0.016	0.022	0.022
	(0.13)	(0.06)	(0.61)	(0.60)	(0.94)	(0.97)
Polygamous HH in 1997 (=1)	-0.014	-0.015	-0.026	-0.025	-0.000	-0.000
	(1.53)	(1.60)	(1.83)	(1.78)	(0.01)	(0.04)
# of Male Adults in 1997	0.005	0.005	0.008	0.009	0.003	0.003
	(1.95)	(1.99)*	(1.92)	(1.88)	(0.87)	(0.97)
# of Female Adults in 1997	0.002	0.002	0.003	0.002	0.001	0.001
	(0.73)	(0.73)	(0.54)	(0.52)	(0.27)	(0.43)
Land Tenure in 1997 (=1)	-0.010	-0.010	-0.009	-0.008	-0.012	-0.012
	(2.12)*	(2.05)*	(1.17)	(1.19)	(2.07)*	(2.22)*
ln (Landholding in 1997)	0.005	0.006	0.004	0.004	0.007	0.006
	(1.21)	(1.29)	(0.54)	(0.53)	(1.29)	(1.42)
ln (Asset Value in 1997)	0.000	0.000	0.004	0.004	-0.002	-0.002
	(0.44)	(0.52)	(2.36)*	(2.65)**	(2.34)*	(2.42)*
Distance to Bus Stop in 1997	-0.001	-0.001	-0.003	-0.003	0.001	0.001
	(0.71)	(0.75)	(1.28)	(1.35)	(0.44)	(0.36)
Distance to Piped Water in 1997	-0.000	-0.000	-0.000	-0.000	-0.001	-0.001
	(1.07)	(0.97)	(0.17)	(0.15)	(1.37)	(1.37)
Year 2002 (=1)	-0.033	-0.034	N.A.	N.A.	N.A.	N.A.
	(4.38)**	(4.50)**				
Province Dummies						
Nyanza province	0.026	0.024	0.029	0.030	0.006	0.008
	(1.37)	(1.27)	(0.93)	(0.95)	(0.36)	(0.46)
Coastal province	0.018	0.019	0.045	0.045	-0.007	-0.007
	(1.30)	(1.35)	(1.72)	(1.90)	(0.56)	(0.57)
Western province	-0.001	-0.001	-0.002	-0.002	-0.005	-0.004
	(0.07)	(0.08)	(0.12)	(0.13)	(0.44)	(0.40)
Central province	0.004	0.004	0.003	0.003	0.001	0.001
	(0.25)	(0.27)	(0.11)	(0.13)	(0.10)	(0.09)
Rift Valley province	-0.008	-0.009	-0.016	-0.016	-0.005	-0.004
	(0.72)	(0.84)	(0.94)	(0.94)	(0.47)	(0.37)
E[y]	0.042		0.058		0.025	
Number of Households	2,688		1,422		1,266	

Note: Numbers in parentheses are absolute z-scores, calculated using heteroskedasticity robust standard errors clustered for households. ** indicates 1 percent significance level; * indicates 5 percent significance level.
(*a*) Estimated coefficients are marginal changes in probability.

Instead of using the lagged HIV prevalence rate, we have also tried the current HIV prevalence rate from 1995-99 in the 2000 survey model. The size of the estimated coefficient declined from 0.520 (t-stat=4.57) to 0.349 (t-stat =4.43), suggesting that the lagged HIV prevalence rate is better correlated with the probability of experiencing the adult mortality than the current HIV prevalence rate, as would be expected given the long-asymptomatic period of the disease.

Turning to household characteristics, we find that an additional male member in the household increases the probability of experiencing working-age adult mortality by 0.5 percent. By contrast, the number of female members does not significantly affect the probability. Having a title deed for at least part of one's land (land tenure) is negatively associated with adult mortality. While we can only speculate as to why this relationship exists, it is possible that individuals who have made the effort to obtain secure land tenure are relatively risk-averse in a variety of ways. It is also possible that households with secure land tenure tend to spend more time farming and less time traveling away from home where the possibility of risky sexual behavior may be higher. Lastly, the sign of the estimated coefficient of the asset value in 1997 changes from positive in 2000 to negative in 2002. This might indicate a shift in characteristics of deceased persons. It is believe that at the beginning of the AIDS epidemic people with higher education and income are more likely to be infected by HIV because they are more able to attract sexual partners and tend to travel more than people with lower education and income. But as information on HIV/AIDS transmission spreads, people with higher education and income may start protecting themselves better than people with lower education and income (Ainsworth and Semali 1998).

5.3. School Attendance and Working-age Adult Mortality

Working-age adult mortality

Results in Table 6 suggest that children experiencing working-age adult mortality since the first survey in 1997 to the last survey in 2002 are about 3 percent less likely to attend school than children who did not experience adult mortality during the same period, although this result is not precisely estimated (t-statistic = 1.27). The results are stable even when using the inverse probability weighting method to control for potential attrition biases.

However, when we disaggregate the adult mortality variable into three variables, we find that children who experienced adult mortality in the preceding survey interval are about 20 percent less likely to attend school than children from non-afflicted households (column C and D).[10]

We find, however, that the results are sensitive to gender and wealth status of households. When we stratify the sample by gender, we find a larger negative impact of the future-mortality variable on girls than on boys (column A and B in Table 7). Table 7 presents results corrected for attrition bias by using the inverse probability weighting method. As with other models in this paper, the results change very little when we do not correct for sample attrition). Results in Table 7 suggest

[10] We have tried to disaggregate working-age adult mortality into two groups: core members (heads and spouses) and non-core members. The results were similar to the ones presented in Table 6 and 7 but with less precision. When we disaggregate adult mortality, we have relatively few observations on adult mortality for each group. Because of this problem, we do not disaggregate the adult mortality variable.

Table 6. Child School Attendance, Pooled Data (Probit[a])

	Basic Model		Mortality Variable Disaggregated	
		corrected for attrition		corrected for attrition
	(A)	(B)	(C)	(D)
Lagged HIV Prevalence Rate				
Lagged Ratio of *HIV*+ pregnant women	-0.202	-0.207	-0.206	-0.211
	(1.66)	(1.70)	(1.71)	(1.75)
Working-age Adult Mortality				
D^A: Any time in 1997-2002	-0.029	-0.027		
	(1.34)	(1.27)		
D^{LP}: in the Long-Past Period			-0.057	-0.054
(in between 3 to 6 years)			(1.69)	(1.62)
D^{IP}: in the Immediate-Past Period			0.022	0.023
(within the last 3 years)			(1.04)	(1.09)
D^F: in the Future Period			-0.200	-0.199
(in the next 3 years)			(2.72)**	(2.75)**
Child Characteristics				
Age in years	0.130	0.129	0.129	0.128
	(7.50)**	(7.50)**	(7.52)**	(7.52)**
Age squared	-0.006	-0.006	-0.006	-0.006
	(6.90)**	(6.91)**	(6.91)**	(6.93)**
Girl (=1)	0.009	0.009	0.009	0.009
	(1.00)	(0.96)	(1.01)	(0.96)
Grand child, Nephew, Niece (=1)	-0.010	-0.011	-0.011	-0.011
	(0.76)	(0.80)	(0.83)	(0.86)
Non Relative (=1)	-0.469	-0.459	-0.470	-0.460
	(10.79)**	(10.65)**	(10.79)**	(10.68)**
Household Characteristics in 1997				
Male: Some Primary (=1)	-0.007	-0.005	-0.007	-0.005
	(0.25)	(0.19)	(0.25)	(0.19)
Male: Primary Finished (=1)	0.011	0.013	0.011	0.014
	(0.40)	(0.48)	(0.43)	(0.50)
Female: Some Primary (=1)	0.008	0.006	0.006	0.004
	(0.46)	(0.35)	(0.35)	(0.24)
Female: Primary Finished (=1)	0.025	0.023	0.022	0.021
	(1.33)	(1.21)	(1.18)	(1.08)
Female Headed in 1997 (=1)	0.048	0.047	0.051	0.049
	(1.96)	(1.87)	(2.13)*	(2.01)*
# of Male Adults in 1997	0.015	0.016	0.015	0.017
	(0.97)	(1.08)	(1.01)	(1.12)
# of Female Adults in 1997	-0.002	-0.002	-0.001	-0.002
	(0.30)	(0.46)	(0.29)	(0.46)
Polygamous HH in 1997 (=1)	-0.004	-0.004	-0.004	-0.004
	(0.67)	(0.65)	(0.62)	(0.61)
Land Tenure in 1997 (=1)	0.012	0.013	0.013	0.013
	(1.07)	(1.11)	(1.17)	(1.20)
ln (Landholding in 1997)	-0.003	-0.003	-0.003	-0.003
	(0.45)	(0.43)	(0.44)	(0.42)
ln (Asset Value in 1997)	0.005	0.005	0.005	0.005
	(3.67)**	(3.56)**	(3.57)**	(3.48)**
Distance to Bus Stop in 1997	-0.001	-0.001	-0.001	-0.001
	(1.48)	(1.75)	(1.46)	(1.73)
Distance to Piped Water in 1997	0.003	0.003	0.003	0.003
	(1.25)	(1.29)	(1.26)	(1.29)
Year 2002 (=1)	-0.010	-0.009	-0.009	-0.008
	(1.19)	(1.06)	(1.07)	(0.95)
Province Dummies Included	YES	YES	YES	YES
Number of children	4,672			

Note: Coefficients are marginal changes in probability and numbers in parentheses are absolute robust z-scores. ** and * indicate 1 and 5 percent significance levels, respectively.

Table 7. Child School Attendance by Gender and Wealth, Pooled Data (Probit[a]) [Corrected for Attrition]

	All		Poor		Non-Poor	
	Boys	Girls	Boys	Girls	Boys	Girls
	(A)	(B)	(C)	(D)	(E)	(F)
Lagged HIV Prevalence Rate						
Lagged Ratio of *HIV*+ pregnant women	-0.305	-0.054	-0.302	-0.258	-0.155	0.098
	(2.17)*	(0.32)	(1.50)	(1.01)	(0.79)	(0.45)
Working-age Adult Mortality						
D^{LP}: in the Long-Past Period	-0.003	0.063	-0.133	0.070	N.A.	N.A.
(in between 3 to 6 years)	(0.04)	(1.70)	(1.60)	(1.06)		
D^{IP}: in the Immediate-Past Period	-0.040	0.018	-0.091	-0.016	-0.009	0.023
(within the last 3 years)	(1.34)	(0.61)	(1.77)	(0.31)	(0.29)	(0.83)
D^{F}: in the Future Period	-0.173	-0.249	-0.326	-0.331	0.040	-0.141
(in the next 3 years)	(2.01)*	(2.73)**	(2.69)**	(2.72)**	(0.64)	(1.28)
Child Characteristics						
Age in years	0.154	0.096	0.181	0.137	0.107	0.070
	(6.59)**	(3.87)**	(5.36)**	(3.50)**	(3.68)**	(2.73)**
Age squared	-0.007	-0.004	-0.008	-0.006	-0.005	-0.003
	(6.09)**	(3.48)**	(4.81)**	(3.11)**	(3.53)**	(2.58)**
Grand child, Nephew, Niece (=1)	-0.021	-0.006	-0.020	0.036	-0.033	-0.044
	(1.32)	(0.35)	(0.86)	(1.19)	(1.55)	(2.73)**
Non Relative (=1)	-0.676	-0.185	-0.757	-0.120	-0.666	-0.199
	(11.38)**	(3.75)**	(7.09)**	(1.76)	(9.95)**	(3.63)**
Household Characteristics in 1997						
Male: Some Primary (=1)	-0.012	-0.002	-0.016	0.002	-0.015	-0.021
	(0.37)	(0.06)	(0.36)	(0.04)	(0.38)	(0.53)
Male: Primary Finished (=1)	0.001	0.035	0.040	0.062	-0.040	-0.017
	(0.05)	(0.93)	(0.93)	(1.20)	(1.39)	(0.54)
Female: Some Primary (=1)	0.009	-0.001	0.017	0.005	-0.002	-0.009
	(0.43)	(0.04)	(0.63)	(0.15)	(0.08)	(0.29)
Female: Primary Finished (=1)	0.034	0.009	0.026	-0.005	0.031	0.006
	(1.51)	(0.38)	(0.87)	(0.13)	(0.94)	(0.22)
Female Headed in 1997 (=1)	0.040	0.057	0.052	0.090	-0.019	-0.011
	(1.23)	(1.78)	(1.19)	(1.74)	(0.31)	(0.22)
Polygamous HH in 1997 (=1)	0.008	0.022	-0.026	0.063	0.032	-0.003
	(0.40)	(1.15)	(0.83)	(1.99)*	(1.54)	(0.16)
# of Male Adults in 1997	0.003	-0.008	-0.002	-0.016	0.005	-0.002
	(0.42)	(1.26)	(0.18)	(1.37)	(0.63)	(0.32)
# of Female Adults in 1997	-0.008	-0.001	-0.015	-0.000	-0.002	-0.003
	(1.11)	(0.07)	(1.66)	(0.01)	(0.19)	(0.46)
Land Tenure in 1997 (=1)	0.021	0.004	0.019	-0.005	0.013	-0.000
	(1.51)	(0.27)	(0.96)	(0.21)	(0.85)	(0.01)
ln (Landholding in 1997)	-0.001	0.000	0.015	0.024	-0.007	-0.007
	(0.13)	(0.03)	(1.04)	(1.54)	(0.81)	(0.91)
ln (Asset Value in 1997)	0.004	0.005	0.004	0.004	0.003	0.008
	(2.14)*	(2.91)**	(1.14)	(0.91)	(0.52)	(1.57)
Distance to Bus Stop in 1997	-0.000	-0.001	0.002	-0.001	-0.001	-0.001
	(0.93)	(1.42)	(1.63)	(0.71)	(2.35)*	(1.72)
Distance to Piped Water in 1997	0.007	-0.001	0.004	0.005	0.007	-0.002
	(2.38)*	(0.22)	(1.00)	(0.92)	(2.47)*	(0.75)
Year 2002 (=1)	-0.010	-0.022	-0.022	-0.025	0.004	-0.014
	(0.90)	(2.07)*	(1.24)	(1.40)	(0.32)	(1.33)
Province Dummies Included	YES	YES	YES	YES	YES	YES
Number of children	2,421	2,251	1,197	1,094	1,224	1,157

Note: Numbers in parentheses are absolute z-scores, calculated using heteroskedasticity robust standard errors clustered for households. ** indicates 1 percent significance level; * indicates 5 percent significance level.
(a) Estimated coefficients are marginal changes in probability.

that girls are likely to be taking care of the sick members in the households prior to their deaths. Furthermore, we find much stronger negative impacts among children in poor households when we rank households by initial 1997 value of productive assets and stratify the sample into two groups: the relatively poor and non-poor. The results indicate that both girls and boys in the relatively poor households are about 33 percent less likely to be in school in the one-two year period prior to adult mortality. Boys in relatively poor households are also 9 percent less likely to be in school some time after experiencing adult mortality. This could be because boys are compensating for the loss of family labor in farm production or other income generating activities, or that such households can no longer afford school fees after incurring the typical increase in expenditures on health care and funeral expenses. It is possible that the recent 2003 decision of the new Kenyan government to provide universal free access to education for children under 12 may help to mitigate the impact of adult mortality on child education.

The lagged HIV prevalence rate and general equilibrium effects

The results in Table 7 indicate that the lagged HIV prevalence rate at the nearest sentinel site is negatively associated with boys', but not girls', school attendance. The HIV prevalence rate may be capturing the indirect or general equilibrium effects of the HIV/AIDS epidemic apart from the direct household-level effects of adult mortality on schooling because we already include information on working-age adult mortality in the model. For instance, it is possible that labor shortages in areas with high HIV prevalence rates have driven up agricultural wage rates, which might give boys an incentive to work instead of going to school. It is, however, difficult to identify the pathway in which the HIV prevalence rate is associated with boys' schooling.

Fixed effects models

Next, we estimate Equation (5) with household fixed effects models to examine if the results presented in Table 6 and 7 are robust when we purge the potential correlation between working-age adult mortality variables and unobserved household fixed effects (Table 8). We present the results from both conditional logit and linear probability models with household dummies. To estimate conditional logit models, only observations with variations in the dependent variable within a group can be included for estimations. This is why there are fewer observations in the conditional logit models. Because of this restriction, we also estimate the model with linear probability models. Because linear probability models suffer from heteroskedasticity problems, use heteroskedasticity robust standard errors to calculate absolute t-values.[11]

The results in Table 8 indicate that, even after controlling for the household fixed effects, the future mortality variable stays negative and significant in the linear probability models, especially for girls. This suggests that girls in households afflicted by adult mortality are less likely to be in school, possibly because of the need to take care of sick members.

[11] We have also estimated child fixed effects models. The results are similar to the ones presented in Table 8.

Table 8. Child School Attendance – Household Fixed Effects Models

	Conditional Logit	Liner Probability Model		
	Pooled	Pooled	Boys	Girls
	(A)	(B)	(C)	(D)
Lagged HIV Prevalence Rate				
Ratio of *HIV+* pregnant women	-16.313	-0.659	-1.292	-0.218
	(1.53)	(1.01)	(1.48)	(0.23)
Working-age Adult Mortality				
D^{IP}: in the Immediate-Past Period	0.792	0.064	0.090	-0.001
(within the last 3 years)	(1.42)	(1.64)	(1.52)	(0.01)
D^{F}: in the Future Period	-0.820	-0.170	-0.088	-0.356
(in the next 3 years)	(0.88)	(1.74)	(0.59)	(2.60)**
Child Characteristics				
Age in years	1.933	0.150	0.116	0.142
	(6.25)**	(6.78)**	(3.39)**	(4.07)**
Age squared	-0.083	-0.006	-0.005	-0.006
	(5.63)**	(6.35)**	(2.98)**	(3.94)**
Girl (=1)	0.086	0.015	N.A.	N.A.
	(0.57)	(1.36)		
Grand child, Nephew, Niece (=1)	0.067	0.011	0.020	0.003
	(0.21)	(0.50)	(0.52)	(0.09)
Non Relative (=1)	-3.195	-0.456	-0.644	-0.259
	(8.74)**	(9.16)**	(9.56)**	(2.74)**
Constant	N.A.	-0.090	-0.090	0.613
		(0.18)	(0.18)	(2.59)*
R-squared		0.50	0.65	0.61
Number of Households		1,167	916	870
Number of observations	1,394	4,672	2,421	2,251

Note: Numbers in parentheses are absolute z-scores in columns A and B and absolute t-statistics, calculated by using heteroskedasticity robust standard errors, in columns C and D. ** indicates 1 percent significance level; * indicates 5 percent significance level.

The results also indicate that, within the household, there is no significant difference in schooling between household heads' own children and grandchildren, nephews, and nieces. The only group of children who are less likely to be in school compared with the household heads' own children in the same household is "non-relative" children who have a relatively distant relationship or no relationship to the household head. These results are similar to those found by Case, Paxson, and Aleidinger (2002) and suggest that children are treated equally as long as they are living with their close relatives. Unfortunately, because we do not know the orphan-status of all sampled children, we cannot determine whether orphans are treated differently than non-orphans in their relatives' households. However, there is likely to be some overlap between the "non-relative" category and those children in the sample who are orphans.

5.4. The School Advancement Model

As indicated earlier, we estimate school advancement models to gauge the robustness of our findings as to the effects of working-age adult mortality on school attendance. The results from the school advancement model support the main results in the school attendance model. The child-fixed effects model indicates that working-age adult mortality slows childrens' school advancement prior to the death (Table 9). When children are stratified into poor and non-poor groups, we find a very strong negative impact among children in poor households, but not in non-poor households. This is consistent with the results in Table 7 on school attendance.

Table 9. School Advancement Model, Controlling for Household Fixed Effects

	Pooled (A)	Boys (B)	Girls (C)
Lagged HIV Prevalence Rate			
Ratio of *HIV+* pregnant women	-7.158	-7.657	-6.646
	(13.68)**	(11.57)**	(8.35)**
Working-age Adult Mortality			
D^{IP}: in the Immediate-Past Period	0.059	0.008	0.163
(within the last 3 years)	(1.19)	(0.13)	(2.16)*
D^{F}: in the Future Period	0.052	-0.066	0.174
(in the next 3 years)	(0.52)	(0.54)	(1.18)
Child Characteristics			
Age in years	-0.166	-0.086	-0.237
	(5.93)**	(2.29)*	(5.33)**
Age squared	0.007	0.004	0.010
	(5.57)**	(2.15)*	(4.91)**
Girl (=1)	0.040	N.A.	N.A.
	(2.94)**		
Grand child, Nephew, Niece (=1)	0.024	N.A.	N.A.
	(0.83)		
Non Relative (=1)	-0.216	N.A.	N.A.
	(5.19)**		
Constant	2.500	2.126	2.878
	(15.83)**	(10.29)**	(11.57)**
Number of households	1,167		
Number of children		3,091	1,492
Number of observations	4,672	4,672	2,240

Note: Numbers in parentheses are absolute z-scores. ** indicates 1 percent significance level; * indicates 5 percent significance level. The dependent variable is a ratio of the *achieved* highest grade over the *projected* highest grade under the normal school progress, measured as Grade / (Age - 6).

Adult mortality also has a generally adverse effect on grade advancement after the death, but in all cases this relationship is imprecisely measured.

6. Conclusions

A rapidly increasing mortality rate among adults due to the AIDS epidemic in sub-Saharan Africa has raised concerns about intergenerational effects, including child education. Using a panel of 1,422 households in rural Kenya, we first estimated the determinants of primary school- aged children experiencing working-age adult mortality. We find a high correlation between lagged HIV-prevalence at nearby urban sentinel surveillance sites and the probability that a household in our sample experiences the death of a working-age adult. We also find that the effects of working-age adult mortality on child school attendance are sensitive to the sex of the child, the initial wealth of the household, and the timing of the death. For example, only 54.7% of girls in initially poor households experiencing working-age adult mortality between 2000 and 2002 were predicted to be in school in 2000, compared to 88.4% of girls in relatively poor households not experiencing adult mortality. However, no effects on girls' school attendance were found *after* the death occurred, whereas these effects were quite severe on boys from relatively poor households (the predicted probability of remaining in school after the death declines 24 percentage points).

By contrast, no clear effects were found among children in households in the top half of the asset distribution in the initial survey, either before or after the timing of the death.

The fact that we find larger impacts among the poor could be because households closer to the edge of economic survival are forced to take more extreme measures to adjust to major shocks to their livelihoods, even at the expense of long-run human capital development, than their less poor neighbors. Because they have fewer options, poor households appear more likely to reallocate their children's time from schooling to care-giving for sick adults or to providing labor to compensate for the lost labor of the sick adult. However, because this study spans only a few years, it is unclear whether the effects on schooling as measured in this study are long-term or only temporary.

Because of limited capacity of local hospitals compared with the overwhelming number of AIDS patients, most AIDS patients are taken care of at their homes in rural areas. Although home-based care should be promoted to ease the burden on the medical system in rural areas, children appear to be bearing part of the burden of taking care of the sick. Policies to reduce the burden of taking care of the sick at home, such as improved community health care systems, may have an added advantage of helping afflicted households keep their children in school.

In Kenya, rural household incomes and primary school enrollment rates are high compared with neighboring countries. Thus, it is quite possible that we would find the impacts of working-age adult mortality on primary school attendance to be even greater in other African countries. There is a need to deepen our understanding on the impacts of adult mortality on child schooling in countries that are suffering from the AIDS epidemic so that governments and donor institutions can determine the best use of limited financial and human resources to mitigate the impacts of HIV/AIDS. Otherwise, there is the risk that the AIDS epidemic may produce as yet unanticipated intergenerational consequences on human capital development that might have been mitigated if understood and addressed earlier.

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Appendix Table A1. Descriptive Statistics, Pooled Data

	Mean	Std. Error	Min	Max
Schooling				
School Attendance (=1)	0.891	0.311	0	1
HIV Prevalence rate				
Ratio of *HIV*+ pregnant women	0.129	0.053	0.036	0.250
Working-age Adult Mortality				
D_t: Working-age Adult Death at *t* (=1)	0.043	0.202	0	1
D^A: Working-age in 1997-2002 (=1)	0.036	0.187	0	1
D^F: Pre-death Period (=1)	0.009	0.097	0	1
D^P: Post-death Period (=1)	0.027	0.161	0	1
Child Characteristics				
Age in years	10.71	2.247	7	14
Girl (=1)	0.482	0.500	0	1
Grand child, nephew, niece (=1)	0.226	0.418	0	1
No relative child (=1)	0.040	0.196	0	1
Household Characteristics				
Male: Some Primary School (=1)	0.171	0.376	0	1
Male: Primary Finished (=1)	0.719	0.449	0	1
Female: Some Primary School (=1)	0.251	0.434	0	1
Female: Primary Finished (=1)	0.619	0.486	0	1
Female headed (=1)	0.057	0.232	0	1
Number of male adults	1.824	1.214	0	8
Number of female adults	1.850	1.171	0	8
Land tenure (=1)	0.383	0.486	0	1
ln (Landholding size in acres)	1.511	0.902	0	4.615
ln (Asset value, Shillings)	5.695	4.359	0	15.85
Distance to Bus Stop (km)	2.313	2.626	0	20
Distance to Piped Water (km)	9.054	11.94	0	70
Year 2002 (=1)	0.451	0.498	0	1
Number of observations		4,672		

Source: Tegemeo Institute (Nairobi)/Michigan State University Agricultural Monitoring and Policy Analysis Household Surveys in 1997, 2000, and 2002.

Appendix Table A2. Three-Year Average HIV-Prevalence Rates Among Pregnant Women At Urban Sentinel Surveillance Sites

City	1990-94	1992-96	1995-99	2000 only
Busia	20.4	25.0	28.6	22.0
Kakamega	11.2	12.0	12.0	12.0
Kisii	3.6	6.4	12.6	16.0
Kisumu	21.6	24.4	28.2	35.0
Kitale	9.8	12.4	12.6	17.0
Kitui	7.2	7.6	6.6	14.0
Meru	5.3	9.5	18.4	35.0
Mombasa	13.2	13.4	15.3	12.0
Nairobi	14.3	17.6	20.0	n.a.
Nakuru	14.8	15.7	22.0	11.0
Nyeri	5.0	6.8	11.0	14.0

Source: NASCOP (2001).