Cash Transfers and Gender Differentials in Child Schooling and Labor: Evidence from the Lesotho Child Grants Programme

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Family and child allowances constitute about 16 percent of total spending on cash transfers (CTs) worldwide (Honorati, Gentilini, and Yemtsov 2015). These programs often focus on increasing investments in children's human capital, particularly in nutrition and schooling, with the goal of reducing the intergenerational transmission of poverty. Other old-age social pension programs and poverty-targeted CTs have similarly targeted human capital investment objectives. To this end, the impacts of CTs on child welfare outcomes have been widely studied (De Hoop and Rosati 2014), showing overall positive results on schooling and in some cases a reduction in child labor.[1] The bulk of such evidence on both conditional and unconditional CTs shows that they have substantial impacts on child enrollment and attendance, particularly in secondary schooling, where attendance tends to be lower in poor households (World Bank 2014).

A remaining important question about CTs, both conditional and unconditional, is whether their impacts on human capital investments are equitable between boys and girls vis-à-vis the use of their labor. The bulk of studies available on CTs show no consistency on whether impacts in education benefit girls or boys more. Gender differences in the impacts of CTs on child labor also remain inconclusive. Differences in outcomes by gender, given household access to CT programs, have important implications for gender equality in human capital accumulation and economic growth. Therefore, when designing CT programs to mitigate constraints faced by households in investing in children, public policy must also consider factors leading to unequal investments by parents and caretakers in children based on sex-specific preferences.

This article contributes to the literature on child well-being by examining gender-differentiated impacts on child schooling, labor, and time use by comparing impacts on outcomes for boys and girls across married male-headed households (MHHs) and *de jure* unmarried female-headed households (FHHs), and by the sex of cash transfer recipients. For the empirical analysis, we use impact evaluation data from the Child Grants Programme (CGP) in Lesotho, which consists of a CT provided to poor and vulnerable rural households with children. As in many sub-Saharan African countries, most rural households in Lesotho are engaged in agriculture, and the vast majority of their children are employed in crop and livestock production activities; this engagement is an important determinant of school enrollment and schooling outcomes (Kimane 2006). Rural households tend to rely on family labor and face more constraints when allocating the time that children dedicate to agricultural activities, household chores, and schooling. The context of Lesotho is also characterized by the HIV pandemic, which has affected the structure of households significantly, reducing its adult labor capacity and potentially further constraining children's time in school.

CTs, child schooling and labor

Conditional CTs (CCTs) mandate child school attendance (among other requirements) for qualification. There is clear evidence that CCTs, including large programs like Brazil's *Bolsa Familia* and Mexico's *Progresa* (Bourguignon, Ferreira, and Leite 2003; Cardoso and Souza 2004; Handa et al. 2009; Skoufias et al. 2001), have positive impacts on children's schooling, especially among students in secondary school (see Martorano and Sanfilippo [2012] for the case of Chile).

There is also evidence that social pensions and unconditional CTs (UCTs) improve child schooling. Edmonds (2006) analyzed pensions for the elderly in South Africa, finding significant increases in schooling and declines in labor participation for children, mostly for boys. Examining a monthly UCT for the ultra-poor in Malawi, Miller and Tsoka (2012) found improved education and reduced labor among children in beneficiary households. More recently, Akresh, de Walque, and Kazianga (2013) found increased school attendance rates as a result of participation in a UCT in Burkina Faso, and Handa et al. (2016) found increased school enrollment, particularly among older children, and decreased child wage labor in a UCT in Zambia.

A number of studies have suggested that education and labor outcomes are influenced by parental expectations of future labor market outcomes relative to the current opportunity cost of boys' and girls' time (World Bank 2014). It is therefore plausible that these factors also influence decisions on how to use CTs, particularly UCTs. In addition, for agricultural households facing nonseparable production and consumption decisions, the impact of CTs on household production—and therefore on labor decisions of both adults and children—are expected to be jointly determined with other outcomes such as schooling investment decisions (Benjamin 1992; Bardhan and Udry 1999; Handa et al. 2010). In Lesotho, child labor is about 23 percent—usually young men involved in the task of livestock rearing. Although no data are available for young children, the Lesotho Demographic and Health Survey (LDHS) shows that in 2009, about 76 percent of boys aged 15–19 participated in agricultural activities, whereas women of the same age-cohort worked in agriculture at a lower rate (36 percent).

In relation to household decision making in child investment by sex, child preference also plays a role in the use of CTs in child investments. Since the seminal work of Becker (1965; 1981), economists have built on his theory of choice framework to analyze intrahousehold and intergenerational resource transmission. The findings of Emerson, Souza, and Portela (2002) in Brazil provide strong evidence that parental child preferences may generate a gender bias in child human capital investments. They find that while both father's and mother's schooling had strong impacts on sons' education and labor, only mother's schooling affected the probability that a daughter works. In addition, nonlabor income (transfers) for either parent had an impact on sons' school attendance, but not on that of daughters. A strand of the literature investigating the impacts of gender-based program features remains inconclusive on the policy implications. For example, Mexico's *Progresa* provided larger transfers to households with girls to reduce the gender gap in schooling enrollment (Handa et al., 2009). However, empirical evidence has not confirmed whether the observed larger impacts on girls derived from lower initial enrollment rates for girls or from the higher payments made to them.

Various studies have already shown that child welfare is improved when women have control of a greater share of household resources, either through income (Thomas, Strauss, and Henriques 1990; Quisumbing and de la Brière 2000) or dowry (Quisumbing and Maluccio 2003), thus making the case for women to be designated cash recipients. However, there is scarce evidence comparing outcomes by sex of transfer recipient. The little existing research on child sex–differentiated impacts by sex of household recipient has in some cases suggested prevalent gender bias in intrahousehold resource allocation (Duflo 2003; Akresh, de Walque, and Kazianga 2013). More recently, a randomized controlled trial on male and female cash recipients of an education grant in Morocco found that girls had marginally higher schooling outcomes when mothers received the transfer instead of fathers. However, this difference was not observed within a UCT applied in the context of the same experiment (Benhassine et al. 2015). Others studies have made the case for a strong association between cash given to mothers and child schooling, nutrition, and general welfare (Behrman and Hoddinott 2005; Manley, Gitter, and Slavchevska 2012; De Brauw et al. 2014). Most of these studies, though, failed to compare these outcomes to a scenario with male cash recipients.

The role of household structure on differences in human capital investments on girls and boys as a result of CTs has not been widely studied either. Constraints derived from lower labor capacity are higher for agricultural households, as they tend to rely on family labor, including that of children. An important question is whether FHHs are more likely to contribute to the intergenerational transmission of poverty, as they face higher constraints for substituting child labor for child education. Empirical evidence on this question and on the social and economic factors that mitigate these poverty dynamics is essential for sub-Saharan Africa, where 26 percent of households are estimated to be headed by a woman and their prevalence has increased since the 1990s (Milazzo and van de Walle 2015), due to its changing population structure,. The extent to which FHHs are disadvantaged relative to MHHs in terms of poverty, labor capacity, access to land and livestock, and lower credit and education varies greatly across studies and contexts (Kossoudji and Mueller 1983; Handa 1996; Quisumbing 1996; Buvinić and Rao Gupta 1997). These factors can also vary between *de jure* FHHs, which are run by single, widowed, divorced, or separated women, and *de facto* FHHs, in which a husband is temporarily absent—for instance, because he is working and living abroad. Further, qualitative evidence from Lesotho shows that women enact social and caregiving roles within a gendered family and household context (Harrison, Short, and Tuoane-Nkhasi 2014).

In the context of sub-Saharan Africa, the age of the head of the household is very relevant, as (due to the HIV pandemic) FHHs sometimes consist of elderly women caring for their grandchildren. In Lesotho, households face particular constraints caused by the HIV pandemic. Starting in the 1990s, the pandemic reduced life expectancy at birth from 59 years in 1990 to 48 years in 2000 (World Bank n.d.), and life expectancy has not yet recovered to pre-1990 levels. According to the LDHS, in 2009, about 23 percent of adults were infected with HIV; in the same year, the proportion of rural households with foster and orphan children reached 47 percent. It is common for grandmothers to take charge of orphan children from their kin, also bringing additional foster children into their care. Lesotho's CGP, the CT analyzed in this article, aims to help households (and children in particular) who have been hit the most by these circumstances. About half of our sample of CGP households are headed by a woman.

Lesotho's CGP

The CGP in Lesotho is a UCT that targets poor rural households with orphans and vulnerable children. Its primary objective is to improve the living standards of such children—to reduce malnutrition, improve their health status, and increase their schooling. At the beginning of the program in 2009, the transfer value was set at a flat rate of LSL120 (US$12) per month per household and was disbursed every quarter. This amount corresponded to around 19 percent of the median consumption of an eligible household. Since April 2013, the size of the transfer was increased and indexed to the number of children, ranging from LSL120 to LSL250 (US$25) per month. Program beneficiaries are selected through a combination of proxy means testing and community validation and are registered in the National Information System for Social Assistance (NISSA) (Pellerano et al. 2014). As of December 2017, 26,600 households were benefitting from the CGP, making it the second largest social protection intervention of the Government of Lesotho after the old-age pension. Phase 1, Round 2 of the program was evaluated through a randomized experiment. A detailed questionnaire was administered to control and treatment households in July–August 2011 (baseline) and during the same months in 2013 (follow-up), so as to avoid seasonality issues. More details on the evaluation design can be found in Pellerano et al. (2014).

Although the CGP is unconditional, program messaging did affect child schooling (Pace et al. 2018). Further, an impact evaluation study carried out by Oxford Policy Management (2014) found that the CGP had a large effect on the proportion of children aged 6–19 who were attending school. The impact was driven mainly by a large decline in enrollment among older boys aged 13–17 in the control group. Enrollment for 13–17-year-old boys was 6–10 percent higher among beneficiaries. Impacts of the CGP on girls' schooling outcomes were not statistically significant but followed a trend similar to that for boys.

A qualitative study of the CGP found that children are commonly taken out of school to engage in labor activities, including farm work for boys and washing and child care for girls, especially in households engaged in agricultural activities (Oxford Policy Management 2014). Our analysis extends existing studies in several ways: (1) it investigates the impact of the CGP on gender inequality in schooling; (2) it investigates whether household characteristics, in terms of the sex of the household head and labor capacity, affect the impact of CGP on gender inequality in schooling; and (3) it investigates whether these impacts are shaped by the sex of the cash transfer recipient.[2]

Empirical framework and testable hypotheses

Our study aims to test four hypotheses. First, we tested whether the CGP has positive effects on child investment in schooling. Positive impacts would be manifested in an increase in children's time in school and a decrease in children's time in labor, particularly among older children (those aged 13–17). Such children tend to be more vulnerable to disinvestment, due to their higher labor value (e.g., in agriculture) as well as their higher schooling costs: primary schooling in Lesotho is free, while secondary schooling is not. Household decisions to invest in child education depend on marginal costs (forgone earnings from child labor and direct educational costs) and marginal benefits (higher expected earnings as an adult as they enter the labor market). CTs may reduce the marginal costs of education by reducing the relative value of children's time in work and leisure compared with schooling. The agricultural household model (Benjamin 1992; Bardhan and Udry 1999) predicts that by alleviating household credit constraints, an exogenous increase in income provided by CTs may affect simultaneously both adult and child labor. If CTs increase labor demand (say through greater employment opportunities on the farm), an increase or a decrease in child labor are both possible, depending on the elasticities of adult and child farm labor with respect to income. However, if child and adult labor are imperfect substitutes, then a decrease in child labor is to be expected. Further, if CTs increase adult participation in wage labor off of the farm, then child labor could increase or decrease, depending on the income effect of the transfer and a household's propensity to hire outside labor. Although the CGP is a UCT, the program included strong messaging about spending money on the needs of children. Hence, we expected to observe an increase in child-specific investments, particularly in education, and a decrease in child participation in agricultural and household labor.

The second hypothesis to be tested was whether the CGP reduced gender inequalities in schooling in Lesotho by generating higher impact in schooling among boys vis-à-vis girls, as boys tend to be at a disadvantage with respect to schooling. The unconditional nature of the transfer, coupled with the vulnerability of recipient agricultural households, could lead households to prioritize different needs over investing in all children equally. Therefore, we expected to observe sex differences in the outcomes of time invested in schooling and time invested in both agricultural labor and household chores by child sex and age.

If parents expect higher lifetime wages and better employment opportunities for boys than girls, then the marginal benefit of one extra year of education for boys is higher than for girls, all else held equal. If this were the case, we would expect to find CTs having a larger impact on boys than on girls. However, if the marginal costs of child education in terms of forgone earnings remain relatively higher for boys than for girls despite the transfer, then we would conclude that girls benefit more from the transfer than boys. Baseline differences between boys and girls in our sample from Lesotho showed that secondary school–aged boys are more likely to miss and repeat school and are vastly more likely to participate in crop and livestock activities than are girls. Boys aged 13–17 spend on average one additional hour on a typical day working (mostly) on farm activities or household chores compared with girls, which is consistent with the national-level data presented earlier. Among poor households participating in the CGP, boys appear to be more disadvantaged than girls with respect to educational prospects due to their participation in income-generating activities.

A third hypothesis was that household composition determines investments in children's schooling, by the sex of the household head and by the household's labor capacity. We thus would expect to find gender differences in child investment impacts due to differences in the value of human capital relative to the cost of present forgone earnings for boys and girls, by household structure. We would also expect to observe more gender-equitable outcomes among MHHs—positive impacts in child schooling, as well as higher impacts among the more disadvantaged children at baseline—as these households tend to be less labor-constrained and, with an increase in income by the CT, are more able to substitute for child labor in favor of more time for schooling. FHHs, on the other hand, tend to be more labor-constrained and, in the context of Lesotho, to be formed by one adult female, usually elderly, and children, usually their grandchildren or foster children.

The fourth hypothesis was that the control provided by assigning a CT recipient influences decision making on households' child schooling investments. In addition to examining differences in investments in boys and girls by household structure, we analyzed heterogeneous impacts by the sex of the CT recipient, who could be the father, mother, grandmother, or caretaker. To test the assumption of unitary household decision making, we compared child outcomes by the sex of the transfer recipient within married MHHs only, in which intrahousehold resource allocation decisions can be made solely or jointly. While we expected to observe gender differences in schooling outcomes according to sex of the recipient, we did not expect to see marked gender preferences. A global study on family preferences based on demographic and health survey (DHS) data suggests that in the case of Lesotho there was no statistically significant difference in girl-boy preference (18–19 percent each) and the vast majority (57 percent) prefer a balanced family in terms of girls and boys (Fuse 2010).

TABLE 1 Sample size

	Baseline	Follow-up
Total households	1,486	1,406
	[8,294]	[8,146]
Households with children aged 13–17	874	862
	[1,186]	[1,209]
Married MHHs with children 13–17	298	326
	[412]	[464]
De jure unmarried FHHs with children 13–17	482	435
	[654]	[614]
Final study sample	780	761
	[1,066]	[1,078]

NOTE: Numbers of individuals are in brackets. FHH and MHH stand for female-headed households and male-headed households.

Data and empirical strategy

Data

The empirical analysis used both baseline and 24-month follow-up data. These surveys were representative of Phase 1 (second round) of the CGP pilot, which covered five districts—Qacha's Nek, Maseru, Leribe, Berea, and Mafeteng—in 10 community councils made up of 96 electoral divisions. Electoral divisions were split equally into treatment and control groups through public lottery events in each community council. Two criteria were used to determine households' eligibility for CGP: having at least one resident child aged 0–17, and being among the poorest households in the community.[3]

Our sample is represented by the cohort of children 13–17 years of age living in *de jure* unmarried FHHs and married MHHs. It included children from both panel and attrition households, and (given the two years' lag between baseline and follow-up) half of the sample consisted of children appearing in both rounds, including those from split households.[4] Overall, the final sample was 2,144 children, 1,066 from the baseline and 1,078 from the follow-up. As shown in Table 1, which describes the sample size and the selection process, approximately 60 percent of the households in the original study had at least one child between 13 and 17 years of age. Among these households, the vast majority (around 90 percent) were either married MHHs or *de jure* unmarried FHHs.

Baseline household summary statistics

Table 2 presents summary statistics at baseline in 2011, across treatment and control households. Given the restriction of the sample to unmarried FHHs and married MHHs with children of secondary school age, some differences between the treated and control groups are to be expected,

TABLE 2 Baseline household summary statistics, by treatment arm

Variable	Control	Treatment	p-value
FHH	0.637	0.599	0.271
Age of household head	54.464	54.627	0.865
Years of education of household head	4.130	4.015	0.584
No. of members in household	5.953	6.444	0.003
Household members ≤5 years old	0.679	0.812	0.035
Household members between ≥6 and ≤12 years old	1.127	1.190	0.395
Household members between ≥13 and ≤17 years old	1.373	1.360	0.767
Household members ≥18 years old	2.775	3.081	0.004
Maseru District	0.181	0.198	0.554
Leribe District	0.207	0.244	0.225
Berea District	0.316	0.269	0.149
Mafeteng District	0.233	0.251	0.556
Qacha's Nek District	0.062	0.038	0.123
Household in crop production	0.741	0.807	0.027
No. of goods produced by household, including fruits and vegetables	1.355	1.596	0.002
No. of crops produced by household	0.777	0.873	0.119
Household owned/herded any livestock in last 12 months	0.588	0.617	0.414
Total livestock owned by household	2.673	2.909	0.452
Household operated nonfarm business in last 12 months	0.197	0.213	0.573
Individual worked in paid work outside household in last 12 months	0.451	0.411	0.265
No. of observations	386	394	

despite the randomized nature of the original design. Household composition for adult members over 18 and members aged 0–5 differed by 0.31 and 0.13 members, respectively, between the treatment and control households, and this difference was statistically significant. As a result, treatment households overall had 0.49 more household members than control households. Controlling for differences in household composition is likely to be important for measuring the impact of CTs on child outcomes, as this reflects labor composition. We also found a significant difference across treatment groups in household engagement in crop production, with control households being 6.6 percentage points less likely to participate and producing on average 0.24 fewer goods, including crops, fruits, and vegetables. Both crop production and livestock rearing were important household economic activities for the poor and vulnerable households sampled in Lesotho, with 74–80 percent and 59–62 percent engaged in crop production and livestock rearing, respectively.[5]

Table 3 compares the samples of *de jure* unmarried FHHs to married MHHs and shows statistically significant differences in characteristics of the

TABLE 3 Baseline household summary statistics, by household structure

Variable	MHH	FHH	p-value
Age of household head	52.030	56.102	0.000
Years of education of household head	2.946	4.768	0.000
No. of members in household	6.980	5.720	0.000
Household members ≤5 yrs old	0.839	0.689	0.021
Household members between ≥6 and ≤12 yrs old	1.372	1.027	0.000
Household members between ≥13 and ≤17 yrs old	1.383	1.357	0.559
Household members ≥18 years old	3.386	2.647	0.000
Maseru District	0.178	0.197	0.506
Leribe District	0.221	0.228	0.827
Berea District	0.275	0.303	0.409
Mafeteng District	0.255	0.234	0.515
Qacha's Nek District	0.070	0.037	0.039
Household in crop production	0.826	0.743	0.007
No. of goods produced by household, including fruits/vegetables	1.661	1.363	0.000
No. of crops produced by household	0.993	0.722	0.000
Household owned/herded any livestock in last 12 months	0.721	0.529	0.000
Total livestock owned by household (TLU)	4.121	1.970	0.000
Household operated nonfarm business in last 12 months	0.191	0.214	0.452
Individual worked in paid work outside household in last 12 months	0.460	0.413	0.199
No. of observations	298	482	

NOTE: While baseline treatment and control groups are not balanced across some variables, using a propensity score matched (PSM) sample does not change the main results of the analysis, suggesting that controlling for observables mitigates differences between control and treatment group.

household head and in household attributes. Household heads in FHHs are on average four years older than those in MHHs. FHH heads are also more educated and have 1.8 years more schooling than MHH heads. Other significant differences include larger households, with more members over age 18 in MHHs than in FHHs. Further, MHHs are relatively more engaged in crop production and livestock rearing, produce fractionally more fruits and vegetables and owning more livestock than FHHs. No significant differences were observed at baseline with respect to nonfarm business operations and engagement in wage labor by family members. Within household structure groups, differences between treatment and control groups were minor. (Descriptive statistics not reported here are available upon request.)

Baseline child summary statistics by gender

In Table 4, we compare how girls and boys differed before CGP payments started, particularly in the outcome variables of interest with respect to children aged 13–17. About 52 percent of girls were enrolled in the last three

TABLE 4 Baseline children summary statistics by sex

Variables	Boys	No.	Girls	No.	p-value
Family characteristics					
Age in years	14.9	544	14.9	522	0.806
Child is son/daughter of head	0.515	544	0.542	522	0.370
Child is grandchild of head	0.392	544	0.358	522	0.262
Household member is disabled	0.044	544	0.031	522	0.248
FHH	0.634	544	0.592	522	0.157
Current level of education					
Primary (1–2)	0.012	410	0.005	417	0.246
Primary (3–4)	0.137	410	0.038	417	0.000
Primary (5–7)	0.585	410	0.516	417	0.044
Secondary—junior (forms A–C)	0.256	410	0.396	417	0.000
Secondary—high (forms D–E) or higher	0.010	410	0.046	417	0.002
Brother enrolled in school	0.386	544	0.427	522	0.172
Sister enrolled in school	0.456	544	0.385	522	0.019
Dependent variables					
Schooling					
Household member ever repeated school year	0.711	519	0.596	503	0.000
Household member enrolled this year in school	0.769	536	0.820	511	0.040
Household member missed day of school in last month	0.389	519	0.323	495	0.028
No. of days of school household member missed in last 30 days	0.845	394	0.750	404	0.650
Labor (crop and livestock in last 7 days)					
Individual worked on own crops/livestock production	0.358	544	0.079	522	0.000
No. of days worked in last week, crops and livestock	2.007	544	0.362	522	0.000
Time use (minutes/day)					
Doing chores	60.9	487	95.6	460	0.000
In farm activities	74.7	488	8.9	464	0.000
In school	307.7	502	328.2	488	0.092
Doing homework	34.3	502	46.7	488	0.000

grades of primary school (years 5 to 7), compared with 58.5 percent of boys in the same age group. However, in the same age category, 39 percent of girls were in secondary school, compared with 25 percent of boys. At baseline, 71 percent of boys aged 13–17 had ever repeated a grade in school (12 percentage points more than girls), and 38.9 percent of boys had missed school in the 30 days prior to the baseline survey (6.5 percentage points more than girls). Hence, schooling among older boys appeared to be more volatile and less favored than for girls. For rural households, especially those engaged in agriculture, this implies that for a large share of older boys, the value of their current earnings relative to the opportunity cost of schooling may be considered greater than the value of their future earnings, resulting in a lower share of boys in school. In addition, researchers have observed that boys in Lesotho have lower school enrollment rates than girls and that in the context of the HIV pandemic there has been growing pressure for boys to support households economically (Nyabanyaba 2008).

In terms of labor and time use, 35.8 percent of boys participated in their own crop or livestock activities in the week prior to the survey, compared with only 7.9 percent of girls. In addition, boys in this age-group spent on average two days per week on such activities, while girls spent just 0.36 days. However, though girls aged 13–17 spent roughly 95 minutes on a typical day engaged in household chores, boys devoted roughly 35 minutes less on such activities. This confirms well-established gender roles in rural households among secondary school–aged boys and girls, which is seen not only in Lesotho but in many rural settings.

Supporting this dichotomy of gender roles, on a typical day, boys participated in farm activities and household chores on average nearly one hour more than girls did. This difference was statistically significant and would add up to a large difference between secondary school–aged boys' and girls' participation within a week. Hence, older boys were typically more disadvantaged than girls among poor rural households in Lesotho, in relative time spent on nonleisure and nonschooling activities, and to a less extent in schooling participation.[6]

Baseline child summary statistics by household structure

We next examined differences in observed child characteristics across FHHs and MHHs. For secondary school–aged children (Table 5), there was a stark contrast in terms of their relationship to the household head. Specifically, 67.7 percent of children in MHHs were the sons or daughters of the household head, while only 43.4 percent in FHHs had this relationship. Further, only 19.4 percent of boys and girls in MHHs were the grandchildren of the head, as opposed to 48.9 percent of grandchildren in FHHs. Grandmothers may view the value of the human capital relative to the opportunity cost of time differently than mothers and fathers. Moreover, households headed by a female elder may face very different constraints in terms of labor capacity and access to assets and services than households headed by younger males.

We did not observe meaningful differences in educational outcomes between MHHs and FHHs for these children. Only 32.5 percent of secondary school–aged children were enrolled in junior secondary school (forms A–C), while most of them (54–56 percent) were enrolled in primary school (years

TABLE 5 Baseline children outcomes, by household structure

Variables	MHH	No.	FHH	No.	p-value
Family characteristics					
Age in years	14.9	412	15.0	654	0.313
Child is son/daughter of household head	0.677	412	0.434	654	0.000
Child is grandchild of household head	0.194	412	0.489	654	0.000
Household member is disabled	0.029	412	0.043	654	0.253
Current level of education					
Primary (1–2)	0.019	311	0.002	516	0.008
Primary (3–4)	0.103	311	0.078	516	0.210
Primary (5–7)	0.540	311	0.556	516	0.654
Secondary—junior (forms A–C)	0.325	311	0.328	516	0.935
Secondary—high (forms D–E) or higher	0.013	311	0.037	516	0.042
Brother enrolled in school	0.403	412	0.408	654	0.863
Sister enrolled in school	0.478	412	0.385	654	0.003
Dependent variables					
Schooling					
Household member ever repeated school year	0.670	388	0.645	634	0.415
Household member enrolled this year in school	0.771	406	0.808	641	0.148
Household member missed day of school in last month	0.360	394	0.355	620	0.857
No. of days of school household member missed in last 30 days	0.900	301	0.734	497	0.443
Labor (crop and livestock in last 7 days)					
Individual worked on own crops/livestock production	0.262	412	0.196	654	0.011
No. of days worked in last week, crops and livestock	1.442	412	1.050	654	0.012
Time use (minutes/day)					
Doing chores	69.5	372	83.1	575	0.023
In farm activities	60.3	374	31.2	578	0.000
In school	313.0	382	320.8	608	0.531
Doing homework	37.3	382	42.4	608	0.132

5–7), below the school grade that they should be in given their age. This indicates a lack of resources to remain in school for children in this age-group, most likely due to household economic constraints and a high level of grade repetition. Further, there were no significant differences across MHHs and FHHs in regard to other key schooling indicators, either in the likelihood of repeating school (69 percent vs. 67 percent) or in the likelihood of having missed school days in the month prior the baseline survey (36 vs. 35.5 percent). Interestingly, boys and girls from MHHs were nine percentage points more likely than those from FHHs to have a sister enrolled in school in the current year, a statistically significant difference.

Consistent with the above, we also observed a large and significant difference in the likelihood of secondary school–aged children participating in farm labor in the seven days prior to the survey (26 percent in MHHs, vs. 19.6 percent in FHHs). On a typical day, such children in MHHs spent 60 minutes on farm activities, while those in FHHs spent just 31 minutes. However, the same children in FHHs spent on average 83 minutes on chores, while those from MHHs spent 69 minutes on them. Children from MHHs also spent less time at school and doing homework than did those from FHHs. Most of these differences were significant, suggesting that farming activities take precedence in MHHs, where livestock rearing is more prevalent, and take up more time among male children. From the summary statistics, labor activities in MHHs for this group of children were likely to lead to greater substitution away from schooling relative to FHHs. Children from FHHs spent more time on household chores, most likely because children in FHHs were less likely to engage in livestock rearing and more likely to substitute time on chores, including fetching water, sibling care, cleaning, cooking, washing, and shopping.

Impact of cash transfers on child schooling, time use, and labor

Our empirical framework was based on two fundamental assumptions: (1) differences between the treatment group (eligible cash recipients) and the control group (eligible but not cash recipients) can be mitigated by conditioning on observables, at the community, household, and individual levels; and (2) unobservable differences for individuals are time-invariant and can be controlled for through individual fixed effects. We recovered the average treatment effect on the treated of the CT on child-level outcomes by estimating the following equation:

$$Y_{iht} = \alpha_0 + \alpha_1 Treat_h * Post_t + \alpha_2 Post_t + \boldsymbol{\theta X}_{iht} + \boldsymbol{\pi Z}_{ht,2011} + \boldsymbol{\rho Q}_{ct} + \delta_i + \varepsilon_{iht} \quad (1)$$

where i indexes individual, h household, c community, and t survey year (t = 2011 or 2013). The dependent variable Y was characterized by outcomes for youth labor, schooling, and time use. $Treat_h$ was an indicator variable set to 1 if the household is a CT beneficiary, and $Post_t$ was an indicator denoting the follow-up period. We denoted by $\boldsymbol{X}_{iht}$ a vector of individual control variables, which include age of the child, whether the child is the son/daughter of the head, whether the child is the grandson/granddaughter of the head, whether the child is disabled, and whether the child has a sister/brother enrolled in school. Similarly, $\boldsymbol{Z}_{ht,\,2011}$ and $\boldsymbol{Q}_{ct}$ were household-level covariates evaluated at baseline and community controls, respectively. Household covariates included age of head, education of head, household size, and household composition, by age-group and sex (to control for potential differences in labor constraints and sex ratio), while community variables consisted of price, wage, and shock indicators. Individual fixed effects (δ_i) were used to control for time-invariant individual characteristics. Our parameter of interest was α_1, the difference in differences estimator, which measures the impact of the CGP on child outcomes.

To estimate the discrete impacts of CTs on gender-bias in child outcomes further, we ran the following fixed-effects regression:

$$Y_{iht} = \beta_0 + \beta_1 Treat_h * Post_t * Girl_i + \beta_2 Treat_h * Post_t + \beta_3 Post_t * Girl_i + \beta_4 Post_t + \boldsymbol{\theta X}_{iht} + \boldsymbol{\pi Z}_{ht,2011} + \boldsymbol{\rho Q}_{ct} + \delta_i + \varepsilon_{iht} \quad (2)$$

The above equation differed from the first equation only in its incorporation of the $Girl_i$ indicator, denoting if the sample individual is a girl.[7] Here, we were interested in the coefficient β_2, which represented the impact of the program on individual-level outcomes (schooling, labor, and time use) for boys, and coefficient β_3, which measured the differential impact in outcomes for girls with respect to boys. In our impact estimates table, we also reported for simplicity $\beta_2+\beta_3$, which represents the impact for girls.

Similar to equation (2), we examined the impacts of CTs on child outcomes by household structure:

$$Y_{iht} = \gamma_0 + \gamma_1 Treat_h * Post_t * FemHead_h + \gamma_2 Treat_h * Post_t + \gamma_3 Post_t * FemHead_h + \gamma_4 Post_t + \boldsymbol{\theta X}_{iht} + \boldsymbol{\pi Z}_{ht,2011} + \boldsymbol{\rho Q}_{ct} + \delta_i + \varepsilon_{iht} \quad (3)$$

where $FemHead_h$ was set to 1 if a household was a *de jure* unmarried female-headed one.

Lastly, for the married MHH sample, we estimated the impact of the gender of the CT recipient:

$$Y_{iht} = \varphi_0 + \varphi_1 Treat_{1h} * Post_t + \varphi_2 Treat_{2h} * Post_t + \varphi_3 Post_t + \boldsymbol{\theta X}_{iht} + \boldsymbol{\pi Z}_{ht,2011} + \boldsymbol{\rho Q}_{ct} + \delta_i + \varepsilon_{iht} \quad (4)$$

where $Treat_{1h}$ was set to 1 if the household received a treatment and the gender of the recipient was female. Similarly, $Treat_{2h}$ was set to 1 if the household treatment recipient was male. For equations (3) and (4), a potential threat to identification stems from the fact that the household structure (FHH–MHH) and the gender of the recipient within the MHH were potentially endogenous and systematically correlated with observed household characteristics, as well as other unobservable factors. To mitigate such concerns, we controlled for observable household characteristics and utilized individual fixed effects, which should minimize time-invariant individual and household differences. These observable household characteristics were

TABLE 6 Impact of CGP on schooling outcomes, overall and by sex, children aged 13–17

Panel A: Schooling outcomes	**(1) Ever repeated school year**	**(2) Enrolled in school**	**(3) Missed any day of school (0/1)**	**(4) No. of days of school missed**
Treat*Post	−0.095*	0.088**	−0.146**	−0.215
	[0.05]	[0.04]	[0.06]	[0.57]
Observations	2,076	2,111	2,063	1,647
R-squared	0.129	0.267	0.189	0.22

Panel B: Heterogeneous impacts by sex	**(1) Ever repeated school year**	**(2) Enrolled in school**	**(3) Missed any day of school (0/1)**	**(4) No. of days of school missed**
ITT boys = β_2	−0.109	0.063	−0.14*	−0.637
	[0.07]	[0.05]	[0.07]	[0.74]
ITT girls = $\beta_1+\beta_2$	−0.074	0.113*	−0.154*	0.244
	[0.07]	[0.06]	[0.09]	[0.57]
Differential effect for girls = β_1	0.035 [0.09]	0.05 [0.08]	−0.014 [0.11]	0.881 [0.65]
Observations	2,076	2,111	2,063	1,647
R-squared	0.13	0.269	0.189	0.224

*$p<0.10$. **$p<0.05$. ***$p<0.01$.
NOTE: Standard errors clustered at the community level in brackets. ITT = intention to treat effect.

measured at baseline, to avoid any bias caused by the inclusion of a covariate that was affected by the treatment.[8]

Results

Gender-differentiated impacts on child investment

Panel A of Table 6 presents the results from the estimation of equation (1) on the impact of the CGP on children. We observed an overall reduction in the likelihood of repeating school years (9.5 percentage points), though the effect was statistically significant only at $p<0.10$. Further, we found that children aged 13–17 were 8.8 percentage points more likely to be enrolled in school and 14.6 percentage points less likely to have missed any days of schooling in the last 30 days (columns 2 and 3, respectively). Both results were significant at $p<0.05$ and seemed driven by girls, as shown in Panel B of Table 6, which shows the heterogeneous impacts by gender obtained by estimating equation (2). Magnitudes for girls were slightly higher (in absolute terms), though the β_1 coefficient (the interaction term) was always statistically nonsignificant.

Looking at the impact of CGP on the time use of girls and boys (Table 7, Panel A), the CGP caused a reduction by 22 minutes in the time spent on household chores for children on a typical day. This represents a large reduction time-wise relative to the baseline average. In addition, from

TABLE 7 Impact of CGP on child time use and labor outcomes, overall and by sex, children aged 13–17

	Time use (minutes/day)			Farm labor (last 7 days)		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Time use and labor outcomes	**Chores**	**Family labor**	**At school**	**Homework/ study**	**Worked (0/1)**	**>No. of days worked**
Treat*Post	−22.082*	2.169	41.447**	5.895	−0.082	−0.612**
	[12.61]	[16.93]	[20.55]	[7.25]	[0.05]	[0.27]
Observations	1,976	1,983	2,048	2,048	2,141	2,141
R-squared	0.169	0.159	0.182	0.192	0.162	0.15

	Time use (minutes/day)			Farm labor (last 7 days)		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel B: Heterogeneous impacts by sex	**Chores**	**Family labor**	**At school**	**Homework/ study**	**Worked (0/1)**	**No. of days worked**
ITT boys $= \beta_2$	−1.092	6.966	19.98	−1.219	−0.117*	−0.824**
	[14.68]	[24.77]	[28.98]	[8.26]	[0.07]	[0.38]
ITT girls $= \beta_1 + \beta_2$	−48.658**	0.494	62.606**	14.535	-0.041	−0.337
	[19.87]	[17.35]	[29.81]	[11.05]	[0.07]	[0.40]
differential effect for girls $= \beta_1$	−47.566**	−6.472	42.625	15.754	0.076	0.487
	[23.40]	[27.13]	[41.72]	[12.69]	[0.09]	[0.57]
Observations	1,976	1,983	2,048	2,048	2,141	2,141
R-squared	0.18	0.166	0.19	0.195	0.163	0.152

$^{*}p < 0.10$. $^{**}p < 0.05$. $^{***}p < 0.01$.
NOTE: Chores include helping at home, fetching water, sibling care, cleaning, cooking, washing, and shopping. Family labor includes family farming/herding and other family business. Standard errors clustered at the community level in brackets.

column (3), children also spent approximately 41 minutes more time at school on a typical day. These changes were statistically significant gains for poor and vulnerable households gaining access to the CGP. Further, as seen in column (6), children were likely to have worked 0.6 significantly fewer days on the farm in the past week. The results on time use and farm labor for complement the results observed in Table 6 on schooling.

Panel B of Table 7 provides estimates of the heterogeneous impacts of CGP on time use by sex. From column (1), we observe that as a result of the CGP, secondary school–aged girls spent significantly less time than boys on household chores—almost 48 minutes per day less. While the result for the difference between girls and boys on time spent at school was nonsignificant (column 3), girls in treatment areas spent 62 minutes more on a typical day at school than did girls in control areas. Boys also had a lower participation rate (–11.7 percentage points) and a smaller number of days per week (–0.8) in farm labor. This is not unusual, as a larger proportion of older boys engage in livestock herding and crop production in Lesotho, while girls typically spend more time on household chores. However, in terms of time allocation, older girls benefited more from the CGP, spending more time in school and less time on household activities.

Overall, the results on child schooling, time use, and labor impacts of Lesotho's CGP suggest gender differences in outcomes among agricultural households favoring 13–17 year-old girls, with this group being less likely to miss school, being likely to spend more time at school, and experiencing less of a time burden in engaging in household chores. Despite the positive results, overall program outcomes seemed not to be working toward a reduction in the existing inequalities between girls' and boys' education among agricultural households in Lesotho.

Gender-differentiated impacts by household structure

Following equation (3), we examined whether the observed impacts differed by household structure. Our results on child schooling in Table 8 showed that children aged 13–17 in FHHs were more likely to repeat a school year than were children in MHHs (a difference of 23.6 percentage points). This difference stems from a statistically significant 23-percentage-point reduction in the likelihood of ever repeating a school year in MHHs. In addition, from the samples stratified by sex of children (columns 2 and 3), we find that the impact by household structure on school repetition is driven by girls, as they are 30 percentage points less likely to ever repeat a school year.

We did not find any differential school enrollment effect for FHHs for the full sample. However, looking at the stratified samples revealed two strikingly different results: first, girls were significantly much more likely to be enrolled in school in MHH treatment areas (23.8 percentage points), and

TABLE 8 Impact of CGP on child schooling, by household structure, children aged 13–17

	Ever repeated school year			Enrolled in school		
	(1) All	(2) Girls	(3) Boys	(4) All	(5) Girls	(6) Boys
ITT MHH = γ_2	−0.23***	−0.302***	−0.151	0.083	0.238**	−0.173**
	[0.08]	[0.11]	[0.12]	[0.05]	[0.10]	[0.08]
ITT FHH = $\gamma_1+\gamma_2$	0.006	0.022	0.041	0.079*	−0.047	0.133**
	[0.06]	[0.11]	[0.08]	[0.05]	[0.07]	[0.07]
Differential effect for FHH = γ_1	0.236**	0.324**	0.193	−0.004	−0.285***	0.306***
	[0.10]	[0.16]	[0.13]	[0.07]	[0.10]	[0.10]
Observations	2,076	1,014	1,062	2,111	1,028	1,083
R-squared	0.14	0.244	0.253	0.273	0.365	0.442

	Missed any day (0/1)			No. of days of school missed		
	(7) All	(8) Girls	(9) Boys	(10) All	(11) Girls	(12) Boys
ITT MHH = γ_2	−0.119	−0.292**	0.075	−0.394	0.942*	−0.11
	[0.08]	[0.13]	[0.12]	[0.59]	[0.50]	[1.07]
ITT FHH = $\gamma_1+\gamma_2$	−0.147**	−0.097	−0.205**	−0.232	0.259	−0.388
	[0.07]	[0.11]	[0.09]	[0.56]	[0.44]	[0.73]
Differential effect for FHH = γ_1	−0.028	0.196	−0.28**	0.162	−0.683	−0.278
	[0.09]	[0.15]	[0.14]	[0.55]	[0.54]	[1.31]
Observations	2,063	1,006	1,057	1,647	842	805
R-squared	0.196	0.318	0.32	0.23	0.375	0.465

*p <0.10. **p <0.05. ***p <0.01.
NOTE: Standard errors clustered at the community level in brackets. ITT = intention to treat effect. MHH = married male-headed households; FHH = *de jure* unmarried female-headed households

while the decline for FHHs was not statistically significant, this difference translates into an effect of –28.5 percentage points. The opposite result was observed for boys: those in MHHs were 17 percentage points less likely to be enrolled, while those in FHHs were 13 percentage points more likely. Thus, there was a 30.6-percentage-point differential impact in FHHs among boys. A similar story emerged if we considered as a dependent variable whether children had missed any school day in the last 30 days. Overall, the impact was significant only for children in FHHs. However, when we looked at the samples by the sex of the child, girls in MHHs were much less likely to miss school (a difference of 29.2 percentage points), while boys in FFHs were more likely to do so (a difference of –20.5 percentage points). The differential effect was statistically significant only for the sample of boys. The results on schooling indicate that girls in MHHs are likely to gain from access to the CGP in Lesotho. However, in FHHs, we observed some benefits to the CGP concentrated among boys. That is, in MHHs, the CGP results in a bias that favors girls, while boys in FHHs are more likely to attain positive school enrollment outcomes as a result of the transfer.

TABLE 9 Impact of CGP on time use (minutes/day), by household structure, children aged 13–17

	Chores			Family labor		
	(1) All	**(2) Girls**	**(3) Boys**	**(4) All**	**(5) Girls**	**(6) Boys**
ITT MHH = γ_2	−34.47*	−78.646**	−7.542	14.013	5.329	−23.76
	[18.84]	[35.23]	[22.72]	[24.80]	[14.94]	[43.67]
ITT FHH = $\gamma_1+\gamma_2$	−7.613	−12.312	−4.648	−6.582	5.42	−20.12
	[15.00]	[26.69]	[14.41]	[18.94]	[10.49]	[35.32]
Differential effect for FHH = γ_1	26.856	66.335*	2.894	−20.595	0.09	3.64
	[22.20]	[34.82]	[24.80]	[28.36]	[17.14]	[46.79]
Observations	1,976	960	1,016	1,983	967	1,016
R-squared	0.183	0.341	0.305	0.162	0.263	0.274
	At school			**Homework**		
	(7) All	**(8) Girls**	**(9) Boys**	**(10) All**	**(11) Girls**	**(12) Boys**
ITT MHH = γ_2	34.517	90.617*	−62.136	17.435	30.747*	−1.736
	[28.19]	[51.69]	[40.17]	[10.79]	[16.70]	[12.30]
ITT FHH = $\gamma_1+\gamma_2$	39.657	52.511	1.629	−2.955	−8.818	−1.225
	[28.53]	[42.98]	[39.22]	[9.10]	[15.49]	[10.19]
Differential effect for FHH = γ_1	5.139	−38.106	63.765	−20.39	−39.565*	0.512
	[39.00]	[64.46]	[45.95]	[13.56]	[21.48]	[14.64]
Observations	2,048	1,004	1,044	2,048	1,004	1,044
R-squared	0.187	0.304	0.325	0.198	0.371	0.271

$^{*}p<0.10$. $^{**}p<0.05$. $^{***}p<0.01$.
NOTE: Standard errors clustered at the community level in brackets. ITT = intention to treat effect. MHH = married male-headed households, FHH = *de jure* unmarried female-headed households

Similar results can be seen in Table 9, where we analyzed the impacts of the CGP on girls' and boys' time use by household structure. Overall, we found no differential impacts across FHHs and MHHs. However, the coefficients from the stratified samples indicated that girls aged 13–17 in MHHs were less likely to engage in household chores, by almost 80 minutes in a typical day. This translated into a 66-minute difference in time spent by girls in FHHs in doing chores when compared with those in MHHs. Despite the lack of a significant difference in the pooled regression, in MHHs girls spent more time at school (90 minutes) and doing homework (30 minutes), though only the latter outcome shows a significant negative differential effect for FHHs (–39 minutes, but significant only at p<0.10). These results complement the impacts of CGP on schooling outcomes across household structures observed for secondary school aged–children.

Finally, looking at the impacts of the CGP on participation in farm labor by household structure (Table 10), we could find no differential impact on FHHs, either in the pooled regressions or in the stratified samples. A significant reduction in the number of days worked in MHHs was driven by

TABLE 10 Heterogeneous impacts on farm labor (last 7 days), by household structure, children aged 13–17

	Worked (0/1)			No. of days worked		
	(1) All	(2) Girls	(3) Boys	(4) All	(5) Girls	(6) Boys
ITT MHH = γ_2	−0.089	−0.007	−0.153	−0.707*	−0.034	−1.146*
	[0.07]	[0.10]	[0.13]	[0.41]	[0.45]	[0.68]
ITT FHH = $\gamma_1+\gamma_2$	−0.09	−0.104	−0.117	−0.547	−0.209	−0.874
	[0.06]	[0.09]	[0.09]	[0.36]	[0.49]	[0.57]
Differential effect for FHH = γ_1	−0.001	−0.097	0.036	0.16	−0.175	0.272
	[0.09]	[0.11]	[0.15]	[0.54]	[0.54]	[0.90]
Observations	2,141	1,045	1,096	2,141	1,045	1,096
R-squared	0.167	0.271	0.235	0.151	0.196	0.241

$^{*}p<0.10$. $^{**}p<0.05$. $^{***}p<0.01$.
NOTE: Standard errors clustered at the community level in brackets. ITT = intention to treat effect. MHH = married male-headed households, FHH = *de jure* unmarried female-headed households

the sample of boys, who at the baseline were in a disadvantaged position compared with girls.

Gender-differentiated impacts in MHHs, by sex of transfer recipient

To assess the influence of potential gender bias by the transfer recipient toward boys or girls, we analyzed the impacts of the CGP within the subsample of married MHHs.[9] In Table 11, ϕ_1 is the coefficient associated with the interaction between the indicator for a female recipient and the time dummy (*Treat1∗Post*) and denotes the CGP impact for households with female recipients; ϕ_2 is the coefficient for the interaction between the indicator for a male recipient and the time dummy (*Treat2∗Post*) and isolates the CGP impact for households with male recipients.[10] From columns (1) to (3), both girls and boys in recipient households were significantly less likely to have ever repeated a school year, though the impact was greater for girls in households with a female recipient and for boys in households with a male recipient. With respect to school enrollment (columns 4 to 6), the negative impact observed for boys in Table 8 was magnified when the cash recipient was a male. Interestingly, the positive impact observed for girls was statistically significant only when the recipient was a male. This latter finding is mirrored by a similar result in the probability of missing any school day in the last 30 days, as the reduction was significant only for the male recipient interaction term.

Table 12 distinguishes the impact of CTs on child time use outcomes by sex of the recipient within the subsample of married MHHs. Columns (1) to (3) indicate that participation in household chores was significantly reduced among children aged 13–17 in households with a female recipient. However, this reduction was highly significant only for the sample of boys

TABLE 11 Heterogeneous impacts on schooling, by sex of the recipient in MHHs, children aged 13–17

	Ever repeated a school year			Enrolled in school		
	(1) All	(2) Girls	(3) Boys	(4) All	(5) Girls	(6) Boys
ϕ_1 (female recipient)	−0.285***	−0.513***	−0.341**	0.022	0.079	−0.353***
	[0.10]	[0.15]	[0.17]	[0.08]	[0.17]	[0.12]
ϕ_2 (male recipient)	−0.338***	−0.42***	−0.44**	0.152**	0.254*	−0.496***
	[0.09]	[0.10]	[0.21]	[0.07]	[0.14]	[0.14]
Observations	757	406	351	777	415	362
R-squared	0.422	0.55	0.658	0.409	0.559	0.736

	Missed any day (0/1)			No. of days of school missed		
	(7) All	(8) Girls	(9) Boys	(10) All	(11) Girls	(12) Boys
ϕ_1 (female recipient)	−0.165*	−0.324	0.026	−0.903**	0.662	−7.522***
	[0.09]	[0.22]	[0.12]	[0.43]	[0.79]	[0.00]
ϕ_2 (male recipient)	−0.204**	−0.4**	0.278*	−1.529**	−0.269	−11.418***
	[0.09]	[0.17]	[0.17]	[0.66]	[1.38]	[0.00]
Observations	763	407	356	597	337	260
R-squared	0.325	0.523	0.69	0.401	0.514	0.973

$^{*}p < 0.10$. $^{**}p < 0.05$. $^{***}p < 0.01$.
NOTE: Standard errors clustered at the community level in brackets.

TABLE 12 Heterogeneous impacts on time use (minutes/day), by sex of recipient in MHHs, children aged 13–17

	Chores			Family labor		
	(1)	(2) Girls	(3) Boys	(4) All	(5) Girls	(6) Boys
ϕ_1 (female recipient)	−35.089*	−23.994	−72.879**	30.461	19.325	45.969
	[20.82]	[29.14]	[29.40]	[24.48]	[17.95]	[59.74]
ϕ_2 (male recipient)	−24.399	−53.054*	−15.634	-9.977	0.844	-19.206
	[16.92]	[29.19]	[19.06]	[26.73]	[14.12]	[60.29]
Observations	732	391	341	733	393	340
R-squared	0.21	0.282	0.305	0.237	0.224	0.276

	At school			Homework		
	(7) All	(8) Girls	(9) Boys	(10) All	(11) Girls	(12) Boys
ϕ_1 (female recipient)	−17.645	−30.412	−17.985	14.385	20.916	4.53
	[30.46]	[52.05]	[46.24]	[10.50]	[17.00]	[14.87]
ϕ_2 (male recipient)	28.395	22.342	6.166	20.799*	4.316	28.321**
	[26.83]	[49.09]	[33.55]	[10.64]	[15.18]	[12.68]
Observations	749	403	346	749	403	346
R-squared	0.261	0.335	0.395	0.235	0.305	0.379

$^{*}p < 0.10$. $^{**}p < 0.05$. $^{***}p < 0.01$.
NOTE: Standard errors clustered at the community level in brackets.

TABLE 13 Heterogeneous impacts on farm labor, by sex of recipient in MHHs, children aged 13–17

	worked (0/1)			# days worked		
	(1) All	(2) Girls	(3) Boys	(4) All	(5) Girls	(6) Boys
ϕ_1 (female recipient)	−0.056	−0.039	0.028	−0.541	−0.11	−0.511
	[0.09]	[0.10]	[0.15]	[0.50]	[0.49]	[0.90]
ϕ_2 (male recipient)	−0.065	0.053	−0.193	−0.307	0.459	−1.03
	[0.09]	[0.10]	[0.15]	[0.50]	[0.45]	[0.88]
Observations	784	418	366	784	418	366
R-squared	0.297	0.199	0.221	0.302	0.206	0.204

$^{*}p<0.10$. $^{**}p<0.05$. $^{***}p<0.01$.
NOTE: Standard errors clustered at the community level in brackets.

(–72 minutes). We also found that girls in MHH spent less time (53 minutes) in chores when the CT recipient was male. Further results distinguishing between male and female CT recipients were weak for most of the other time use outcomes. In columns 10 and 12, we found that the increase in time spent doing homework was concentrated in households where men were the recipients of CTs. Lastly, Table 13 showed no significant impact of the sex of the cash recipient on aspects of farm labor.

Discussion and conclusion

In this article, we investigated gender differences in household child investment behavior using data from a randomized controlled trial aimed at measuring the impacts of the Lesotho Child Grants Program, a CT program directed to poor households with children. The analysis focused on households with children of secondary school age, as in Lesotho access to primary school is almost universal. In addition to observing impacts between boys and girls, we sought to examine whether gender-differentiated impacts varied according household structure. This exercise furthered understanding on the different constraints experienced by different types of households and their decision making as a result of cash transfers. We therefore analyzed impacts in child investments by married MHHs and *de jure* unmarried FHHs, the latter having lower labor capacity constrained by older status of the head and a higher household dependency ratio. Finally, we explored the relationship between gender-differentiated impacts and potential sex bias, as determined by who in the household receives the cash. This aimed to deepen the evidence for or against the idea of child preference by sex, particularly in schooling. Our hypothesis was that in the case of Lesotho's CGP, household structure—and therefore the household's capacities and constraints—plays a role in influencing observed gender-differentiated impacts on child investments, rather than gender bias.

First, we found that the CGP increased existing secondary school–aged girls' advantage, as these girls in treatment households benefited significantly more from increased schooling enrollment, fewer missed days of school, and time use activities shifting away from household chores. Among secondary school–age boys, however, we noted a significant decline of almost one day less spent working in crop production or livestock activities as a result of the transfer. These results are not unusual, as a larger proportion of older boys engage in livestock herding and crop production in Lesotho, while girls typically spend more time on household chores. However, these girls benefited more than boys from the CGP, considering that girls were already in an advantaged position before the introduction of the transfer.

From a theoretical perspective, this could first suggest that parental preferences favor the allocation of resources toward girls. It could also be that the current opportunity cost of boys' time, despite the CT, is perceived as being higher than the future benefit of human capital accumulation, and that this difference for boys exceeds that of girls. If households rely more on boys for sustaining current agricultural incomes (which is suggested in other work in Lesotho), it is plausible that the size of the transfer was not large enough to increase secondary school–aged boys' participation in schooling but was sufficiently large to incentivize girls to attend school.

Second, we found that in *de jure* unmarried FHHs, the CGP improved schooling outcomes for secondary school–aged boys. For girls in this cohort, the treatment impacts were not as strong as among boys. We hypothesize that heads of FHHs, where a larger proportion of the children are grandchildren as opposed to sons or daughters of the head, may respond differently when accessing additional income through the CT and may have different preferences on gender and child education. One hypothesis is potential bias toward males reflected by positive impacts in boys' education vis-à-vis future prospects. Given the *ex ante* disadvantages for boys, this is a positive outcome of the program. Adding to this, investment in agricultural activities by FHHs—brought about by a reduction in liquidity constraints—may have led to an increase in adult agricultural labor participation. In this situation, girls' time may still be required for household chores.

In MHHs, girls benefited much more in terms of schooling and time use outcomes than did boys in the treatment evaluation. This would suggest that in MHHs, where a larger proportion of secondary school–aged boys engaged in crop production and livestock activities, the opportunity cost of boys' time may be still too high relative to that of girls, despite their access to the transfer.

Finally, much of the empirical and theoretical literature supports targeting women as transfer recipients to improve household well-being, such as children's health and educational outcomes. Our research showing that women spent in more "family-friendly" ways is primarily based on the assumption that women systematically differ from men in their preferences for types of expenditures or for the welfare of particular family members. However, when transfers are allocated to women rather than to men in a household, other factors beyond preferences and incentives determine whether differences in outcomes related to well-being will actually be realized (Yoong, Rabinovich, and Diepeveen 2012). These include, for example, differences in bargaining power over allocation of resources, under the assumption that intrahousehold bargaining is not fully cooperative, or differences in income-generating ability, for which women may face many other constraints (such as social restrictions on occupational type, or a relative lack of training) that result in lower returns for the transfer. Our analysis suggests that child investment, particularly for girls, may not be driven by the sex of the transfer recipient, contrary to previous literature. More plausibly, rather than male or female preference, it is the household structure and constraints that determine these differentiated effects.

For program design, our findings suggest that an undifferentiated CT for different types of households should at least include gender-specific messaging to promote boys' and girls' equal benefit in schooling. In addition, higher transfer levels and other mechanisms that could facilitate household access to agricultural labor would be required for children to be able to spend more time at school and increase their educational level.

Notes

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1 With the term "child labor," international organizations often define work that deprives children of their childhood,

their potential, and their dignity and that is harmful to their physical and mental development. Engagement of children in labor activities can be difficult and demanding, hazardous, and even morally reprehensible. With the available survey instrument used to collect the data for this study, it is impossible to disentangle the many kinds of work children do. For this reason, in this study, we use interchangeably such terms as child labor, child work, or engagement of children in family farming or wage labor.

2 The results for primary school–aged children (6–12) are not presented in this article, due in part to space constraints and to the lack of strong statistical significance in results for children in this age-group.

3 For more details about the identification process of the poorest households, see Pellerano et al. 2014.

4 The purpose of the survey was to track children. In some cases, the children of one household from baseline may have split into multiple households by the time of follow-up. (Additional details and discussion can be found in Pellerano et al. 2014.)

5 The results presented here do not use Propensity Score Matching (PSM) techniques, like reweighting for the propensity score, since impact estimates are virtually unchanged. This suggests that controlling for observables is sufficient to mitigate differences between control and treatment groups. Results are available from the authors on request.

6 A prior draft of this article included baseline statistics and all results for younger children aged 6–12. However, due to space limitations and the lack of strong results for this age-group, these results were not included in this version of the analysis. These results are available upon request from the authors.

7 Note that the "Girl" indicator and the "FemHead" indicator (from the subsequent equation), as well as the treatment indicator, were omitted due to individual fixed effects.

8 Controlling for time-varying household characteristics brings about only minor and negligible changes to impact estimates. Further, we did Hausman tests for overidentifying restrictions, and fixed-effects models were always preferred to random-effects models.

9 We cannot test this potential gender bias by the gender of cash recipients in FHHs, since we selected those households where women were *de jure* single. Hence, the overwhelming majority of women received the cash. There were few spare cases in which a man in these households received the transfer (a brother or a child), though the econometric analysis would lack statistical power.

10 The sample size in the heterogeneity analysis by the gender of the recipient is smaller than the overall MHH sample, as 10 percent of the data on sex of cash recipients are missing.

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