Can Conditional Cash Transfers Compensate for a Father's Absence?

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Abstract: This paper investigates how the departure of the father from the household, that results in his permanent absence, affects children's school enrollment and work participation in rural Colombia. The results show that departure of the father decreases children's school enrollment by around 5 percentage points and increases child labor by 3 percentage points. The paper exploits the roll-out of a conditional cash transfer program during the period of study and shows that it counteracts these adverse effects. This, and other pieces of evidence provided, strongly suggests that the channel through which the father's departure affects children is through reducing the income of very poor households, which is tightening their liquidity constraints. It also highlights the important safety net role played by such welfare programs, in particular for disadvantaged households, which are unlikely to find formal or informal ways of insuring themselves against such vagaries.

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A major disruption to family life can have serious consequences for children. A particularly traumatic event is the departure of the father from the household on a

permanent basis. There are at least three different channels through which this can affect children's human capital accumulation, and in particular their school and work participation (more discussion of the following points is to be found in Case, Paxson, and Ableidinger 2004 and Gertler, Levine, and Ames 2004). First, it is likely to involve a substantial income loss and this may be important for school choices in the presence of credit and insurance market failures. Second, the balance of decision-making power within the household may change, with the preferences of remaining adults gaining increased importance, which may have important consequences for children. Third, the loss of a parent can have significant emotional and psychological consequences for children. The importance of the first and third channels was highlighted in a World Bank Development Outreach report (Bell, Bruhns, and Gersbach 2006):

"if parents sicken and die while their children are still young, then all the means needed to raise the children so that they can become productive and capable citizens will be greatly reduced. The affected families' lifetime income will shrink, and hence also the means to finance the children's education, whether in the form of school fees or taxes. On a parent's death, moreover, the children will lose the love, knowledge and guidance which complement formal education."

Some countries, particularly in Africa, have put in place policies to provide education and health support to children who have lost one or both parents. These policies appear to be a response to the increase in HIV-associated mortality, which has resulted in millions of children losing parents to AIDS. Yet the absence of the father from the household whilst a child is still young is a pervasive phenomenon. Despite this, there is surprisingly little evidence on how children are affected by the long-term departure of

one or more parents (exceptions are referred to below) and on how policies may protect them against such adversities. In this paper, we first investigate how the departure of the father from the household - that results in his *permanent* absence - affects children's school enrollment and work participation in Colombia.[1] We are interested in the effects on children's school and work participation because of their importance for human capital accumulation; moreover, child work also affects family income and current poverty, which is indeed the reason why we may expect it to increase in order to compensate for income reductions. We then exploit the roll-out of a conditional cash transfer (CCT) program, *Familias en Acción*, the purpose of which is to increase schooling of children from poor backgrounds, and examine the extent to which receiving the cash transfer program mitigates the adverse effects we find.

Departure of the father is a relatively rare occurrence amongst our households. In order to focus on more permanent reductions in income, which are more difficult to insure against than transitory ones, we consider only departures due to death and divorce, which we can be confident are permanent. A central concern is that divorce or widowhood is not exogenous with respect to other determinants of child outcomes (see van de Walle 2011 for related selection issues). Previous work has attempted to exploit exogenous variation to overcome this problem, for instance in divorce laws (Gruber 2004) and child sex composition (Dahl and Moretti 2008). In this paper, we provide several pieces of evidence which, taken together, help build confidence in the quasi-random nature of the departure of the father. First, we show that observable characteristics (*before* the departure happened) of households in which the father did and did not subsequently depart are quite similar. Although reassuring, the concern remains

that unobserved heterogeneity may differ between these two types of households. We deal with time-invariant unobserved heterogeneity by allowing for household fixed effects in a three-year panel of households. So, in line with related literature (for instance De Janvry et al. 2006), our empirical method assumes common trends across both types of household. Note that this is conditional on a set of covariates, including transitory income shocks, making it more credible. We assess the plausibility of this common trends assumption by looking at pre-departure trends in children's schooling and per capita income, across households where the father does and does not subsequently depart, and are reassured by the fact that they do not differ significantly. To deal with endogeneity concerns due to potential correlation with time-varying shocks, we check whether divorce is correlated with recent significant time-varying shocks, including crop losses, business losses, and illnesses, and find that it is not. Finally, to further build confidence in the quasi-random nature of the departure, we carry out a falsification exercise by checking whether current child activities are correlated with future departure of the head: the idea here is that *future* departure should not lead to a significant effect on *current* activities if departure is effectively quasi-random. We find, reassuringly, no evidence that it does.

Our main finding is that, in the setting we consider in Colombia, the father's permanent absence from the household affects adversely the schooling of both boys and girls, and it increases their participation in paid and unpaid work. These findings are particularly pronounced for the relatively less well-off, who are likely to face the more severe liquidity constraints, and are consistent with the father's absence affecting activities through the income reduction associated with it. A second key finding of the

paper is that the conditional cash transfer program *Familias en Acción* helps protect children against the vagaries of the event: it protects their schooling and offsets the increased child labor after the father's departure. The fact that the CCT program acts as a safety net suggests that the main impact of father's departure is through the income loss associated with it.

The paper is structured as follows. In section I, we provide a brief overview of the related literature. Section II describes the data that we use in this research. We discuss identification issues in section III and present the empirical methodology and main results in section IV. Section V considers whether the CCT program introduced in the environment we consider has cushioned the poor households in our sample against these effects and section VI concludes.

I. Related Literature

Our work fits into a number of strands of literature. First, it is related to the growing literature in developing countries on parental deaths and children's education. This literature investigates the importance of different channels in explaining the observed impacts (Case, Paxson, and Ableidinger 2004; Gertler, Levine, and Ames 2004; Yamano and Jayne 2005; Beegle, De Weerdt, and Dercon 2006, 2010; Evans and Miguel 2007; van de Walle 2011). It generally finds adverse effects on schooling, particularly on primary school participation. This literature generally does not consider the effects on child labor, however - clearly an important economic activity amongst children in developing countries and one which may be particularly responsive to an event that induces a substantial income reduction. We consider permanent departure of the father

through either death or divorce, which are the two events in our data resulting in substantial income reductions for households. Whilst the channels through which both may affect outcomes may differ,[2] our empirical work suggests that it is the income reduction entailed that is the main driver of observed effects.

More generally, our work is related to the literature that considers the causal impacts of a family disruption on child outcomes, mainly covering the US or other developed countries. Whilst the literature has mainly focused on a permanent absence of a parent due to either death or divorce, recent work has also highlighted negative effects of parental migration on children's schooling, in particular that of the father since it is males who migrate in most contexts (see Antman 2012 for a survey, Giannelli and Mangiavacchi 2010, and Lahaie et al. 2009). A key recognition throughout this literature is that the absence of a parent is likely to be correlated with unobserved factors, which may also explain the poorer outcomes of the children. Several methods have been used to account for the influence of such factors. For example, some have used sibling-difference (household fixed effects) models (Case, Lin, and McLanahan 2001; Ermisch and Francesconi 2001; Ginther and Pollak 2004; Gennetian 2005) taking account of the fixed unobservable endowments that are shared by siblings from the same family or the same mother; others compare children's outcomes before and after the divorce of their parents (Cherlin et al. 1991; Painter and Levine, 2000), assuming that pre-existing disadvantages of the family or the child are captured by child fixed effects. Finally, quasi-experimental studies have either considered parental death as an exogenous source of parental absence (Biblarz and Gottainer 2000; Corak 2001; Lang and Zagorsky, 2001) or exploited

exogenous variation in separation rates due to differences in divorce laws either across states (Gruber 2004) or over time (Piketty 2003).

Our work also fits into the literature that considers the relationship between children's work participation and negative income shocks in developing countries, such as labor market shocks (Parker and Skoufias 2006) and/or crop losses (Jacoby and Skoufias 1997; Guarcello, Mealli, and Rosati 2003; Dehejia and Gatti 2005; Beegle, Dehejia, and Gatti 2006; Dammert 2007; Duryea, Lam, and Levison 2007; Gubert and Robilliard 2008). In line with this literature, our results are consistent with the presence of credit and insurance market failures in rural Colombia.

The second part of the paper, which provides evidence of CCT programs attenuating the negative income effects entailed by permanent absence of the father on children's activities, fits into a growing literature on the role of CCTs as safety nets. Indeed, CCT programs are a fast-growing part of safety net policy, and there is evidence that they provide households with protection against short-term shocks, both systemic and idiosyncratic. For instance, De Janvry et al. (2006) show that the Mexican PROGRESA program fully protected children's schooling from shocks due to unemployment and illness of the household head, as well as natural disasters in the community. Maluccio (2005) shows that the Nicaragua *Red de Protección Social* protected households' total and food expenses and children's school attendance against the effect of the Central America coffee crisis in 2000–01. More recently, Gitter, Manley, and Barham (2011) provide evidence of CCT programs mitigating the effects of negative shocks on physical development in early childhood. Our results are very much in line with these papers, suggesting that CCT programs provide a safety net against income

losses. A distinctive feature of our work is that we consider income losses that are likely to be permanent and that are thus even more difficult to insure against than transitory reductions in income.

II. DATA

In this section, we discuss the data used in the paper and present some key descriptive statistics relating to our sample of interest.

Background

We use three years of panel data from a survey of households and individuals in rural Colombia. These data have been collected to evaluate the large-scale welfare program *Familias en Acción*, which has been in place in some rural areas of Colombia since 2002 and which has since expanded to cover urban areas. The program aims to alleviate poverty by fostering human capital accumulation among the poorest households through conditional subsidies for investments in education, nutrition, and health.

The first wave of data collection for the evaluation of the program took place in 2002, when around 11,500 households were interviewed. We refer to this as the baseline survey. A year later, after the program started, a second wave of data was collected, and a third wave was collected in 2006. We refer to these as the first and second follow-up surveys respectively. In this paper, we estimate the effects of the father's permanent absence on children's outcomes.[3] The socio-economic data are rich, reflecting face-to-face interviews that lasted on average 3.5 hours.

Descriptive Statistics

We follow the school and work status of the children in households with at least one child aged 7–14 at the baseline across the first follow-up survey (1 year later) and the second follow-up (3.5 years after the baseline), up until they are at most 17 years of age. As we are considering the effects of departure of the father since baseline, we restrict the sample to households in which both parents are present at baseline.[4]

Outcomes. We consider two outcomes - school enrollment, which relates to whether the individual is enrolled in school at the time of the survey, and work participation, which includes all types of paid and unpaid economic activities, as well as looking for work as a main activity.[5] Looking at the proportions of our sample enrolled in school and participating in work, by age and gender (table 1), we see that school participation rates are high amongst children aged 7–11, corresponding to primary school.[6] The first substantial drop in school enrollment is observed at age 12, at the transition from primary to secondary school. Another point worth noting is that school enrollment of females is higher than that of males. Engagement in work is around twice as high for males as for females, and is very low for both before the age of 12 (participation in work is not recorded for individuals under age 10).

{Table 1 about here}

Permanent absence of the father. To capture a potentially very important disruption to family life and a long-term reduction in income, we focus on the departure of the child's father from the household since the baseline, and in particular departure that results in his subsequent *permanent* absence from the household.[7] Divorce and death are the two reasons for permanent absences that are identifiable from the data. As they

are relatively rare events, we pool them in order to improve statistical precision.[8] One might be concerned that they result in different levels of transfers to the household; however, we checked this in our data and found that the amount of transfers received by the household is very similar in magnitude after death and divorce (table A1 in the appendix).

To measure the incidence of divorce, we combine information on marital status of the child's mother at times *t*–1 and *t*, and the status of the father at time *t*. In particular, if her marital status at time *t* is divorced and at time *t*–1 is married, and if his status at time *t* is 'no longer in the household', we consider this to be a divorce. Deaths, on the other hand, are coded directly in the survey. Departure of the father due to death or divorce has occurred in 5.6% of our sample of households (i.e. those with at least one 7- to 14-year-old at baseline). Divorce accounts for 82% of such departures and death for 18% of them.[9]

The average age of fathers who leave the household is 43 at baseline, and it results in a substantial income reduction: 90% of them were working at the baseline. To give some idea as to the extent of the income loss associated with the departure, we compare total household labor earnings across households with and without an absent father. Total labor earnings are around 22% lower, controlling for household composition (number of male adults, number of female adults, number of children aged 0–6, and number of children aged 7–17).[10] We also compare total household consumption, a more direct indicator of welfare of the households in our sample. It is lower by around 13% in households in which the father subsequently departed than in households in which he did

not, controlling for household composition as above. Both differences are statistically significant at the 1% level.

Whether such events can be fully anticipated or not, it is unlikely that the households in our sample have ways of fully insuring against the income losses they entail, in particular as they live in rural municipalities where credit and insurance markets are typically thin (Edmonds 2006). We also checked in our data that, although monetary transfers and transfers-in-kind increase significantly after the departure of the father, the magnitude is very small compared with the income losses it entails.[11] Under these conditions, we expect paternal absence to affect the decisions to send children to school and/or work. In addition to this, paternal absence is likely to have a number of other important repercussions (see for example Gertler, Levine, and Ames 2004 for a discussion of these). First, the father is likely to be one of the key decision-makers in the household, so his departure may bring about changes in bargaining power and decision-making within the household, which may affect children's education and work. Second, the father can be an important figurehead for children. Though we cannot disentangle these channels with the available data, we note in anticipation of our results that the evidence we find is strongly consistent with income loss being the key factor affecting children's activities.

III. IDENTIFICATION

Two issues that arise in identification relate to the potential endogeneity of parental absence and attrition from the sample over time. In this section, we discuss each of these issues in turn.

Endogeneity

An important concern with paternal absence, and indeed one that has received much attention in the related literature (see for example Gruber 2004), is that it may not be exogenous to the outcomes of interest - children's work and schooling. For instance, couples may split up due to having different preferences over investment in children, in which case we may be picking up the effects of preferences rather than divorce per se.[12] Whilst it is reassuring that pre-departure (i.e. baseline) observed characteristics of households that do and do not go on to experience departure of the father are quite similar, as shown in table 2 (differences are mainly in relation to education of the head of household), it is clearly important to address endogeneity concerns.

{Table 2 about here}

In the empirical work, we deal with these concerns in two ways. First, we control for time-invariant unobserved confounding factors through household fixed effects. Accordingly, our identification strategy relies on the assumption that the time trends are the same in households where the father does and does not depart, the plausibility of which we look at in detail below. Second, to address the concern that there may be time-varying factors correlated with father's departure and child outcomes, we assess the relationship between divorce and observed time-varying shocks, including crop losses, business losses, and illnesses, and also control for such time-varying shocks in the analysis to improve the conditional exogeneity of paternal departure.

To examine the plausibility of the common trends assumption, we look at trends in two key variables. First, we look at whether trends in children's schooling were the same in both types of household *before* the father departed. We have two periods of

school enrollment data before the departure - at baseline (2002) and the year before (collected retrospectively at baseline). We cannot reject that schooling trends are the same in both types of household, as shown by the statistically insignificant coefficient on the interaction between the type of household ("Absence") and the year dummy in the upper panel of table 3. As a second check for common trends, we compare trends in household per capita income in both types of household, before the father departed. We have three periods of income data before the departure, all collected retrospectively at baseline. The evolution of per capita household labor income in the years 1999, 2000, and 2001 is very similar across both types of household prior to the departure (lower panel of table 3). This gives us no reason to believe they would not have been so if departure had not occurred.

{Table 3 about here}

It is also worth noting that the signs of the estimates in table 3 point, if anything, to a positive selection of the families with absent fathers in terms of schooling and income trends (although they are not statistically significant).[13] In this worst-case scenario, we would in fact underestimate the magnitude of the "true" impacts of departure: to pre-empt our main results, the detrimental impacts of the permanent absence of the father on school enrollment and work would be even more pronounced than the ones we find.

As an additional exercise to build more confidence in the quasi-random nature of the absence of the father, we check whether current child activities are correlated with future absence of the father: *future* absence should not lead to a significant effect on *current* activities if departure is effectively random. To do this, we regress current

children's activities (schooling/work at time *t*, for $t = 1,2$) on future absence (at time $t+1$) and, reassuringly, find extremely small and statistically insignificant correlations between them (0.009 for schooling, –0.009 for work, both with p-values of 0.5).

Whilst all of the evidence above is reassuring, it does not address concerns that there may be unobserved time-varying shocks affecting both the father's permanent departure and child outcomes. For instance, a temporary shock to income such as a crop/business loss or illness may affect the quality of the marital relationship and the likelihood of divorce, as well as affecting child outcomes. We can gauge the importance of this to some extent, as households report the most important shocks in the year prior to the survey, including crop loss, illness, and business loss: when we check whether such shocks in period $t-1$ are correlated with divorce in period *t*, we find that they are not (table 4), which is reassuring. We also note that we control for these shocks in the empirical work and our point estimates of the main coefficient of interest are very similar with and without them.

{Table 4 about here}

Taken together, the above evidence helps build confidence in the quasi-random nature of father's absence. We also reiterate that we control for time-invariant unobserved household-level characteristics and time-varying observed ones (including, importantly, shocks) throughout the empirical analysis.

Attrition

Overall, around 5% of households left the sample between the baseline survey and the first follow-up and an additional 8.5% of households left between the first and

second follow-ups (3.5 years after baseline).[14] Although this attrition rate is relatively low,[15] it is a concern if the reason for leaving the sample is related to the behavior being modeled, as might be the case if households from which the father departs are more likely to drop out of the sample. To address this, we compare baseline characteristics of households that did and did not subsequently leave the sample (table 5). As expected, households that own a house are significantly less likely to attrit than those that do not; and those living at relatively high altitudes are more likely to attrit. Other than that, attrition is not systematically related to any of the variables considered in the table. Whilst this is reassuring, potential selection biases on the basis of unobserved characteristics cannot be ruled out, which we account for in our empirical work. The methods we use to correct for this are discussed in section IV and all results presented take into account this possible selection problem, although it makes little difference to the effects we estimate.

{Table 5 about here}

IV. Effects of Permanent Absence of the Father on Schooling and Work

In this section, we present the empirical specification used to estimate the effects of permanent absence of the father on children's schooling and work outcomes. We then show the empirical findings.

Main Specification

To estimate the effects of the permanent absence of the father on children's school and work participation, we estimate the following model:

$$(1) \qquad y_{ijt} = \alpha_1 + \alpha_2 D_{jt} + X'_{ijt}\alpha_3 + I'_{jt-1}\alpha_4 + f_j + \delta_t + u_{ijt}$$

where i denotes child, j denotes household, and t denotes time, $t = 1$ (baseline), 2 (first follow-up), and 3 (second follow-up), y_{ijt} is a discrete indicator for participation in school or work, and D_{jt} is an indicator that takes the value 1 if the father is absent from the household permanently and 0 otherwise. Note that by definition, $D_{j1} = 0$.[16] If the father departed the household between baseline and first follow-up, then $D_{j2} = 1$ and $D_{j3} = 1$; if the father departed between first and second follow-ups, then $D_{j2} = 0$ and $D_{j3} = 1$. X_{ijt} is a vector of observed time-varying child and household characteristics including a cubic in the age of the child, number of siblings of different age categories (0–6, 7–12, 13–17, 18+), I_{jt-1} is a vector of time-varying shocks that occurred in the year prior to the survey, including dummies for crop losses, business losses, and illnesses, f_j is a household fixed effect capturing the effects of unobserved time-invariant household characteristics, δ_t is a survey round dummy, and u_{ijt} is an error term that we assume to be *iid*. The coefficient of interest is α_2, the effect of absence of the father on the outcome (school or work participation).

We estimate equation (1) using a linear probability model (LPM) and cluster the standard errors at the municipality level to adjust for potential correlations of household decisions within the same municipalities. Although the dependent variable is discrete, in our case the main advantage of the linear model over discrete choice models is that it is considerably easier to incorporate fixed effects. Another point to note is that, in our application, most of the explanatory variables are discrete and take on only a few values, strengthening the case for the LPM (Wooldridge 2002, chapter 15). Though a potential limitation of the LPM is that it can yield predicted probabilities outside the unit interval, in our case this is not a big concern as less than 3% of predictions lie outside the unit

interval. Note also that we checked for robustness of our results to this linear specification, by estimating a fixed effects logit model (Honoré 2002). The estimates, though less precisely estimated as they are based on the subset of children who changed their activity over time, point to the same patterns of coefficients as are discussed in the main text on the basis of LPMs and are shown in table A2 in the appendix.

As discussed in section III, an important issue is that our variable of interest, father's permanent absence, may be correlated with unobserved household characteristics that have a direct effect on children's schooling and work. To net out the effects of unobserved characteristics that are fixed over time and may lead to spurious correlations between father's permanent absence and children's outcomes, we use a household fixed effects model. We also control for important time-varying shocks to mitigate concerns that shocks may be determining both the paternal absence and the child's outcomes.

A second issue, also discussed in section III, is that non-random attrition, if present, will yield inconsistent parameter estimates. We use a standard correction in a two-step sample selection model (Heckman 1979) and estimate the probability that the individual does not leave the survey using a probit model:

$$(2) \qquad \Pr(S_{ijt} = 1) = \beta_1 + \beta_2 Z_{jt-1} + X'_{ijt-1}\beta_3 + \eta_j + \delta_t + v_{ijt}$$

where S_{ijt} takes the value 1 if child i from household j does not leave the survey between wave t–1 and wave t, and 0 otherwise, Z_{jt-1} are the instruments used for identification, discussed below, X_{ijt-1} are individual and household characteristics at wave t–1, δ_t is a survey round dummy, η_i is a household-level fixed effect, which may be correlated with f_j in equation (1), and v_{ijt} is an error term.

The instrument set Z_{jt-1} includes characteristics of the previous interview - its date (day of the month) and whether the survey respondent was the household head or spouse. Both may affect the overall experience of the interview and thus willingness to be re-interviewed but are unlikely to affect the outcomes of interest since they relate to the previous interview, which took place at least a year earlier.[17] The estimates from equation (2) are shown in table A3 in the appendix. The instruments are jointly statistically significant at the 1% level. We use these estimates to construct the inverse Mills ratio, which is appended to the set of control variables in equation (1). The selection correction term turns out not to be statistically significant at conventional levels in most cases and the estimates change very little when it is included in equation (1). Nonetheless, all reported results take into account this selection correction.

Results

We next turn to the estimates from our equation of interest, equation (1), which are shown in table 6. As we observe work (schooling) for children aged 10 (7) and above (see section II), we include an additional column containing estimates for schooling for the subsample aged 10 and above, to be able to make meaningful comparisons between the estimates for work and schooling outcomes.[18] We see from column 3 that the permanent absence of the father from the household significantly increases participation in work, by around 3 percentage points. Interestingly, we see from column 2 that the increase in work comes entirely from schooling (and not leisure) since the absence of the father has a significant negative effect on school enrollment, of close to 5 percentage points.[19] Note that the effects on schooling for the full sample, shown in column 1, are

very similar to those for the restricted sample. The estimated effects are not significantly different by gender (columns 4–6).

{Table 6 about here}

An important reason why these negative effects on schooling and positive effects on work may be expected, discussed in section II, is that households in which the father left permanently incur a substantial income reduction. To investigate the extent to which the income loss associated with the absence of the father underlies the estimated impacts, we interact it with education of the head (as at baseline, i.e. pre-departure), a proxy for household income. On the one hand, households with relatively low-educated heads have less to lose from a departure through an "income effect".[20] On the other hand, the relatively less well-off are more likely to face credit constraints and insurance market failures, and to have fewer formal ways to mitigate the impacts of income losses, such that they are likely to suffer more from father's absence. Accordingly, the interaction effect can go in both directions and we test it empirically in columns 7–9. We see that the detrimental effects of father's absence on schooling and child labor are driven by relatively less well-educated households. This highlights the importance of liquidity constraints for these households, which dominates the effect entailed by their relatively lower loss of income in the case of departure.

Finally, we checked whether the effects vary depending on the reason for the father's absence, by allowing the effects of death and divorce to be different. A caveat is that the incidence of death is very low, affecting just 1% of our sample of households (compared with 4.6% for divorce), resulting in its effects being imprecisely estimated. The results (available upon request) show that the impacts appear driven mainly by

divorce, though we cannot reject that the coefficient estimates are statistically the same. In what follows, we continue to pool these events as we are interested in events that we are fairly confident induce permanent income reductions. Another reason for combining them is to maintain statistical power given the rarity of the events.

V. Do Conditional Cash Transfers Help Protect Children?

In this section we investigate whether the effects of the father's permanent absence on children's outcomes differ depending on whether or not a CCT program is in place. We start off by describing the CCT program. We then go on to check whether the absence of the father due to divorce has itself been affected by the CCTs, and find no evidence that it has. We then estimate whether the CCTs mitigate the adverse effects of permanent absence of the father.

The CCT Program

In order to evaluate the impacts of the *Familias en Acción* CCT program, a representative stratified sample of municipalities was selected; strata were defined in terms of region and an index of infrastructure relating to health and education. Some municipalities from the same strata that were excluded from receiving the CCTs, but that were as similar as possible to eligible municipalities in terms of population, area, and an index of quality of life, were chosen as controls.[21] A total of 122 municipalities were chosen for the evaluation, of which 70, "eligible for CCTs", received the CCTs, which were phased in during the period we are considering: 26 received CCTs by the time of the baseline survey ("early-treat"), 31 by first follow-up ("mid-treat"), and 13 by second follow-up ("late-treat"). The final evaluation sample comprised approximately 100

households randomly selected in each of these 122 municipalities. Attanasio et al. (2010) provide an evaluation of the program's main impacts.

The CCT Program and Divorce

Before studying the interaction between the CCTs and permanent absence of the father, we address the potential concern that absence of the father - particularly in the case of divorce - may itself be affected by the CCTs. Indeed, there is direct evidence of positive effects of the PROGRESA CCTs on divorce in Mexico (Bobonis 2011) and indirect evidence that the *Familias en Acción* CCTs may have increased women's bargaining power (Attanasio, Battistin, and Mesnard 2012). Given this, one might expect women receiving CCTs to transit more readily out of relationships.

To estimate the effect of the CCTs on divorce, we use data from the first and second follow-ups only (as there is no variation in the outcome, divorce, at baseline - see endnote 3).[22] We estimate the following regression at the household level on our sample of households:

$$(3) \qquad y_{jt} = \alpha_0 + \alpha_1 T_{jt} + X'_{jt}\alpha_2 + I'_{jt-1}\alpha_3 + f_j + \delta_t + u_{jt}$$

pooling $t = 2$ and $t = 3$, where y_{jt} is a dummy variable indicating whether the parents living in household j divorced between periods t–1 and t, and T_{jt} is an indicator equal to 1 if household j lives in a municipality that is receiving CCTs at time t and 0 otherwise. Note that $T_{jt} = 1$ for {(early-treat = 1 or mid-treat = 1) and $t = 2,3$} and for {late-treat = 1 and $t = 3$}. X_{jt} are time-varying measures of the composition of children in the household in period t, I_{jt-1} is a vector of dummies indicating whether the household experienced a

crop loss, business loss, or illness in period t–1, f_j is a household fixed effect, δ_t is a survey round dummy, and u_{jt} is an error term.

We cannot reject the hypothesis that the cash transfers have had no statistically significant effect on divorce (table 7).[23] It is thus unlikely that such an effect underlies the results we discuss next, which show that receiving CCTs compensates for a father's absence.

{Table 7 about here}

Interaction Effects

There is a growing literature on the safety net role played by CCTs in the presence of income shocks but, to our knowledge, no work has been done studying the case of risk entailed by permanent loss of income. To investigate whether the effects of the father's permanent absence on children's outcomes differ depending on whether or not CCTs are in place, we augment equation (1) to include an interaction between our variable of interest, father's permanent absence, and receiving the CCTs. We thus estimate the following model:

$$(4) \qquad y_{ijt} = \alpha_0 + \alpha_1 D_{jt} + \alpha_2 D_{jt} * T_{jt} + \alpha_3 T_{jt} + X'_{ijt}\alpha_5 + I'_{jt-1}\alpha_6 + f_j + \delta_t + u_{ijt}$$

where T_{jt} is equal to 1 if household j lives in a municipality that is receiving CCTs at time t and 0 otherwise, and all other notation is as defined in equation (1). As before, T_{jt} reflects the gradual roll-out of the program, so $T_{jt} = 1$ for {early-treat = 1 and t = 1,2,3}, {mid-treat = 1 and t = 2,3}, and {late-treat = 1 and t = 3}.

In equation (4), the coefficient of interest, α_2, measures the extent to which receiving CCTs mitigates the effect of the permanent absence of the father, α_1.[24] Note that the above specification also implicitly controls for pre-program differences in

outcomes across municipalities that are and are not eligible for the CCTs (through fixed effects), which is potentially important given the quasi-experimental setting.

In municipalities not receiving the CCTs, the permanent absence of the father reduces school enrollment and increases child labor, particularly amongst the relatively less-educated households (left-hand columns of table 8): this is picked up by the coefficient α_1 displayed in the first row, which estimates the effect of departure in the absence of CCTs. Added to this, the second row, α_2, shows that when CCTs are in place, these adverse effects are offset (as shown by $\alpha_1+\alpha_2$ - which is close to zero and not significantly different from zero as shown by the p-values of the test).[25]

{Table 8 about here}

Finally, as a robustness check, we restrict the comparison to households living in municipalities eligible for the CCTs falling within the common support, i.e. the region over which treated individuals have a counterpart in the group of controls (according to the propensity score). In line with Attanasio et al. (2010), we do this by matching treatment and control observations using kernel-weighted propensity score matching, and imposing common support by dropping 10% of the treatment observations at which the propensity score density of the control observations is the lowest. The results are qualitatively similar and shown in table A4 of the appendix.

The fact that the welfare program provides insurance to protect the very poor children from the adverse consequences of a father's permanent absence is, perhaps, not very surprising to the extent that the CCTs received represent a sizeable share of income for these households - more than 20% of their monthly total consumption on average (see

Mesnard 2009) - and that the drop in household labor earnings entailed by father's departure is of a similar magnitude. Moreover, the welfare program is in place on a permanent basis, which gives some credence that the insurance it provides will continue as long as the child is enrolled in school. Interestingly, this result is somewhat distinct from that of De Janvry et al. (2006), who show that PROGRESA did not prevent children from working more following shocks due to unemployment and illness of the household head, as well as natural disasters in the community, though it fully protected their schooling.

Taken together, our results point towards the existence of credit and insurance market imperfections, with adverse implications for children, who play an important role in cushioning the household against the income losses entailed by departure of fathers. Whilst one cannot rule out the psychological impacts of a parent departing playing a role too, we believe they are of secondary importance to the income loss channel. In particular, we have no reason to believe that psychological impacts would be stronger amongst the less well-educated and they do not lend themselves easily to explaining why the CCTs would help mitigate such effects.

VI. CONCLUSION

This paper has investigated the link between the permanent absence of the father from the household and the school enrollment and work participation of children in rural Colombia. We find that absence of the father decreases schooling, by around 5 percentage points, and increases participation in work by around 3 percentage points. We provide evidence that these effects are mainly driven by households with relatively less-

educated heads, which, of the indigent households in our sample, are the very poorest. We show that receiving conditional cash transfers offsets these adverse consequences, offering the children a form of insurance when the father leaves the household for good. This also suggests that the income reduction associated with paternal absence, which is tightening liquidity constraints of already very poor households, is the main mechanism at play.

Our results have a number of important policy implications. First, they suggest that credit and insurance market failures are potentially important in the context of rural Colombia and can contribute to lower human capital accumulation of children. Second, an event such as the permanent departure of the father has potentially important consequences for the schooling and work of children, in particular those with relatively low levels of education, who are particularly vulnerable to permanent income losses given insurance market failures. Third, such adverse effects can be offset by well-designed conditional cash transfer programs targeted at very poor households, which, in the case of Colombia, represent on average more than 20% of total household consumption and are in place as long as the child is enrolled in school.

The last finding is the first of this kind, and offers an important agenda for future work. An important question is whether it also holds for investments other than schooling (such as children's health and nutrition) and in other contexts and environments. Another question is whether this should be taken into account in the design of safety nets and their targeting to lone parents, as it may also have the unintended consequence of promoting single parenthood. A final thought is on the particular relevance of these findings for sub-Saharan Africa, which has seen a dramatic rise in orphanhood due to the prevalence of

HIV/AIDS, with estimates suggesting 12% of all children are orphaned (UNICEF 2006). Families and communities have been sharing the burden of this, and it may be time for government support to be put in place to help households cope.

APPENDIX

{Tables A.1, A.2, A.3 & A.4 about here}

REFERENCES

Antman, Francisca M. 2012. "The Impact of Migration on Family Left Behind." Forthcoming in A. Constant and K.F. Zimmermann, eds., *The International Handbook on the Economics of Migration*. Edward Elgar, forthcoming 2013.

Attanasio, Orazio, Emla Fitzsimons, Ana Gomez, Maria I. Gutiérrez, Costas Meghir, and Alice Mesnard. 2010. "Child Education and Work Choices in the Presence of a Conditional Cash Transfer Program in Rural Colombia." *Economic Development and Cultural Change* 58 (2): 181–210.

———, Erich Battistin, and Alice Mesnard. 2012. "Food and Cash Transfers: Evidence from Colombia." *Economic Journal* 122 (559): 92–124.

Beegle, Kathleen, Rajeev Dehejia, and Roberta Gatti. 2006. "Child Labor and Agricultural Shocks." *Journal of Development Economics* 81 (1): 80–96.

———, Joachim De Weerdt, and Stefan Dercon. 2006. "Orphanhood and the Long-Run Impact on Children." *American Journal of Agricultural Economics* 88 (5): 1266–72.

———, Joachim De Weerdt, and Stefan Dercon. 2010. “Orphanhood and Human Capital Destruction: Is There Persistence into Adulthood?” *Demography*, 47 (1): 163–80.

Bell, Clive, Ramona Bruhns, and Hans Gersbach. 2006. “Economic Growth, Education, and AIDS in Kenya: A Long-Run Analysis.” World Bank Policy Research Working Paper 4025. Washington, D.C.: World Bank.

Biblarz, Timothy J., and Greg Gottainer. 2000. “Family Structure and Children’s Success: A Comparison of Widowed and Divorced Single-Mother Families.” *Journal of Marriage and Family* 62 (2): 533–48.

Bobonis, Gustavo J. 2011. “The Impact of Conditional Cash Transfers on Marriage and Divorce.” *Economic Development and Cultural Change* 59 (2): 281–312.

Case, Anne, I-Fen Lin, and Sara McLanahan. 2001. “Educational Attainment of Siblings in Stepfamilies.” *Evolution and Human Behavior* 22 (4): 269–89.

———, Christina Paxson, and Joseph Ableidinger. 2004. “Orphans in Africa: Parental Death, Poverty and School Enrolment.” *Demography* 41 (3): 483–508.

Cherlin, Andrew J., Frank F. Furstenberg, P.Lindsay Chase-Lansdale, Kathleen E. Kiernan, and others. 1991. “Longitudinal Studies of Effects of Divorce on Children in Great Britain and the United States.” *Science* 252 (5011): 1386–9.

Corak, Miles 2001. “Death and Divorce: The Long-Term Consequences of Parental Loss on Adolescents.” *Journal of Labor Economics* 19 (3): 682–715.

Dahl, Gordon, and Enrico Moretti. 2008. “The Demand for Sons.” *Review of Economic Studies* 75 (4): 1085–120.

Dammert, Ana. 2007. “Child Labor and Schooling Response to Changes in Cocoa Production in Rural Peru.” IZA Discussion Paper 2869. Bonn: IZA.

Dehejia, Rajeev, and Roberta Gatti. 2005. “Child Labor and Agricultural Shocks.” *Journal of Development Economics* 81 (1): 80–96.

De Janvry, Alain, Frederico Finan, Elizabeth Sadoulet, and Renos Vakis. 2006. “Can Conditional Cash Transfer Programs Serve as *Safety-Nets* in Keeping Children at School and from Working when Exposed to Shocks?” *Journal of Development Economics* 79 (2): 349–73.

Duryea, Suzanne, David Lam, and Deborah Levison. 2007. “Effects of Economic Shocks on Children’s Employment and Schooling in Brazil.” *Journal of Development Economics* 84 (1): 188–214.

Edmonds, Eric. 2006. “Child Labor and Schooling Responses to Anticipated Income in South Africa.” *Journal of Development Economics* 81 (2): 386–414.

Ermisch, John F., and Marco Francesconi. 2001. “Family Structure and Children’s Achievements.” *Journal of Population Economics* 14 (2): 249–70.

Evans, David, and Edward Miguel. 2007. “Orphans and Schooling in Africa: A Longitudinal Analysis.” *Demography* 44 (1): 35–57.

Fiszbein, Ariel, Norbert Schady, Francesco Ferreira, Margaret Grosh, Niall Keleher, Pedro Olinto, and Emmanuel Skoufias. 2009. *Conditional Cash Transfers: Reducing Present and Future Poverty*. Washington, D.C.: World Bank.

Gennetian, Lisa. 2005. “One or Two Parents? Half or Step Siblings? The Effect of Family Composition on Young Children.” *Journal of Population Economics* 18 (3): 415–36.

Gertler, Paul, David Levine, and Minnie Ames. 2004. “Schooling and Parental Death.” *Review of Economics and Statistics* 86 (1): 211–25.

Giannelli, Gianna, and Lucia Mangiavacchi. 2010. “Children’s Schooling and Parental Migration: Empirical Evidence on the ‘Left-Behind’ Generation in Albania.” *Labour* 24 (S1): 76–92.

Ginther, Donna, and Robert Pollak. 2004. “Family Structure and Children’s Educational Outcomes: Blended Families, Stylized Facts, and Descriptive Regressions.” *Demography* 41 (4): 671–96.

Gitter, Seth, James Manley, and Brad Barham. 2011. “The Coffee Crisis, Early Childhood Development, and Conditional Cash Transfers.” IDB Working Paper 245. Washington, D.C.: Inter-American Development Bank.

Gruber, Jonathan. 2004. “Is Making Divorce Easier Bad for Children? The Long-Run Implications of Unilateral Divorce.” *Journal of Labor Economics* 22 (4): 799–833.

Guarcello, Lorenzo, Fabrizia Mealli, and Furio Rosati. 2003. “Household Vulnerability and Child Labor: The Effects of Shocks, Credit Rationing, and Insurance.” In *Understanding Children’s Work*. Inter Agency Research Cooperation Project: ILO, UNICEF and World Bank Group.

Gubert, Flore, and Anne Sophie Robilliard. 2008. “Risk and Schooling Decisions in Rural Madagascar: A Panel Data-Analysis.” *Journal of African Economies* 17 (2): 207–38.

Heckman, James. 1979 “Sample selection bias as a specification error” *Econometrica: Journal of the econometric society*, 153-161.

Honoré, Bo E. 2002. “Nonlinear models with panel data” *Portuguese Economic Journal*, 1: 163–179.

Jacoby, Hanan, and Emmanuel Skoufias. 1997. “Risk, Financial Markets, and Human Capital in a Developing Country.” *Review of Economic Studies* 64 (3): 311–35.

Lahaie, Claudia, Jeffrey Hayes, Tinka Piper, and Jody Heymann. 2009. “Work and Family Divided across Borders: The Impact of Parental Migration on Mexican Children in Transnational Families.” *Community, Work and Family* 12 (3): 299–312.

Lang, Kevin, and Jay Zagorsky. 2001, “Does Growing Up with a Parent Absent Really Hurt?” *Journal of Human Resources* 36 (2): 253–73.

Maluccio, John. 2005. “Coping with the ‘Coffee Crisis’ in Central America: The Role of the Nicaraguan *Red de Protección Social*.” FCND Discussion Paper 188. Washington, D.C.: International Food Policy Research Institute (IFPRI).

Mesnard, Alice. 2009. “Migration, Violence and Welfare Programs in Rural Colombia.” IFS Working Paper 09/19. London: Institute for Fiscal Studies.

Painter, Gary, and David Levine. 2000. “Family Structure and Youths’ Outcomes: Which Correlations are Causal?” *Journal of Human Resources* 35 (3): 524–49.

Parker, Susan W., and Emmanuel Skoufias. 2006. “Job Loss and Family Adjustments in Work and Schooling during the Mexican Peso Crisis.” *Journal of Population Economics* 19 (1): 163–81.

Piketty, Thomas. 2003. “The Impact of Divorce on School Performance: Evidence from France, 1968–2002.” CEPR Discussion Paper 4146. London: Centre for Economic Policy Research.

Ravallion, Martin, and Quentin Wodon. 2000. “Does Child Labour Displace Schooling? Evidence on Behavioural Responses to an Enrollment Subsidy.” *Economic Journal* 110 (462): C158–75.

UNICEF. 2006. *Africa’s Orphaned and Vulnerable Generations: Children Affected by AIDS*. New York, NY: UNICEF.

Van de Walle, Dominique. 2011. “Lasting Welfare Effects of Widowhood in a Poor Country.” World Bank Policy Research Working Paper 5734. Washington, D.C.: World Bank.

Wooldridge, Jeffrey 2002. *Econometric Analysis of Cross Section and Panel Data*. Cambridge, MA: MIT Press.

Yamano, Takashi, and Thomas S. Jayne. 2005. “Working-Age Adult Mortality and Primary School Attendance in Rural Kenya.” *Economic Development and Cultural Change* 53 (3): 619–54.

TABLE 1. School and work participation, by age, survey and gender

	Males			Females		
Age at baseline	Baseline	First (+1 yr)	Second (+3.5 yrs)	Baseline	First (+1 yr)	Second (+3.5 yrs)
		School enrolment **%**			School enrolment **%**	
7	0.904	0.928	0.963	0.922	0.953	0.970
8	0.935	0.951	0.933	0.961	0.959	0.947
9	0.952	0.943	0.895	0.966	0.960	0.918
10	0.932	0.907	0.813	0.958	0.950	0.867
11	0.917	0.884	0.764	0.935	0.901	0.835
12	0.856	0.782	***0.675***	0.897	0.859	***0.786***
13	0.791	0.755	***0.577***	0.832	0.791	***0.633***
14	0.660	***0.620***	***0.457***	0.740	***0.728***	***0.536***
N	6090	5726	5033	5589	5266	4482
		Work participation **%**			Work participation %	
10	0.021	0.019	0.083	0.010	0.006	0.045
11	0.031	0.050	0.132	0.012	0.025	0.057
12	0.057	0.093	***0.209***	0.029	0.042	***0.117***
13	0.109	0.148	***0.284***	0.057	0.086	***0.143***
14	0.213	***0.285***	***0.371***	0.091	***0.169***	***0.203***
N	3672	4233	5022	3265	3870	4480

Notes: Work includes full-time paid and unpaid activities and look for work as a main activity. Figures in bold (italics) denote ages corresponding to post-compulsory schooling. Note that +1 yr (3.5 yrs) means 1 yr (3.5 yrs) after the baseline survey. N denotes the number of individuals (aged 7-14 at baseline) present in the survey listed at top of column. Schooling observed for children aged ≥7; work observed for children aged ≥10.

TABLE 2. Comparison of baseline (pre-departure) characteristics across households that do and do not experience subsequent departure

Characteristic, Baseline ↓	Permanent absence of father (D)		
	D=1	D=0	p-value
Age of household head	42.88	42.21	0.182
Age of spouse	37.81	37.14	0.110
Education of head			
None	0.282	0.230	**0.015**
Some (complete/incomplete primary)	0.535	0.638	**0.000**
High (incomplete secondary or more)	0.181	0.132	**0.004**
Education of spouse			
None	0.199	0.195	0.846
Some (complete/incomplete primary)	0.633	0.661	0.238
High (Incomplete secondary or more)	0.168	0.144	0.171
Household composition			
Ave # of kids ≤ 6	0.388	0.467	**0.011**
Ave # of boys 7-11	0.727	0.738	0.775
Ave # of girls 7-11	0.718	0.684	0.366
Ave # of boys 12-17	0.635	0.641	0.890
Ave # of girls 12-17	0.581	0.590	0.809
Ave # of female adults	1.232	1.244	0.708
Ave # of male adults	1.366	1.396	0.432
School enrolment rate of 7-14 yr olds in household	0.924	0.899	0.057
Household monthly consumption	421286	441994	0.085
Program area	0.700	0.682	0.453
Altitude	574.45	601.90	0.451
N	426	5720	

Notes: Sample consists of households where both parents are present at baseline and there is a 7-14 year old. N = number of households at baseline. P-values are based on standard errors clustered at the municipality level. Figures in bold in column (4) indicate that the figures in columns (2) and (3) are significantly from each other at the 5 per cent level or less.

TABLE 3. Common Trends: Schooling and Income

	School enrolment[A]
Year = 2002	0.0343**
	(0.005)
Absence * Year=2002	0.0238
	(0.0143)
N^1	11679

	Per capita income[B]
Year = 2000	0.498**
	(0.111)
Year = 2001	1.1019**
	(0.1432)
Absence * Year = 2000	-0.0845
	(0.4641)
Absence * Year = 2001	0.6144
	(0.6028)
N^2	5066

Notes: [A] Dependent variable is school enrolment. Estimates from household fixed effects model; also control for quadratic in child age, gender (female=1). Reference year = 2001. [B] Dependent variable is per capita household labour income. Estimates from household fixed effects model. Reference year = 1999. N^1 is the number of children in the sample at baseline with non-missing school enrolment data. N^2 is the number of households in the sample at baseline that report income retrospectively for 1999, 2000 and 2001. Standard errors, clustered at municipality level, in parentheses.

TABLE 4. Correlation between divorce and shocks in previous period

	Divorce $_{(t)}$
Crop loss $_{(t-1)}$	-0.0033
	(0.0045)
Business loss $_{(t-1)}$	0.0205
	(0.0181)
Illness $_{(t-1)}$	-0.0061
	(0.0071)
P-value for joint significance	0.55
N	5796

Notes: Dependent variable is divorce. Reference year=2001. Estimates from household fixed effects model pooling first and second follow-ups; also control for child composition and time dummies. N is the number of households remaining in the sample by first follow-up. Standard errors, clustered at the municipality level, in parentheses.

TABLE 5. Comparison of characteristics across households that do and do not attrit at any time after baseline

Baseline Characteristics	Did not attrit	Did attrit	p-value difference
Age of head	42.1123	42.3617	0.5057
Age of spouse	37.1413	37.4586	0.2987
Head no education	0.2309	0.2497	0.2263
Spouse no education	0.1927	0.2120	0.1889
Head some education	0.6321	0.6208	0.5251
Spouse some education	0.6579	0.6655	0.6668
Head high education	0.1365	0.1260	0.4047
Spouse high education	0.1493	0.1225	**0.0398**
Treated area	0.6844	0.6756	0.6064
Altitude	577.69	726.43	**0.0000**
Crop loss at first survey	0.1339	0.1249	0.4708
Owns house	0.6466	0.5321	**0.0000**
N	5289	857	

Notes: Sample consists of households where both parents are present at baseline and there is a 7-14 year old. N = number of households at baseline. P-values are based on standard errors clustered at the municipality level. Figures in bold in column (4) indicate that the figures in columns (2) and (3) are significantly from each other at the 5 per cent level or less.

TABLE 6. Marginal effects of the father's absence on children's schooling and work

	School Overall	School Restricted	Work Restricted	School Overall	School Restricted	Work Restricted	School Overall	School Restricted	Work Restricted
Permanent Absence	-0.0412*	-0.0484*	0.0301+	-0.0422+	-0.0486+	0.0361	-0.0556**	-0.0641**	0.0360+
	(0.0172)	(0.0216)	(0.0167)	(0.0233)	(0.0289)	(0.0233)	(0.0193)	(0.0234)	(0.0183)
Permanent Absence * Girl				0.0021	0.0005	-0.0122			
				(0.0241)	(0.0275)	(0.0245)			
Permanent Absence * High Educated Head							0.0742*	0.0851*	-0.0319
							(0.0313)	(0.0367)	(0.0295)
N	32186	24531	24531	32186	24531	24531	32186	24531	24531

Notes: Marginal effects from a fixed effects linear probability model reported (equation (1)). Also control for absence of father from household for unknown reason, absence of both parents, time dummies, cubic in child age, sibling composition, dummies for crop, illness and business shocks, inverse mills ratio computed as in equation (2) (see Table A3). High educated = 1 if incomplete secondary or more at baseline, 0 otherwise. N is the number of children in the sample pooled across three waves. 6146 households are in our initial sample, from which 426 fathers have subsequently departed. Schooling observed for all children in sample, i.e. ≥7 ('overall' sample); work observed for children ≥10 ('restricted' sample). Robust standard errors clustered at municipality level in parentheses. + significant at 10%; *significant at 5%; ** significant at 1%.

TABLE 7. Marginal Effects of CCTs on Divorce

	Pr (Divorce = 1)
CCTs	0.0024
	(0.0082)
Time = 2	-0.0286
	(0.0039)
N	5796

Notes: Marginal effects from equation from a fixed effects linear probability model reported (equation (3)). Pools first and second follow-ups. Also control for child composition, dummies for crop, illness and business shocks, inverse mills ratio computed as in equation (2) (see Table A3). N is the number of households that have not attrited by first follow-up. Robust standard errors clustered at municipality level in parentheses.

TABLE 8. Cushioning effects of CCTs: Marginal effects on schooling and work

	Low Ed			All		
	School	School	Work	School	School	Work
	Overall	Restricted		Overall	Restricted	
Permanent Absence (α_1)	-0.103**	-0.114*	0.0782*	-0.0801**	-0.0919*	0.0694*
	(0.0341)	(0.0437)	(0.0355)	(0.0284)	(0.0376)	(0.0300)
Permanent Absence * CCTs (α_2)	0.0818*	0.0833+	-0.0687+	0.0566+	0.0615	-0.0542+
	(0.0381)	(0.0472)	(0.0356)	(0.0330)	(0.0420)	(0.0307)
CCTs (α_3)	0.0143	0.0057	-0.0262**	0.0121	0.0053	-0.0257**
	(0.0123)	(0.0148)	(0.0096)	(0.0112)	(0.0134)	(0.00902)
Time = 2	0.1080	0.1520	-0.2010*	0.0783	0.1180	-0.1500+
	(0.0957)	(0.105)	(0.0847)	(0.0872)	(0.0964)	(0.0773)
Time = 3	0.0261	0.0689	-0.148+	-0.0025	0.0347	-0.1010
	(0.1010)	(0.1100)	(0.0889)	(0.0927)	(0.1020)	(0.0809)
P-value ($\alpha_1 + \alpha_2$)= 0	0.3498	0.2636	0.6224	0.2633	0.2379	0.3854
N	28027	21464	21464	32186	24531	24531

Notes: Marginal effects from a fixed effects linear probability model reported (equation (4)). Also control for absence of father for unknown reason, absence of both parents, cubic in child age, sibling composition, dummies for crop, illness and business shocks, inverse mills ratio computed as in equation (2) (see Table A3). CCTs indicates whether the household lives in a municipality that is receiving CCTs at time of survey. N is the number of children in the sample pooled across three waves. 6146 households are in initial sample, from which 426 fathers have subsequently departed. Schooling observed for all children in sample, i.e≥7 ('overall' sample); work observed for children ≥10 ('restricted' sample). Robust standard errors clustered at municipality level in parentheses. + significant at 10%; *significant at 5%; ** significant at 1%.

TABLE A1. Marginal effects of paternal death and divorce on transfers received by household

	Institutional	Monetary	In-Kind
Death	-1,317	16,809*	22,596
	(8680)	(7512)	(14788)
Divorce	1,512	17,704**	8,484
	(4294)	(4106)	(7967)
N		6,069	

Notes: N is number of households at baseline for which we observe transfers. Complete data on transfers missing for 77 of the sample of 6146 households. Pools baseline, first and second follow-ups. We trim the top 1% of outliers in each period. Each column represents a separate regression. Also control for household fixed effects, absence of father for unknown reason, absence of both parents, time dummies, household child composition, dummies for crop, illness and business shocks, inverse mills ratio computed as in equation (2) (see Table A3). Robust standard errors clustered at municipality level in parentheses. Robust standard errors clustered at municipality level in parentheses. + significant at 10%; *significant at 5%; ** significant at 1%.

TABLE A2. Marginal effects of the father's absence on children's schooling and work
Estimates from Conditional Logit Model

	School	School	Work	School	School	Work	School	School	Work
	Overall	Restricted		Overall	Restricted		Overall	Restricted	
Permanent Absence	-0.507*	-0.596*	0.550+	-0.488*	-0.579*	0.661+	-0.555*	-0.662**	0.515
	(0.223)	(0.248)	(0.309)	(0.245)	(0.277)	(0.349)	(0.231)	(0.250)	(0.317)
Permanent Absence * Girl				-0.0423	-0.0367	-0.241			
				(0.267)	(0.300)	(0.376)			
Permanent Absence * High Educated Head							0.555	0.711	0.471
							(0.847)	(0.855)	(1.007)
N^1	4284	3895	2536	4284	3895	2604	4284	3895	2604
N^2	1829	1661	1064	1829	1661	1075	1829	1661	1075

Notes: Marginal effects from a conditional logit model reported (equation (1)) with household fixed effects. Additional controls include control for absence of father for unknown reason, absence of both parents, time dummies, cubic in child age, sibling composition, dummies for crop, illness, and business shocks, inverse mills ratio computed as in equation (2) (see Table A3). High Educated is equal to 1 if incomplete secondary or more at baseline, 0 otherwise. N^1 (N^2) is the number of children (households) in the sample that (contain a child that) switch outcome status at least once, pooled across three waves. Non-switcher children drop out of the conditional likelihood function. Schooling observed for all children in sample, i.e≥7 ('overall' sample); work observed for children ≥10 ('restricted' sample) . Robust standard errors clustered at municipality level in parentheses. + significant at 10%; *significant at 5%; ** significant at 1%.

TABLE A3. Probability of not leaving the sample, marginal effects

	Dep vble=1 if stay in sample, 0 otherwise
Female	0.0057+
	(0.0031)
Time = 2	0.0367**
	(0.0071)
Owns house	0.0544*
	(0.0213)
Urban	0.0075
	(0.0069)
Day of month 1	-0.0478+
	(0.0288)
Day of month 2	-0.0385
	(0.0318)
Day of month 3	-0.0583+
	(0.0349)
Day of month 4	-0.0287
	(0.0309)
Day of month 5	-0.0011
	(0.0228)
Day of month 6	-0.0410
	(0.0265)
Day of month 7	-0.0050
	(0.0246)
Day of month 8	-0.0593+
	(0.0331)
Day of month 9	-0.0466
	(0.0351)
Day of month 10	-0.0130
	(0.0263)
Day of month 11	-0.0132
	(0.0227)
Day of month 12	-0.0771*
	(0.0392)
Day of month 13	-0.0645+
	(0.0378)
Day of month 14	-0.0581
	(0.0356)
Day of month 15	-0.0514
	(0.0374)
Day of month 16	-0.0826*
	(0.0361)

Day of month 17	-0.0311
	(0.0335)
Day of month 18	-0.0165
	(0.0290)
Day of month 19	-0.0524
	(0.0366)
Day of month 20	-0.0584
	(0.0414)
Day of month 21	-0.0471
	(0.0361)
Day of month 22	-0.0319
	(0.0323)
Day of month 23	-0.0429
	(0.0341)
Day of month 24	-0.0435
	(0.0322)
Day of month 25	-0.0194
	(0.0234)
Day of month 26	-0.0439
	(0.0364)
Day of month 27	-0.0456
	(0.0308)
Day of month 28	-0.0135
	(0.0249)
Day of month 29	-0.0034
	(0.0234)
Day of month 30	0.0078
	(0.0224)
Respondent = head	0.0132
	(0.0188)
Respondent = spouse	0.0597*
	(0.0272)
p-value of joint significance of instruments	0.0000
N	11679

Notes: N is the number of children in the sample at baseline with non-missing school enrolment data. Day of month = dummy variables for day baseline interview took place. Robust standard errors clustered at municipality level in parentheses. + significant at 10%; *significant at 5%; ** significant at 1%.

TABLE A4. Marginal effects of the father's absence on children's schooling and work, common support only

	School		Work
	Overall	Restricted	
Permanent Absence (α_1)	-0.102**	-0.106*	0.0682+
	(0.0373)	(0.0477)	(0.0367)
Permanent Absence * CCTs[1] (α_2)	0.0881*	0.0829	-0.0658+
	(0.0398)	(0.0506)	(0.0368)
CCTs[1] (α_3)	0.0145	0.006	-0.0224*
	(0.0125)	(0.0146)	(0.0103)
Time = 2	0.121	0.17	-0.222*
	(0.101)	(0.11)	(0.087)
Time = 3	0.0371	0.0856	-0.167+
	(0.106)	(0.114)	(0.091)
N	24982	18972	18972

Notes: See notes to Table 8. Note further that we match treatment and control observations using kernel-weighted propensity score matching, and impose common support by dropping 10% of the treatment observations at which the propensity score density of the control observations is the lowest.

NOTES

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1 Note that departure of the mother is also an important issue and may have different effects from those stressed in this paper. However, there is insufficient variation in the data to allow us to look at this.

2 An absent but living father can visit and influence the children's upbringing in a way that a deceased father obviously cannot. On the other hand, relations with the absent parent's family might also be very different in the two cases, perhaps more supportive in cases of early death of the father than in cases of acrimonious separation. Moreover, transfers from the father or in-laws may compensate in different ways depending on the

reason for departure. However, in our data, transfers are very similar regardless of whether departure is due to death or divorce, as table A1 later attests to.

3 We start with a sample of households where both parents are married so, by definition, fathers are all present at baseline.

4 This sample selection criterion means that we retain 9,187 out of 11,502 households. The reason we do not keep mono-parental households is that the departure of the father (if present) in such households would raise additional issues, which would be difficult to disentangle.

5 School enrollment is defined on the basis of whether the child is registered at school in the academic year corresponding to the survey. Work participation is equal to 1 if the child's main activity in the week before the survey is reported to be any of work, household chores (paid and unpaid), or looking for work. We note that our main results are similar if we exclude unpaid household chores.

6 The school system in Colombia operates as follows. Compulsory education is free and lasts for nine years; it consists of basic primary (*educación básica primaria*, five years, ages 7 through 11) and basic secondary (*educación básica secundaria*, four years, ages 12 through 15). The secondary school system also includes the middle secondary cycle (*educación media*, two years, ages 16 and 17). Successful completion of studies leads to the *Bachillerato*. Students must pass an entrance examination for access to universities.

7 Note that absent fathers are not being 'replaced' in households, at least in the 3.5-year span of our surveys: whilst the number of male adults is lower by almost 1 in

households that experience departure, the number of female adults is the same, as is the number of children.

8 We also checked that, considered separately, they do not have significantly different impacts (see section IV).

9 In an additional 1% of households, both the father and mother have left the household for an unknown, possibly temporary, reason; there is also a small percentage (1.2%) of households in which the father has left for an unknown reason, but the mother has remained in the household and reports being married, so we assume that these are temporary departures. These are not the main variables of interest but we control for them throughout the analysis.

10 If we do not control for adult composition, the difference is larger, at around 34%, which we would expect since departure of the father decreases the number of adults in the household. Further, we see this as a lower bound of the magnitude of the departure effect in terms of total household adult earnings, as it includes labor supply responses to it, which are likely to cushion the potential adverse effects on income. This figure excludes earnings from children to mitigate this problem.

11 Table A1 shows that the total value of additional transfers received by the households after the paternal departure (institutional, monetary, and in-kind) is less than 50,000 pesos per annum, compared with an average monthly total household consumption in excess of 420,000 pesos at baseline. Nonetheless, these responses by the extended family or friends may also contribute to explaining why household consumption does not drop by as much as household labor income, as noted earlier.

12 However, if this were the case, then we would rather expect to see it having a positive impact on children, whereas we in fact observe the contrary. It must also be acknowledged that departure of the father due to death may not be a random event, though this is much less of a concern.

13 In the lower panel of table 3, the coefficient associated with the interaction of absence of the father with year of survey 2000 which is negative, is very small in magnitude and has a very large standard error, so is of no concern.

14 Attrition at the individual level is extremely rare, at less than 1%.

15 It is comparable to the attrition rate of 6% between the baseline and follow-up surveys for the evaluation of the *Bono de Desarrollo* CCT program in Ecuador, which is considered "low", and just under the attrition rate of 15% over four years in Nicaragua for the evaluation of the *Red de Protección Social* CCT program, which is considered "reasonably low". It is slightly higher than the rate for the PROGRESA program, which was around 6% over the first three years of the program and considered to be "very low". (Fiszbein et al. 2009)

16 As discussed in section II, our sample is restricted to households in which the father is present at baseline. We only observe departures after baseline.

17 Attrition in our sample is predominantly at the household level. Moreover, very few households (3.7% in the entire sample) have migrated out of their village of residence and additional resources have been invested into tracking them (Mesnard 2009), so attrition is mostly due to non-willingness to answer.

18 We retain estimates for the full sample in order to improve statistical power.

19 This suggests that child labor and schooling are strong substitutes, in contrast to the finding of Ravallion and Wodon (2000) that increases in schooling in Bangladesh following a welfare program only partially come from decreased child labor.

20 Similarly, if paternal education is positively correlated with paternal quality as a figurehead / role model, then one would expect the loss of a high-educated father to involve the loss of a more positive impact on the child's life.

21 In order to be eligible to qualify for the program, municipalities had to satisfy four criteria: (i) have less than 100,000 inhabitants and not be a departmental capital, (ii) have basic education and health infrastructure, (iii) have a bank, and (iv) have relatively up-to-date welfare lists at the municipality administrative office. The evaluation design was carried out by a consortium led by the Institute for Fiscal Studies and that included the authors of this paper.

22 This also means that the identification of the effect of the CCTs on divorce comes from the roll-out of the program to late-treat areas at time 3.

23 As an additional check, we compared the characteristics of households that divorce, across areas eligible for the CCTs and control areas, and found them to be very similar for both types of area.

24 Note that due to the gradual phasing-in of the CCTs, the "early-treat" municipalities do not contribute to identifying α_3, the impact of the CCTs, as there are no pre-program data collected for these municipalities. However, we retain them in the analysis as they do contribute to identifying α_1 and α_2.

25 The table also shows that the effect of the CCTs on children in our sample who are not affected by paternal absence, given by α_3, is to increase school enrollment and reduce child labor. Although the effect on schooling is not significant, this is most likely due to the fact that the early-treat municipalities do not contribute to the identification of the CCT effect, unlike in Attanasio et al. (2010), which contains the general analysis of the impacts of the program.