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

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Abstract

This paper investigates how the permanent departure of the father from a household affects children's school enrollment and work participation in rural Colombia. The results indicate that the permanent departure of the father decreases children's school enrollment by approximately 5 percentage points and increases child labor by 3 percentage points. This paper explores the rollout of a conditional-cash-transfer program during the period of study and shows that this program counteracts these adverse effects. When coupled with other evidence, this finding strongly suggests that the channel through which the father's departure most affects children is by reducing the income of very poor households, which tightens their liquidity constraints. This finding also highlights the important safety-net role played by welfare programs with respect to disadvantaged households, particularly because these households are unlikely to have formal or informal mechanisms with which to insure 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 father's permanent departure from the household. There are at least three different channels through which this event can affect children's human capital accumulation, particularly their school and work participation rates (for further discussion of the following points, see Case, Paxson, and Ableidinger [2004] and Gertler, Levine, and Ames [2004]). First, this event is likely to involve substantial income loss that may particularly affect school choices coupled with credit and insurance market failures. Second, the balance of decision-making power in the household may change, with the preferences of the remaining adults gaining increased importance, which may have important consequences for children. Third, the loss of a parent may 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 established policies to provide education and healthcare support to children who have lost one or both parents. These policies may have been implemented in response to the increase in HIV-associated mortality, which has resulted in millions of children losing parents to AIDS. However, a

father who is absent from the household while a child is young is a pervasive global phenomenon, although there is surprisingly little evidence that shows how children are affected by the long-term departure of one or more parents and that recommends policies to protect children against such adversities. In this paper, we first investigate how the father's departure 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 the importance of these institutions for human capital accumulation. Moreover, child labor affects family income and current poverty, which is the reason we may expect it to increase to compensate for income reductions. We then explore the rollout of a conditional cash transfer (CCT) program, *Familias en Acción* (the purpose of which is to increase schooling for children from poor backgrounds), and we examine the extent to which the effects of the CCT program mitigate the adverse effects that are caused by the father's permanent absence.

The departure of the father from the family household is a relatively rare occurrence among households in rural Colombia. To focus on more permanent reductions in income, which are more difficult to insure against than transitory departures, 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 that, taken together, build confidence in the quasi-random nature of the father's departure. First, we

show that important observable characteristics of households (before the departure occurred) in which the father did and did not subsequently depart are similar. Although reassuring, the concern remains that unobserved heterogeneities might distinguish these types of households. Thus, we address time-invariant unobserved heterogeneity by allowing for household-fixed effects over a three-year period with a panel of households. Consistent with the related literature (see, for example, De Janvry et al. [2006]), our empirical method assumes common trends across both types of households. Note that this is conditional on a set of covariates, including transitory income shocks, which makes our approach more credible. We assess the plausibility of this assumption of common trends by examining predeparture trends in children's schooling and per capita income across households in which the father does and does not subsequently depart. We are reassured by the fact that these trends do not differ significantly. To address endogeneity concerns caused by potential correlation with time-varying shocks, we determine whether divorce is correlated with recent significant time-varying shocks, including crop losses, business losses, and illnesses. We find that it is not correlated in our sample. Finally, to build more confidence in the quasi-random nature of the departure, we perform a falsification exercise by determining whether current child activities are correlated with future departure of the father (that is, if departure is quasi-random, then future departure should not lead to a significant effect on current activities). Reassuringly, we find no evidence that it does.

In the setting we consider in Colombia, our main finding is that the father's permanent absence from the household adversely affects the schooling of both boys and girls and increases their participation in paid and unpaid work. These findings are

particularly pronounced for poorer households, which are likely to face more severe liquidity constraints, and are consistent with the father's absence affecting activities directly through the associated income reduction. A second key finding of the paper is that the CCT program *Familias en Acción* helps protect children against the vagaries of the event by protecting their schooling and offsetting increased child labor after the father's departure. Because the CCT program effectively acts as a safety net, it appears that the main impact of the father's departure is felt through its associated income loss.

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 into the environment we consider has cushioned poor households in our sample against these effects, and section VI concludes.

<<A>>*I. Related Literature*

Our work fits into several strands of the literature. First, it is related to the growing literature in developing countries on parental death and children's education. This literature investigates the importance of different channels in explaining the observed effects (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). The literature generally finds adverse effects on schooling, particularly on primary school participation. However, this literature generally does not consider effects on child labor; clearly, this is an important economic activity among

children in developing countries and one that may be particularly responsive to an event that induces a substantial income reduction. We consider the permanent departure of the father through either death or divorce, which are the two events in our data that consistently result in substantial income reduction for households. Although the channels through which these events might affect outcomes may differ,[2] our empirical work suggests that income reduction is the main driver of the observed effects.

More generally, our work is related to the literature that considers the causal effects of family disruption on child outcomes, which has concentrated on the United States and other developed countries. Although this literature has mainly focused on the permanent absence of a parent resulting from death or divorce, recent work has also highlighted the negative effects of parental migration on children's schooling—particularly the migration of the father because it is males who migrate in most contexts (see Antman 2013 for a survey; Giannelli and Mangiavacchi 2010; Lahaie et al. 2009). A key finding 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 studies have used sibling-difference (household fixed effects) models (Case, Lin, and McLanahan 2001; Ermisch and Francesconi 2001; Ginther and Pollak 2004; Gennetian 2005) that account for the fixed and unobservable endowments that are shared by siblings from the same family or the same mother, whereas others have compared children's outcomes before and after the divorce of their parents (Cherlin et al. 1991; Painter and Levine, 2000), which assumes that preexisting 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 based on differences in divorce laws across states (Gruber 2004) or over time (Piketty 2003).

Our work also dovetails with 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) or crop losses (Jacoby and Skoufias 1997; Dehejia and Gatti 2005; Beegle, Dehejia, and Gatti 2006; Dammert 2007; Duryea, Lam, and Levison 2007; Gubert and Robilliard 2008; Guarcello, Mealli, and Rosati 2010). As described in this strand of the 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 that the CCT programs attenuated the negative income effects entailed by the permanent absence of the father on children's activities, is consistent with the growing literature on the role of CCTs as safety nets. Indeed, CCT programs are a rapidly growing part of social safety-net policies, and there is evidence that they provide households with protection against both systemic and idiosyncratic short-term shocks. For instance, De Janvry et al. (2006) show that the Mexican PROGRESA program fully protected children's schooling from shocks caused by unemployment and illness of the head of the household and by natural disasters in the community. Maluccio (2005) shows that the Nicaraguan *Red de Protección Social* protected families' total expenses (including food expenses) and children's school attendance against the effects of the Central American coffee crisis in 2000–2001. More recently, Gitter, Manley, and Barham (2011) offer evidence that CCT programs have

mitigated the effects of negative shocks on physical development in early childhood. Our results are consistent with these papers, suggesting that CCT programs do provide a safety net against income losses for poor households. A distinctive feature of our work is that we consider income losses that are likely to be permanent and that are therefore even more difficult to insure against than transitory reductions in income.

<<A>>*II. DATA*

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

<<B>>*Background*

We use three years of panel data from a survey of households and individuals in rural Colombia. These data were collected to evaluate the large-scale welfare program *Familias en Acción*, which has been in place in several rural areas of Colombia since 2002 and 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 occurred in 2002, when approximately 11,500 households were interviewed. We refer to this as the baseline survey. A year later, after the program began, 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 socioeconomic data are rich and reflect face-to-face interviews that lasted 3.5 hours on average.

<<B>>*Descriptive Statistics*

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

<<C>>*Outcomes.* We consider two outcomes: school enrollment, which relates to whether the child is enrolled in school at the time of the survey, and work participation, which includes all types of paid and unpaid economic activities and includes looking for work as a main activity.[5] Table 1 shows the proportions of our sample enrolled in school and participating in work by age and gender (table 1). The table indicates that school participation rates are high among children aged 7–11 years, which corresponds to primary school.[6] The first substantial drop in school enrollment is observed at age 12, which is the transition from primary to secondary school. Another point worth noting is that the school enrollment of females is higher than that of males. Engagement in work is approximately 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 10 years of age).

<<C>>*Permanent absence of the father.* To capture a potentially important disruption to family life and long-term reduction in income, we focus on the departure of the father from the household since the baseline. In particular, we focus on departure that results in the father's subsequent permanent absence from the household.[7] Divorce and death are the two primary reasons for permanent absences that are identifiable from the data. Because these are relatively rare events, we pool them to improve statistical

precision.[8] There may be some concern that these events result in different levels of transfers to the household. However, in our data, we found that the amount of transfers received by the household is very similar in magnitude after death and divorce. This is shown in table S1 in the appendix.

To measure the incidence of divorce, we combine information on the 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 the mother's marital status at time t is divorced and her status at time $t - 1$ is married and if the father's status at time t is "no longer in the household", we consider this a divorce. Deaths, in contrast, are coded directly into the survey. The departure of the father as the result of death or divorce occurred in 5.6 percent of our sample of households (that is, those with at least one 7- to 14-year-old at baseline). Divorce accounts for 82 percent of such departures, and death accounts for 18 percent.[9]

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

not depart, after controlling for household composition as above. Both differences were statistically significant at the 1 percent level.

Regardless of whether such events can be fully anticipated, it is unlikely that the households in our sample have mechanisms to fully insure against the income losses they entail, particularly because these households are in rural municipalities in which credit and insurance markets are thin (Edmonds 2006). Our data confirm that although monetary transfers and in-kind transfers increase significantly after the departure of the father, the magnitude is small compared to the income loss the departure entails.[11] Under these conditions, we expect paternal absence to affect decisions about sending children to school or work. Furthermore, paternal absence is likely to have a number of other important repercussions (see, for example, Gertler, Levine, and Ames [2004] for a discussion). First, the father is likely to have been one of the key decision makers in the household, and his departure may bring about changes in bargaining power and decision making within the household that may affect decisions about children's education and work. Second, the father may be an important figurehead for children. Although we cannot disentangle these channels in the available data, in anticipation of our results, we note that the evidence is strongly consistent with income loss as the key factor that affects children's activities.

<<A>>*III. IDENTIFICATION*

Two issues that arise in identification are related 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.

<<B>>*Endogeneity*

An important concern about paternal absence, and one that has received considerable attention in the related literature (see, for example, Gruber [2004]), is that it may not be exogenous to the outcomes of interest (that is, children's work and schooling). For instance, couples may divorce because they have different preferences for investment in children, in which case we may be identifying the effects of preferences rather than divorce per se.[12] Although it is reassuring that the predeparture (that is, baseline) observed characteristics of households that do and do not experience the subsequent departure of the father are similar (the chief difference mainly relates to the education level of the head of household, as shown in table 2), it is important to address endogeneity concerns.

In our empirical work, we address 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 in which the father does and does not depart. We examine the plausibility of this assumption in detail below. Second, to address the concern that there may be time-varying factors correlated with the father's departure and child outcomes, we assess the relationship between divorce and observed time-varying shocks, including crop losses, business losses, and illnesses. We then 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 investigate trends in two key variables. First, we examine whether trends in children's schooling were the same in both types of households before the father departed. We have two

periods of school enrollment data before the departure, the baseline (2002) and the year before (collected retrospectively at the baseline). We cannot reject the possibility that schooling trends are the same in both types of households, as indicated by the statistically insignificant coefficient of 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 households before the father departed. We have three periods of income data before the departure, all collected retrospectively at the baseline. The evolution of per capita household labor income in the years 1999, 2000, and 2001 is similar across both types of households prior to the father's departure (lower panel of table 3). This finding gives us no reason to believe that they would have differed if departure had not occurred.

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

As an additional exercise to build confidence in the quasi-random nature of the father's absence, we determine whether current child activities are correlated with the future\e absence of the father; that is, *future* absence should not lead to a significant effect on current activities if departure is random. To do so, we regress current children's activities (schooling/work at time t for $t = 1,2$) on future absence (at time $t + 1$).

Reassuringly, we find extremely small and statistically insignificant correlations between these variables (0.009 for schooling, –0.009 for work, both with P values of 0.5).

Although all of the evidence above is reassuring, it does not address concerns that there may be unobserved time-varying shocks that affect both the father's permanent departure and the children's 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 in addition to affecting children's outcomes. We can gauge the importance of this type of shock to some extent because 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 no correlation (table 4), which is reassuring. When we control for the point estimates of the main coefficients of interest with and without such shocks in our empirical work, we note that they are very similar.

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

<<**B**>>*Attrition*

Overall, approximately 5 percent of households left the sample between the baseline survey and the first follow-up, and an additional 8.5 percent of households left between the first and second follow-ups (3.5 years after the 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 concern, we compare the baseline characteristics of households that did and did not subsequently leave the sample (table 5). As expected, households that own a house or that do not live at high altitudes are significantly less likely to abandon the study than those who do not own a house or who live at high altitudes. Other than these factors, attrition is not systematically related to any of the variables considered in the table. Although this is reassuring, potential selection biases on the basis of unobserved characteristics cannot be ruled out, for which we account in our empirical work. The methods we use to correct for this are discussed in section IV. All results presented take into account this possible selection problem, although it makes little difference to the effects that we estimate.

<<A>>*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 the permanent absence of the father on children's schooling and work outcomes. We then present our empirical findings.

<<B>>*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 the child; j denotes the household; t denotes time, where $t = 1$ denotes baseline, $t = 2$ denotes the first follow-up, and $t = 3$ denotes the 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 one if the father is absent from the household permanently and zero otherwise.

Note that $D_{j1} = 0$, by definition.[16] If the father departed the household between the baseline and the first follow-up, then $D_{j2} = 1$ and $D_{j3} = 1$; if the father departed between the 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 and in the number of siblings in 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 independent and identically distributed. The coefficient of interest is α_2, which is the effect of the absence of the father on the outcome (school or work participation).

We estimate equation (1) using a linear probability model 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, the main advantage of the linear model over discrete choice models in our case is that it is considerably easier to incorporate fixed effects. Most of the explanatory variables are discrete and take on only a few values in our application, which strengthens the case for the linear probability model (Wooldridge 2002, chapter 15). Although a potential limitation of the linear probability model is that it may yield predicted probabilities outside the unit interval, this is not a concern in our case because less than 3 percent of predictions lie outside the unit interval. We verify the robustness of our results to this linear specification by estimating a fixed-effects logit model (Honoré 2002). Although the estimates are less precisely estimated because they are based on the subset of children

who changed their activity over time, they indicate the same patterns of coefficients that are discussed in the main text on the basis of linear probability models. These are shown in table S2 in the appendix.

As discussed in section III, an important issue is that our variable of interest, the father's permanent absence, may be correlated with unobserved household characteristics that have a direct effect on children's schooling and work. To determine the effects of unobserved characteristics that are fixed over time and that may lead to spurious correlations between the 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 determine both paternal absence and the child outcomes.

A second issue, which is also discussed in section III, is that if nonrandom attrition is present, it 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 will not leave the survey by 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 one if child i from household j does not leave the survey between wave $t-1$ and wave t and zero 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, including the date (day of the month) of the interview and whether the survey respondent was the household head or spouse. Both may affect the overall experience of the interview and the respondent's willingness to be reinterviewed, but they are unlikely to affect the outcomes of interest because they relate to the previous interview, which occurred at least a year earlier.[17] The estimates from equation (2) are shown in table S3 in the appendix. The instruments are jointly statistically significant at the 1 percent 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 is not 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.

<<B>>*Results*

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

sample, which are shown in column 1, are similar to those for the restricted sample. The estimated effects are not significantly different by gender (columns 4–6).

An important reason why the negative effects on schooling and positive effects on work may be expected, which is discussed in section II, is that households in which the father leaves 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 the education of the household head (at the baseline, that is, predeparture), which is a proxy for household income. On the one hand, households with relatively less 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 are more likely to have fewer formal ways to mitigate the effects of income loss; thus, they are likely to suffer more from the father's absence. Accordingly, the interaction effect may run in both directions. We test it empirically in columns 7–9. We observe that the detrimental effects of the father's absence on schooling and child labor are driven by relatively less educated households. This finding highlights the importance of liquidity constraints for these households, which dominates the effect of the loss of relatively less income with the father's departure.

Finally, we determine 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 percent of our sample of households (compared with 4.6 percent for divorce), which results in imprecise estimations of its effects. The results (available upon request) show that the effects appear

to be driven primarily by divorce, although we cannot reject that the coefficient estimates are statistically the same. In this study, we continue to pool these events because we are interested in events that induce permanent income reductions. Another reason for combining these events is to maintain statistical power given their rarity.

<<A>>*V. Do CCTs 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 a CCT program is in effect. We begin by describing the CCT program. We determine whether the absence of the father due to divorce is affected by the CCTs and find no evidence that it is. We then estimate whether CCTs mitigate the adverse effects of the father's permanent absence.

<<B>>*The CCT Program*

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 by using an index of infrastructure relating to health and education. Municipalities from the same strata that were excluded from receiving CCTs (but that were similar to eligible municipalities in terms of population, area, and a quality-of-life index) were chosen as controls.[21] A total of 122 municipalities were chosen for the evaluation. Of these, 70 that were eligible for CCTs received the CCTs, which were phased in during the period that we are considering; 26 received CCTs by the time of the baseline survey ("early-treat"); 31 received CCTs by the first follow-up ("mid-treat"); and 13 received CCTs by the second follow-up ("late-treat"). The final evaluation sample comprised approximately 100 households that were randomly selected in each of the 122

municipalities. Attanasio et al. (2010) provide an evaluation of the program's main impacts.

<<B>>*The CCT Program and Divorce*

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

To estimate the effect of CCTs on divorce, we use data from the first and second follow-ups only (because there is no variation in the outcome of divorce at the baseline; see note 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 one if household j lives in a municipality that is receiving CCTs at time t and zero 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 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.

<<B>>*Interaction Effects*

There is a growing body of literature on the safety net role played by CCTs in the presence of income shocks. To our knowledge, however, no work has examined the role of CCTs in alleviating the 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 CCTs are in place, we augment equation (1) to include an interaction between our variable of interest, father's permanent absence, and receiving 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 one if household j lives in a municipality that is receiving CCTs at time t and zero otherwise; all other notations are as defined in equation (1). As before, T_{jt} reflects the gradual rollout of the program, such that $T_{jt} = 1$ 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 preprogram differences in

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

In municipalities that are not receiving CCTs, the permanent absence of the father reduces the likelihood of school enrollment and increases the likelihood of child labor, particularly among relatively less educated households (the left-hand columns of table 8). This is indicated by the coefficient α_1 (displayed in the first row), which estimates the effect of departure in the absence of CCTs. The second row, α_2, indicates 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 indicated by the P values of the test).[25]

Finally, as a robustness check, we restrict the comparison to households living in municipalities eligible for the CCTs that fall within the common support (that is, the region in which treated individuals have a counterpart in the group of controls according to the propensity score). Consistent with Attanasio et al. (2010), we do this by matching treatment and control observations using kernel-weighted propensity score matching and by imposing common support by dropping the 10 percent of the treatment observations in which the propensity score density of the control observations is the lowest. The results are qualitatively similar and are shown in table S4 of the appendix.

The fact that the welfare program provides insurance to protect very poor children from the adverse consequences of a father's permanent absence is perhaps not surprising to the extent that the CCTs received represent a sizeable share of income for these households (more than 20 percent of their monthly total consumption on average; see Mesnard [2009]), and the drop in household labor earnings entailed by the father's departure is of a similar magnitude. Moreover, the welfare program is in place on a

permanent basis, which ensures that the insurance it provides will continue as long as the child is enrolled in school. Notably, this result is somewhat different from that of De Janvry et al. (2006), who show that PROGRESA did not prevent children from working more following shocks related to unemployment and illness of the household head, in addition to natural disasters in the community, although it fully protected their schooling.

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

<<A>>*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 the absence of the father decreases schooling by approximately 5 percentage points and increases participation in work by approximately 3 percentage points. We provide evidence that these effects are mainly driven by households with relatively less educated heads, which are the very poorest of the indigent households in our sample. We show that receiving CCTs offsets these adverse consequences and offers children a form of insurance when the father leaves the household permanently. This

finding also suggests that the income reduction associated with paternal absence, which tightens the 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 reducing the human capital accumulation of children. Second, an event such as the permanent departure of the father has potentially important consequences for the schooling and employment of children. This is particularly true for households with relatively low levels of education that are vulnerable to permanent income losses because of insurance market failures. Third, these adverse effects can be offset by well-designed CCT programs targeted at very poor households, which, in the case of Colombia, represent more than 20 percent of total household consumption on average and are in place as long as the child is enrolled in school.

The last finding is the first of its type and offers an important agenda for future work. An important question is whether the same mechanism holds for investments other than schooling (such as children's health and nutrition) and in other contexts and environments. Another question is whether this mechanism should be taken into account in the design of safety nets and in targeting them at single parents. Such a policy might have the unintended consequence of promoting single parenthood. A final thought involves the particular relevance of these findings for sub-Saharan Africa, which has witnessed a dramatic rise in orphanhood as the result of the prevalence of HIV/AIDS, with estimates suggesting that 12 percent of all children are orphaned (UNICEF 2006).

Families and communities have shared this burden, and it might be time to institute government support to help households cope.

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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 looking 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 years.

TABLE 2. Comparison of Baseline (Predeparture) 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			
Avg. # of children ≤ 6 years	0.388	0.467	**0.011**
Avg. # of boys 7–11 years	0.727	0.738	0.775
Avg. # of girls 7–11 years	0.718	0.684	0.366
Avg. # of boys 12–17 years	0.635	0.641	0.890
Avg. # of girls 12–17 years	0.581	0.590	0.809
Avg. # of female adults	1.232	1.244	0.708
Avg. # of male adults	1.366	1.396	0.432
School enrollment rate of 7- to 14-year-olds in household	0.924	0.899	0.057
Household monthly consumption	421,286	441,994	0.085
Program area	0.700	0.682	0.453
Altitude	574.45	601.90	0.451
N	426	5,720	

Source: Familias en Accion Evaluation Data.

Note: Sample consists of households in which both parents are present at the baseline and there is at least one 7- to 14-year-old child. N denotes the 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 correlation of the figures in columns (2) and (3) are significant at the 5 percent level or less.

TABLE 3. Common Trends: Schooling and Income

	School enrollment[a]
Year = 2002	0.0343**
	(0.005)
Absence * Year = 2002	0.0238
	(0.0143)
N^1	11,679

	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	5,066

Source: Familias en Accion Evaluation Data.

Note: N^1 is the number of children in the sample at the baseline with nonmissing school enrollment 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 the municipality level, in parentheses.

[a] Dependent variable is school enrollment. Estimates are from a household fixed effects model, controlled for quadratic in child age and gender (female = 1). Reference year is 2001.

[b] Dependent variable is per capita household labor income. Estimates are from a household fixed-effects model. Reference year is 1999.

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	5,796

Source: Familias en Accion Evaluation Data.

Note: Dependent variable is divorce. Reference year is 2001. Estimates are from a household fixed-effects model pooling first and second follow-ups, controlled 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, are in parentheses.

TABLE 5. Comparison of Characteristics across Households That Do and Do Not Abandon the Study at Any Time after Baseline

Baseline Characteristics	Did not abandon	Did abandon	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
			0.0000
Owns house	0.6466	0.5321	
N	5,289	857	

Source: Familias en Accion Evaluation Data.

Note: Sample consists of households in which both parents are present at the baseline and there is at least one 7- to 14-year-old child. N denotes the 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 correlation of the figures in columns (2) and (3) are significant at the 5 percent level or less.

TABLE 6. Marginal Effects of the Father's Absence on Children's Schooling and Work

	School	School	Work	School	School	Work	School	School	Work
	Overall	Restricted		Overall	Restricted		Overall	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 * Head High Education							0.0742*	0.0851*	−0.0319
							(0.0313)	(0.0367)	(0.0295)
N	32,186	24,531	24,531	32,186	24,531	24,531	32,186	24,531	24,531

Source: Familias en Accion Evaluation Data.

Note: Marginal effects are presented from a fixed-effects linear-probability model (equation (1)), controlling for absence of father from household for unknown reason; absence of both parents; time dummies; cubic in child age; sibling composition; dummies for crop losses, illness, and business shocks; and the inverse mills ratio, as computed in equation (2) (see table S3). High education is one for incomplete secondary or more at baseline, zero otherwise. N is the number of children in the sample pooled across three waves. We included 6,146 households in our initial sample, from which 426 fathers have subsequently departed. Schooling is observed for all children in sample ≥7 years (overall sample); work is observed for children ≥10 years (restricted sample). Robust standard errors clustered at the municipality level are in parentheses. + significant at 10 percent; *significant at 5 percent; ** significant at 1 percent.

TABLE 7. Marginal Effects of CCTs on Divorce

	Pr (Divorce = 1)
CCTs	0.0024
	(0.0082)
Time = 2	−0.0286
	(0.0039)
N	5,796

Source: Familias en Accion Evaluation Data.

Note: Marginal effects are presented from equation from a fixed-effects linear-probability model (equation (3)), pooling first and second follow-ups and controlling for child composition; dummies for crop losses, illness, and business shocks; and the inverse mills ratio, as computed in equation (2) (see table S3). N is the number of households that did not abandon the study by the first follow-up. Robust standard errors clustered at the municipality level are in parentheses.

TABLE 8. Cushioning Effects of CCTs: Marginal Effects on Schooling and Work

	Low Education			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	28,027	21,464	21,464	32,186	24,531	24,531

Source: Familias en Accion Evaluation Data.

Note: Marginal effects are presented from a fixed-effects linear-probability model (equation (4)), controlling for absence of father for unknown reason; absence of both parents; cubic in child age; sibling composition; dummies for crop losses, illness, and business shocks; and the inverse mills ratio, as computed in equation (2) (see table S3). "CCTs" indicates whether the household lived in a municipality that was receiving CCTs at the time of survey. N is the number of children in the sample pooled across three waves. We included 6,146 households in our initial sample, from which 426 fathers have subsequently departed. Schooling is observed for all children in sample ≥7 years ('overall' sample); work is observed for children ≥10 years ('restricted' sample). Robust standard errors clustered at the municipality level are in parentheses. + significant at 10 percent; *significant at 5 percent; ** significant at 1 percent.

APPENDIX

TABLE S1. 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 S2. 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 S3. 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 S4. 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

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 examine this issue.

2 An absent but living father can visit and influence the children's upbringing in a way that a deceased father obviously cannot. However, relations with the absent parent's family may be very different in the two cases and may be more supportive in cases of the 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 the departure. However, in our data, transfers are very similar regardless of whether departure is caused by death or divorce, as shown in table S1 in the appendix.

3 We begin with a sample of households in which both parents are married; thus, 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 single-parent households is that the departure of the father (if present) in such households would raise additional issues that 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 one if the child's main activity in the week before the survey is reported to be any 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 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. Whereas the number of male adults is lower by almost one in households that experience departure, the number of female adults and children is the same.

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

9 In an additional 1 percent 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 percent) 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, approximately 34 percent, which we would expect because the departure of the father

decreases the number of adults in the household. We see this as a lower bound of the magnitude of the departure effect in terms of total household adult earnings because it includes labor supply responses, which are likely to cushion the potential adverse effects on income. This figure excludes earnings from children to mitigate this problem.

11 Table S1 in the appendix shows that the total value of additional transfers received by the households after paternal departure (institutional, monetary, and in-kind) is less than 50,000 pesos per annum compared to an average monthly total household consumption in excess of 420,000 pesos at the baseline. Nonetheless, these responses by extended family or friends may contribute to explaining why household consumption does not drop as much as household labor income, as noted above.

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

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

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

15 This is comparable to the attrition rate of 6 percent 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 percent 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 approximately 6 percent over the first three years of the program and is considered “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 the baseline. We only observe departures after the baseline.

17 Attrition in our sample is predominantly at the household level. Moreover, very few households (3.7 percent in the entire sample) have migrated out of their village of residence, and additional resources have been invested to track them (Mesnard 2009). Attrition is thus mostly due to nonwillingness to answer.

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

19 This suggests that child labor and schooling are strong substitutes, in contrast to the findings of Ravallion and Wodon (2000) indicating that increases in schooling in Bangladesh following a welfare program are only partially due to 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 highly educated father to involve the loss of a more positive impact on the child’s life.

21 To be eligible to qualify for the program, municipalities had to satisfy the following four criteria: (i) have fewer 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 undertaken by a consortium led by the Institute for Fiscal Studies and included the authors of this paper.

22 This also means that the identification of the effect of the CCTs on divorce comes from the rollout of the program to late-treat areas at $t = 3$.

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

24 Note that because of the gradual phasing in of the CCTs, the early-treat municipalities do not contribute to identifying α_3, the impact of the CCTs, because there are no preprogram data for these municipalities. However, we retain them in the analysis because they 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 because the early-treat municipalities do not contribute to the identification of the CCT effect, as opposed to the findings of Attanasio et al. (2010), which provide a general analysis of the impacts of the program.