Published in final edited form as:
Soc Sci Med. 2009 October ; 69(8): 1281–1286. doi:10.1016/j.socscimed.2009.07.030.

Father Absence due to Migration and Child Illness in Rural Mexico

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Abstract

Little research to date has assessed the importance of the presence of fathers in the household for protecting child health, particularly in developing country contexts. Although divorce and nonmarital childbearing are low in many developing countries, migration is a potentially important source of father absence that has yet to be studied in relation to child health. This study utilizes prospective, longitudinal data from Mexico to assess whether father absence due to migration is associated with increased child illness in poor, rural communities. Rural Mexico provides a setting where child illness is related to more serious health problems, and where migration is an important source of father absence. Both state- and individual-level fixed effects regression analyses are used to estimate the relationship between father absence due to migration and child illness while controlling for unobserved contextual and individual characteristics. The state-level models illustrate that the odds of children being ill are 39% higher for any illness and 51% higher for diarrhea when fathers are absent compared with when fathers are present in the household. The individual-level fixed effects models support these findings, indicating that, in the context of rural Mexico, fathers may be important sources of support for ensuring the healthy development of young children.

Keywords

Mexico; father absence; migration; child illness

Introduction

Child illness reduces societal and individual well-being around the world. In developing countries, diarrhea and respiratory infections coupled with malnutrition lead to roughly 40% of deaths in children under five (Black, Morris, & Bryce, 2003). Even when not leading to death, illness constrains children’s play, learning, and social interactions, as well as parental work time. Repeated and/or severe illness can affect children’s cognitive skills (Caughy,

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1996), physical development, and susceptibility to disease (Martorell & Ho, 1984), resulting in a lifetime of disadvantage (Case, Fertig, & Paxson, 2005).

Developing country households, particularly those in poor, rural areas, face a complex set of challenges to ensuring their children's healthy development. This study addresses one of these challenges: the temporary loss fathers from the household due to migration. To date, we know little about the role of fathers in protecting child health in developing country contexts (Bianchi, 1998), and even less about the implications of father absence due to migration (as opposed to divorce or nonmarital childbearing) for child health. Although migration is a relatively common phenomenon in developing countries, little research has addressed this source of family change.

This study uses longitudinal data to assess the relationship between father absence due to migration and child illness in rural Mexico. It is important that we understand how fathers, and particularly their absences, may affect child illness in settings such as rural Mexico where child illness and male migration rates are high. The findings from this study not only inform our understanding of the relationship between father absence and child illness in rural Mexico, but also highlight paternal migration as a source of change in family structure that may have significant implications for child health throughout the developing world.

Background

Mexico is a middle-income country, with 40% of the population living below the national poverty line (Central Intelligence Agency, 2005). Poverty inevitably leads to poor health conditions for children. Moderate and severe stunting (low height-for-age), which reflects chronic malnutrition and illness, currently occurs in about 18% of Mexican children under 5, compared with 16% in the Latin American region as a whole (UNICEF, 2005). Both child poverty and poor health are concentrated in rural areas. In the late 1990s it was estimated that almost 32% of children under five suffered from malnutrition in rural Mexico compared with 12% in urban areas (INSP, 1999); and, in the early 2000's the child mortality rate was two times higher in rural than urban areas (World Health Organization, 2006).

Father absence has not been widely studied in this setting, and we know even less about the role of fathers in taking care of children. Research from other countries suggests that residential fathers may be important in ensuring the health development of children through their economic and social roles. The loss of a father, whether due to migration, divorce, or death, may reduce economic and social resources, producing stressful organizational changes in the family, and decreasing emotional support to the child and his/her mother. These may, in turn, work to increase a child's susceptibility to disease, as well as decrease the resources needed to prevent and treat child illness (see Figure 1).

Empirical literature in the U.S. (Dawson, 1991; Guttmann, Dick, & To, 2004; Heiland & Liu, 2006; Page & Stevens, 2004; Wen, 2008) and developing countries (Bronte-Tinkew & DeJong, 2004; Bronte-Tinkew & Dejong, 2005; Carvalhaes, Benicio, & Barros, 2005; Gage, 1997) suggests that children have better health outcomes in two-parent families than in single mother families. Although in the U.S. economic resource loss is a main pathway through which father absence affects child well-being being (Clarke-Stewart, Vandell, McCartney, Owen, & Booth, 2000; Hango & Houseknfcht, 2005), there is evidence in developing countries that social support and time received from residential fathers are key factors in protecting child health as well (Carter, 2004; Carvalhaes et al., 2005; Gage, 1997; Gertler, Martinez, Levine, & Bertozzi, 2003; Tran, 2008).

The above studies focus on father absence due to nonmarital childbearing, divorce, or death. In settings such as Mexico, however, migration is an important source of change in family structure, as poor households routinely use migration to gain access to economic resources (Douglas S. Massey, Durand, & Malone, 2002). Mexico has a long history of international migration to the U.S. (Bean, Corona, Tuirán, & Woodrow-Lafield, 1998), and the net annual international migration rate for the country is currently estimated at −4.57 migrant(s)/1,000 population (Central Intelligence Agency, 2005). The high rate of out-migration includes substantial repeat migration, since the majority of migrants return home within one year (D.S. Massey, Alarcon, Durand, & Gonzalez, 1987; Reyes, 2001).

In terms of father absence due to migration, a recent unpublished paper estimated that 7% of Mexican children under age 15 currently live with their parents in union and their father absent; and, on average, 17% of children are likely to experience father absence due to migration at least once during their first 14 years of life (Nobles, 2006). Whether such absences may be associated with child health has not been adequately studied to date.

Father absence due to migration may have both positive and negative effects on child well-being, since it is often spurred by fathers aiming to provide for their families rather than separating from their spouses. Empirical research on migration suggests that household migration experience benefits infant health in sending households (Frank & Hummer, 2002; Kanaiaupuni & Donato, 1999; McKenzie & Hildebrandt, 2005), and that much of the effect is due to remittances (money sent back during absences). Family organizational changes that come with paternal migration may also improve child health if mothers are better able to engage sources of support and direct household spending when fathers are absent (Villarreal & Shin, 2008). This may be beneficial if, as research suggests, mothers, more than fathers, prioritize spending on child health (Case & Paxson, 2001; Maitra, 2004; Schmeer, 2005).

There is also evidence of negative effects of migration on child health. One study suggests that, although infant mortality declines and birth weight goes up in migrant households, young children in these households tend to receive fewer immunizations and less time breastfeeding (McKenzie & Hildebrandt, 2005). This suggests negative effects of father absence on child health after the first year of life, and that losses in parental time may be a particularly critical pathway through which father absence due to migration affects child health.

However, most migration studies focus on the implications of having a migrant in the household rather than on father migration per say. The one study to consider paternal migration finds no relationship between paternal migration experience and length of time breastfeeding, immunizations and food allocation among children in five rural Mexican villages (Fernandez, 1998). The study uses retrospective job histories to measure paternal migration experience rather than focusing on prospective father absence due to migration. It may be that while parental or household migration experience may have either a positive or null effect on children, their health may suffer during the time when fathers are migrating.

In sum, based on current research, we know little about whether fathers are important for child health in Mexico, and how short term father absences due to migration may be related to child illness. If fathers are important sources of social support and supervision, their children's health may suffer when they are migrating. On the other hand, if women are better able to provide for their children when fathers are absent (due to increases in decision making power and economic resources), the results may be the reverse.

To address this gap in the literature, this study investigates the relationship between father absence due to migration and child illness using prospective, longitudinal data from rural Mexico. One methodological challenge to assessing the importance of father absence to child illness is the potential selectivity of absent fathers. I address this issue using statistical analyses (individual-level fixed effects models) that reduce coefficient biases due to time-invariant unobserved differences among children and their fathers. The findings from this study provide insight into an understudied source of father absence (migration) in a setting that lacks sufficient research on family structure and child well-being (Mexico); while also providing less biased assessments of the relationship between father absence and child health than in many past studies.

Data

The data utilized here were collected by the Mexican government as part of the PROGRESA project. In the late 1990's the government created the Education, Health and Nutrition Program (*Programa de Educación, Salud y Alimentación,* or PROGRESA) to fight rural poverty, and with it began a longitudinal study of all households in 506 poor, rural communities in the states of Guerrero, Hidalgo, Michoacán, Puebla, Querétaro, San Luis Potosí, and Veracruz (World Bank, 2004). In implementing this large-scale welfare program, eligible (i.e., poor) households were enrolled incrementally based on a random selection of "treatment" and "control" communities. Households in treatment communities (60% of the communities) were enrolled in fall 1998, while eligible households in control communities were delayed entry until 2000 (after the final wave of data used here). All households from the 506 communities were included in four follow up surveys (fall 1998, spring 1999, fall 1999 and fall 2000), where child health and family structure data were collected. The data from these surveys were merged to produce the longitudinal dataset used here.

The main strength of these data for considering the question of how father absence affects child illness is the ability to assess the residential status of fathers and child illness prospectively. The close spacing of the survey waves allows for short term father absences to be assessed, and the large sample size and geographic coverage provide both statistical power and the ability to generalize the results to households in poor communities across rural Mexico. Finally, the longitudinal data allow for individual-level fixed effects modeling that reduces bias in assessing the relationship between father absence and child illness.

The sample used in this study consists of children aged 0–5 living in the surveyed households during 1998–2000. This age group closely corresponds to the age group utilized by official child mortality and morbidity measures (usually, 0–4). The study sample is restricted to children with at least one valid sick measure across the four waves and to those children living with their parents in union (marital or consensual) throughout the study period. This avoids the problematic selection bias that comes with comparing single-mother or divorced households with married couple households in estimating the effects of father absence on child well-being (Aughinbaugh, Pierret, & Rothstein, 2005; Fomby & Cherlin, 2007).

Table 1 provides the sample size and descriptive statistics for the dependent variables, father absence, and control variables used in this study (described in detail below). As Table 1 shows, the percent of children who are ill declines over time, due in part to the aging of the sample (older children are less likely to be ill) and perhaps due to improvements made in child health with the implementation of PROGRESA. The number of children per wave also declines over time (the sample size in wave 1=14,481, wave 2=12,447, wave 3=11,094, and wave 4= 8,848) as they age out of the sample or if their families are missing from the survey wave. In the analysis, the data are converted into child-years to conduct state- and individual-level fixed effects analyses, where individual children contribute multiple measures to the sample over time. The sample size for *any illness* totals 46,870 child years (17,914 children observed, on average, 2.6 times) and for *diarrhea* totals 46,594 cases (17,767 children observed, on average, 2.6 times). The slightly smaller sample size for the *diarrhea* dependent variable is due to missing values on the classification of illness. As Table 1 indicates, *illness* was reported in 21% of the child-years, and *diarrhea* occurred in 2.5% of the child-years.

Table 1 also shows the descriptive characteristics of the sample, with mean values for all variables (dummy variable means are equivalent to percentages) and standard deviations for the continuous variables. Consistent with the rural setting, this is a sample of largely economically disadvantaged children with, on average, fewer than 2 household assets (which includes kitchen appliances, washer dryer, vehicles, televisions, etc.). Only 55% of the sample mothers are literate and 37% are part of an ethnic minority group (speak a language other than Spanish in the household).

Methods

Measures

This study utilizes two dependent variables: *any illness* and *diarrhea*. The dependent variable *any illness* is measured by whether the respondent (the child’s mother) reported the child as being ill in the four weeks prior to the survey. This measure has been substantiated in the literature as being highly correlated with objective measures of child illness (Rousham, Northrop-Clewes, & Lunn, 1998). 38% of the sample children were ill at least once across the four survey waves.

The *diarrhea* dependent variable was assessed based on the mother’s classification of an illness in the past four weeks as diarrhea. This type of illness was chosen as a measure of a potentially serious illness—in 2002 diarrhea was the third leading cause of death among children under five in the developing world (World Health Organization, 2003). Diarrhea illnesses represent 11% of the child-year illnesses, and 6% of the sample children were reported as having had diarrhea at least once across the survey waves.

The main independent variable of interest is *father absence due to migration*, defined prospectively as a father who is in union (formally married or consensual union) with the mother and is not currently living in the household because he is studying or working, or for an undefined reason. This is a dichotomous variable, and the comparison group is fathers in union with the mother who are present in the household. Using this household roster approach provides a direct, prospective assessment of father absence due to migration; however, a limitation is the lack of other information about the absence (i.e., we don't know where the father went, whether he has sent money back, etc). In this sample of poor, rural communities, about 10% of the fathers have migrated at some point during this two year period. This is consistent with the estimate of 7% of married fathers absent due to migration in a 2002 nationally-representative sample of Mexican children under 15 (Nobles, 2006). The percentage of child-years with father absence (2.5%) is lower than the percent of children with fathers ever absent across the four waves (10%) because most of these fathers are absent during only one wave.

Statistical Methods

The first set of analyses utilizes logistic regression with state dummy variables (creating a state-level fixed effects models) to assess the relationship between father absence and child illness (any illness and diarrhea). The state dummy variables reduce potential biases due to unobserved state-level differences among children and their families (e.g., social networks, social norms, access to health care, access to information). Including community dummy variables did not change the results and thus the state-level fixed effects models were preferred for their parsimony.

The models utilize maximum likelihood to estimate the log odds of being ill or of having had diarrhea in the past four weeks (Long, 1997). The standard errors are adjusted for heteroskedasticity and to account for clustering of the data from the state down to the individual level using the "robust cluster" command in Stata (Angeles, Guilkey, & Mroz, 2005). The following variables, potential confounders identified in the literature, are included as controls: *child sex* (male=1 female=0); *child age* (in years, with spline to allow for different slopes for 2 and under and those over age 2); *mother's age* (in years); *mother's literacy* (literate=1, illiterate=0); *mother's ethnicity* (ethnic=1 non-ethnic=0); *household assets* (total # of household assets out of 14 possible); *number of permanent adults in the household* (all children in the sample have their father and mother listed as household members so this varies only by the number of other adults who are household members); *number of children in the household*; and *survey wave dummy variables* (to control for time trends and seasonal effects). Since there is some evidence that the social welfare program PROGRESA has impacted male migration (Stecklov, Winters, Stampini, & Davis, 2005) and child health (Gertler, 2004), I control for both whether the household qualified in terms of poverty status (*poor*) and being in the treatment community (*treat*) in 1998. The

PROGRESA program effect, then, is a time invariant variable made up of the interaction between *poor* and *treatment* variables (PROGRESA=1 if both poor=1 and treat=1, 0 otherwise).

It is possible that there are unobserved differences at the individual level that may produce bias in the results in the state-level fixed effects model. To address this, I provide a second set of analyses using conditional logit estimation, which is equivalent to individual-level fixed effects estimates for binary outcomes (Chamberlain, 1980). The conditional logit models assess the effects of father absence on the log odds of the child being ill conditional on the child's illness status in the previous wave (see Chamberlain 1980 for details on how this is done). Since this model estimates coefficients based on within individual variation, the model can only be estimated for children who have changed illness status across waves. This drops the pooled sample size to roughly 22000 child-years for *illness* and 3600 for *diarrhea*. Thus, one caveat to using this method is that the results are based on a sample that includes only those children who have ever been ill (possibly the sickest children). The main benefit of the conditional logit models is that they produce estimates based on variation within individual children's measures over time, effectively controlling for time-invariant unobserved individual and family differences (e.g., parental ability, father commitment, child's genetic susceptibility to disease) (Chamberlain, 1980; Wooldridge, 2000). The following control variables that vary over time are also included in the individual-level fixed effects models: *household assets*, *number of adults in the household*, *number of children in the household*, and *survey wave dummies*.

The state- and individual-level fixed effects analyses are conducted for both dependent variables—*any illness* and *diarrhea*—and the odds ratios (exponentiated coefficients) are presented in the tables (an odds ratio greater than one indicates a positive relationship and a ratio less than one indicates a negative relationship). I assess the significance of the coefficients through two-tailed tests of $p<.05$. Potential differences in the association between father absence and illness by sex of the child, enrollment in PROGRESA, and maternal characteristics are tested with interaction terms (statistical significance also evaluated at $p<.05$).

Results

Table 2 below shows the results from the state-level fixed effects logit models for *any illness* and *diarrhea*. The interaction terms tested were not statistically significant at $p<.05$, thus they were not included in the final models. Model 1 shows that the odds of being ill are 39% higher when fathers are absent compared with when fathers are living in the household regularly, holding all else constant (see Table 2). The control variables illustrate, that the odds of illness go down with child age, mother's ethnicity, the number of adults and children in the household, and over time. Maternal literacy and household assets slightly increase the chance of illness, but the effect is small and might represent reporting biases. As would be expected, the PROGRESA program has a significant negative effect on child illness. Several of the state dummy variables are statistically significant, and are jointly significant as a group ($p<.01$), indicating important state differences in the odds of child illness (see Table 2).

Model 2, the state-level fixed effects model for diarrhea, indicates that children have 51% higher odds of having had a diarrhea episode when their fathers are absent compared to present. This is an important association, since diarrhea is the third leading cause of child mortality in Mexico and the developing world as a whole (World Health Organization, 2006). The control variables operate similarly as in the general illness model, although several have lost significance.

Table 3 provides the results from the individual-level fixed effects (i.e., conditional logit) models for *any illness* (Model 3) and *diarrhea* (Model 4). In these models, father absence due to migration is significantly associated with a 27% increase in the odds of being ill, and a 79% increase in the odds of being ill with diarrhea. Time- varying controls were included, of which only the wave dummies are statistically significant (see Table 3). The individual-level fixed effects results illustrate support for the idea that married father absence is associated with higher odds of being ill even when accounting for unobserved time-invariant differences among children and their families. These results are consistent with the state-level fixed effects models in terms of direction and significance of the odds ratios on father absence. The small difference in magnitude of the father absence effects suggests that there may be some bias in the state-level models due to unobserved heterogeneity or in the individual-level models from using a sample of potentially sicker children.

Discussion

Married father absence due to migration is a relatively understudied aspect of children's family contexts that has the potential to improve or hinder child health. Family structure research suggests that father absence (due to non-marital childbearing, divorce or death) may increase child illness through the loss of economic and social resources, increased social disorganization, and less emotional support available to children when fathers leave. However, married father absence due to migration may be a distinct form of father absence. The migration literature suggests that, especially in poor communities, fathers may provide more economic resources to children by moving away. Mothers may engage social networks (family and friends) to support their households and may have more power to direct scarce resources towards child health when fathers are gone.

Given the lack of knowledge about father absence due to migration and the effects on children, the research question posed here was whether father absence due to migration has an overall positive, negative or no association with child illness in rural Mexico. Using data from the large scale PROGRESA social welfare program allowed for a prospective, longitudinal assessment of father absence due to migration and child illness (any illness and diarrhea) among children under age 6.

The results provide consistent evidence that, in the context of poor, rural Mexican communities, father absence due to migration is associated with an increased risk of illness (both any illness and diarrhea) among young children. These effects were significant in individual-level fixed effects models, suggesting that the positive association between father absence due to migration and child illness is not due to unobserved time-invariant differences among children or their families. The effects were stronger for diarrhea than for the general illness outcome. This may be due, in part, to the parental time and supervision needed to prevent diarrhea by boiling water and having children wash their hands regularly. With controls for household economic conditions included in the models, the loss of parental time and supervision seems to be a plausible explanation for how short term father absence may be related to child illness in rural Mexico.

When viewed within the context of children's lives in these poor, rural household, the association of father absence with increased risk of illness is important, since an illness episode (and particularly diarrhea) may result in malnutrition and subsequent illnesses, especially among young children (Black et al., 2003; Martorell & Ho, 1984). Furthermore, if a father migrates repeatedly, a child may be subject to repeated illnesses, further exacerbating their precarious health situation.

There are several limitations of this study that should be considered. First, the child illness measures depend on mothers' recall of illness in the past four weeks, and in the case of diarrhea, the mother's own interpretation of what constitutes an episode of diarrhea. The measures likely include some recall bias, possibly underreporting of diarrhea, and have the further disadvantage of only obtaining reports on only one illness in the past four weeks rather than the number of times a child was ill. Other studies should build on these findings using more objective measures of child illness where possible.

A second limitation is the lack of clear temporal ordering of the father absence and child illness events. The possibility that father absence due to migration may be the consequence rather than cause of child illness cannot be ruled out by the findings from this study. Although testing the effects of lagged (previous wave) father absence would be one way to address this, virtually all fathers have returned by the subsequent wave. Since the focus of this study is on how children fare while their fathers are absent, the analytical technique uses father absence and child illness assessed concurrently. Further, the lack of specificity in the data regarding the timing of both child illness and father absence preclude even a prospective (one wave to the next) analysis from disentangling the sequence of events. Because of this, the results reflect associations rather than causal relationships between father absence due to migration and child illness.

A third limitation of the findings presented here is that they reflect only short term father absence and the association with child illness in the short term. It cannot be concluded here whether longer absences (when fathers may have more time to send back money and families may adjust to their absence) would have similar deleterious effects; or, how father absence due to migration is associated with child health in the long run. It may be that migration by fathers has distinct effects, and possible benefits, for child health at the community or national levels; and, that short term reductions in their own children's health are made up for by longer term gains in community infrastructure, health information, or education that may be possible when sufficient numbers of adults migrate and invest in their origin communities. It cannot, therefore, be concluded from this study whether Mexican migration, as a whole, is bad for child health. Further, due to the sample limitations, the findings are generalizable only to 2-parent households in poor, rural Mexican communities.

Finally, the large dataset and relatively few cases of father absence due to migration (2.5% of the cases and 10% of the sample children) require caution in interpreting the statistical significance of the results.

These limitations notwithstanding, this study provides further evidence of the importance of fathers and their presence in the household for child well-being. This is particularly strong evidence, given that the models compare married (in union) father present and married father absent households, and that the relationships remained significant even when controlling for time-invariant unobserved individual differences. Although mechanisms were not explicitly tested here, it is interesting to note that controls for economic status and welfare enrollment did not account for the association between married father absence due migration and child illness. This indirectly indicates further support for the idea that fathers provide social support and time resources that may be important for child health. It is hoped that future research can build on these findings to further explore the role of fathers in protecting the health of children in developing countries.

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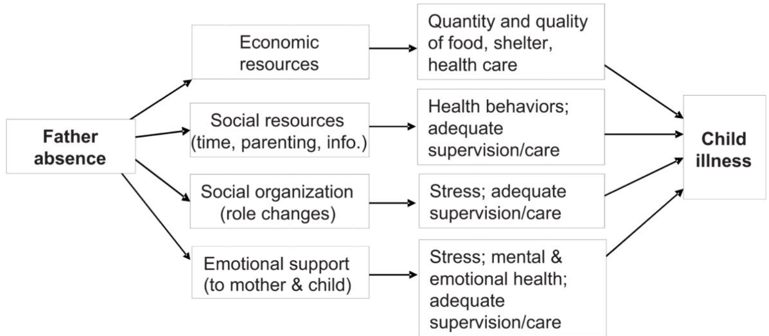

Figure 1.
Theoretical pathways through which father absence may affect child illness

Table 1

Descriptive Statistics of Sample Children Age 0 – 5

	Mean	Std. Dev.[1]
Reported any illness		
Average across all child-years	0.21	
Wave 1	0.25	--
Wave 2	0.20	--
Wave 3	0.21	--
Wave 4	0.16	--
Reported diarrhea		
Average across all child-years	0.024	
Wave 1	0.034	--
Wave 2	0.032	--
Wave 3	0.015	--
Wave 4	0.010	--
Father absent		--
Average across all child-years	0.025	
Children with father ever absent	0.10	
Control variables		
Child male	0.50	--
Child age	2.9	1.6
Mom age	29.8	7
Mom literate	0.55	--
Mom ethnic	0.37	--
# household assets	1.8	1.9
# of adults in household	2.5	1.9
# of children in household	2.8	1.6
Enrolled in PROGRESA	0.55	--

[1]Provided for non-dummy variables.

Table 2

State-Level Fixed Effects Results

	(1) Anny Illness	(2) Diarrhea
Father absent (ref: father present)	1.39*** (0.093)	1.51*** (0.12)
Control Variables		
Child male	1.02 (0.020)	1.01 (0.078)
Child age spline <=2 years old	0.95 (0.035)	0.94 (0.072)
Child age spline >2 years old	0.85*** (0.0088)	0.70*** (0.029)
Mom age at baseline	1.00*** (0.0012)	0.99 (0.0035)
Mom ethnic	0.85*** (0.018)	1.01 (0.17)
Mom literate	1.06** (0.029)	0.95 (0.050)
# household assets	1.03*** (0.013)	1.00 (0.023)
# of adults in household	0.95*** (0.018)	0.98 (0.014)
# of children in household	0.90*** (0.013)	0.96* (0.019)
HH enrolled in PROGRESA (poor x treat)	0.86** (0.058)	0.85 (0.15)
HH qualified for welfare (poor)	0.97 (0.033)	0.91 (0.087)
HH in treatment community (treat)	1.06 (0.069)	0.98 (0.19)
Survey wave 2 (spring 1999)[1]	0.68*** (0.064)	0.90 (0.14)
Survey wave 3 (fall 1999)[1]	0.74** (0.088)	0.44*** (0.065)
Survey wave 4 (fall 2000)[1]	0.50*** (0.054)	0.22*** (0.018)
State dummy variables[2]		
State 4	0.94*** (0.0067)	0.84*** (0.011)
State 5	0.87*** (0.0044)	0.84*** (0.018)
State 6	1.00 (0.013)	1.50*** (0.067)
State 12	0.78*** (0.013)	0.99 (0.064)
State 27	1.01 (0.021)	0.91 (0.081)
State 28	0.89*** (0.023)	1.31*** (0.11)
Constant	0.88 (0.087)	0.095*** (0.024)
Number of child-years[3]	46870	46594
Number of children	17914	17767
Log likelihood	−23517	−5015

Robust standard errors in parentheses.

*** $p<0.01$,

** $p<0.05$

[1]Reference group=Wave 1 (fall 1998)

[2]Reference group=State 3. State names not identified in the data.

[3]All cases over the 4 waves of data.

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Table 3

Individual-Level Fixed Effects Results (conditional logit models)

	(3) Any Illness	(4) Diarrhea
Father absent (ref: father present)	1.27** (0.12)	1.79** (0.42)
Time-varying controls		
# household assets	1.00 (0.016)	0.97 (0.043)
# of adults in household	0.98 (0.023)	0.98 (0.071)
# of children in household	1.00 (0.019)	1.05 (0.053)
Survey wave 2 (spring 1999)[1]	0.71*** (0.026)	0.85 (0.075)
Survey wave 3 (fall 1999)[1]	0.75*** (0.030)	0.40*** (0.044)
Survey wave 4 (fall 2000)[1]	0.40*** (0.030)	0.14*** (0.035)
Number of child-years[2]	22623	3669
Number of children	6838	1092
Log Likelihood	−8133	−1134

Standard errors in parentheses.

*** $p<0.01$,

** $p<0.05$

[1] Reference group=Wave 1 (fall 1998)

[2] All cases over the 4 waves of data for children who vary on the dependent variable over time.