Parental Health and Child Schooling*

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June 13, 2011

Abstract

Evidence on the role of parental health on child schooling is surprisingly thin. We explore this issue by estimating the short-run effects of parents' illness on child school enrollment. The analysis is based on household panel data from Bosnia-Herzegovina, a country whose health and educational systems underwent extensive destruction during the 1992-1995 war. Using child fixed effects to correct for potential endogeneity bias, we find that — contrary to the common wisdom that shocks to the primary household earner should bear more negative consequences for child education — it is especially maternal health that makes a difference as far as child schooling is concerned. Children whose mothers self-reported having poor health are about 7 percent points less likely to be enrolled in full-time education at ages 15-24. These results are robust to considering alternative indicators of parental health status such as the presence of limitations in activities of daily living and depression symptoms. Mothers' health shocks have more negative consequences on younger children and sons.

JEL. I21 O15

Keywords. Bosnia and Herzegovina, children, education, parents, school, self-reported health

*Comments are welcome. Simone Lombardini provided excellent research assistance. Financial support from Centro Studi Luca d'Agliano is gratefully acknowledged. Massimiliano Bratti also gratefully acknowledges funding from the Seventh Framework Programme's project 'Growing inequalities' impacts' (GINI). We thank participants to presentations given at the University of Milan for useful comments. The usual disclaimer applies.

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1 Introduction

It is widely recognized that in the absence of an adequate system of social protection, illness can take a large and unexpected toll on household well-being (Hamoudi and Sachs 1999, Dercon and Krishnan 2000, Wagstaff 2007). Indeed, health shocks typically have an impact on labor supply while also squeezing resources for consumption, due to forgone income and higher health care expenses (Gertler and Gruber 2002, Wagstaff 2007). This is especially true in developing countries where many individuals are not covered by formal health or disability insurance while out-of-pocket payments are the most important means of financing health care (World Bank 1993; 1995). Adverse health events may be even more costly in terms of well-being and growth if their economic and non-economic consequences are transferred to future generations. Yet the extent to which parental health shocks may affect children's human capital accumulation has received very little empirical attention.

This paper addresses this gap by examining the impact of parental health on children's human capital acquisition in such as setting as Bosnia and Herzegovina, where the long-term human capital costs in the population of the 1992-1995 war, in terms of both reduced education and health, could have been particularly severe.

There is a rich literature showing that exposure to income shocks is detrimental to child education when households cannot rely on formal or informal mechanisms to smooth out negative events (Jensen 2000, Edmonds 2006, Dureya et al. 2007). In particular, households facing adverse shocks may divert child time away from education and towards labor in order to substitute adult work or generate immediate income (Jacoby and Skoufias 1997, Beegle et al. 2006, Kruger 2007). Yet, while evidence has been concentrated on the income effect of shocks in market production (e.g. agricultural income shocks, crop losses, etc.), less attention has been paid to the direct and indirect consequences of household members' health on children's schooling decisions. As emphasized by Morduch (1995), income changes may be the result of ex-ante smoothing strategies, which is not the case for the type of large and unpredictable shocks that are represented by changes in health status. Moreover, parental health status may have direct non-pecuniary effects on the child's schooling, over and above the pecuniary effects mediated by lower household income and higher health expenditures. These influences may act through the reduction of the quantity and/or quality of parental time inputs into child rearing, or the emotional distress caused to children.

We examine these issues by using a detailed household panel survey collected in Bosnia and Herzegovina between 2001 and 2004, one of the few panel datasets available for transition countries. In 1990 Bosnia and Herzegovina enjoyed the economy, health status and health care of a middle-income country, but the war from 1992 to 1995 left the country's physical and human resources devastated. Health services, especially those supporting women and children, were severely disrupted, with over 35% of facilities destroyed or heavily damaged (DFID 2003). Half of the country's schools were destroyed during the conflict, decreasing access to education (World Bank 2005). Thus, due to the pervasive destruction of both the health and the education systems, the effect of parental health on child schooling is of particular concern.

A major challenge in the evaluation of the causal impact of parental health on child school enrollment is to disentangle spurious correlation, due to unobserved heterogeneity, from causality. Parents with poor health, and presumably high intertemporal discount rates, for instance, may also be those who want to invest less in their children's education. In order to address this endogeneity concern, we employ longitudinal data and child fixed effects (FE, hereafter). The main identifying assumption is that conditional on a wide range of observable characteristics and child fixed effects, parental health *shocks* triggering a poor health status are random.

Our preferred FE estimates show that children with only mothers with poor health are 7 percent points less likely to be enrolled in full-time education at ages 15-24 compared to children with healthy parents, while we do find much lower and statistically insignificant effects of paternal illness. Thus we find that — contrary to the common wisdom that shocks to the primary household earner should bear more negative consequences for child education — it is especially maternal health that makes a difference as far as child schooling is concerned. These results are robust to considering alternative indicators of parental poor health such as the presence of limitations in activities of daily living and depression symptoms, which should be less prone to measurement error.

The structure of the paper is as follows. Section 2 discusses the role of parental health in determining children's human capital acquisition, as explored in the existing literature. Section 3 describes the Bosnia and Herzegovina's context. Section 4 presents the econometric strategy and challenges to identification. Section 5 describes the data and reports some descriptive statistics. Results using parental self-reported poor health status are presented in Section 6. Section 7 checks the sensitivity of our findings against the use of alternative measures of parental health such as limitations in activities of daily living or depression symptoms. Section 8 concludes.

2 Background literature

Illness is one of the most sizable and least predictable shocks to household well-being, leaving little scope for ex-ante income smoothing strategies (Morduch 1995). Adverse health events impose to households' members current pecuniary costs, both direct, i.e. the price of accessing health care, and indirect, i.e. the loss of income associated with reduced labor supply and productivity. As a result, having a major health shock may make a family experience both a short-term income fall and a prolonged poverty trap (Wagstaff 2007, Sun and Yao 2010). Based on the theory of full insurance, Gertler and Gruber (2002) test and reject the hypothesis of consumption smoothing in the context of Indonesia, showing that households significantly reduce both labor supply and consumption patterns when hit by an adverse health event. Similarly, Asfaw and von Braun (2004) show that in Ethiopia illness has a significant negative impact on the stability and the level of household consumption. Focusing on the direct monetary costs of health, instead, Wagstaff (2007) finds evidence that the financial implications of ill health in Vietnam can be catastrophic, being associated with a significant reduction of consumption in households with no access to insurance (see also Dercon and Krishnan 2000, Baeza and Packard 2005, Bredenkamp et al. 2010).

In countries with poor systems of social protection, though, ill health may have significant economic consequences for both current and future generations (Hamoudi and Sachs 1999, Wagstaff 2007). Drawing from the economic theory of the household, if families with ill members are not able to access formal insurance markets — as it is likely to be the case in less developed or poor contexts — they may be compelled to rely on other coping mechanisms such as trading the future welfare of all or some of their members against current access to health care or forgone income for one of them. This is to say that when hit by an adverse health event, households may increase their use of child labor, by having children substitute adult labor supply and decrease school attendance.

Furthermore, parents' illness may also have non-pecuniary, e.g. psychological, costs on children, which negatively impact on their school achievement. Last but not least, as parents not only contribute monetary inputs but also time inputs into the "production" of child quality, their poor health status may reduce both the quantity and the quality of their time contributions, and affect negatively a child's quality, in our specific case education.

In the conclusion to their well known survey on the determinants of children's attainments Haveman and Wolfe (1995) mentioned information on the health status of both parents and children as one of the most pressing data needs in this area of research.

Surprisingly enough, more than 15 years later works on the effects of parental health on children's educational achievement can still be counted on the fingers of one hand. Indeed, most studies investigate the effect of that extreme form of health "shock" which is parental death. Gertler et al. (2004) use three repeated cross-sections of household data from Indonesia to test how the loss of a parent affects investment in children. They find that a parent's recent death has a large effect on the child's school enrollment, irrespective of the gender of the child and of the parent who dies. On the other hand, using longitudinal data Case and Ardington (2006) and Chen et al. (2009) present strong evidence that maternal death has a much larger impact on child education than paternal death in sub-Saharan Africa and Taiwan, respectively. Adda et al. (2011) find for Sweden that mothers are somewhat more important for children's cognitive skills and fathers for noncognitive ones. All the above mentioned papers, though, recognize that if important health problems predate parental death, the treatment effects of parental death on child school enrollment might be seriously biased. To put it in other words, parental health is considered as a confounding factor.

However, from a policy perspective, the international community is increasingly concerned about the impact of better health care (or effective risk protection) on well-being and development (World Bank 2007). This is even more relevant if ill health has (intergenerational) implications in terms of intra-household resource allocation and investments in children's human capital. Yet, as mentioned above, there is little empirical evidence pointing explicitly at the effect of parental health on child schooling. Sun and Yao (2010) investigate the consequences of household adults' health shocks on a child's likelihood of entering and finishing middle school using Chinese panel data. They find that primary school-age children are the most vulnerable to severe health shocks, measured by health expenditures larger than 5000 yuan per year, and that girls are more susceptible than boys to the damage of these shocks. Choi (2010) analyzes the long-run effects of parental self-reported poor health on children's probability of having completed at least 15 years of schooling in Russia. Her results show that a father's poor health status is a significant predictor of lower daughter's educational attainment and working probability during adulthood. Morefield (2010) investigates the effect of poor parental health, proxied by health conditions which limit an individual's daily activities or ability to work, on children's cognitive and non-cognitive skills. Cognitive skills are measured by the Revised Woodcock-Johnson (WJ-R) applied problem achievement test.[1] His results indicate that

[1]The WJ-R applied problem test evaluates a child's ability to solve practical mathematical questions and is a measure of quantitative knowledge, while non-cognitive skills refer to behavioral problems and are measured using the Behavior Problems Index (BPI) developed by Peterson and Zill (1986).

parental health is determinant only for non-cognitive skills, that are health shocks related to a vascular or cancerous condition which bear the more negative consequences, and that sons are more negatively affected than daughters.

We add to the existing literature in several respects. As to the identification strategy, unlike Sun and Yao (2010) and Choi (2010) who do not address potential endogeneity issues generated by household's or child's unobservable characteristics, we take into account child unobserved heterogeneity using child fixed effects (see section 4). Our strategy differs also from the one proposed by Morefield (2010), who seeks to account for unobserved heterogeneity by including in the child's attainment equation lagged parental inputs and past educational achievements.

Another major difference with respect to the existing studies is that we use multiple measures of parental health. We employ parental self-reported health status like Choi (2010), but also self-reported limitations in activities of daily living, like Morefield (2010), and depression scales. So doing we are able to check the robustness of our results to different indicators of parental health.[2]

We depart from the previous literature also with respect to the specific outcome variable considered. Due to the limited length of our panel data, like Gertler et al. (2004) we focus on the short-run effects of parental health, namely on *current school enrollment*. As we already stated, Choi (2010) focuses on long-term effects in terms of higher education achievement, while Morefield (2010) considers attainment in a standardized test. Both short-run and long-run effects are of interest. When considering the short-run effects of parents' health shocks, it may be argued that a child's school drop-out may be only temporary, since individuals could go back to education when parents' health improve. As we have panel data, we do account for this potential issue unless the time elapsed between dropping out and re-entering education is very long, in which case we claim drop-out can be considered as a particularly negative outcome. It should be noted, though, that health shocks producing a sudden deterioration (or improvement) of health status — the kind of shocks we consider in our analysis after conditioning on child fixed effects — are unlikely to be very temporary, and we expect the negative effects to persist overtime. Moreover, the fact that an individual dropped-out from school *per se* may make the option of re-entering the educational system less attractive, as the future benefits of schooling fall with age, while the costs, especially the psychological ones, are likely

[2] Moreover, our study presents some specific advantages with respect to Sun and Yao (2010) whose study relies on a retrospective survey to identify major health shocks and may be affected by a severe recall bias (i.e. in 2003 individuals were asked about major illnesses happened to any of the family members during 1987-2002).

to increase with age and time spent out of education. Considering short-run effects also gives the advantage that it is easier to keep under control potential confounding factors in the analysis, while in studies of long-term effects it is very difficult to account for all potential events intervened between the time parental health worsened and the time children's outcomes are observed. Last but not least, considering current school enrollment rather than achievement in a standardized test of quantitative knowledge, in which an individual's innate ability probably plays a stronger role, is likely to give different information with respect to Morefield (2010) and to complement his findings.

3 The country context

Formerly one of the six federal units constituting the Socialist Federal Republic of Yugoslavia, Bosnia and Herzegovina (Bosnia, hereafter) gained its independence during the Yugoslav wars of the 1990s and it is now transforming its economy into a market-oriented system. With a population among the youngest in the European region, Bosnia is a country where health and education levels lag substantially behind neighboring countries. Prior to the war, Bosnia was a country with a GDP of US$11 billion, a per capita income of US$2,400 and a sophisticated health system.[3] Primary and secondary schools were free, with primary education (for those aged 7-15) compulsory so that the completion of the first nine years of schooling was virtually universal. The war, though, destroyed much of the country's infrastructure and economy and the toll on the population was extremely severe. It left an estimated 250,000 people dead, 240,000 wounded, and 25,000 permanently disabled. Some 50,000 children were wounded. About 50,000 people still require rehabilitation and a 15% of the population suffer from post traumatic stress disorder. There are estimated 800,000 externally displaced people still refugees abroad, plus about 1 million internally displaced people (DFID 1999). By 1995, GDP declined to US$2 billion, and the per capita income to US$500. The unemployment was estimated to have risen to 80%. With the support for reconstruction provided by The World Bank, European Commission (EC), and a broad coalition of donors, by the end of 2000 macroeconomic stability has been achieved despite extremely unfavorable conditions. Annual economic growth has averaged about 40% in real terms since 1995, and GDP reached US$7 billion in 2001, with per capita income approaching US$1,800 (DFID 2003).

The Bosnian health system was devastated by the war. One third of all health infras-

[3]The provider network was publicly owned and financed through a para-state insurance system that provided health insurance, social security and disability insurance.

tructures were totally destroyed. About 30% of the doctors and nurses left the country or were killed in the conflict. Government financing of the service is no longer in place. There are two health systems, one for the Federacija Bosna i Hercegovina (FBiH) and one for the Republika Srpska (RS). Both Ministries of Health lack the necessary financial resources and are highly dependent on external funding and humanitarian aid. Before the war, health care services were covered 100 per cent by the social system, which collapsed during the war. In the FBiH it was replaced by an insurance fund that merged with the Federal Ministry of Health. In RS an insurance fund operating from Mostar came into operation. However, in reality the health system is funded through a diversity of sources (DFID 2003) and is still far from being able to provide financial protection against adverse health events, with only 60% of the population covered by health insurance. As other countries in the region, the major reconstruction process is now focused on enhancing the containment and efficiency of public spending. In the health sector, the main challenge is to make progress with respect to the population health status while providing protection against the short- and long-term costs of illness in terms of human capital levels and growth. Indeed, health outcomes in Bosnia lag behind those found in other countries of the region. Some key outcome indicators raise concerns: the incidence of tuberculosis is four times higher than the EU average; disability, post traumatic stress, depression and chronic diseases rank highly on the burden of diseases. Accidents and injuries are at a high level and appear to be rising. The incidence of high-cost diseases of the heart and circulatory system, stroke and cancer is above the European averages (World Bank 2005).

The war hampered access to education as well. Many school buildings were damaged, destroyed or forced to be converted into refugee centers and hospitals during the war (Mazowiecki 1994, Swee 2009). Reliable enrollment data during conflict are very rare but it has been estimated that 50 percent of schools in Bosnia required repair or reconstruction after the conflict (World Bank 2005). Furthermore, teachers also became a scarce resource due to out-migration, decreasing access to education even further. Even though several reports suggest that remaining teachers strove to share energy and resources in order to informally organize classes in occasional locations, this was easier for primary education but more difficult for secondary education and above (due to more specialized curricula). Overall, education access has suffered seriously as a result of conflict, leaving a lasting impact and developmental lag. However, primary schooling enrollment in school recovered rapidly following the conflict. By contrast, secondary and tertiary education display less consistent patterns of resilience, although they suffered equal or greater damage during conflict, and the gross enrollment rates started from a much lower level (World Bank 2005). Overall, forty percent of students do not acquire basic skills and knowledge by the end of fourth grade, while many students enrolled in costly vocational schools receive insufficient general education and are ill-equipped to meet the challenges of today's labor markets. Pre-primary education enrollment rates are the lowest in the region. While primary education enrollment rates remain high at about 93 percent, Bosnia has the lowest rate of net secondary enrollment (73 percent overall, with only 57 percent of the poor attending) of all transition countries for which data are available (World Bank 2005).

The post-conflict transition posed major challenges also to employment and labor force participation in the new labor market. The 2007 Labor Force Study estimates the overall unemployment rate in Bosnia at 31.8%, with youth unemployment much higher than for adults (up to 60%). This is among the highest in the region, and according to a recent State commission's study on Youth Issues, Bosnia unemployment rate is about 4 times the EU average (CCYI 2008).

4 Econometric issues and identification strategy

Our aim is to estimate a child's school enrollment equation in which parents' self-reported health status appears as a regressor. By assuming a linear specification (linear probability model, LPM hereafter):

$$s_{ict} = \alpha_0 + \alpha_{1m} PM_{it} + \alpha_{1f} PF_{it} + \alpha_{1mf} PMF_{it} + \alpha_2 \mathbf{x}_{it} + \alpha_3 \mathbf{w}_{it} + \delta_c + \delta_t + v_{it} \quad (1)$$

where i, c, t are subscripts for individuals, cities of current residence and calendar years, respectively, s_{ict} is child's education and the PM_{it}, PF_{it} and PMF_{it} are indicators of poor health status *self-reported* by parents. We have included three different indicators, PM_{it} takes on value one if only the child i's mother reported poor health at time t, PF_{it} equals one if only the child's father reported poor health, and PMF_{it} equals one if both parents reported poor health. This way of specifying the child schooling equation — instead of including mother's and father's poor health as separate regressors — has two main advantages: (i) it reduces potential multicollinearity problems between mother's and father's health status as the three health indicators are mutually exclusive, (ii) it is more general, in that it allows for non-linearities in the effect of parental health and relaxes the additivity assumption.[4] $\mathbf{x}_{it}$ is a vector of individual time-varying and time-

[4]More in detail, the effect of having two parents with poor health is no longer equal to the sum of the effects of having each parent in bad health conditions.

invarying characteristics such as age and sex, $\mathbf{w}_{it}$ a vector of child i's parental time-varying and time-invarying characteristics such as education and age, δ_c and δ_t are city and calendar year fixed effects, respectively, which capture city specific time invariant unobservables and nationwide time trends or macroeconomic conditions, and v_{it} is a time varying individual error term. The α's are parameters to be estimated.

Potential endogeneity of parental health status must be tackled when pursuing the task of estimating the causal effect of parental health on a child's schooling. Some parents' unobservables are likely to determine both parental health and child schooling and enter the error term v_{it} causing an endogeneity problem. One possible example of such unobservables are parents' intertemporal discount rates: parents with low discount rates will invest more both in their health and in their children's education. OLS estimates of equation (1) are likely to be affected by this source of bias. In case endogeneity is only generated by time-invariant individual unobservable characteristics, a possibility to get rid of the bias is by using a fixed effect (FE) estimator. Let us assume that the individual error term in (1) is additive and consists of a time invariant part (u_i), which may be correlated with the regressors included, and a white-noise (ϵ_{it}), i.e. $v_{it} = u_i + \epsilon_{it}$. Then equation (1) can be rewritten as:

$$s_{ict} = \alpha_0 + \alpha_{1m}PM_{it} + \alpha_{1f}PF_{it} + \alpha_{1mf}PMF_{it} + \alpha_2\mathbf{x}_{it} + \alpha_3\mathbf{w}_{it} + \delta_c + \delta_t + u_i + \epsilon_{it} \quad (2)$$

and the FE estimator with *child fixed effects* will deliver consistent estimates of the treatment effects of interest (α_{1m}, α_{1f} and α_{1mf}). As the reader will notice, only one source of potential endogeneity remains unaddressed by the fixed effects estimator above: the one coming from unobservable determinants (or correlates of) *time-variant shocks* to parental health that are also correlated with factors directly affecting a child's education (ϵ_{it}). It is hard to think what these shocks may be, and in any case they are unlikely to be very frequent. For instance, they may be serious accidents which involve parents and children, causing both a deterioration of the health status of both parents and children and a reduction in the school attendance of the latter. To avoid these odd cases, we will check the sensitivity or our estimates to including in empirical specifications also children's (and their siblings') health status. Were common shocks the main responsible for the correlation between parental health and child schooling, we would expect the coefficient on parental health to significantly decline after including the additional regressors.

How does our strategy compare with the identification strategies employed in the past literature on the effects of parental health or death? There are mainly three types of studies. A first group makes an attempt to address endogeneity by simply including in the estimation children's observable characteristics (Gertler et al. 2004, Sun and Yao

2010, Choi 2010) or using matching techniques to compare 'similar' individuals (Gertler et al. 2004), relying in both cases on a 'selection on observables' assumption.

A second group of studies seeks to improve identification by using *family fixed effects*, and exploiting differences in educational achievement between siblings (Chen et al. 2009, Adda et al. 2011). Chen et al. (2009), for instance, identify the causal effect of parental death estimating differences in college enrollment between siblings who are orphaned by an unexpected accident before versus after the age of 18. The main identification assumption, there and in similar studies, is that the best control group for an individual are his/her siblings. However, this is not necessarily the case, as children of the same parents may differ by their ability levels or non-cognitive attributes, and parents may vary their monetary and time inputs across children using compensatory or reinforcing policies (Ermisch and Francesconi 2000).[5] Our identification strategy explicitly accounts for this by using *child fixed effects*. Similar in spirit is Morefield (2010), which however does not use child fixed effects but a 'value-added plus' model (Todd and Wolpin 2006) including in the current child attainment equation both lagged attainment and lagged parental inputs.

Finally, a third stream of studies makes an attempt to address endogeneity, or to bound the endogeneity bias, by using 'most exogenous' sources of parental death or some assumptions about the time-pattern of endogeneity, respectively (Adda et al. 2011).[6]

5 Data and descriptive statistics

Our empirical analysis is based on the Bosnian Living Standards Measurement Surveys (LSMS), a panel survey conducted by the World Bank in four consecutive years (2001, 2002, 2003, and 2004). The 2001 survey is nationally representative and contains over 5,400 households and more than 9,000 individuals, half of which have been re-interviewed for the panel in the following years. The attrition rate across waves is around 5 percent, which is relatively low compared to other national panels. As other LSMS, the survey contains detailed information on individual health status (both self-reported health and

[5]Other potential weaknesses of the family fixed effects estimator are stressed by Adda et al. (2011) and concern the fact that they implicitly use for identification only children in families with two or more children, and in the case of Chen et al. (2009) with a certain spacing between births. These subpopulations, and the treatment effect there estimated, may not necessarily be representative of the general population.

[6]Adda et al. (2011) assume that the amount of endogeneity is constant or decreasing during childhood. To put it in other words, they assume that parental deaths at early ages are more likely to be endogenous than at later ages (p. 10).

physical disabilities) and educational levels along with detailed demographic characteristics of household members, household asset endowments and wealth position, ethnicity, area of residence. Consumption and income aggregates are available only in the 2001 and 2004 waves, while self-reported health status was asked in 2002, 2003 and 2004. Hence, we restrict our analysis to the last three waves and our population of interest are children aged 15-24 living in families with both parents currently alive. 15-24 is the age between the end of compulsory (universal) schooling and the age at which the hazard of being in education decreases sharply or is almost zero. In order to focus on the effect of parental health only, and to avoid that its effect being confounded with those of parental absence and parents' deaths, we exclude single parents' households and parental deaths. Moreover, since we need information on parental data, we necessarily have to focus on individuals who reside with their parents. Among individuals aged 15-24 in the LSMS, 83% are living with their parents. The corresponding percentages are 90% in the 15-19, and 74% in the 20-24 age group. Co-habitation may introduce a sample selection bias mainly on older children. Indeed, we may expect the latter to be more likely to coreside with parents in bad health conditions to offer them daily assistance. This could also be negatively correlated with children's school enrollment and generate an upward bias in our estimates. However, on the grounds that cohabitation is more likely when parents suffer from long-term, i.e. permanent, health impairments, child fixed effects are likely to attenuate the severity of this selection bias.

The sample selection criteria are detailed in Table 1. The final sample is an unbalanced panel of 786 individuals and 2,061 observations.

Current school enrollment of children aged 15-24 is the outcome of interest, which we measure with a dummy variable equal to one if the individual is enrolled in full-time education. This variable allows us to estimate both the probability to drop-out and of not enrolling in the next level of education. We do not distinguish between the two, since this would require modeling also past student status (dynamic panel), but we do not have enough waves. We do control for the (time varying) highest diploma achieved by the individual, though. This is done as to capture the fact that some individuals stop studying because they already achieved their desired level of education (e.g. many individuals may stop at the end of secondary schooling irrespective of parental health).[7] Hence, after controlling for the highest diploma hold by individuals, we are able to estimate if parental health has a *contemporaneous effect* over and above the level of education

[7]By including past educational achievement, our model resembles a value-added model (Todd and Wolpin 2003).

already achieved.

In order to measure a parent's poor health status, i.e. a major illness having potentially severe consequences for the rest of the family, we use a dummy variable equal to one if the individual reported her/his health condition over the last fourteen months as 'Very poor' (compared to the other categories provided by the survey question, that are (i) Excellent, (ii) Good, (iii), Fair and (iv) Poor).[8] In what follows, we will refer to these parents as those with 'poor health' or 'ill'. As we mentioned above though, self-reported health status may contain a considerable amount of noise, and we will also consider in section 7 information on more objective indicators of physical disabilities and mental health.

In our sample, for about 9.9%, 10% and 9.6% of observations only mother's, only father's and both parents' health status is poor, respectively. Tabulations from the population of 15-24 years old children indicate that 56 percent of those living with healthy parents are students, while the enrollment rate drops to 35 percent if both parents report poor health. Interestingly though, the enrollment rate is 36 percent if only the mother reports poor health and 49 percent if only the father is ill (see Table 2). The same pattern holds if we split the sample according to child age (i.e. if we look at secondary and tertiary education ages, separately).

6 Results

6.1 OLS, random effects and fixed effects estimates

In this subsection, we report the estimates obtained with OLS, random effects (RE) and FE models.

6.1.1 The baseline specification

Table 4 illustrates the estimates for the child school enrollment equation on the full sample of 15-24 years old children. In the first three columns we include the following standard controls: a child's age, sex, ethnic group and (time variant) highest educational qualification; the mother's and father's age and the (time variant) highest educational qualification; a dummy for the household owing a farm; the number of children in the

[8]In 2004, the survey question is: 'Please think back over the last fourteen months about your health has been. Compared to other people of your own age would you say that your health has been on the whole', with the possible answers reported in the main text. In 2002 and 2003 the question refers to the last twelve months.

household, household size, the number of sons aged 0-6, the number of daughters aged 0-6, the number of sons aged 7-15 and the number of daughters aged 7-15 in the household to account for family composition; some indicators of household wealth (house ownership, logarithm of the number of rooms, availability of water, telephone and house connected to sewer); city of current residence and calendar year fixed effects.[9] Table 3 reports sample summary statistics. Parents' attributes, such as age and education are included as they are likely to be correlated with both their health and investments in their progeny; the child's age and highest educational qualification achieved are included as the likelihood of school enrollment are included as the the likelihood of school enrollment tends to decrease with both these variables; proxies of household wealth are included as they affect both the health status and the schooling level of household members; time and city fixed effects are included to capture macroeconomic and local conditions, such as the local provision of health services. Columns (4)-(6) also include the child's and his/her siblings' poor health status, and are mainly intended as robustness checks. In particular, after conditioning on child fixed effects, these additional controls may capture time-variant household common health shocks. We do not include work-related variables, such as parents' labor force status, working hours or wages, as these variables are likely to be affected by parents' health status. So doing, we will be estimating the overall effect of parents' illness, including both pecuniary and non-pecuniary effects.

Column (1) shows the estimates obtained with OLS, from which we can see that children in families in which only mothers have poor health status are 14 percent points (p.p., hereafter) less likely to be enrolled in education than children of healthier parents. The probability of school enrollment of children with both parents in bad health conditions is about 7.6 p.p. lower. The effect of father's illness turns out to be much smaller in magnitude and statistically insignificant.

For the sake of completeness, although they are affected by the very same weaknesses of the OLS estimates, in column (2) we have reported the RE estimates, which show a reduction in the coefficient of mother's poor health. Column (3) reports our preferred specification. From the FE estimator we obtain that the mother's poor health has a negative effect on child school enrollment of about -7 p.p., statistically significant at the 1% level.

Columns (4)-(6), as we anticipated, report some robustness checks. Indeed, the FE

[9] Swee (2009) studies the effect of the conflict on individual school attainment and include in his analysis some proxies of war destructions, and other controls for the city of residence *before* the war. We cannot do the same, as these variables are time-invariant and perfectly collinear to individual fixed effects. For the same reasons, other time-invariant variables are excluded (e.g., birth order).

estimator is not consistent when, for instance, time varying shocks to parental health are correlated with child's health shock (household common shocks), which are in turn correlated to child education. We already took the example of accidents involving the whole family, another one may be viral diseases. For this reason we have included the child's and her/his siblings' poor health status in the regression but the estimates of parental health effects remain unaffected.

Our finding of a stronger effect of mothers than of fathers in line with the literature on parental deaths (Case and Ardington 2006, Chen et al. 2009, Adda et al. 2011). Also the absence of paternal effects is not new to the economic literature. Chen et al. (2009) finds, for instance, that after conditioning on family fixed effects paternal deaths have very small and statistically insignificant effects on children's college going behavior. Some of the migration literature reports similar findings. Cortes (2010) shows, for instance, that children with migrant mothers are more likely to lag behind in school compared to children with migrant fathers, supporting the fact that mother's absence has a stronger detrimental effect on child achievement than father's absence.

Generally, all these studies tend to explain the asymmetric effects of mothers and fathers with the higher importance of time over pecuniary inputs into the production of child quality, and the fact that mothers tend to be primarily responsible for childcare. Hence, a mother's poor health condition is likely to determine a large fall in a crucial input into childrearing and have a more negative impact on children's outcomes than a father's illness whose consequences on child quality are mainly of a pecuniary nature. Unfortunately, not having time diary data or other information on household members' time allocation, we are not able to directly test this hypothesis. However, we can provide some evidence consistent with it. Our results seem to suggest that pecuniary costs related to parental poor health (both health expenditures and foregone earnings) should only have a minor role for a child's school enrollment. We implemented a more direct test of this hypothesis in Table 5, which reports the estimates of the child's schooling equation also including mother's and father's monthly salaries.[10] We considered the two measures of salary available in all waves of LSMS: the last paid monthly salary or earning and the usual monthly net salary or earning (expressed in hundreds of convertible marks, KM).[11] In both cases the estimate of mother's poor health is not affected. Although salaries are not significant in the FE models, in OLS and RE models child's school enrollment seems to be more sensitive to the mother's salary.

All these pieces of evidence taken together suggest that the main causal pathway for

[10]Since for some working parents salaries are not available we included a missing value dummy.

[11]Salaries were deflated using the GDP deflator and are expressed at 1996 value.

the negative effect of mother's illness could be non-pecunary and be a reduction of the quantity and/or the quality of parenting time.

6.1.2 Heterogeneous effects by child age and gender

Up to now, we have assumed homogeneous effects of parental health on children of different age and gender. Yet according to the literature changes in household's economic conditions may differentially affect the 'treatment effects' of children according to their attributes (Becker 1981). This is so as marginal productivities of monetary resources and parental time are likely to change according to child maturity and gender. This is also the case when households are hit by parental health shocks.

Estimating heterogeneous effects of parental health shocks by child age for instance, Sun and Yao (2010) find that children are more sensitive to shocks happening during their primary school age, while those during middle school age bear no effect.[12] As for child gender, Adda et al. (2011) report a penalty from maternal death of -0.10 and -0.16 years of schooling for daughters and sons, respectively.[13] Moreover, a related literature on elderly parents' health and adult women labor force participation shows that daughters are more likely to take care of ill parents (Pagani and Marenzi 2008, Choi 2010).[14] Similarly, while the mother can take care of the father when the latter is ill, when the mother is in poor health the burden of care may be disproportionately borne by daughters.

Thus, exploring age and gender patterns in the effects of parental poor health may also shed more light into causal pathways hidden behind the stronger negative effect of mother's health we found above.

Columns (1)-(3) of Table 6 show the estimates on the sample of 15-19 years old children. The OLS estimator gives significant and negative effects for households in which only the mother is in poor health and for those in which both parents are in poor health, of -14.5 and -8.5 p.p., respectively. However, when we switch to panel estimators (RE and FE models) the latter effect falls and looses statistical significance, and only children in households where the mother (only) is in poor health turn out to have a significant penalty (-9.3 p.p.) in the probability of school enrollment. Columns (4)-(5),

[12] Their explanation is that children have already been screened during primary education, which reduces their likelihood of dropping out due to poor parental health.

[13] These are the estimates corrected for potential endogeneity of parental death.

[14] Evidence of differential effects by child gender can be also found in the migration literature. Gindling and Poggio (2010) report that children separated from parents due to migration are more likely to be high school drop-outs than children who migrated with parents, and that this effect is larger for sons (higher relatively probability to drop-out).

reporting OLS and RE estimates, show that apparently also older children are affected by mother's poor health status. However, the FE estimator indicates that this negative association is likely to reflect a spurious correlation, and disappears when we control for unobserved individual time-invariant characteristics which in the FE estimator are allowed to be correlated with both child's schooling and parental health. Under some specific assumptions, the estimates in columns (4)-(6) can also be considered as a sort of falsification check. Indeed, if older children's schooling decisions are less affected by their parents' decisions or conditions ('life-course hypothesis', Shavit and Blossfeld 1993), finding an effect of parental health in the age group 20-24 as large as the one in the 15-19 age group could be considered as a symptom that we are catching a spurious correlation due to unobserved heterogeneity. This is not the only possible interpretation of course. For instance, older children in general have a higher earnings potential and therefore their human capital investment decisions may be more sensitive to parental health shocks if they mainly involve pecuniary costs to the household. However, the fact that for both age groups the higher coefficient is attracted by mother's poor health, which is also the one more often statistically significant, makes us to propend for a greater relative importance of non-pecuniary parental inputs rather than pecuniary ones, since mothers are less likely to be working (or to find a job) in the labor market and earn an income.[15]

Table 7 reports the results of the estimates split by child gender. Columns (1)-(3) list the results for daughters and columns (4)-(6) for sons. For daughters, coefficient estimates from RE and FE estimators are much lower than those obtained using OLS, suggesting that unobserved heterogeneity is partly responsible for the negative association between parental health and their schooling. By contrast, for sons the estimates of OLS, RE and FE are much more stable. Interestingly enough, when using the FE estimator sons show a 8.4 p.p. lower probability of school enrollment, statistically significant at the 1% level. Our results of a larger impact of mothers on sons are consistent with some of the results reported in the literature, and point to a potentially more relevant role of mothers' investment on sons than on daughters.

7 Sensitivity to alternative measures of parental health

The index of self-reported parental poor health status that we have used until now may be affected by a reporting bias. Indeed, reported health status may contain a measurement

[15] In the estimation sample, 31% of child observations refer to children with working mothers as compared to 73% for children with working fathers.

error due to differences in individual reference points, e.g. more optimistic individuals may systematically overstate their health status. Unfortunately, we cannot correct for response scale bias using vignettes, as they are not available in the dataset (see Kapteyn et al. 2007). Thus it is important to use alternative measures of poor parental health and check the robustness of our results. In this section, we use as alternative proxies of parental poor health parents' self-reported ability to physically perform activities of daily living (ADLs, hereafter). Similar health indexes have been already used in the economic literature by Strauss et al. (1993), Gertler and Gruber (2002), Gertler et al. (2004) and Morefield (2010), among others. ADLs indexes are often considered more objective than self-reported health status, and less likely to be affected by differences in individual response scales, as they represent answers to very specific questions in which the interviewer asks for the ability to perform certain daily activities. These measures have been validated both in the US and in East Asian countries (Andrews et al. 1986, Guralnik et al. 1989, Ju and Jones 1989). In 2003 and 2004, the LSMS asked individuals the following questions:[16] (i) Has your health limited your ability to perform vigorous activities such as lifting heavy objects, running, or participating in strenuous sports?; (ii) Has your health limited your walking uphill?; (iii) Has your health prevented you from bending, lifting, or stooping?. The possible answers are: 'No', 'Yes, less than 3 months' and 'Yes, more than 3 months'. Scores of 1, 2 and 3 are given to the first, second and third answers, respectively. The scores to the single questions can be aggregated into a single health indicator, which we label the *ADLs score*. This last variable can then be included as a continuous indicator of parental health in the child schooling equation.[17] The ADLs score can be also dichotomized to build an indicator of poor health status. As questions on ADLs were not asked in the 2002 wave, and in order not to restrict too much the number of parents with poor health, we fixed the threshold of the dichotomous variable at an arbitrary level of 6 (corresponding, for instance, to individuals not being able to perform all three activities for less than 3 months, or only two activities for more than three months).[18]

Table 8 reports the estimates on the whole sample. The FE estimator shows that a one standard deviation increase (about 2 points) in the ADLs score of the mother

[16] The English translation is ours as original survey questions were asked in the local language.

[17] In this case, as the ADLs score is continuous, we do not include mutually exclusive indicators of mother's and father's health, but include an interaction term in the regression.

[18] We also replicated the analysis with cut points of 5 and 7 without significant differences in the results. As older parents are more likely to be affected by ADLs limitations, it is crucial to control in the child schooling equation for parents' ages, which may also have direct effects on child's attainments.

(meaning worse health) is associated with a 6.4 p.p. penalty in the likelihood of child school enrollment. The effect turns out to be statistically significant only at the 10% level. Effects are more precisely estimated when we use the dichotomous version of the indicator. Children with mothers with poor health (ADLs score ≥ 6) are about 10 p.p. less likely to be enrolled in school. The overall picture is very consistent with the results of the previous section, showing a stronger effect of mothers.

Table 9 reports the results obtained using the dichotomous health indicator on subsamples defined by child age and gender. The subsamples are quite small, and the estimates are likely to be less precise. Columns (1)-(6) of Table 9 reports the estimates in the two subsamples defined by child age. Mother's poor health is significant only for children aged 15-19, and in the FE is associated with a -22.9 lower probability of school enrollment. Columns (7)-(12) show the estimates for daughters and sons. Consistently with the previous section, point estimates seem to be larger for sons, and in the FE model mother's poor health turns out to be significant at the 10% level, showing a -13.3 p.p. penalty in the probability of school enrollment. Findings are in line with those in Tables 6 and 7.

While the proxy of poor health considered in the previous section encompasses both physical and mental health, the one based on the ADLs score refers to physical disability only. However, the LSMS also reports questions that can be used to assess individuals' mental health. In particular, waves 2003 and 2004 of LSMS provide a battery of questions that can be used to compute depression scales. Despite being subjective, as they ask respondents about their internal states and associated behaviors, these scales have been validated in the medical literature. In particular, Radloff (1977) Center of Epidemiological Studies Depression (CES-D) Scale was administered to LSMS respondents.[19] This scale has been subject to a specific validation for Bosnia and Herzegovina (Kapetanovic 2009). In the current study, we use the following items that are present in both the 2003 and the 2004 waves: (i) For the next few questions please look at Showcard C And tell me if during the last week you felt low in energy, slowed down?; (ii) During the last week did you accuse yourself for different things?; (iii) During the last week did you have problems falling asleep or sleeping?; (iv) During the last week did you feel hopeless in terms of the future?; (v) During the last week did you feel melancholic?; (vi) During the last week did you feel that you worried too much about different things?; (vii) During the last week did you feel that everything was an effort?. The possible answers are 'Not at all', 'A little', 'Quite a bit', 'Extremely often', which are attributed scores 0, 1, 2 and

[19] For more information see Do and Iyer (2009).

3, respectively. Scores in single questions can be summed to obtain an aggregate score ranging between a minimum of 0 (no depression symptoms) and a maximum of 21 (very severe depression symptoms). Panel (a) of Table 10 reports the estimates of the child schooling equation including the mother's and the father's CES-D scales along with their interaction terms. Raising by one point the CES-D score of the mother is associated to a lower 3 p.p. probability of child school enrollment when fathers show no symptoms of depression (i.e., their score is zero). Also these results are pointing to a higher role of maternal health for child's schooling achievement.

8 Concluding remarks

A major parental illness is one of the most sizeable and least predictable shocks to household's welfare, with potential long-lasting consequences if investment in children is affected. Unlike the effect of parents' death on children's outcomes, the role of parental illness on investments in children's human capital has been rarely investigated by the economic literature. Yet lack of access to both health care and insurance mechanisms is increasingly perceived by policymakers as a crucial hurdle for household well-being and economic development.

We explore this issue by estimating the short-run effects of parental illness on child school enrollment, using a detailed longitudinal panel dataset from Bosnia and Herzegovina. The latter is a transition country where the 1992-1995 conflict left both health and schooling infrastructures in a very poor state and where the levels of educational and health achievements in the population are relatively lowe compared to neighbouring countries.

Methodologically — to be best of our knowledge — this is the first paper which seeks to address the potential endogeneity of parental health status with respect to a child's education using *child fixed effects* instead of the family fixed effects widely used in cross-section studies. Unlike the latter, child fixed effects allow unobserved heterogeneity not to be the same for all children within the same household. This is important, for instance, when parents allocate different amounts of financial and time resources to their children, according to their ability or non-cognitive skills. Moreover, child fixed effects allow to focus on the whole sample of children and not only on those with at least one sibling.

Our findings show that, contrary to the common wisdom that shocks to the primary household earner bear more negative consequences for child education, it is especially maternal health that makes a difference as far as child school enrollment is concerned. If the mother self-reported to be in poor health conditions, our FE model suggests that her child is 7 p.p. less likely to be enrolled in full-time education at ages 15-24. This finding is robust to considering other — presumably more objective — measures of parental health, such as limitations in activities of daily living and depression scales, which have been validated in the medical literature. Last but not least, we find an interesting heterogeneous pattern of parental effects by child age and gender: younger children (aged 15-19) and sons seem to be more negatively affected.

By finding that negative effects of parental health shocks are stronger when the mother is ill, our analysis has important policy implications. Women's access to health care services is likely to be particularly difficult in developing and transition countries (see, for instance, Oster 2009). For this reason, especially in those countries, the implementation of an adequate system of social protection, better prevention and improved women's access to health care might contribute to greatly reducing the intergenerational cost of adults' illness.

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Table 1: Sample selection criteria

sample selection criteria	dropped	sample size
children		12,426
age≥15 & age≤24	7847	4579
cohabiting with mother and father	794	3785
self-reported health status asked in the panel wave (2002-2004)	1024	2761
parental self-reported health status non-missing	242	2519
all regressors non-missing	304	2215
at least 2 time observations in the panel	154	2061

Note. The table shows the initial sample size (number of observations) and the observations lost applying our sample selection criteria.

Table 2: Children's school attendance by parents' health status

No parent ill	Mother ill only	Father ill only	Both parents ill	Total
		Full sample		
0.56	0.36	0.49	0.35	0 .51
(0.50)	(0.48)	(0.50)	(0.48)	(0.49)
		15-19 age group		
0.76	0.6	0.74	0.65	0.73
(0.43)	(0.50)	(0.44)	(0.49)	(0.44)
		20-24 age group		
0.53	0.34	0.46	0.32	0.48
(0.50)	(0.47)	(0.50)	(0.47)	(0.49)

Note. Standard deviations in parentheses.

Table 3: Sample summary statistics

Variable	n. obs.	mean	SD
enrolled in full-time education	2061	0.522	0.500
mother only with poor health (PM)	2061	0.098	0.298
father only with poor health (PF)	2061	0.100	0.301
both parents with poor health (PMF)	2061	0.096	0.294
poor health child	2061	0.035	0.184
poor health siblings	2061	0.016	0.126
ADLs score mother (3-9)	1285	4.332	1.972
ADLs score father (3-9)	1284	4.245	1.968
interaction ADLs scores mother and father	1284	20.099	17.619
ADLs score ≥ 6 mother only	1285	0.084	0.278
ADLs score ≥ 6 father only	1284	0.067	0.250
ADLs score ≥ 6 both parents	1284	0.107	0.309
CES-D scale child (0-21)	1282	2.515	2.768
CES-D scale mother (0-21)	1282	5.566	3.328
CES-D scale father (0-21)	1281	4.628	3.312
interaction CES-D scales father and mother	1279	33.051	35.168
age	2061	19.557	2.597
male	2061	0.553	0.497
ethnic group (Bosniak)			
Serbian	2061	0.402	0.490
Croat	2061	0.083	0.276
other	2061	0.024	0.152
highest education child (primary)			
secondary	2061	0.535	0.499
tertiary	2061	0.005	0.073
age father	2061	48.740	5.683
age mother	2061	45.201	5.579
highest education father (none)			
primary	2061	0.306	0.461
secondary	2061	0.623	0.485
tertiary	2061	0.063	0.242

Continued on next page

Table 3 – continued from previous page

Variable	n. obs.	mean	SD
highest education mother (none)			
primary	2061	0.533	0.499
secondary	2061	0.399	0.490
tertiary	2061	0.022	0.148
household owns a farm	2061	0.077	0.266
number of children	2061	2.243	0.950
household size	2061	4.471	1.180
number of sons 0-6	2061	0.017	0.147
number of daughters 0-6	2061	0.038	0.216
number of sons 7-15	2061	0.216	0.466
number of daughters 7-15	2061	0.205	0.480
dwelling not appropriate	2061	0.153	0.360
house owned	2061	0.852	0.355
availability of water	2061	0.888	0.315
log number of rooms	2061	0.986	0.435
telephone	2061	0.770	0.421
house connected to sewer	2061	0.853	0.354
last real monthly salary mother (,00 KM)$^{(a)}$	545	2.253	1.508
last real monthly salary father (,00 KM)$^{(a)}$	1284	2.791	2.615
usual real net monthly salary mother (,00 KM)$^{(a)}$	557	2.210	1.262
usual real net monthly salary father (,00 KM)$^{(a)}$	1296	2.769	2.142

$^{(a)}$ Means and standard deviations refer only to the samples with positive salaries. Salaries are expressed in hundreds of convertible marks (KM) at the 1996 value.

Note. Summary statistics are reported for the estimation sample in Table 4. We also reported in this table summary statistics for other variables used in this paper based on the non-missing observations in the same sample. Reference categories for categorical variables are shown in parenthesis and the other categories in italics.

Table 4: Effect of parental health on child school enrollment

	OLS (1)	RE (2)	FE (3)	OLS (4)	RE (5)	FE (6)
mother only with poor health	-0.141***	-0.092***	-0.069***	-0.141***	-0.091***	-0.068**
	(0.032)	(0.024)	(0.027)	(0.032)	(0.024)	(0.026)
father only with poor health	-0.053	-0.032	-0.017	-0.053	-0.031	-0.015
	(0.033)	(0.024)	(0.025)	(0.033)	(0.024)	(0.025)
both parents with poor health	-0.076**	-0.021	0.001	-0.076**	-0.020	0.002
	(0.035)	(0.025)	(0.026)	(0.035)	(0.025)	(0.026)
child with poor health				0.000	-0.029	-0.046
				(0.047)	(0.040)	(0.042)
at least one sibling with poor health				0.015	-0.008	-0.023
				(0.068)	(0.063)	(0.073)
N. observations	2061	2061	2061	2061	2061	2061
N. children		786	786		786	786

*, **, *** statistically significant at the 10%, 5% and 1% level, respectively.

Note. The dependent variable is a dichotomous indicator for being enrolled in full-time education. The samples include individuals aged 15-24 in the Bosnian LSMS (2002-2004) cohabiting with both their parents. The table reports OLS, child random effects (RE) and child fixed effects (FE) estimates of the effect of parents' self-reported poor health on the probability of child school enrollment using a linear probability model. All models also control for the variables listed in section 6.1.1. Heteroskedasticity robust standard errors in parentheses. OLS standard errors are clustered by child.

Table 5: Effect of parental health on child school enrollment controlling for parents' salaries

	OLS (1)	RE (2)	FE (3)	OLS (4)	RE (5)	FE (6)
mother only with poor health	-0.134***	-0.088***	-0.065**	-0.134***	-0.089***	-0.069***
	(0.032)	(0.024)	(0.027)	(0.032)	(0.024)	(0.027)
father only with poor health	-0.047	-0.029	-0.015	-0.044	-0.027	-0.014
	(0.033)	(0.024)	(0.025)	(0.033)	(0.024)	(0.025)
both parents with poor health	-0.064*	-0.014	0.004	-0.059	-0.013	0.004
	(0.036)	(0.025)	(0.026)	(0.036)	(0.025)	(0.026)
last monthly wage mother	0.026**	0.016	0.006			
	(0.013)	(0.014)	(0.017)			
last monthly wage father	0.004	0.003	0.001			
	(0.005)	(0.003)	(0.004)			
usual net monthly salary mother				0.033**	0.036**	0.035
				(0.015)	(0.016)	(0.024)
usual net monthly salary father				0.011*	0.007	0.003
				(0.006)	(0.005)	(0.005)
N. observations	2061	2061	2061	2061	2061	2061
N. children		786	786		786	786

*, **, *** statistically significant at the 10%, 5% and 1% level, respectively.

Note. The dependent variable is a dichotomous indicator for being enrolled in full-time education. The samples include individuals aged 15-24 in the Bosnian LSMS (2002-2004) cohabiting with both their parents. The table reports OLS, child random effects (RE) and child fixed effects (FE) estimates of the effect of parents' self-reported poor health on the probability of child school enrollment using a linear probability model. All models also control for the variables listed in section 6.1.1. Heteroskedasticity robust standard errors in parentheses. OLS standard errors are clustered by child.

Table 6: Effect of parental health on child school enrollment by child age

	age 15-19			age 20-24		
	OLS	RE	FE	OLS	RE	FE
	(1)	(2)	(3)	(4)	(5)	(6)
mother only with poor health	-0.145***	-0.125***	-0.093**	-0.193***	-0.086***	-0.047
	(0.043)	(0.039)	(0.043)	(0.045)	(0.031)	(0.034)
father only with poor health	-0.018	-0.028	-0.028	-0.082	-0.022	0.001
	(0.036)	(0.032)	(0.034)	(0.051)	(0.038)	(0.040)
both parents with poor health	-0.085*	-0.055	-0.017	-0.050	0.004	0.029
	(0.048)	(0.041)	(0.046)	(0.041)	(0.030)	(0.032)
N. observations	979	979	979	998	998	998
N. children		370	370		374	374

*, **, *** statistically significant at the 10%, 5% and 1% level, respectively.

Note. The dependent variable is a dichotomous indicator for being enrolled in full-time education. The estimation samples include individuals aged 15-19 and 20-24, respectively, in the Bosnian LSMS (2002-2004) cohabiting with both their parents. The table reports OLS, child random effects (RE) and child fixed effects (FE) estimates of the effect of parents' self-reported poor health on the probability of child school enrollment using a linear probability model. All models also control for the variables listed in section 6.1.1. Heteroskedasticity robust standard errors in parentheses. OLS standard errors are clustered by child. The sum of the two subsample sizes is different from the one reported in Table since there are some individuals who turn 20 and change subsample during the 2002-2004 period, and they are dropped from the subsamples if they have less than 2 time observations.

Table 7: Effect of parental health on child school enrollment by child gender

	daughters			sons		
	OLS	RE	FE	OLS	RE	FE
	(1)	(2)	(3)	(4)	(5)	(6)
mother only with poor health	-0.140***	-0.073**	-0.040	-0.097**	-0.088***	-0.084**
	(0.052)	(0.035)	(0.037)	(0.038)	(0.034)	(0.041)
father only with poor health	-0.068	-0.049	-0.039	-0.000	-0.007	0.010
	(0.046)	(0.031)	(0.033)	(0.044)	(0.035)	(0.038)
both parents with poor health	0.048	0.065	0.062	-0.122***	-0.065**	-0.026
	(0.059)	(0.044)	(0.047)	(0.039)	(0.030)	(0.033)
N. observations	922	922	922	1139	1139	1139
N. children		353	353		433	433

*, **, *** statistically significant at the 10%, 5% and 1% level, respectively.

Note. The dependent variable is a dichotomous indicator for being enrolled in full-time education. The estimation samples include individuals aged 15-24 in the Bosnian LSMS (2002-2004) cohabiting with both their parents. The table reports OLS, child random effects (RE) and child fixed effects (FE) estimates of the effect of parents' self-reported poor health on the probability of child school enrollment using a linear probability model. All models also control for the variables listed in section 6.1.1. Heteroskedasticity robust standard errors in parentheses. OLS standard errors are clustered by child.

Table 8: Sensitivity analysis: limitations in activities of daily living (ADLs) – full sample

	OLS (1)	RE (2)	FE (3)	OLS (4)	RE (5)	FE (6)
ADLs score mother	-0.021	-0.028*	-0.032*			
	(0.019)	(0.015)	(0.017)			
ADLs score father	-0.022	-0.021	-0.013			
	(0.020)	(0.015)	(0.018)			
ADLs score mother × DAL score father	0.004	0.004*	0.004			
	(0.003)	(0.003)	(0.003)			
ADLs score only mother ≥ 6				-0.057	-0.088**	-0.104**
				(0.058)	(0.042)	(0.046)
ADLs score only father ≥ 6				-0.079	-0.030	0.039
				(0.060)	(0.054)	(0.066)
ADLs score for both parents ≥ 6				0.006	-0.020	-0.034
				(0.045)	(0.033)	(0.034)
N. observations	948	948	948	948	948	948
N. children		474	474		474	474

*, **, *** statistically significant at the 10%, 5% and 1% level, respectively.

Note. The dependent variable is a dichotomous indicator for being enrolled in full-time education. The estimation samples include individuals aged 15-24 in the Bosnian LSMS (2003 and 2004) cohabiting with both their parents. The table reports OLS, child random effects (RE) and child fixed effects (FE) estimates of the effect of parents' having reported limitations in ADLs (both continuous and dichotomized) on the probability of child school enrollment using a linear probability model. The continuous ADLs score ranges between 3 (no limitation) and 9 (all three limitations listed in section 7 for more than 3 months). All models also control for the variables listed in section 6.1.1. Heteroskedasticity robust standard errors in parentheses. OLS standard errors are clustered by child.

Table 9: Sensitivity analysis: limitations in activities of daily living (ADLs) – by child's age and gender

	age						sex					
	15-19			20-24			daughters			sons		
	OLS	RE	FE	OLS	RE	FE	OLS	RE	FE	OLS	RE	FE
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
ADLs score only mother >= 6	-0.225***	-0.226***	-0.229***	0.060	0.012	-0.003	-0.036	-0.069	-0.075	-0.062	-0.095	-0.133*
	(0.070)	(0.067)	(0.086)	(0.075)	(0.026)	(0.020)	(0.076)	(0.064)	(0.071)	(0.076)	(0.059)	(0.070)
ADLs score only father >= 6	-0.111	-0.083	0.006	0.027	0.119	0.153	0.001	0.042	0.100	-0.128	-0.085	-0.013
	(0.080)	(0.076)	(0.087)	(0.103)	(0.089)	(0.103)	(0.095)	(0.079)	(0.086)	(0.080)	(0.079)	(0.108)
ADLs score for both parents >= 6	0.071	0.064	0.042	-0.015	-0.072	-0.082	0.083	0.026	0.008	-0.024	-0.037	-0.056
	(0.075)	(0.066)	(0.056)	(0.062)	(0.049)	(0.051)	(0.073)	(0.058)	(0.062)	(0.055)	(0.040)	(0.041)
N. observations	384	384	384	456	456	456	418	418	418	528	528	528
N. children		192	192		228	228		209	209		264	264

*, **, *** statistically significant at the 10%, 5% and 1% level, respectively.

Note. The dependent variable is a dichotomous indicator for being enrolled in full-time education. The estimation samples include individuals aged 15-24 in the Bosnian LSMS (2003 and 2004) cohabiting with both their parents. The table reports OLS, child random effects (RE) and child fixed effects (FE) estimates of the effect of parents' having reported an ADLs score greater than 6 on the probability of child school enrollment using a linear probability model. All models also control for the variables listed in section 6.1.1. Heteroskedasticity robust standard errors in parentheses. OLS standard errors are clustered by child.

Table 10: Sensitivity analysis: CES-D depression scale – full sample

	OLS (1)	RE (2)	FE (3)
CES-D mother	-0.018**	-0.025***	-0.030***
	(0.008)	(0.007)	(0.009)
CES-D father	0.007	-0.002	-0.002
	(0.011)	(0.008)	(0.010)
CES-D mother × CES-D father	0.000	0.002*	0.003**
	(0.001)	(0.001)	(0.001)
N. observations	940	940	940
N. children		470	470

*, **, *** statistically significant at the 10%, 5% and 1% level, respectively.

Note. The dependent variable is a dichotomous indicator for being enrolled in full-time education. The estimation samples include individuals aged 15-24 in the Bosnian LSMS (2003 and 2004) cohabiting with both their parents. The table reports OLS, child random effects (RE) and child fixed effects (FE) estimates of the effect of parents' Center of Epidemiological Studies Depression (CES-D) scale on the probability of child school enrollment using a linear probability model. The CES-D scale we consider ranges between 0 (no depression symptoms) and 21 (maximum depression symptoms). All models also control for the variables listed in section 6.1.1. Heteroskedasticity robust standard errors in parentheses. OLS standard errors are clustered by child.