Parental Health and Child Schooling*

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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. Our 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 have 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 percentage points less likely to be enrolled in education at ages 15–24. These results are robust to considering alternative indicators of parental health status such as the presence of limitations in the activities of daily living and depression symptoms. Moreover, we find that mothers' health shocks have more negative consequences on younger children and sons.

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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 labour 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 aims at filling this gap by examining the impact of parental health on children's human capital acquisition in a setting such as Bosnia and Herzegovina, where the long-term human capital costs to 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 labour in order to substitute for 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 adult 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 high intertemporal discount rates, for instance, are likely to engage in health-damaging behaviour, have worse health, and at the same time invest less in their children's human capital. In order to address this endogeneity concern, we employ longitudinal data and, for the first time in this literature, a child fixed effects estimator.

Our preferred estimates show that children with only mothers with poor health are 7 percentage points less likely to be enrolled in education at ages 15–24 compared to children with healthier parents, while we do find much lower and statistically insignificant effects of paternal illness. Thus, it appears 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 the presence of limitations in the activities of daily living and depression symptoms, which should be less prone to measurement error. We also find that in general younger children and sons suffer the most from their mother's illness.

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 context of Bosnia and Herzegovina. 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 to the use of alternative measures of parental health. Section 8 presents heterogeneous results by child's attributes. Section 9 concludes.

2 Background literature and our main contribution

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 household's members current pecuniary costs, both direct, in terms of the price of accessing health care, and indirect, in terms of the loss of income associated with reduced labour 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 labour 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 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 labour, by having children substitute for adult labour supply, thus decreasing school attendance.

Furthermore, parents' illness may 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 papers on the effects of parental health on children's educational achievement can still be counted on the fingers of one hand. In a related literature, some influential 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 might be seriously biased. To put it in other words, parental health is considered as a confounding factor.

From a policy perspective, though, the international community is increasingly concerned about the growth-dampening effects of low levels of human capital on the one hand, and about the impact of better health care (and effective risk protection) on well-being and development on the other (e.g., the Millenium Development Goal; 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) analyses 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 probability of working 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 parental health is determinant only for non-cognitive skills, that health shocks related to a vascular or cancerous condition bear more negative consequences, and that sons are more negatively affected than daughters.

Our study adds 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). To the best of our knowledge this is the first paper using such an identification strategy in the context of evaluating the effects of parental health on child schooling. Indeed, former studies have generally used cross-section data (Chen et al. 2009), or in spite of using longitudinal data they have omitted individual fixed effects (Sun and Yao 2010) or have focused on a measure of schooling observed only at one single point in time, such as the highest level of schooling achieved by adults (Choi 2010). Our strategy differs also from the one proposed by Morefield (2010), who uses cross-section data but seeks to account for unobserved heterogeneity by including lagged parental inputs and past educational achievements in the child's attainment equation.

Another major difference from 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 indicators of mental health ('depression scales'). Hence we are able to check the robustness of our results to various indicators of parental health, which are differentially affected by measurement error.[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, we focus on the short-run effects of parental health, namely on current school enrollment.[3] As mentioned above, Choi (2010) focuses on the 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

[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 behavioural problems and are measured using the Behavior Problems Index (BPI) developed by Peterson and Zill (1986).

[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 of any family members during 1987–2002).

[3] The same is done by Gertler et al. (2004) in their study of parental deaths.

shocks, it may be argued that a child's school drop-out may be only temporary, since individuals could go back to education when the parents' health improves. 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. 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 to increase with age and time spent out of education. Considering short-run effects also gives us 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 events which have potentially 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 is likely to play a stronger role, gives different information than Morefield (2010) and complements 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 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 are substantially below those of neighbouring 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.[4] 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 and 20,000 to 50,000 women were raped and tortured. About 50,000 people still require rehabilitation and 15% of the population suffer from post traumatic stress disorder. There are an es-

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

timated 800,000 externally displaced people still refugees abroad, plus about 1 million internally displaced people (DFID 1999). By 1995, GDP had declined to US$2 billion, and the per capita income to US$500. Unemployment was estimated to have risen to 80%. With the support for reconstruction provided by The World Bank, the European Commission (EC), and a broad coalition of donors, by the end of 2000 macroeconomic stability had been achieved despite extremely unfavourable 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 infrastructures 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% 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 with other countries in the region, the major reconstruction process is now focused on enhancing the efficiency of public spending. In the health sector, the main challenge is to make progress in the population's 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 are below 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 high 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 centres and hospitals during the war (Mazowiecki 1994, Swee 2009). Reliable enrollment data during the conflict is very rare but it has been estimated that 50% of the 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 the 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 all due to the 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 the conflict, and the gross enrollment rates started from a much lower level (World Bank 2005). Overall, 40% of the 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 labour markets. Pre-primary education enrollment rates are the lowest in the region. While primary education enrollment rates remain high at about 93%, Bosnia has the lowest rate of net secondary enrollment (73% overall, with only 57% 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 labour force participation in the new labour 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's unemployment rate is about four 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. We adopt 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 the child's education and 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 the 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 the mother's and father's poor health as separate regressors—has two main advantages: (i) it reduces potential multicollinearity problems between the 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.[5] $\mathbf{x}_{it}$ is a vector of individual time-varying and time-invariant characteristics such as age and sex, $\mathbf{w}_{it}$ a vector of child i's parental time-varying and time-invariant 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 αs are parameters to be estimated.

The 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. If endogeneity is only generated by time-invariant individual unobservable characteristics, one possibility for getting rid of the bias is by using a fixed effect (FE) estimator.

Let us assume that the individual error term in equation (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:

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

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-varying 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 might be, and in any case they are unlikely to be very frequent. For instance, they might 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

[5] In more 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.

will check the sensitivity of our estimates to including in the empirical specifications also the children's (and their siblings') health status. Were common shocks the main cause 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.

It is important to stress what are our main source of identification and identifying assumption. In the child fixed effects estimator, identification comes from time-variation within the same individual in parental poor health status, i.e., by 'health shocks' (either positive or negative) which trigger parental poor health. Accordingly, the main identifying assumption is that, conditional on child and parent observables, such shocks are exogenous. The idea is that children living in certain families may be systematically (i.e., in each period) more likely to live with ill parents, but that the timing of the deterioration or improvement of their parents' 'poor' health status is substantially random after controlling for a large set of observables.

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, to identify the causal effect of parental death by 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).[6] Our identification strategy explicitly accounts for children's specificities by using child fixed effects. Similar in spirit is Morefield (2010),

[6]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.

who makes an attempt to address endogeneity by using a 'value-added plus' model (Todd and Wolpin 2006), i.e., 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).[7]

5 Data and descriptive statistics

Our empirical analysis is based on Bosnia's 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 were re-interviewed for the panel in the following years. The attrition rate across waves is around 5%, 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 physical disabilities) and educational levels along with detailed demographic characteristics of household members, household asset endowments, and wealth position, ethnicity, and area of residence. Consumption and income aggregates are available only for 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 comprises 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 parental health only, and to avoid its effect being confounded with those of parental absence and parents' deaths, we exclude single parent 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, the vast majority (83%) live with their parents. The corresponding percentages are 90% in the 15–19, and 74% in the 20–24 age group. Co-habitation could introduce a sample selection bias, if any, but this is a real possibility only for 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

[7] 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).

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 785 individuals and 2,060 observations.[8]

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 education.[9] This variable allows us to estimate both the probability of dropping 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 so as to capture the fact that some individuals do not go on in education 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).[10] Hence, after controlling for the highest diploma held by individuals, we are able to estimate whether parental health has a contemporaneous effect over and above the level of education already achieved. In the whole sample the school enrollment rate is 52.2% , 73.6% in the age group 15–19, and 48% in the age group 20–24. Boys' (girls') school enrollment rate is 43.1% (63.4)%. In the age group 15–19 (20–24) 69.6% (37.8%) of boys and 77.8% (60%) of girls are enrolled in education. Thus, school enrollment generally appears to be more frequent among girls, especially after age 19.

In order to measure a parent's poor health status, i.e., a major illness's 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).[11] 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

[8]Due the sample selection we made before the analysis, we did not use survey weights as they no longer reflect population proportions. LSMS only provides cross-sectional weights.

[9]The survey's question is: 'Are you presently attending school?'.

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

[11]In 2004, the survey question is: 'Please think back over the last fourteen months about how 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.

Section 7 information on more objective indicators of physical disabilities and mental health. In our sample, for about 9.8%, 10% and 9.6% of observations only the mother's, only the father's, and both parents' health status is poor, respectively.

As in our preferred child fixed specification the effect of parental health is identified by 'switchers', i.e., children for whom parental poor health status changes at least once during the observation period, it may be important to assess for how many individuals this happens. In our main estimation sample, the mother's poor health changes for 147 observations, the father's poor health for 175 observations, and both parents' poor health for 152 observations. Health 'shocks' are almost evenly split between bad shocks (i.e., changes in poor health from 0 to 1) and good shocks (changes from 1 to 0). Indeed, in the sample positive shocks for mothers, fathers, and both parents are 70, 83 and 79, respectively. The corresponding figures for negative shocks are 77, 92 and 73. From these statistics, it appears that the incidence of both positive and negative shocks is quite similar for mothers and fathers.

Tabulations for the population of 15–24 years old children indicate that 56% of those living with 'healthy' parents are students, while the enrollment rate drops to 35% if both parents report poor health (see Table 2). Interestingly though, the enrollment rate is 36% if only the mother reports poor health and 49% if only the father is ill. 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 OLS, random effects and fixed effects estimates

In this subsection, we report the estimates obtained with OLS, random effects (RE), and fixed effects (FE) models. Table 4 illustrates the estimates for the child school enrollment equation on the full sample of 15–24 years old children. In all OLS specifications 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 owning a farm; variables related to the household's demographic structure; a set of 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.[12] Table 3 reports sample summary statistics. Parents' attributes, such as age and

[12]Swee (2009) studies the effect of the conflict on individual school attainment and includes in his analysis some proxies of war destruction, 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

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 the likelihood of school enrollment tends to decrease with both these variables; proxies of household wealth and demographic structure are included as they affect both the health status and the schooling level of household members; time and city fixed effects capture macroeconomic and local conditions (such as the local provision of health services), respectively. Columns (4)–(6) of Table 4 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 exclude for the moment work-related variables, such as parents' labour force status, working hours or wages, as they are likely to be affected by parents' health status. Hence, we will be estimating the overall effect of parents' illness, including both pecuniary and non-pecuniary effects. Yet, as a robustness test, we will also present later estimates with controls for parents' salaries.

Column (1) of Table 4 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 percentage 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 instead 7.2 p.p. lower. The effect of the 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 as the OLS estimates, in column (2) we have reported the RE estimates, which show a reduction in the coefficient of the 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. On the other hand, when using panel estimators the effect of having both parents ill loses statistical significance with respect to OLS, suggesting that unobserved heterogeneity may be partly responsible for spurious correlation.[13]

Columns (4)–(6), as we anticipated, report some robustness checks. Indeed, the FE estimator is not consistent when, for instance, time varying shocks to parental health are correlated with child's health shocks (household common shocks), which are in turn correlated to child education. We already took the example of accidents involving the

effects. For the same reasons, other time-invariant variables are excluded (e.g., birth order).

[13] A possible reading of this result is that while bad health shocks that hit only one parent are approximately random, the same is not true when these shocks hit simultaneously both parents, reflecting a sort of 'negative assortative mating' in terms of health.

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 is in line with most of 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 a very small and statistically insignificant effect on children's going to college. 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 the mother's absence has a stronger detrimental effect on child achievement than the father's absence.

According to all of these studies, the asymmetric effects of mothers and fathers may be explained by the higher importance of time over pecuniary inputs into the production of child quality combined with the mother's traditional time-use pattern of rearing a child which is complementary to the child's education.[14] In our estimation sample, for instance, only 31% of mothers work for pay (i.e., are employees, independent workers, or seasonal workers), compared to 73% of fathers (see Table 3). Hence, a mother's poor health condition in Bosnia is likely to involve a reduction of the quantity and/or quality of parenting time and have a more negative impact on children's outcomes than paternal illness, whose consequences may be mainly pecuniary in nature.[15] Even though we are not able to directly test the intra-household time allocation hypothesis (our survey does not contain time diary data or other information on household members' time use) we can still provide some evidence consistent with it. In Table 5 we report the estimates of the child's schooling equation with controls for the mother's and father's monthly salaries. In so doing we test whether pecuniary costs related to parental poor health (in particular earnings losses) play only a minor role for a child's school enrollment. We consider two measures of salary available in all LSMS waves: the last paid monthly salary or earning and the usual monthly net salary or earning (converted into hundreds of 'convertible marks', KM).[16] As expected, in both cases the coefficient of the mother's poor health

[14]Several studies tend to stress the lower importance of current income as a determinant of human capital, compared to other non-pecuniary, especially early, parental inputs (see Cameron and Heckman 1998, Carneiro and Heckman 2002, , among others).

[15]In principle, poor health may also imply a larger amount of time spent with children for working parents, but this is unlikely to be the case for Bosnian mothers, given their low level of labour force participation.

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

remains unaffected by the inclusion of parental salaries. Although the latter are not significant in the FE models, in OLS and RE models the child's school enrollment seems to be more sensitive to the mother's salary. All of these pieces of evidence taken together suggest that the main causal pathway for the negative effect of the mother's illness is likely to be non-pecunary and related to a reduction of the quantity and/or the quality of parenting time.

It may argued, though, that results in this section reflect higher misclassification in self-assessment of health by fathers (men) than by mothers (women), hence producing an attenuation bias for the former. This may happen if, for instance, fathers are for any reason more likely to wrongly report poor health than mothers. Under the assumption that the incidence of true poor health was similar across mothers and fathers (or higher for fathers), we would expect a higher incidence of self-reported poor health for the latter, which however does not seem to be the case observing the raw stastitics in Section 5. As we can observe in the same section, the incidence of self-reported 'bad' and 'good' health shocks is very similar across mothers and fathers too.

However, for a more complete assessment of the relevance of the potential issue of measurement error, in the next section we use alternative indicators of poor health which are generally considered as less subject to report bias. Furtheremore, a direct test of potential gender bias in the self-assessment of health is reported in the Appendix.

7 Sensitivity to alternative measures of parental health

The index of self-reported parental poor health status we used so far may be affected by a reporting bias. Indeed, reported health status may contain a measurement 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 to check the robustness of our results. In this section, we use parents' self-reported ability to physically perform the activities of daily living (ADLs, hereafter) as an alternative proxy of parental poor health. Similar health indicators have been already used in the economic literature by Strauss et al. (1993), Gertler and Gruber (2002), Gertler et al. (2004), and

not available for some working parents, we included a missing value dummy. Last salaries are missing for 4.3% of mother-time observations and 10.6% of father-time observations, while usual salaries for 1.9% of mother-time observations and 5.1% of father-time observations. We could not include health expenditures among the regressors as in the LSMS they are only available in 2001 and 2004.

Morefield (2010), among others. ADLs indicators 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) among others. In 2003 and 2004 only, the LSMS asked individuals the following questions:[17] (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 three months' and 'Yes, more than three 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* and which increases with the severity of the disability. This last variable can then be included as a continuous indicator of parental health in the child schooling equation.[18] 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 three months, or only two activities for more than three months, etc.).[19]

Table 6 reports the estimates on the 2002–2003 sample. The FE estimator shows that a one standard deviation increase (about two points, see Table 3) in the mother's ADLs score (meaning worse health) is associated with a 6.2 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 9 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' poor health. Columns (7)–(9) report the estimates using self-reported poor health status, which show

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

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

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

that the different estimation sample with respect to the analysis in Table 4 is not causing remarkable differences in the estimated effects.

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 disabilities only. Hence, we turn to indicators of mental health provided by LSMS. In particular, waves 2003 and 2004 answer 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 behaviour, 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.[20] This scale has been subjected to a specific validation for Bosnia and Herzegovina (Kapetanovic 2009). In the current study, we use the following seven 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 blame 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 in Showcard C are 'Not at all', 'A little', 'Quite a bit', and 'Extremely often', which are assigned scores of 0, 1, 2, and 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). Higher CES-D scores means worse mental health.

Table 7 reports the estimates of the child schooling equation including the mother's and the father's CES-D scales along with their interaction terms. Results show that the mother's mental health is more important than the father's mental health in explaining children's school enrollment. Raising by one standard deviation (3.265) the CES-D score of the mother is associated to an about 11 p.p. lower probability of child school enrollment when fathers show no symptoms of depression (i.e., their score is zero). Consistent effects are found using the CES-D dichotomous variable. As an indicator of poor mental health status we adopted the threshold of 5.6, which was set by reporting on our 21-point (and 7-item) scale the 16 score threshold suggested by Radloff (1977) on a 60-points scale including 20 items. The results are shown in columns (4)–(6): the mother's depression is associated with a 9 p.p. negative premium in the child's school enrollment. Table 3

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

also suggests that depression is relatively more widespread among mothers than fathers. Estimates using self-reported poor health status in the same sample are reported in colums (7)–(9) respectively.

Overall, the findings in this section are robust to considering alternative indicators of parental health status and confirm a primary role of maternal health shocks in negatively affecting child school achievement.

8 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 the household's conditions may differentially affect the 'treatment effects' of children according to their attributes (Becker 1981). This is so as the marginal productivities of monetary resources and parental time, which are stretched by health shocks, are likely to change according to child maturity and gender.

Child age may lead to heterogenous treatment effects as, for instance, older children are likely to be more (financially) independent from their parents and less sensitive to shocks hitting them. Furthermore, if they have been already screened in the educational system, parents who are aware of their level of ability will not withdraw them from school in case of adverse health events (see Sun and Yao 2010). Heterogenous effects may be also expected with respect to child gender. The latter may depend on the degree to which returns to schooling are different between boys and girls, and whether sons and daughters play different (economic) roles in the household—by taking care of ill parents, for instance.[21] In addition, parents may have different preferences for the quality of female and male education based either on social and cultural norms or on the child rearing technology (Thomas 1994, Alderman and Gertler 1997).[22]

[21]Human capital theory argues that the underlying costs and benefits of schooling may vary with gender for several reasons (Schultz 1988). Gendered labour markets may provide lower wages for educated females than males, whilst girls' unpaid domestic labour is often a significant component of family production (Smock 1981, McMahon 1999). Marriage markets may take a girls' future income out of her family and transfer the benefits of education to the husband's family (Boserup 1995).

[22]While there are differences in the allocation of household resources depending on the gender of the child, these differences even vary with the gender of the parent. This may reflect either the role of preferences or differences in the technology of child rearing, e.g., it may be efficient for mothers to spend more time with daughters and fathers with sons. Thus, if the mother (father) has a preference for daughters (sons), she (he) will invest marginal resources in sons (daughters); correspondingly, sons (daughters) will lose more from a mother's shock (see also Duflo 2003, Gertler et al. 2004).

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

Tables 8 and 9 investigate heterogenous effects using self-reported poor health status as it allows for larger sub-samples and hence more precise estimates.[23] Columns (1)–(3) of Table 8 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 loses 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), reporting OLS and RE estimates, show that apparently also older children are affected by the 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, as we already said, the fact that for both age groups the higher coefficient is attached to the mother's poor health, which is also the one more often statistically significant, makes us 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 labour market and earn an income.

Table 9 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, the coefficient estimates from RE and FE estimators are much lower than those obtained using OLS,

[23] We also ran regressions using alternative indicators of parental health status such as the ADLs and CES-D scores. Due to the smaller sub-samples, the estimates are less precise but are overall consistent with the results reported in the section.

and insignificant in the FE model, 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 and when using the FE estimator sons show an 8.4 p.p. lower probability of school enrollment, statistically significant at the 1% level. This seems to suggest that according to the several intra-household allocation hypotheses, either the mother spends more time with daughters and a reduction in the quantity of her parenting time due to illness is especially damaging to sons (for whom the marginal productivity of time is higher) or, assuming that boys and girls are equally dependent on maternal time, boys have a higher opportunity cost of education when mothers are ill.[24] These results are robust to including parental wages (results are not reported here).

Overall, findings in this section seem to suggest that younger children and sons are likely to be more sensitive to maternal illness and pay the highest toll in terms of low education achievement. However, a thorough analysis of this heteregeneity would require larger samples.

9 Concluding remarks

A major parental illness is one of the most sizeable and least predictable shocks to household welfare, with potentially long-lasting consequences if investment in children is affected. Unlike the effect of parental 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 low compared to neighbouring countries.

[24] We made an attempt to distinguish between the two hypotheses by estimating a regression using child's labour force participation (LFP) as the dependent variable. In the pooled sample by gender, the mother's poor health is significantly (at the 10%) positively correlated with children's LFP in FE model. When we split the sample by child gender, the mother's effect is very imprecisely estimated. However, for both sons and daughters, the effects show a positive sign for the mother's illness and a negative sign for the father's illness.

Methodologically—to the 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 for 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 of focusing 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-reports to be in poor health, our FE model suggests that her child is 7 p.p. less likely to be enrolled in education at ages 15–24. The results are robust to considering other—presumably more objective—measures of parental physical and mental health, such as limitations in activities of daily living and depression scales, which have been validated in the medical literature. We also find heterogeneous patterns of parental effects by child age and gender: younger children (aged 15–19) and sons seem to be more negatively affected by maternal illness. By finding that the negative effects of parental health shocks are stronger when the mother is ill, our analysis supports the hypothesis that the maternal non-financial support to children is a key input for their school achievement. Furthermore, our findings point to 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 may contribute to greatly reducing the intergenerational cost of low levels of human capital.

Appendix. Differential self-report bias by gender?

We have noted in the main text that the larger effect of self-reported health on child's schooling for mothers may be determined by a larger measurement error in father's self-reported health, which may cause an attenuation bias. Although in section 7 we indirectly address this issue by considering less subjective measures of individual health, in the present Appendix we provide a more direct assessment of differences in report bias by gender.

As we have noted, the LSMS does not provide vignettes, however, the latter are available in the World Health Survey (WHS) administered in 2002 by the World Health Organization (WHO). Vignettes are hypothetical situations in which individuals are asked to rate the health of a third person, whose health conditions are carefully described. An example is the following:

> [Anton] does not exercise. He cannot climb stairs or do other physical activities because he is obese. He is able to carry the groceries and do some light household work.
>
> Q2105 Overall in the last 30 days, how much of a problem did [name of person] have with moving around? 1. None 2. Mild 3. Moderate 4. Severe 5. Extreme/Cannot do
> Q2106 In the last 30 days, how much difficulty did [name of person] have in vigorous activities, such as running 3 km (or equivalent) or cycling? 1. None 2. Mild 3. Moderate 4. Severe 5. Extreme/Cannot do

In the questions thefirst name of the vignette's person matches the respondent's gender. The idea is that differences in individual responses completely stem from differences in reporting styles, as the true health of the persons described in the vignettes cannot depend on the respondents' characteristics. In what follows, we present estimates referring to the twenty vignettes of Set-A ('mobility and affect') for which data are available for 262 individuals. The full set of vignettes is described in the WHS website.

We test gender differential in self-report bias in two ways. First, we estimate an ordered probit model where the cut-points depend on individual observable characteristics, among which gender. The latent health index for the vignette's person is:

$$Y^*_{vi} = \theta_v + \epsilon_{vi} \tag{3}$$

where the indexes v and i stand for the vignette and the individual, respectively. θ_v's are vignettes' specific fixed effects. The error term ϵ_{vi} is assumed to be standard normally distributed. The observational rule for rated health is instead:

$$Y_{vi} = j \quad \text{if} \quad \tau_i^{j-1} < Y^*_{vi} \leq \tau_i^j \quad \text{for} \quad j = 1...5 \tag{4}$$

where $j = 1...5$ are the points of the Likert scale (none, mild, moderate, severe, extreme). The cut-points are modelled as:

$$\tau_i^j = a_j + \mathbf{x'}_i\gamma. \tag{5}$$

In the vector of controls affecting report bias ($\mathbf{x}_i$) we include an individual's gender, age and years of completed education. In this model, individual characteristics of the respondents produce the same shift in all cut-points, implying a reporting bias which affects extreme categories only. For instance, a positive coefficient on the female variable would mean that women are relatively more likely than men to report the 'none' category and less likely to report the 'extreme' category of the Likert scale. Given the linearity of the cut-points, this model can be estimated using Maximum Likelihood by simply including in an ordered probit specification the additional variables affecting the thresholds. The results are reported in column (1) of Table 10 and show that gender does not affect report bias. The only statistically significant variables turn out to be the vignettes' fixed effects (we do not report the coefficients but the Wald test for their joint exclusion).

Second, we estimated a model allowing for self-report bias to affect the last two cut-points only, which are the most important for reporting 'poor health' in our analysis based on the LSMS. In this case, we modelled the cut-points following Pradhan and van Soest (1995):

$$\begin{aligned} \tau_i^j &= a_j && \text{for} \quad j = 1, 2 \\ \tau_i^j &= \tau_i^{j-1} + \exp(\mathbf{x}'_i \gamma^j) && \text{for} \quad j = 3, 4. \end{aligned} \tag{6}$$

This model allows for individual characteristics to have *different* effects on the last two cut-points. The results are reported in columns (2)-(3) of Table 10, and also in this case there is no evidence of a self-report bias.

We also tried with a model allowing for a greater flexibility, by making all cut-points depend on individual characteristics as in Kapteyn et al. (2007), but the Maximum Likelihood estimation did not achieve convergence.

Overall, all these additional checks suggest the absence of a differential self-report bias by the respondent's gender.

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Table 1: Sample selection criteria

sample selection criteria	dropped	sample size
children		12,426
age$\geq$15 & age$\leq$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	305	2214
at least 2 time observations in the panel	154	2060

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		
0.57	0.36	0.52	0.37	0 .52
(0.50)	(0.48)	(0.50)	(0.49)	(0.50)
		15-19 age group		
0.76	0.60	0.74	0.63	0.73
(0.43)	(0.50)	(0.44)	(0.49)	(0.44)
		20-24 age group		
0.52	0.31	0.48	0.34	0.48
(0.50)	(0.46)	(0.50)	(0.47)	(0.50)

Note. Standard deviations in parentheses. The sample includes 785 individuals and 2060 observations.

Table 3: Sample summary statistics

Variable	n. obs.	mean	SD
child enrolled in education	2060	0.522	0.500
mother only with poor health (PM)	2060	0.099	0.298
father only with poor health (PF)	2060	0.100	0.301
both parents with poor health (PMF)	2060	0.096	0.295
poor health child	2060	0.035	0.184
poor health siblings	2060	0.017	0.129
age	2060	19.564	2.597
male	2060	0.553	0.497
ethnic group (Bosniak)			
Serbian	2060	0.402	0.491
Croat	2060	0.082	0.274
other	2060	0.024	0.152
highest education child (primary)			
secondary	2060	0.535	0.499
tertiary	2060	0.005	0.073
age father	2060	48.732	5.673
age mother	2060	45.196	5.575
highest education father (none)			
primary	2060	0.305	0.460
secondary	2060	0.624	0.485
tertiary	2060	0.063	0.242
highest education mother (none)			
primary	2060	0.533	0.499
secondary	2060	0.399	0.490
tertiary	2060	0.022	0.148
household owns a farm	2060	0.076	0.265
number of children	2060	2.239	0.938
household size	2060	4.465	1.162
number of sons 0-6	2060	0.017	0.147
number of daughters 0-6	2060	0.038	0.216
number of sons 7-15	2060	0.217	0.466
number of daughters 7-15	2060	0.204	0.480
dwelling not appropriate	2060	0.154	0.361
house owned	2060	0.852	0.355
availability of water	2060	0.889	0.314
log number of rooms	2060	0.985	0.435
telephone	2060	0.770	0.421
house connected to sewer	2060	0.853	0.354
last real monthly salary mother (,00 KM)$^{(a)}$	545	2.253	1.508

Continued on next page

Table 3 – continued from previous page

Variable	n. obs.	mean	SD
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
ADLs score mother (3-9)[b]	946	4.276	1.934
ADLs score father (3-9)[b]	946	4.174	1.905
ADLs score interation[b]	946	19.425	16.908
ADLs mother ≥ 6(b	946	0.103	0.304
ADLs father ≥ 6[b]	946	0.091	0.288
ADLs both parents ≥ 6[b]	946	0.099	0.299
CES-D scale mother (0-21)[c]	940	5.555	3.265
CES-D scale father (0-21)[c]	940	4.606	3.278
CES-D mother $\times$ CES-D father[c]	940	32.705	35.124
CES-D mother > 5.6[c]	940	0.180	0.384
CES-D father > 5.6[c]	940	0.052	0.222
CES-D both parents > 5.6[c]	940	0.318	0.466

[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.

[b] Summary statistics refer to the estimation sample used in Table 6.

[c] Summary statistics refer to the estimation sample used in Table 7.

Note. Summary statistics are reported for the estimation sample in Table 4. 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***	-0.140***	-0.091***	-0.068**
	(0.032)	(0.024)	(0.027)	(0.032)	(0.024)	(0.026)
father only with poor health	-0.053	-0.031	-0.017	-0.053	-0.031	-0.016
	(0.033)	(0.024)	(0.025)	(0.033)	(0.024)	(0.025)
both parents with poor health	-0.072**	-0.020	0.001	-0.070*	-0.018	0.003
	(0.036)	(0.025)	(0.026)	(0.036)	(0.025)	(0.026)
child with poor health				0.012	-0.026	-0.048
				(0.048)	(0.040)	(0.042)
at least one sibling with poor health				-0.059	-0.022	-0.008
				(0.068)	(0.038)	(0.041)
N. observations	2060	2060	2060	2060	2060	2060
N. individuals		785	785		785	785

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

Note. The dependent variable is a dichotomous indicator for being enrolled in 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. 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 wage mother				0.033**	0.036**	0.035
				(0.015)	(0.016)	(0.024)
usual net monthly wage father				0.011*	0.007	0.003
				(0.006)	(0.005)	(0.005)
N. observations	2060	2060	2060	2060	2060	2060
N. individuals		785	785		785	785

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

Note. The dependent variable is a dichotomous indicator for being enrolled in 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. Heteroskedasticity-robust standard errors in parentheses. OLS standard errors are clustered by child.

Table 6: Sensitivity analysis: limitations in activities of daily living (ADLs)

	ADLs score (continuous)			ADLs score (dichotomous)			poor health status		
	OLS	RE	FE	OLS	RE	FE	OLS	RE	FE
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
ADLs score mother	-0.020	-0.028*	-0.032*						
	(0.019)	(0.015)	(0.017)						
ADLs score father	-0.021	-0.021	-0.013						
	(0.020)	(0.016)	(0.018)						
ADLs score mother × ADLs score father	0.004	0.005*	0.004						
	(0.003)	(0.003)	(0.003)						
ADLs score mother ≥ 6				-0.046	-0.068*	-0.091**			
				(0.056)	(0.039)	(0.042)			
ADLs score father ≥ 6				-0.031	0.008	0.046			
				(0.052)	(0.041)	(0.045)			
ADLs score both parents ≥ 6				0.010	-0.012	-0.025			
				(0.045)	(0.032)	(0.034)			
mother only with poor health							-0.161***	-0.127***	-0.092**
							(0.047)	(0.041)	(0.047)
father only with poor health							-0.119**	-0.098**	-0.053
							(0.052)	(0.042)	(0.047)
both parents with poor health							-0.081	-0.070	-0.044
							(0.051)	(0.044)	(0.052)
N. observations	946	946	946	946	946	946	946	946	946
N. individuals		473	473		473	473		473	473

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

Note. The dependent variable is a dichotomous indicator for being enrolled in 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. Heteroskedasticity-robust standard errors in parentheses. OLS standard errors are clustered by child.

Table 7: Sensitivity analysis: CES-D depression scale

	mental health (continuous)			mental health (dichotomous)			poor health status		
	OLS	RE	FE	OLS	RE	FE	OLS	RE	FE
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
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)						
CES-D mother $\geq$ 5.6				-0.067*	-0.085**	-0.091**			
				(0.040)	(0.033)	(0.039)			
CES-D father $\geq$ 5.6				-0.030	-0.027	0.017			
				(0.056)	(0.045)	(0.049)			
CES-D both parents $\geq$ 5.6				-0.050	-0.026	0.008			
				(0.038)	(0.026)	(0.028)			
mother only with poor health							-0.161***	-0.127***	-0.092**
							(0.047)	(0.041)	(0.047)
father only with poor health							-0.121**	-0.098**	-0.053
							(0.052)	(0.042)	(0.047)
both parents with poor health							-0.081	-0.070	-0.043
							(0.051)	(0.044)	(0.052)
N. observations	940	940	940	940	940	940	940	940	940
N. individuals		470	470		470	470		470	470

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

Note. The dependent variable is a dichotomous indicator for being enrolled in 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. Heteroskedasticity-robust standard errors in parentheses. OLS standard errors are clustered by child.

Table 8: 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)	(0.044)	(0.045)	(0.031)	(0.034)
father only with poor health	-0.018	-0.028	-0.028	-0.083	-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	-0.046	0.004	0.029
	(0.048)	(0.041)	(0.046)	(0.042)	(0.030)	(0.032)
N. observations	977	977	977	999	999	999
N. individuals		369	369		374	374

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

Note. The dependent variable is a dichotomous indicator for being enrolled in 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. 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 4 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 9: 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	-0.139***	-0.072**	-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	-0.067	-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	0.053	0.066	0.062	-0.122***	-0.065**	-0.026
	(0.059)	(0.044)	(0.047)	(0.039)	(0.030)	(0.033)
N. observations	921	921	921	1139	1139	1139
N. individuals		352	352		433	433

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

Note. The dependent variable is a dichotomous indicator for being enrolled in 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. Heteroskedasticity-robust standard errors in parentheses. OLS standard errors are clustered by child.

Table 10: Evidence on reporting bias in self-reported health

	Ordered probit		Flexible-thresholds ordered probit			
			cut point 3		cut point 4	
female	0.0107	(0.054)	-0.0217	(0.0717)	0.0372	(0.0599)
age	-0.0002	(0.0018)	0.0026	(0.0026)	-0.0008	(0.0022)
education	-0.0006	(0.0069)	0.0078	(0.0116)	-0.0155	(0.0105)
constant			-0.3136*	(0.1904)	0.4297**	(0.1695)
cut point 1	-2.3584***	(0.1663)	-2.3469*** (0.0894)			
cut point 2	-1.3841***	(0.1467)	-1.3730*** (0.0523)			
cut point 3	-0.5036***	(0.1490)	(see above)			
cut point 4	0.8000***	(0.1500)	(see above)			
Log likelihood	-4830		-4823			
Wald test θ_v's ($\chi^2(19)$)	1107.78 [0.00]		1294.59 [0.00]			
Wald test female = 0	0.04 [0.84]		0.38 [0.83]			
N. observations	5238		5238			
N. individuals	262		262			

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

Note. The dependent variable is an ordered categorical variable (1. None; 2. Mild; 3. Moderate; 4. Severe; 5. Extreme) for vignettes for health state descriptions in Set-A ('Mobility and affect') of the Bosnia-Herzegovina's WHS (WHO). Coefficient estimates are reported in the table. Heteroskedasticity-robust standard errors clustered by individual are reported in parentheses and p-values in brackets. For two observations vignettes' responses are missing.