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Ecological Influences on Child and Adolescent Development: Evidence from a Philippine Birth Cohort

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Declaration

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Abstract

The largest number of children and young people in history are alive today, so the costs of them failing to realise their potential for development are high. Most live in low-income and lower-middle-income countries (LLMICs), where they are vulnerable to risks that may compromise their development. Yet many risk factors in LLMICs are not well understood. Moreover, recent studies suggest that in addition to the critical first 1,000 days there are several key periods of development in later childhood and adolescence which have received comparatively little research attention. This work responds to the gaps in the evidence, examining the influence of exposure to risks in the physical and social environment on health, education and development outcomes in a birth cohort of children from the Philippines.

The first chapter provides a brief introduction to the theoretical and empirical evidence on the risks children face in LLMICs as well as a description of the Philippine country context and the birth cohort. The second chapter tests the associations between infant exposure to sanitation risks and subsequent school survival. The third chapter investigates the effects of housing instability in early to middle childhood on cognitive performance at 11 years of age. And, the fourth chapter examines the links between forms of social marginalisation and adolescent mental health and wellbeing. This work's findings suggest infant exposure to faecal contamination in the home environment shortens the overall length of time children later spend at school. Pre-primary-school age children appear to be at risk of developmental deficits and/or delays as a result of changes to their neighbourhood environment. And, adolescents who are excluded or become disengaged from the important socialising institutions of school and the workplace are at increased risk from developing mental disorders,

while among older teens the protective effects associated with being in employment are greater than those linked to being in education.

Impact Statement

The design and implementation of evidence-based interventions to support potential development can make an important difference to a child's chances of thriving. Yet gaps in the evidence base undermine the efforts of policymakers and development practitioners, especially in LLMICs where the ambitious 2030 Agenda for Sustainable Development underscores the urgency of the issue. This research responds to these critical gaps and its findings have clear policy and practice implications.

This work contributes to our understanding of the social determinants of development during middle childhood and adolescence. For example, examining the associations between social marginalisation and adolescent mental health revealed not only the serious burden of health facing young people in developing countries like the Philippines, but also how adolescent health and wellbeing underpins the transition from school to work and early adulthood and therefore guarantees earlier investments in nurturing care and quality education. This research also identified new cross-cutting issues in child health and education. For example, its findings describe the interconnectedness of water, sanitation and hygiene, early childhood development and education, as well as the influence of both education and employment on health outcomes. In addition, this work highlights the potential problems arising from the lack of evidence from LLMICs. For example, a review of the literature uncovered a single study on adolescent mental health and social marginalisation in LLMICs, which found being in school improved teen mental health but no health benefits to being in work. However, the results of this research suggest youth employment may have a strong protective effect on mental health and wellbeing. The social ecology of children varies according to local demographic, economic and cultural contexts, and the paucity of developing country-specific evidence is likely to limit the effectiveness of interventions.

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Chapter 1.

Risks to Child and Adolescent Development in Low-Income and Lower-Middle-Income Countries: A Brief Introduction

1.1 Introduction

Before reaching adulthood, children and adolescents have developed many of the physical, cognitive, emotional and social attributes that will shape their lives. The exposure to developmental risks during this period can therefore have wide-ranging effects on individual, family and community capabilities. In low- and lower-middle-income countries (LLMICs) an estimated 250 million children are at risk of failing to reach their development potential primarily as a result of poverty conditions, malnutrition and inadequate learning opportunities (Lu, Black, and Richter, 2016). Adversity in early life is known to undermine school preparedness and school performance, which in turn affects adult health and earnings, and contributes to the intergenerational transmission of poverty (Engle et al., 2007; Gertler et al., 2014). The developmental trajectories of growing children and adolescents also reflect structural and social determinants such as the lack of good quality education and youth employment in many LLMICs (Bell, Donkin, and Marmot, 2013; Blum and Boyden, 2018).

Yet there remain important gaps in our understanding of the causes of inequalities of child and adolescent development worldwide. Much of what is known about the effects of exposure to risk factors on developmental outcomes comes from studies conducted in high-income country (HIC) contexts, even though the majority of the global population of children live in the developing world and the social and economic contexts across LLMICs are highly heterogenous. Moreover, a large proportion of the research on the determinants of child development in LLMICs has concentrated on early childhood, attracting to it considerable resources and policy attention. Although the valuable evidence gathered as a result of this research focus highlights the importance of investing in the crucial first 1,000 days, it has led to the neglect of other developmentally sensitive periods, notably middle childhood and adolescence. In addition, there are multiple potentially important cross-sectoral linkages that are not well understood, despite the established interdependence of sectors like education and health in LLMICs (Glandon et al., 2018).

No single academic discipline encompasses physical, cognitive and socio-emotional growth from birth to early adulthood. Moreover, the multiple domains of development – physical health, the development of motor, cognitive and language skills, social and emotional functioning at different ages – cannot be fully understood in isolation. The present work therefore follows the example of influential interdisciplinary research on child and adolescent development, such as the 2007 and 2011 *Lancet Series* on child development, the 2007 and 2012 *Series* on adolescent health, and the 2016 *Lancet Commission* on adolescent health and wellbeing, drawing on knowledge and methods from several of the applied health, behavioural and social sciences, including epidemiology, psychology and economics.

In doing so this research aims to respond to several of the outstanding evidence gaps. Specifically, it examines the influence of exposure to risks in the physical and social environment in childhood and adolescence on health, education and development outcomes in an urban birth cohort of children from the Philippines, strengthening our shared knowledge of risk and protective factors and better enabling the successful design and implementation of effective evidence-based policies and programme interventions.

The second section of this introductory chapter is a critical summary of the developmental risk factors facing children and adolescents in LLMICs, including the relationship between poverty and risk exposure. The third section considers these risk factors within the bio-ecological framework (Bronfenbrenner and Morris, 2006). Included is discussion of several implications of the bio-ecological model for this research. The fourth section reviews the demographic, social and economic context of the Philippines over the last 30 years and describes the data used in this work. It concludes with an overview of the four chapters that follow the introduction.

1.2 Risks in the context of child and adolescent development

Developmental risks refer to adverse biological, psychosocial and/or environmental events known to compromise 'the acquisition and growth of the physical, cognitive, social and emotional competencies required to engage fully in family and society' (Abner et al., 1997, p.47; quoted in Boyden, Dercon, and Singh, 2014, p.195). Yet the inherently probabilistic nature of risk means the actual consequences for children and adolescents depend on the combined influence of individual and contextual factors, the timing, duration and severity of the exposure, as well as the role played by protective processes (Wachs and Rahman, 2013). So, although children in households with inadequate water and sanitation are more likely to have compromised physical growth, the outcome is contingent on characteristics of the child, its family and neighbourhood environment, whether exposure occurs during a developmentally sensitive period, and the presence of mitigating factors.

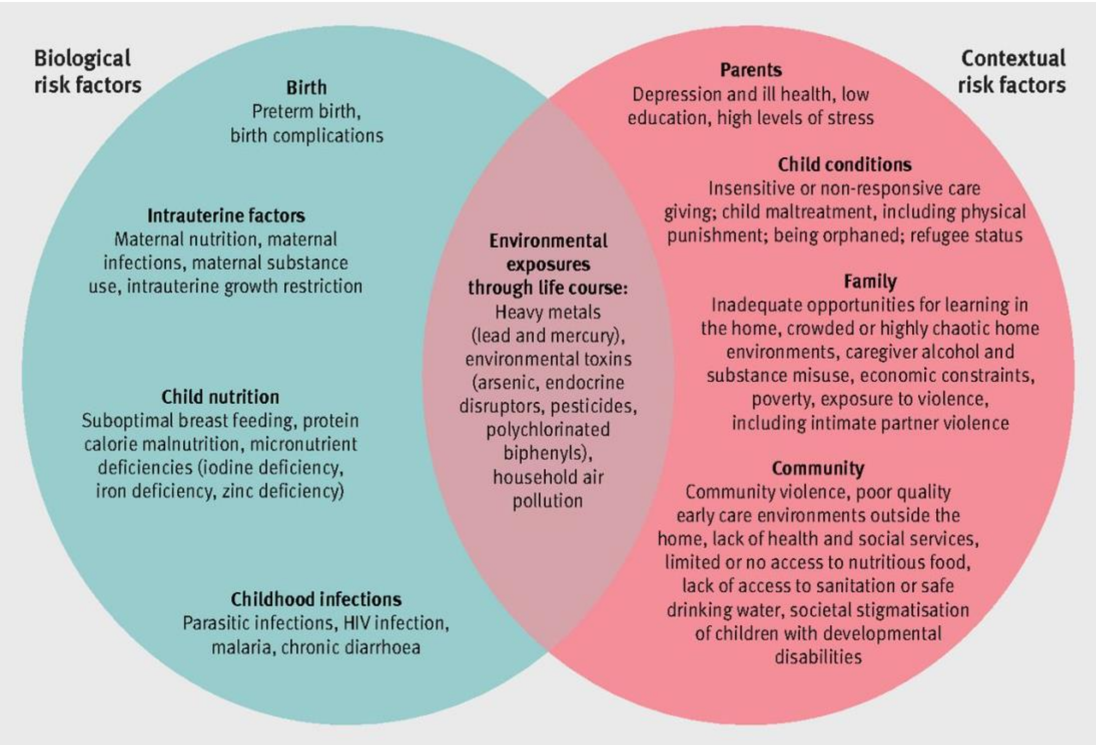
Individual biology is a key influence on the developmental burden of risks. For example, cerebral malaria in childhood is known to be a potentially important cause of neurological deficits in countries where malaria is widespread, but there is evidence that genetic variation provides protection for some children in malaria-endemic areas against the most severe forms of the disease (Brewster, Kwiatkowski, and White, 1990; Greenwood, Marsh, and Snow, 1991; Hill et al., 1991). Other individual characteristics like gender, cognitive ability and psychosocial resources also influence the extent to which children and adolescents are resilient to adversity (Masten et al., 1999). For example, low cognitive test scores and male gender were found to increase the probability of developing psychological difficulties in a group of children from Brazil (Goodman, Fleitlich-Bilyk, Patel, & Goodman, 2007).

Equally, the multiple physical and social contexts in which risks occur also importantly determine their developmental effects. For example, several studies have observed poorer motor, cognitive and socioemotional outcomes in otherwise healthy infants with iron-deficiency anaemia in low-income and lower-middle-income countries (e.g. Grantham-McGregor and Ani, 2001). Yet research conducted in Peru found associations between maternal intelligence and improved offspring iron status – measured using individual haemoglobin and transferrin saturation levels, as well as the nutritional value of reported diet – even after accounting for family socio-economic position (Wachs et al., 2005). Although no direct pathway between mothers' education and infant iron status was established, the same study did find a significant association between years of maternal schooling, duration of breast-feeding and offspring diet. In other words, despite risk factors including low family income and limited food choices, the impact on young children was mitigated by the protective characteristics of the family and the mother in particular.

The effects of risk exposure also depend on timing and whether exposure occurs in a critical or sensitive period of development, during which children may be especially vulnerable (Bornstein, 1989). There are generally recognised to be multiple such periods before children reach adulthood, with some considered more important for certain developmental domains than others (Wachs et al., 2014). For example, researchers examining the impact of iron-deficiency on brain development have focused on infancy since it is characterised by a combination of rapid cerebral change and high nutrient requirements, although animal studies suggest the window during which nutrient deficiencies have the biggest effect on the brain includes the late foetal/neonatal stage (Lozoff et al., 2006). Longitudinal evidence from Costa Rica shows that the effects of iron deficiency in early life are not limited to the first few years, persisting through middle childhood and adolescence and affecting school performance (Lozoff, Jimenez, and Wolf, 1991; Lozoff et al., 2000). However, adolescence has also been identified as a sensitive period for the developing brain (Steinberg, 2005). Adolescence is associated with major physical, psychological and social change, which coincides with changes to specific regions and systems of the brain important for regulating behaviour and emotion (Casey, Jones, and Hare, 2010). For example, this neural development occurs at the same time as young people are exposed to potentially stressful social environments such as new schools and peer-groups, a combination which is thought to contribute to the onset of mental illnesses observed in this age-group (Blakemore and Mills, 2014; Fuhrman, Knoll, and Blakemore, 2015).

The risks children and adolescents contend with can be grouped into two broad categories: biological and psychosocial/contextual (Walker et al., 2007; Walker et al., 2011). Although Figure 1.1 presents a summary of risks factors that can be found across high-, middle- and low-income countries, it is important to recognise many have a higher prevalence and severity in the developing world (unsafe drinking water, inadequate sanitation, and severe childhood diarrhoea), while a small number are almost exclusively found in low- and lower-middle-income country settings (exposure to armed conflict, natural disasters, diseases like cholera and malaria, and banned environmental toxins like polychlorinated biphenyls) (Wachs and Rahman, 2013). For example, although children living in poverty in HICs can experience malnutrition, the 'dosage' (i.e. the frequency, intensity and duration of exposure) is likely to be much higher for poorer children in LLMICs (Wachs, Cueto, and Yao, 2016). Moreover, there are fewer protective factors in LLMICs, such as skilled health workers attending births (Mills, 2014) or quality education (Glewwe and Kremer, 2006).

Figure 1.1 Developmental risk factors found in low- and lower-middle-income countries



Source: Daelmans, B. et al. (2015). Effective interventions and strategies for improving early child development. BMJ, 351: h4029. Reproduced with permission from BMJ Publishing Group Ltd.

1.2.1 Poverty and risks

Many of the risk factors in low- and lower-middle-income countries are linked to poverty conditions (Grantham-McGregor et al., 2007; Walker et al., 2007). Poverty has multiple definitions. For example, a common approach defines poverty using the fixed level of income necessary for households to be able to meet basic needs like food, clothing and shelter (Ravallion, 1992). In dollar-a-day terms, the percentage of the world's population estimated to be living in extreme poverty – those on $1.90 per day or less – has fallen from about 35 percent in 1990 to 10 percent in 2015. Nonetheless, estimates vary substantially across regions, with the most recent ranging from more than 40 percent of the population in Sub-Saharan Africa to just 1.5 percent in Europe (obtained from World Bank website: https://data.worldbank.org/indicator).

Some researchers, notably Amartya Sen (1993), conceptualise poverty as more than a lack of money, arguing it is characterised by broader deprivations (Ruggeri-Laderchi, Saith, and Stewart, 2003). From this perspective, being poor means the failure to achieve essential capabilities; a term referring to 'the ability to satisfy certain crucially important functionings up to certain minimally adequate levels' (Sen, 1993, p.41). Operationalising Sen's 'capabilities' approach Alkire and Foster (2007, 2011) developed an index to measure multidimensional poverty – which combined health, education and standard of living data from LLMICs – and found evidence of distinct patterns of poverty. For example, the probability of being multidimensionally poor but *not* income-poor was estimated at 12 percent in China but 59 percent in Chad (Alkire and Santos, 2010). This indicates that even where households are not in severe income-poverty they may still contend with major deprivations of key services such as primary education, safe drinking water and adequate sanitation facilities.

Notwithstanding its definition poverty is widely recognised as a major influence on a person's chances of thriving (Walker et al., 2011; Patton et al., 2016; Black et al., 2017). An analysis of Mexico's *Oportunidades* – a conditional cash transfer programme – found the cash component was associated with a range of outcomes measured in children aged 24 to 72 months, including physical health and growth, motor and cognitive development and receptive language (Fernald, Gertler, and Neufeld, 2008). In other words, the results showed that children from low-income families who received money achieved better health and development outcomes, regardless of the conditionalities of the programme (participating families were required to attend preventative health care services, agree to regular medical check-ups for children, as well as ensure attendance at school). Conversely, children from poor households with less money did less well.

The study authors suggest several possible mechanisms. Greater purchasing power may have improved quality of diet, indirectly increased cognitive stimulation (as a result of parents buying more toys and books), and/or reduced the risk of exposure to infection. Alternatively, the additional income could have increased the mental health and wellbeing of family members leading to improvements in nurturing care (Fernald, Gertler, and Neufeld, 2008). These explanations suggest poverty conditions increase the likelihood of exposure to other more proximate risks that in turn compromise development (Wachs and Rahman, 2013; Wachs, Cueto, and Yao, 2016). Put another way, although poverty is a major influence on the lives of children and young people in LLMICs, effective policies and intervention programmes aimed at mitigating its effects on the potential for human development must address specific risk factors (Wachs, Cueto, and Yao, 2016).

1.3 The bio-ecological model

Following in the footsteps of major interdisciplinary research on child and adolescent development in LLMICs – for example, the 2007 *Lancet Series* of papers on child development, and the 2016 *Lancet Commission* on adolescent health and wellbeing – this work adopts a conceptual framework broadly based on Bronfenbrenner's bioecological theory (Bronfenbrenner and Morris, 2006).

An ecological perspective is appropriate for several reasons. First, child and adolescent development is a multidimensional concept concerned with an outcome – a child's overall development – which depends on complex interactions between individual children and contextual biological, psychosocial and environmental factors, including maternal-child health, nutrition and education (Britto, Engle, and Super, 2013).

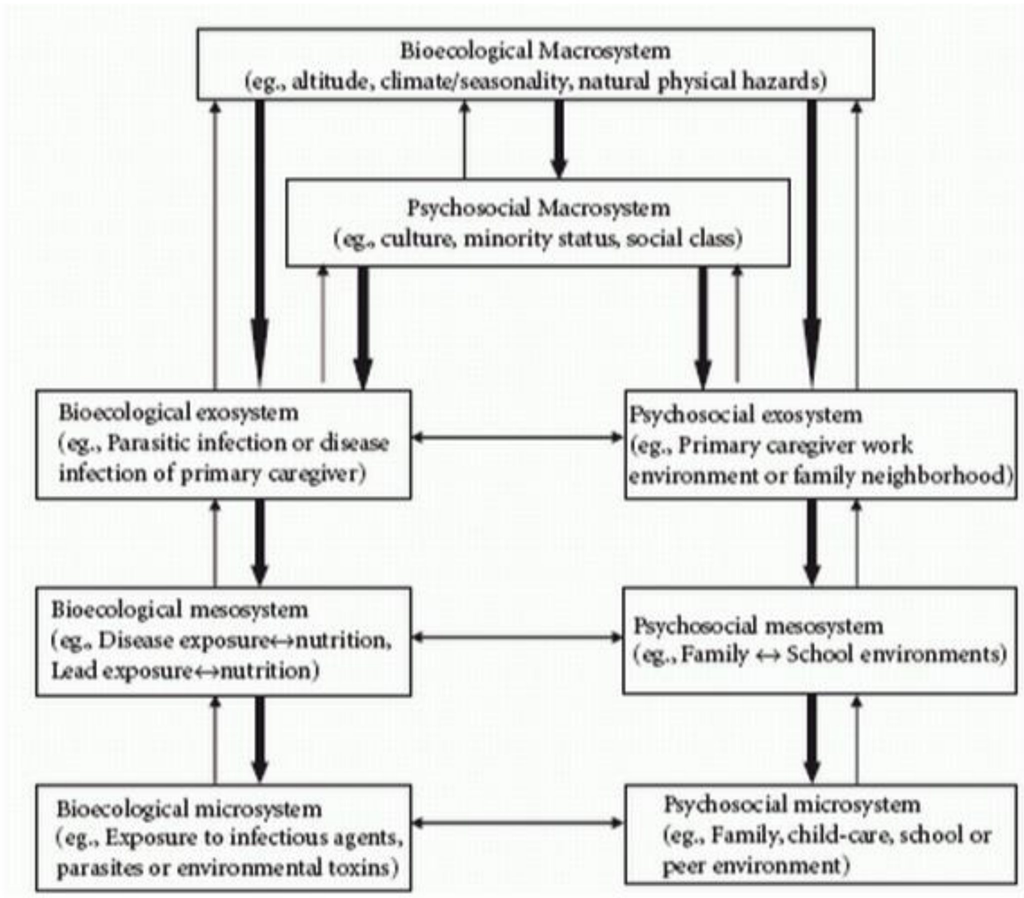
Second, the concept is fundamentally ecological, since it is predicated on a holistic, rights-based view of human development; Article 6 of the Convention on the Rights of the Child establishes the rights of children to 'life, survival and development' (United Nations General Assembly, 1989), while General Comment No. 7 'Implementing Child Rights in Early Childhood' specifies that 'ensuring survival and physical health are priorities, but States parties are reminded that article 6 encompasses *all aspects of development*, and that a young child's health and psychosocial well-being are in many respects interdependent' (United Nations Committee on the Rights of the Child, 2006 p.38; quoted in Bornstein et al., 2012, emphasis added).

Widely applied across multiple fields of study from epidemiology and economics to psychology and public health, as well as in high-, middle- and low-income country settings, Bronfenbrenner's is an ecological perspective which has undergone significant refinement and revision (Tudge, Mokrova, Hatfield, & Karnik, 2009).

At its core is the original ecological framework, according to which human development is the result of relationships between individuals and their ecological environment, 'conceived as a set of nested structures, each inside the other like a set of Russian dolls' (Bronfenbrenner, 1979, p.3). Building on what became known as the bioecological model, ideas regarding the role played by reciprocal interactions between individual biological and psychological characteristics and the immediate physical and social environment were added later (Bronfenbrenner and Ceci, 1994; Bronfenbrenner, 1995). These individual-environment interactions – referred to as proximal processes – were identified as the main drivers of development (Bronfenbrenner and Morris, 1998). The final iteration of the theory was the Process-Person-Context-Time (PPCT) model (Bronfenbrenner and Evans, 2000; Bronfenbrenner and Morris, 2006).

Process in the PPCT model are the activities and interactions (i.e. the proximal processes) between a biologically and psychologically developing person and the people and objects in their immediate surroundings, 'in either direction or both; that is, from the developing person to features of the environment, from features of the environment to developing person, or in both directions, separately or simultaneously.' (Bronfenbrenner and Evans, 2000, p.118). For example, parent-child interactions and playing with toys in the early years or peer-to-peer relations among older children and young adolescents. However, the outcome of these activities and interactions depends on the three remaining elements of the model. *Person* represents the biological, genetic and psychosocial characteristics of the individual, such as age, gender, health and personality which modify development. For example, there is evidence showing the interactive relationships between parents and children (the proximal process), family socio-economic position (a contextual factor) and socioemotional and behavioural outcomes in children are influenced by gender (Ashiabi and O'Neal, 2015). The *context* refers to the bioecological and psychosocial environment, which comprises a hierarchy of four interconnected systems (e.g. Figure 1.2). The first, the microsystem, is generally considered the most influential, since it is in home, school and neighbourhood environments that the most instrumental interactions between children and the people and objects in their surroundings are believed to occur (Wachs, 2010). The second, the mesosystem, represents relations between multiple, different microsystems. For example, the impact of chaotic neighbourhoods may diminish the beneficial effects a stable home environment has on child development (Evans and Wachs, 2010; Shamama-tus-Sabah, Gilani, and Wachs, 2011). Third is the exosystem, which is made up of contexts that children only experience indirectly. In other words, the environmental conditions that shape the environments children inhabit. The link between conditions in the caregivers' work environment and levels of family stress is an example of how these contexts might indirectly affect developmental outcomes (Bronfenbrenner, 1986).

Figure 1.2 An example of an integrated model of bioecological and psychosocial environments.



Source: Wachs, T.D. (2003). Expanding our view of context: The bioecological environment and development. In R. Kail (Ed.), *Advances in child development and behaviour* (Vol. 31, p. 365-411). New York: Academic Press. Reproduced with permission from Elsevier.

The fourth, which refers to characteristics of the natural, socio-cultural, political and economic environment is known as the macrosystem. The influence of these distal contexts on proximal processes could be wide-ranging, especially in LLMICs. For example, low levels of investment in essential services and infrastructure increases the risk to children and families from infectious and parasitic diseases (Freeman et al., 2013); the cultural production of norms and values restricts access to education for girls (Unterhalter et al., 2014); and among the many adverse effects of climate change, poor nutrition and economic disruption reduces school readiness and school performance (Hanna and Oliva, 2016). The final element of the PPCT model is *time* which refers to the timing of exposure of proximate processes to risks (e.g. whether during a critical or sensitive developmental stage), duration, frequency, and exposure intensity (Bronfenbrenner and Evans, 2000). *Time* also includes the influence of wider historical contexts on child development, such as periods of political crisis, economic recession and/or mass unemployment.

Bronfenbrenner's theory extends our understanding of the relationship between risk exposure and compromised development in several ways. For example, the framework identifies the reciprocal activities and interactions between children and young people and their immediate physical and social environment as the main motor of development. And that to be effective these proximal influences – which become more complex as children get older – need to occur predictably and in a stable environment (Bronfenbrenner and Morris, 2006). In doing so, the bioecological perspective offers a more precise definition of developmental risk. That is, viewed from within the bioecological framework a risk factor is any biological or psychosocial exposure that can compromise, directly or indirectly, the effective functioning of such proximal processes.

Testing this model is beyond the scope of this work, but the framework described in this section has several potentially important implications for this research. First, although the following chapters examine the impact of specific risks on child and adolescent development, it is important to recognise that bidirectional links between risk factors may mean exposure to one risk factor leads to the exposure to many (see Figure 1.2), particularly in LLMICs where risks are expected to cluster together

(Wachs and Rahman, 2013). For example, children exposed to poverty have a higher likelihood of experiencing nutritional deficiencies and living in neighbourhoods with inadequate water and sanitation facilities, elevating the risk from infectious diseases like severe childhood diarrhoea at the same time as vulnerability to infection is heightened as a result of malnutrition and lowered immunity, which in turn accelerates the loss of essential nutrients (Abubakar and van de Vijver, 2017). In addition, multiple exposures to developmental risks like these could cause an accumulation of risks over time (Evans and Kim, 2007; Evans, Dongping, and Whipple, 2013).

Second, the bioecological model was heavily influenced by Theodore Wachs and his work on the impact of the physical environment on development (e.g. Wachs, 1979, 1989, 1990). Wachs also emphasised the role played by both the biological and psychosocial dimensions of the *same* risk, which is crucial to understanding the way risk and protective factors determine developmental outcomes in low- and lower-middle-income countries (Wachs, 2010; Wachs and Evans, 2010).

Third, the basic mechanism underlying compromised child development is the disruption of key proximal processes, so the closer the risk exposure is to these processes the greater its influence (Wachs and Evans, 2010). However, specific exposures must be situated within the wider contextual environment, from stressful or unsanitary neighbourhood conditions to patterns of social exclusion linked to high youth unemployment or negative sociocultural attitudes towards girls' education.

1.4 The Philippine context

An East Asian country, the Philippines is an archipelago of over 7,000 islands in three main groups: Luzon, Visayas and Mindanao. The most recent census revealed a population of over 101 million people, which grew between 1980 and 2015 at an average annual rate of 1.5 percent (PSA, 2015). This comparatively high rate of population growth reflects the fact the country has only partially undergone the demographic transition seen regionally from the early 1960s (Mapa, 2015). In other words, instead of moving from a population characterised by high mortality and high fertility to one with low mortality and low fertility the mortality rate in the Philippines declined but fertility remained high. The result of this incomplete transition can be observed in the age structure of the population, with children under 19 years of age making up more than 40 percent of the total number of people (PSA, 2015).

In 2017, the Philippines' Human Development Index (HDI) value put the country in the medium human development category, tied in 113th out of 189 countries and territories with South Africa (UNDP, 2018). The country developed slowly between 1990 and 2017; progress was below the East Asian and Pacific country average and compared poorly to individual countries of similar population size and HDI ranking (UNDP, 2018). Although the percentage of the population in extreme poverty (on less than $1.90 per day) has dropped from more than 25 percent in 1990 to 7.8 percent in 2015, around 30 percent are still living below the low- to lower-middle-income poverty line of $3.20 (figures are obtained from the World Bank http://povertydata.worldbank.org/poverty/country/PHL). Using the Alkire-Foster method (Alkire and Foster, 2007, 2011), 37 percent of overall deprivation in the Philippines is attributable to the education dimension of poverty; 26 percent represents deprivations of health and nutrition; housing, water and sanitation is also 26 percent; while 11 percent is linked to employment (PSA, 2018).

Mothers and young children face considerable health challenges related to poverty and disease, including an under-five mortality rate of 27 deaths in 1000 live births; high levels of stunting (33%) and underweight (22%) among children under 5 years of age; and anaemia in pregnant women and children (25% and 39%, respectively) (UNICEF, 2016). There are also important disparities between urban and rural areas and the regions.

Despite progress, including passing major education legislation, the Philippines only partially achieved its 2015 Education for All goals (UNESCO, 2015). As of 2012, net enrolment in primary education was at more than 95 percent and there was an average annual increase in kindergarten enrolment of 8.5 percent between 2006 and 2012 (net enrolment in kindergarten had reached 77% by 2012-2013) (UNESCO, 2015). However, the net enrolment rate in secondary education was about 65 percent, while completion rates for both primary and secondary levels hovered around 70 percent, highlighting the scale of the challenge the country faces meeting the Sustainable Development Goal for education (SDG4) of universal access to quality primary and secondary education by 2030 (United Nations General Assembly, 2015). One relatively unusual characteristic of education in the Philippines is that boys are less likely to attend either primary or secondary school compared to girls and the gender gap in enrolment increases with age (Maligalig and Albert, 2008; Maligalig et al., 2010). School completion rates are also lower for males than for females, although being from poorer and/or more rural households significantly increases the likelihood of school dropout for children of either sex (Albert et al., 2012).

1.4.1 Cebu Longitudinal Health and Nutrition Survey

The data used in this research are from the Cebu Longitudinal Health and Nutrition Survey (CLHNS), a cohort study conducted in Metro Cebu.[1] Located in the Central Visayas region, the metropolitan area around the city of Cebu has undergone major demographic change since the 1980s; growing at an average annual rate of more than 4 percent until 2010, it has a current population of more than 2.8 million people making it the second most populous city area after Metro Manila (OECD, 2017). Due to rapid urban growth Metro Cebu reflects many of the economic and social opportunities associated with urbanisation as well as its environmental, health and infrastructural challenges, mirroring the rest of the Philippines as well as much of the wider developing world (Cohen, 2005). The study location is also ecologically diverse, comprising densely populated cities (Cebu City, Mandaue and Lapu-Lapu), the surrounding peri-urban areas, as well as smaller rural towns and nearby mountain and island communities (Adair et al., 2010).

The CLHNS used a stratified, single stage cluster-sampling approach, randomly selecting 17 urban and 16 rural *barangays* (a village or district).[2] Approximately 28,000 households were then surveyed for pregnant women and those due to give birth between May 1983 and April 1984 were included in the sample, making it representative of births in Cebu for this period. Altogether 3,327 women were

[1] The data are publicly available at https://dataverse.unc.edu/dataverse/cebu

[2] Stratification was by urban-rural residence and the clusters were *barangays* (Feranil, Gultiano, and Adair, 2008).

interviewed at baseline. Women who had multiple births, a stillbirth or miscarriage, refused to participate further or out-migrated were excluded (Adair et al., 2010). This left 3,080 women who had given birth to a single live infant. Data were collected during the pregnancy, immediately after the birth, and every two months for next 24 months. Although originally designed only as a two-year follow-up, women and children who could be traced were surveyed again in 1991-1992, 1994-1995, 1998-1999, 2002 and 2005 (Feranil, Gultiano, and Adair, 2008).

The sample attrition in the CLHNS required investigation. Specifically, into the extent to which attrition was non-random and whether non-random attrition led to selection bias. Sample selection bias arises from attrition when individual, family or community characteristics are linked to both the likelihood of attrition and the outcome variable of interest (Outes-Leon and Dercon, 2008). In other words, when non-random patterns of attrition distort results (Alderman et al., 2001). These patterns can be selective on both observable and unobservable characteristics but the focus here is attrition on observables. Although 39 percent of the original Cebu cohort had been lost to follow-up by 2005 the annual rate of attrition was just 1.8 percent. This compares favourably to other longitudinal low- and lower-middle-income country surveys of similar duration (Alderman et al., 2001; Outes-Leon and Dercon, 2008). The main cause of attrition was out-migration, which was estimated by the CLHNS team in Cebu to account for more than one-third of women and children lost to follow-up (Feranil, Gultiano, and Adair, 2008).[3]

[3] Infant mortality in the first 24 months was another major factor (Feranil, Gultiano, and Adair, 2008).

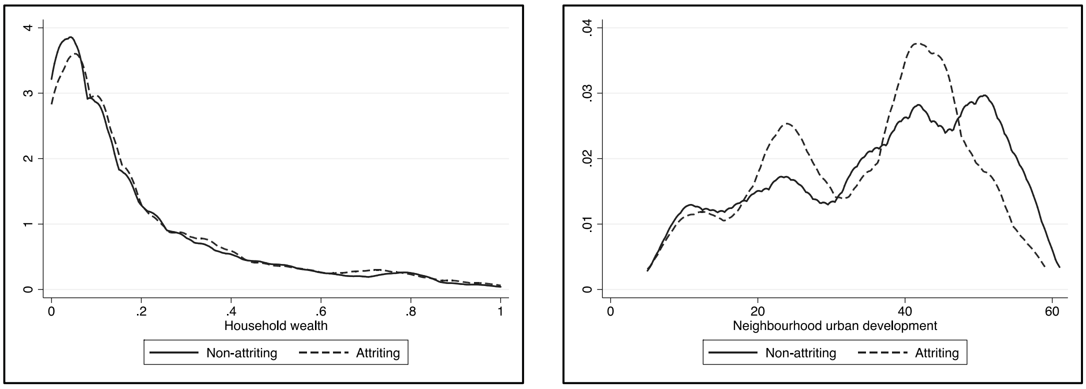


Figure 1.3 Wealth and urban development index kernel densities by attrition status.

The investigation of sample attrition involved searching for non-random patterns in observable individual, family and community characteristics. Figure 1.3 shows attriting households had similar patterns of wealth to non-attriting households and that attrition was primarily urban. More formally, the comparison of mean demographic, social and economic characteristics across attrition and non-attrition groups at baseline revealed attriting families had smaller children, but were better educated, wealthier, lived in less crowded conditions and in more developed urban neighbourhoods than non-attriting families (some differences were only borderline statistically significant) (see Appendix 1A). Table 1.1 reports the results from further tests of sample attrition using probit regression models.

Table 1.1 Probit regression predicting likelihood of attrition in CLHNS cohort.

	Variables tested individually		Variables tested jointly	
	Coef.	p-value	Coef.	p-value
Male	0.032	0.493	-0.017	0.743
Age	0.068	0.599	0.130	0.346
Birth weight	-0.000	0.694	0.000	0.227
Height-for-age	0.009	0.754	-0.019	0.625
Weight-for-height	0.025	0.445	0.037	0.355
Mother's height	0.003	0.661	-0.001	0.904
Mother's age	-0.002	0.663	-0.002	0.797
Mother primary school	0.018	0.826	-0.076	0.456
Mother secondary school	0.084	0.237	-0.089	0.291
Father primary school	0.035	0.637	-0.050	0.536
Father secondary school	0.130	0.003	0.129	0.097
Household income	0.020	0.046	0.001	0.933
Household wealth	0.082	<0.001	0.092	0.003
Household size	0.013	0.173	-0.013	0.503
Household crowding	-0.229	0.138	-0.536	0.061
Neighbourhood urban development	0.007	0.052	0.005	0.172
Constant	-	-	-0.324	0.804
Chi-square test (prob>chi2)	-	-	115.85	<0.001

Dependent variable = loss to follow-up between 1984 and 2005.

Tested individually (i.e. one at a time) paternal education, household income, household wealth and neighbourhood urban development were each predictive of loss to follow-up (although neighbourhood urban development was statistically significant at the lowest threshold). These results were similar to those obtained from tests of the equality of means (Appendix 1A). When the covariates were tested jointly the only coefficient that was significant at 95 percent confidence was for household wealth, although a joint χ^2 test indicated that collectively the covariates were highly predictive of attrition.[4] However, evidence of non-random differences between attriting and non-

[4] Outes-Leon and Dercon (2008) make the point that the pseudo R-squared from the probit regression gives an indication of the attrition that was not explained by the covariates in the model (i.e. the proportion of sample attrition that might be considered non-random). According to the pseudo R-

attriting households does not necessarily imply bias. The evidence collected from additional regression analyses indicates that the determinants of child development outcomes central to this research, such as school enrolment and cognitive test scores, were not statistically different for the full sample and attriting sample (see Appendix 1B and 1C). This suggests attrition on observable characteristics does not lead to biased estimates.

Sample attrition in the CLHNS is to some extent non-random with wealthier, better educated families living in more urban areas more likely to be lost to follow-up. However, there appears to be relatively little evidence to suggest sample selection bias will affect the results of this research.

squared from the probit regression reported in Table 1 the model explained just two percent of sample attrition, suggesting it was a largely random phenomenon.

1.5 Overview

This introductory chapter forms the first part of this work. It includes the research motivation, a summary of both the developmental risk factors faced by children and adolescents in low- and lower-middle-income countries and the theoretical framework for the analysis, as well as description of the country setting and dataset.

The second chapter examines the indirect effects of exposure to sanitation risks in early life on subsequent school survival. Existing research has focused on the impact of inadequate WASH conditions on infant mortality and morbidity in LLMICs, but little attention has been paid to the indirect consequences (i.e. the impact of health effects on later developmental outcomes). This paper contributes to the literature by estimating whether exposure to faecal-contaminated home environments are associated with the length of time children and young people in the Philippines spend in school before stopping or dropping out. The findings suggest the effect of infant exposure to contaminated conditions are not limited to early childhood and by shortening school life may adversely affect adult health and wellbeing.

The third chapter investigates the influence of residential mobility during early to middle childhood on cognitive performance at 11 years of age. Widely recognised as a risk factor in high-income country contexts, little is known about the disruption housing instability in the pre-primary-school period causes children living in low- and lower-middle-income country neighbourhoods. Taking advantage of multiple waves of CLHNS data this paper uses inverse probability weighting to assess the role played by changes in neighbourhood surroundings and the mediating effect of move quality and school moves.

The fourth chapter tests the associations between the adolescent mental health problems and social marginalisation, particularly disengaged youth who are neither in education, employment nor training (NEET). The term NEET was originally used to measure disengagement from major socialising institutions during the transition from school to work in the UK but is now applied widely across HIC and LLMIC contexts. However, while the issue of adolescent mental ill health in developing countries is beginning to attract the research attention it warrants, its relationship with the soaring rates of school dropout and youth unemployment found in many LLMICs is far from established.

Finally, the fifth chapter presents conclusions and highlights some policy recommendations based on the findings of this research.

Chapter 2.

Does Infant Exposure to Sanitation Risks in the Physical Home Environment Predict School Survival?

2.1 Introduction

The link between early childhood experience and the development of children and adolescents in less industrialised countries is well established (Grantham-McGregor et al., 2007; Walker et al., 2011; Black et al., 2017). Child development refers to the highly complex process of acquiring sensory-motor, cognitive, language, social-emotional and self-regulation skills and the first few years of life represent a developmentally critical period (Black, Gove, and Merseth, 2017). For that reason, infants and young children are particularly vulnerable to exposure to biological, psychosocial or environmental risk factors, which can cause developmental deficits and/or delays affecting school readiness and school performance leading to poorer individual health, higher fertility rates and lower household income in adulthood (Engle et al., 2007; Grantham-McGregor et al., 2007). In low-income and lower-middle-income countries (LLMICs) major poverty-related risk factors include linear growth retardation, inadequate cognitive stimulation, iodine deficiency and iron deficiency anaemia (Walker et al., 2007). The evidence also points to the risks posed by intrauterine growth restriction, malaria, lead exposure, HIV, maternal depression and societal violence (Walker et al., 2011). This paper focuses on the associations between conditions in the physical environment in which children are born and raised – specifically, the risk of faecal-oral transmission through contact with excreta in the physical home environment – and duration of child and adolescent schooling.

Infant exposure to sanitation risks may adversely affect subsequent school outcomes by compromising child development in the first 24 months. Linear growth retardation (commonly known as stunting) is a widely used marker of impaired or delayed child development (Leroy and Frongillo, 2019), which prior research has linked to water, sanitation and hygiene (WASH) conditions in the home environment. At two years of age children in a Peruvian birth cohort living with the worst water and sanitation conditions were shorter than children living in the best conditions (Checkley et al., 2004). The same study also found lack of adequate sewage disposal explained a similar growth deficit. Researchers using Young Lives study data for Ethiopian, Indian, Peruvian and Vietnamese children observed access to improved sanitation was more often associated with reduced stunting than access to improved water and that stunting was less common among children who had had access to a flushing toilet at 12 months when children were assessed at five and eight years of age (Dearden et al., 2017). Another study conducted in rural Bangladesh also found children from more sanitary households had higher height-for-age z-scores (HAZ) (Lin et al., 2013). However, the experimental evidence is more mixed. The results of a Cochrane review of five randomised-controlled trials suggest access to improved drinking water sources and increased handwashing only has small effects on HAZ, although the review was unable to identify any sanitation interventions (Dangour et al., 2013). Cluster-randomised trials of several subsequent sanitation interventions have shown exposure to increased sanitation risks is significantly associated with stunting (Hammer and Spears, 2013; Pickering et al., 2015). For example, a community-led intervention in Mali, West Africa, which encouraged communities to build their own toilets, found that after one and a half years there was more latrine use and less open defecation in villages that received the intervention, while children under five years of age were taller and significantly less likely to be stunted than children in the control villages (35% compared to 41%) (Pickering et al., 2015). Yet results from trials of individual and combined nutrition, water, sanitation and hygiene interventions in Bangladesh and rural India were less conclusive, revealing a decreased incidence of childhood diarrhoea but no improvement in linear growth (Clasen et al., 2014; Luby et al., 2018).

Undernutrition as indicated by stunting affects brain and motor development, physical growth and physical activity levels. It is also believed to indirectly influence the quality and quantity of essential interactions between caregiver and child as well as between children and their immediate surroundings (Pollitt et al., 1995; Grantham-McGregor et al., 2007). Recent experimental findings support the existing evidence connecting child nutrition to a number of different developmental domains (Stewart et al., 2018; Tofail, et al., 2018). Stewart and colleagues (2018) tested the effects of separate and combined WASH and nutrition interventions on a range of outcomes in a group of children from Kenya, including language, gross motor and personal social development skills, and found the combined intervention had a protective effect on motor skills. Tofail and colleagues (2018) also observed consistent improvements in communication, gross motor and personal social skills for all children receiving combined WASH and nutrition interventions in Bangladesh. Additionally, evidence from multiple observational studies underscores the association between linear growth in early life and developmental outcomes in older children, including measures of motor and cognitive skills, academic attainment in standardised tests, age at initial enrolment, likelihood of grade repetition as well as school dropout and completion rates (Daniels and Adair, 2004; Victoria et al., 2008; Adair et al., 2013; Crookston et al., 2013; Fink and Rockers, 2014; Sudfeld et al., 2015).

2.1.1 Pathways between WASH conditions and subsequent schooling

The connection between infant exposure to faecal-contaminated environments and subsequent schooling is hypothesised to run indirectly via the effects on child nutrition, since WASH conditions have previously been linked to linear growth, which reliably predicts health and human capital outcomes in older children and young adults (Santos et al., 2008; Adair et al., 2013). Possible mechanisms therefore include child diarrhoeal illnesses and other parasitic infections, given a large share of the infectious disease burden among children under five years of age is attributable to WASH-related risk factors in LLMICs and infection and malnutrition are known to be intricately linked (Black et al., 2008; Katona and Katona-Apte, 2008; Black et al., 2013; Prüss-Üstün et al., 2014).

The evidence suggests repeated bouts of severe childhood diarrhoea are associated with an increased risk of stunting in early childhood as well as with modestly retarded growth over the long term (Adair and Guilkey, 1996; Checkley et al., 2008; Richard et al., 2013). However, the contribution of diarrhoeal diseases to growth retardation and the risk of long-term sequalae is not fully understood. One reason for this is that it is extremely difficult to disentangle relations between undernutrition and the increased frequency, duration and severity of diarrhoeal episodes (Guerrant et al., 1992; Brown, 2003). Another reason is that some randomised-controlled trials of sanitation interventions have observed significant differences in stunting between treated and control group children without corresponding differences in the incidence of diarrhoea (Pickering et al., 2015). An alternative mechanism to childhood diarrhoea is environmental enteropathy, also known as environmental enteric dysfunction (EED) (Humphrey, 2009; Ngure et al., 2014). Like diarrhoea, EED is spread through the faecal-oral transmission of pathogens or parasites from human or animal faeces to an individual's mouth. Unlike diarrhoea, EED is a typically asymptomatic infection that is believed to cause chronic enteric inflammation, leading to reduced nutrient absorption of the intestine and increased intestinal permeability (Keusch et al., 2014; Crane et al., 2015). Research testing for EED is limited but results from a study conducted in Bangladesh observed higher height-for-age z-scores and lower lactulose to mannitol (L:M) ratios in children from less faecal-contaminated environments, which was interpreted as evidence supportive of the EED hypothesis. (Lin et al., 2013). The estimated prevalence also suggests EED is widespread in low-income and lower-middle-income countries (Crane et al., 2015).

In summary, the existing evidence indicates conditions in the WASH environment may compromise early childhood development. It also shows that compromised development measured in terms of retarded linear growth is associated with undesirable outcomes in the short and medium term across a wide array of developmental domains (Leroy and Frongillo, 2019). Nevertheless, there remain several important gaps in our understanding.

First, relatively little is known about the longer-term consequences of WASH-related risks in infancy since there is scarce research evidence on the effects of exposure on outcomes measured in children over five years of age. Second, a small handful of studies have examined the impact on school-age cognition and learning outcomes and found positive effects associated with access to improved water and sanitation (Bhalotra and Venkataramani, 2013; Wulan et al., 2015; Dearden et al., 2017b). Wulan and colleagues (2015) estimated the effect of prenatal exposure to unsafe water and sanitation in Indonesia on cognition when children were followed up nine to twelve years later and found the women who had had access to safe WASH facilities during their pregnancy had children who performed better in testing. Dearden and colleagues (2017) revealed access to improved water and sanitation at 12 months was associated with higher standardised test scores measured at five and eight years old using longitudinal data from the multi-country Young Lives study. Bhalotra and Venkataramani (2013) exploited variation in the risk of waterborne disease caused by a major water reform in Mexico to investigate the impact of reduced exposure to WASH-related risks on cognitive development and academic achievement. The study assessed cognitive functioning when children were between nine and 14 years of age and maths and reading skills at 15 years of age, observing that a reduction in childhood diarrhoea associated with water improvements at 12 months predicted better subsequent test results for girls. However, though valuable these studies relied on limited or one-off measures of child development. Third, almost no consideration has been given to the impact on school enrolment. Enrolment in LLMICs is not only a matter of school performance, but also poverty and gender (Hunt, 2008). For example, the longer girls stay in school the later they marry and the older they tend to be when they first have children (Roest, 2016). The evidence from the two studies found in a review of the published and grey literature to have considered WASH effects on school enrolment is mixed. Results from a cross-sectional study conducted in Brazil implied concurrent WASH conditions were associated with completed years of schooling (Ortiz-Correa, Fiho, and Dinar, 2016), but a randomised-controlled trial of the effects of latrine adoption and ownership in India found young children from intervention villages were no better off in terms of school enrolment ten years later than children from areas that did not receive the intervention (Orgill-Meyer and

Pattanayak, 2017). Nonetheless, neither study tested the effects of exposure to inadequate sanitation in the first 1,000 days – the sample in the India trial comprised households with children under the age of five but age at exposure was unknown – nor were they able to model the fundamentally time dependent nature of enrolment in school. Stopping school can be a long, drawn-out process that fluctuates during important periods of school transition and may involve patterns of temporary dropout and reenrolment (Hunt, 2008).

The purpose of this study was to respond to these evidence gaps by investigating the indirect effects of inadequate WASH conditions in early childhood on subsequent schooling. Specifically, testing the associations between infant exposure to sanitation risks and school survival using longitudinal data collected from a cohort of children from Cebu, Philippines. Prior research has shown sex differences in cognitive and socioemotional development, so child gender was tested for its potential moderating influence on the relationship (McCoy et al., 2016). This study's main hypotheses were: (1) infant exposure to increased sanitation risks is associated with subsequent school stoppage and shorter overall school life, but accounting for family socio-economic position and community characteristics like population density attenuates the associations; (2) the probability of school survival is lower for the exposed group of children during periods of school transition; (3) female children are more vulnerable to the impacts of contaminated home environments compared to males.

2.2 Data

The data used in this analysis were collected as part of the Cebu Longitudinal Health and Nutrition Survey (CLHNS); a cohort study conducted in Metro Cebu, which comprises the second largest city in the Philippines and its surroundings. The study area included densely populated urban centres, the peri-urban periphery, as well as smaller rural towns and nearby mountain and island communities (Adair et al., 2010). A single stage cluster-sampling approach was used to randomly draw 17 urban and 16 rural administrative units (*barangays*). These were surveyed for pregnant women and those due to give birth between May 1983 and April 1984 were selected into the sample, making it representative of births in Cebu for this period. In total 3,327 women were interviewed at baseline. Women who had multiple births, a stillbirth or miscarriage, refused to participate further or out-migrated, were excluded (Adair et al., 2010). This left 3,080 women who had given birth to a single live infant. Data were collected in the prenatal period, immediately after birth, and then bimonthly for 24 months. Subsequently, there were follow up surveys in 1991, 1994, 1998, 2002 and 2005. The CLHNS collected richly detailed sociodemographic and health data for each child and family, in addition to information about the services and infrastructure in the community in which children were born and grew up. It is well suited to analysing the influence of environmental conditions in early childhood on school outcomes, covering each relevant stage of the life course. Data of particular importance to this study included pre- and postnatal environmental characteristics collected at household and neighbourhood levels, linear growth in the first 24 months, school entry in 1991 and school enrolment in all subsequent rounds.

1,888 children were still participating in the survey in 2005, representing 61 percent of the original cohort. The attrition rate of less than two percent per annum was primarily driven by out-migration and compares favourably to rates for similar longitudinal surveys conducted in low-income and lower-middle-income countries (Outes-Leon and Dercon, 2008). Adair and colleagues (2010) confirm that 232 index children are known to have died and that the majority of deaths occurred within the first two years of life. The final sample of 1,664 children (799 females and 865 males) was based on the availability of complete data from birth until 21 years of age.

Schooling outcome (dependent variable). School survival was based on current enrolment in school or enrolment in the most recent school year if the survey was carried out during the summer recess (in the Philippines the school calendar runs June to March). The failure event was therefore the first time an individual was not enrolled in school having previously been enrolled. Time of enrolment was assumed to be the same for everyone (i.e. the beginning of the academic year). This simplifying assumption ignores late enrolment as well as enrolment and irregular attendance. Although measures based on enrolment can be inconsistent with expected age for grade – for example, when children who are enrolled in school have fallen behind compared to their age-group as a result of grade repetition – enrolment in school is expected to be correlated with overall school achievement (Hunt, 2008). For children and adolescents in many LLMICs stopping school early can also lead to substantively different and potentially riskier transitions to adulthood in terms of employment, marriage and fertility (Jensen and Thorton, 2003). The focus of this analysis was new cases of non-enrolment, investigating the environmental determinants of the first incidence of school stoppage. However, patterns of temporary withdrawal from school caused by child or family-member ill health, overdue school fees and seasonal labour requirements are not uncommon in low-income and lower-middle-income country contexts (Hunt, 2008). In the Cebu cohort there were a fairly small number of cases in which dropout was followed by reenrolment at a later stage (9%). In the main analysis temporary withdrawals were censored and subsequent reenrolment in school ignored, although a separate outcome variable for temporary dropout was also created for robustness.

WASH environment (independent variable). The exposure variable was constructed according to (1) the WHO/UNICEF Joint Monitoring Program (JMP) definition of 'unimproved' sanitation, which includes open pit latrines, hanging and bucket toilets, and (2) the practice of open defecation (WHO, 2015: p.52). Shared sanitation services were not included since sharing circumstances can vary widely – for example, toilets can be shared between two families living in the same compound or by a whole community – making the differentiation between hygienic and unhygienic facilities extremely problematic (Evans et al., 2017). The main exposure of interest (sanitation risk) was measured by a binary variable that took one if a family had unimproved sanitation facilities likely to be favourable to faecal-oral transmission and zero otherwise. However, lack of access to improved sanitation alone cannot fully capture sanitation risk. For example, informal settlements with inadequate drainage and informal housing structures are both likely to increase infant exposure to excreta (Rheingans et al., 2012). Two additional exposure variables were therefore created to estimate the combined effects of exposure to sanitation risks and household and neighbourhood vulnerabilities. The first variable measured exposure to unimproved sanitation facilities for infants living in houses constructed using lightweight materials like *nipa* (palm) and bamboo. The second exposure measured lack of access to improved sanitation, such as flush toilets, in households where poor drainage capacity was also observed in the surrounding area. These indicators were used in robustness checks. The variables used reported and interviewer-observed information about the household environment collected as part of the baseline survey (carried out in the sixth or seventh month of the mother's pregnancy).

Other covariates. The choice of variables used to adjust for potential confounding was based on a review of earlier studies, in particular research investigating the early childhood effects of environmental conditions on morbidity and linear growth as well as the link between growth and child development outcomes in the Cebu cohort (VanDerslice, Popkin, and Briscoe, 1994; VanDerslice and Briscoe, 1995; Daniels and Adair, 2004). Individual and parental characteristics that were expected to be associated with conditions in the WASH environment and school outcomes included child sex, age and birthweight, whether they were breastfed, mothers' age, and whether mother and father completed primary and secondary school. The maternal decision to breastfeed has been identified as having an important protective effect against infant morbidity and malnutrition, especially with respect to the risk from diarrhoeal diseases (VanDerslice, Popkin, and Briscoe, 1994). Breastfeeding was a dummy variable that took one if mothers reported that they had breastfed the infant and zero if not. Potential household and barangay confounders included number of people living in the house, household assets and income, as well as neighbourhood urban development. The household assets index measured reported ownership of the house the respondent was living in, an electric fan, furniture, a jeepney (a small van), colour television, cassette recorder and refrigerator. The index was calculated using polychoric principal component analysis (PCA). Kolenikov and Angeles (2004) demonstrated the appropriateness of using PCA to aggregate discrete data into a single measure by estimating the polychoric correlations between several ordinal variables to assess household socio-economic status. Accordingly, the first principal component was treated as an index of household assets or wealth with lower index values indicating lower levels of wealth and higher values corresponding to higher levels. Total household income (in pesos) was a truncated variable derived by assigning the 99th percentile value to outlier values greater than the 99th percentile, and zero to values less than zero (income net of investments in livestock or machinery will sometimes produce negative values). The variable was expressed as a multiple of 100. The neighbourhood urban development index was developed by members of the CLHNS team and components were selected according to the results of an earlier factor analysis (McDade and Adair, 2001; Dahly and Adair, 2007). The measure combined neighbourhood population, population density, communications (cell phones, the internet, cable TV etc.), transportation (the density of paved roads, availability of public transportation), educational facilities, health services, and markets (including grocery stores and gas stations), each of which was assigned values from zero to 10 resulting in a scale between zero and 70.

2.3 Data analysis

This study modelled schooling using discrete time survival analysis. This approach was appropriate because it accounted for the way the time-to-event was recorded and allowed for right censoring, which occurs when individuals have not experienced the event of interest (school stoppage) by the end of the period under observation.

Time is represented discretely in a survival analysis when the exact time the event occurs is unknown. The CLHNS collected information regarding current enrolment status (i.e. enrolment for the most contemporary school year at the time of the survey). However, since there were two or three academic years to every survey round there was no way to determine whether children who did not enrol for the year in which the survey was undertaken only failed to enrol in that school year alone. The solution to the problem was to treat the period between survey rounds as a discrete interval of analysis, restructuring the dataset so that there was one record for every child in each interval.

A single risk model was estimated using logistic regression. The hazard function h_i was the conditional probability that an individual stopped school during interval t given they had been enrolled in all previous periods:

$$h_i(t) = P(y_i(t) = 1 | y_i(t-1) = 0) \tag{1}$$

where $y_i(t)$ was the binary indicator of event occurrence at time t. In this hazard framework, once an individual stopped school for the first time they were censored.

Bounded between zero and one the hazard was related to the covariates using the following logistic function:

$$logit[h_i(t)] = \log\left[\frac{h_i(t)}{1-h_i(t)}\right] = \alpha(t) + \beta x_i(t) \qquad (2)$$

where $x_i(t)$ was a set of fixed or time-invariant factors including WASH exposure in infancy, mother's height and education, as well as time-varying variables like child age and household assets. Variation in $h_i(t)$ over time was captured by $\alpha(t)$ which were interval specific intercepts representing the baseline hazard i.e. the likelihood of experiencing the school failure event in each discrete period or interval when all the covariates were zero. Time was treated as categorical so $\alpha(t)$ was a set of dummy variables for intervals 1991 to 1994, 1994 to 1998, 1998 to 2002 and 2002 to 2005. The effect of exposure on the probability of the failure event was described in terms of the hazard ratio. An estimated hazard ratio of greater than one was interpreted as evidence that exposure to sanitation risks increased the likelihood that a child experienced the failure event compared to an unexposed child. Ratios of less than one indicated exposure reduced the odds of event occurrence. The effect of the covariates is usually assumed to be constant within a given time period – commonly known as the proportional hazards assumption – however, in a discrete time analysis the probability of the event occurring is estimated regardless of the duration of exposure time within each interval so the assumption was not required.

Equation (2) was then extended to incorporate *barangay* or neighbourhood specific random effects to account for between cluster heterogeneity in the hazard of the occurrence of the outcome. This random intercept or shared-frailty model can be written as follows (Goldstein, 2011):

$$logit\left[h_{ij}(t)\right] = \alpha(t) + \beta x_{ij}(t) + u_j \qquad (3)$$

where i were level 1 units (children in households), j were level 2 units (*barangays*) and the random effects were assumed to be normally distributed $u_j \sim N(0, \sigma_u^2)$. Importantly, in this model the interpretation of β is the barangay-specific effect of x or the probability of experiencing the failure event for children exposed to different WASH conditions from the same *barangay* or neighbourhood.

In addition, since two differing outcomes were possible a competing risks model was also run as a robustness check, simultaneously modelling school stoppage and temporary school dropout using multinomial logistic regression. Under this framework, the hazard function h_i in equation (1) becomes the conditional probability that individual i stopped school ($y_i = 1$) or was temporarily withdrawn ($y_i = 2$) so long as neither event had occurred in any period before time t.

2.4 Results

Sample characteristics. The analytic sample was defined by the availability of data, which made it potentially vulnerable to selection bias arising from non-random attrition. To assess the extent of the problem baseline sociodemographic characteristics of children and adolescents with complete data were compared to individuals excluded due to missing or otherwise incomplete information (Table 2.1). Young people with complete data were not significantly different in terms of birth length, height-for-age z-scores at 24 months, mothers' height and age, or the proportion who were male, compared to individuals in the missing data group. However, individuals included in the analysis came from larger families and weighed more at birth, which was expected since infant death was a cause of loss to follow-up (Daniels and Adair, 2004).

Table 2.1 Characteristics of children and families in CLHNS at baseline.

Characteristic	Missing (n=1396)	Complete (n=1664)	*p*-value[1]
	means ± SD		
Birth weight, g	2976.5 (459.8)	3008.4 (433.5)	0.049
Birth length, cm	49.2 (2.2)	49.3 (2.1)	0.192
HAZ	-2.3 (1.1)	-2.3 (1.1)	0.388
Mother's height, cm	150.7 (5.0)	150.6 (5.2)	0.556
Mother's age, y	26.0 (6.1)	26.1 (5.9)	0.455
Mother's education, y	7.7 (3.7)	7.4 (3.7)	0.026
Household income, pesos	285.1 (562.3)	276.8 (479.1)	0.659
Household assets	0.04 (1.1)	-0.03 (1.0)	0.084
Household size	5.7 (2.9)	5.7 (2.8)	0.953
Neighbourhood urban development	31.9 (12.1)	29.5 (12.9)	<0.001
	n (%)		
Male	739 (54)	865 (52)	0.221
Unimproved sanitation	397 (29)	616 (37)	<0.001
Unimproved water	96 (7)	233 (14)	<0.001
No private toilet	520 (38)	716 (43)	0.002

[1] Values of *p* correspond to one-way ANOVA *F*-test statistics (continuous variables) or are based on Pearson's chi-square test (categorical variables).

Mothers in the missing data group had a higher level of education compared to those in the analytic sample and excluded families lived in more developed neighbourhoods according to an index measuring community characteristics including infrastructure and access to health and education services. Families with missing data were also financially better off based on an indicator of household assets, although no significant differences were found in reported household income. Environmental conditions also differed meaningfully across the two groups; access to improved water and sanitation facilities was greater among attriting households than in the non-attriting group. Since households included in this analysis were less educated and had a poorer standard of living compared to those excluded as a result of missing data the effects of exposure to sanitation risks on schooling estimated in this study may be larger than in the population. Nevertheless, the presence of non-random attrition is not necessarily evidence of selection bias and this study’s findings still make an important contribution to our understanding of the role played by risks in the physical environment for long-term education outcomes in lower-middle-income countries like the Philippines.

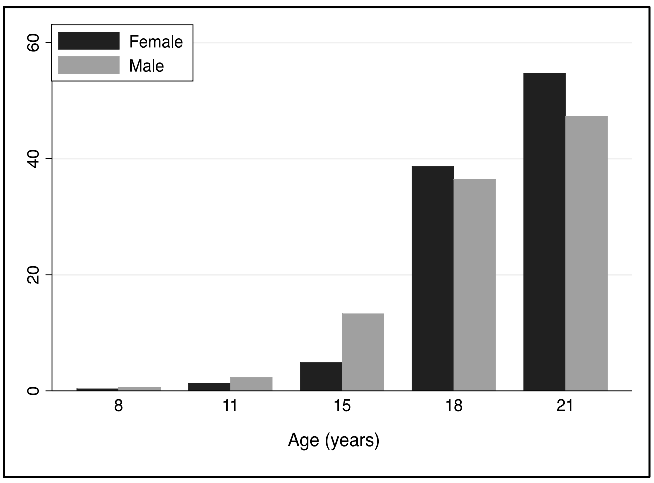


Figure 2.1 Distribution of new cases of school stoppage in males and females in the Cebu cohort.

The distribution of new cases of school stoppage for male and female children is shown in Figure 2.1. The number of cases between eight and 11 years of age was extremely low, which is consistent with the period of compulsory education in the Philippines at the time. Yet mandatory schooling did not include secondary level education, and this was reflected in an increase in the number of school stoppages by the time children were 15 years old. The largest gender gap was also at this age, with the proportion of males stopping school three times greater than the proportion of females (19% and 6% respectively). Peak school stoppage was between 18 and 21 years of age. The number of female cases in those years indicates more young women stayed in school for longer compared to their male cohort peers.

Table 2.2 Results of multilevel logistic regression of school stoppage (odds ratio).

	Model 1			**Model 2**			**Model 3**		
	OR	95% CI		OR	95% CI		OR	95% CI	
Unimproved sanitation	1.638***	1.389	1.932	1.298**	1.094	1.539	1.181*	0.993	1.405
Male	1.230**	1.076	1.405	1.273***	1.112	1.458	1.282***	1.119	1.469
Age	1.330***	1.155	1.532	1.345***	1.166	1.552	1.380***	1.195	1.594
Breastfed	1.215	0.898	1.643	0.905	0.664	1.235	0.843	0.616	1.153
Birth weight	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
Mother’s age	-	-	-	0.980***	0.969	0.991	0.982**	0.971	0.994
Mother primary school	-	-	-	0.843*	0.715	0.995	0.888	0.752	1.049
Mother secondary school	-	-	-	0.578***	0.481	0.695	0.662***	0.548	0.801
Father primary school	-	-	-	0.872	0.735	1.035	0.891	0.750	1.057
Father secondary school	-	-	-	0.717***	0.604	0.850	0.783**	0.658	0.932
Urbanicity	-	-	-	0.996	0.989	1.003	0.997	0.990	1.004
Household size	-	-	-	1.024*	0.999	1.050	1.033**	1.007	1.060
Household assets	-	-	-	-	-	-	0.800***	0.743	0.862
Household income	-	-	-	-	-	-	0.986	0.970	1.003
Between Barangay	0.103**	0.047	0.224	0.063*	0.027	0.151	0.061*	0.026	0.145

*** $p < 0.001$, ** $p < 0.05$, * $p < 0.1$.

Survival analysis. *Hypothesis 1* (infant exposure to increased sanitation risks is associated with subsequent school stoppage and shorter overall school life, but accounting for family socio-economic position and community characteristics like population density attenuates the associations): Results from the multilevel logistic regression presented in Table 2.1 show that exposure to unsanitary conditions in early childhood was associated with poorer school outcomes. Specifically, for children from households without access to improved sanitation the chances of stopping school and a shorter overall school life were higher compared to children born and raised with access to improved WASH facilities. The results from Model 1 show that exposure to uncovered latrines and/or open defecation from birth was associated with significantly higher odds of stopping school although adjusting for characteristics including parental education and level of neighbourhood urban development in Model 2 weakened the associations. In the fully adjusted model (Table 2.1, Model 3), which also accounted for household assets and income, the association between unimproved sanitation and school stoppage was attenuated further (and was only borderline statistically significant). Yet exposure to sanitation risks still had an odds or hazard ratio of 1.18 (95% CI 0.99-1.41, p =0.06). In other words, for two children who were the same with respect to the observed covariates in the model but differed in terms of sanitation exposure, the odds of not being enrolled in school were higher for the child exposed to faecal-contaminated home and play environments than for the child whose family had exclusive or shared access to a flush toilet. Re-estimating the model using measures of sanitation risk which incorporated specific household and neighbourhood vulnerabilities underscored these findings. These checks of robustness showed that exposure to unimproved sanitation and/or open defecation combined with informal housing structures or inadequate drainage modestly increased the negative effect on subsequent schooling (see Appendix 2A).

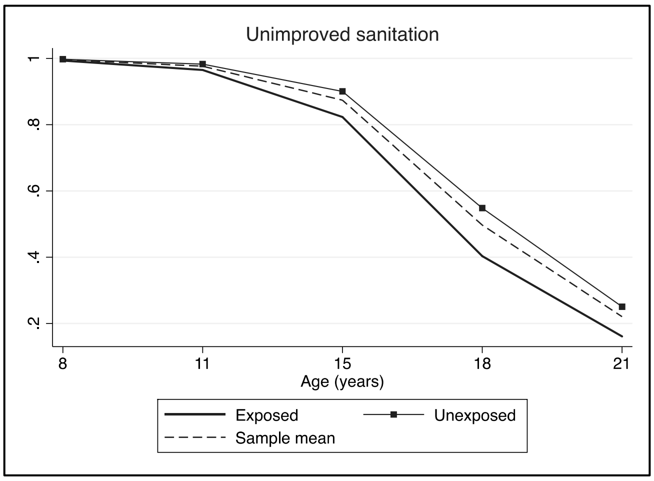


Figure 2.2 Predicted school survival for children exposed to unimproved sanitation in infancy.

Hypothesis 2 (the probability of school survival is lower for the exposed group of children during periods of school transition): Figure 2.2 shows infants who experienced exposure to faecal-contaminated conditions at home had lower predicted probabilities of being in school from 11 years of age. However, the gap in school survival increased between 15 and 18 years of age, suggesting children exposed to adverse physical environments in early life may have been less prepared or able to negotiate the post-secondary school transition.

Figure 2.3 Predicted school survival for male and female children exposed to unimproved sanitation in infancy.

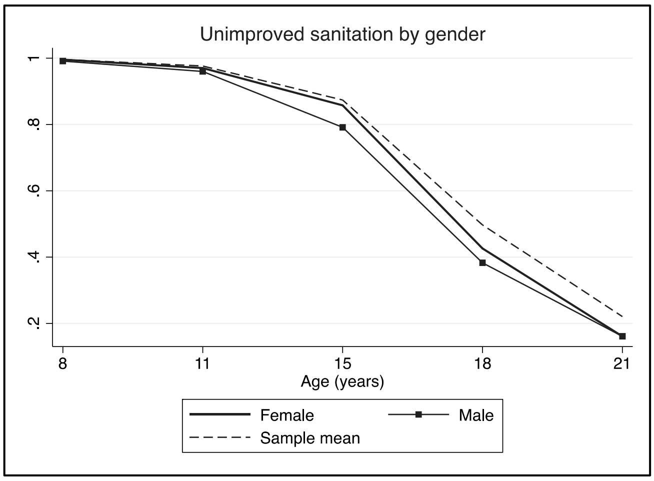


Figure 2.3 shows mean probabilities of school survival were lower for exposed males than females between 11 and 21 years of age or from the end of compulsory primary-level education. This is consistent with the pattern of male and female school enrolment in both the Cebu cohort (see Figure 2.1) and the Philippines nationally, which shows enrolment rates among boys and young men below those of girls and young women (Maligalig and Albert, 2008). The predicted school survival of female children exposed to unsanitary conditions was not much lower than the sample mean during primary- and secondary-level education (the period ending when female children were 15 years of age). However, between 15 and 18 years of age female survival trended more sharply downwards until it was similar to exposed male survival. This suggests risks in the physical environment in early childhood may have levelled some of the educational advantages young women otherwise had over their male peers.

Table 2.3 Results of multilevel logistic regression of school stoppage by gender group (odds ratio).

	Females			Males		
	OR	95% CI		OR	95% CI	
Unimproved sanitation	1.512***	1.162	1.968	0.932	0.734	1.183
Male	-	-	-	-	-	-
Age	1.067	0.823	1.384	1.081	0.862	1.355
Breastfed	0.481**	0.309	0.748	1.504	0.913	2.476
Birth weight	1.000	1.000	1.000	1.000	1.000	1.000
Mother's age	0.987	0.970	1.005	0.977**	0.963	0.992
Mother primary school	0.900	0.696	1.164	0.850	0.679	1.064
Mother secondary school	0.532***	0.394	0.718	0.779*	0.607	1.001
Father primary school	0.833	0.637	1.088	0.940	0.747	1.183
Father secondary school	0.729*	0.556	0.957	0.850	0.675	1.072
Urbanicity	1.000	0.989	1.010	0.991	0.981	1.002
Household size	1.034*	0.995	1.076	1.035*	1.000	1.071
Household assets	0.838**	0.745	0.944	0.772***	0.700	0.851
Household income	0.990	0.964	1.017	0.986	0.965	1.009
Between Barangay	0.070	0.020	0.247	0.192**	0.088	0.418

*** p < 0.001, ** p < 0.05, * p < 0.1.

Hypothesis 3 (female children are more vulnerable to the impacts of contaminated home environments compared to males): Gender differences were also observed when the fully adjusted model was run separately for males and females. The results presented in Table 2.3 suggest that among female children exposure to sanitation risks was associated with a higher chance of school stoppage, while the equivalent odds of stopping school in the male group were lower than the odds of continuing (although the point estimates for males should be interpreted with caution in view of their imprecision). Wald tests confirmed the presence of significant differences between estimates obtained for each gender group ($\chi^2(18)$=116.5, *p*<.001). In addition, results from two-way and three-way interactions models (sanitation risk by gender; gender by stunting; sanitation risk by gender by stunting) also suggest sanitation risks in early childhood may have been worse for female children and for stunted infant girls in particular (Appendix 2B).

The results presented thus far were from the single risks model, which estimated the probability of the first incidence of school stoppage with temporary dropout events censored. However, for robustness the two events were also estimated jointly in a competing risks model, allowing the hazard of the occurrence of school stoppage and dropout to be correlated with the aim of assessing the extent to which the results obtained from the single risks model were sensitive to model specification. Appendix 2C shows estimates from a multinomial logistic regression (i.e. the competing risks model) were similar to those obtained via the single risks model (RR=1.24, 95% CI 1.06-1.46, p=.007) suggesting the main modelling approach taken in this study was appropriate.

2.5 Discussion

Using longitudinal data collected from a cohort of children living in Metropolitan Cebu, the second-largest urban area in the Philippines, this study examined the effects of infant exposure to sanitation risks on subsequent school survival. Children who had no access to improved sanitation in infancy were found to have higher odds of stopping school and a shorter overall school life than children from families with improved facilities, such as flushing toilets or well-constructed latrines, although estimates were attenuated after accounting for household and community characteristics. Alternative measures of sanitation risk, incorporating household and neighbourhood vulnerabilities like informal housing construction materials and inadequate drainage, also indicated exposed infants faced increased odds of school stoppage. Despite the pattern of male and female enrolment in the cohort as well as nationally, which shows girls are generally enrolled in school for longer than boys, young women exposed to contaminated conditions as children were only as likely as their male peers to be in school by the time they were 18 years old. This suggests exposure to sanitation risks increases the likelihood that young women discontinue post-secondary education and do not receive the social and economic benefits of further and/or higher education. Results from the subgroup analysis also indicated differential effects for males and females. Remarkably few studies have addressed the long-term consequences of conditions in the physical home environment in low-income and lower-middle-income countries. So, these findings have potentially important implications for policymakers and development practitioners, highlighting how the possible effects of faecal contamination extend beyond early childhood and may influence the development of competencies necessary to take full advantage of the learning opportunities children and adolescents have at school.

This study found that children exposed to inadequate WASH environments had poorer school outcomes compared to children from households with access to improved facilities, consistent with the results from one of the two previous studies known to have examined the indirect influence of WASH on school enrolment. In line with the present study's findings, Ortiz-Correa, Filho and Dinar (2015) found access to improved water and sanitation services in Brazil had a positive effect on completed years of schooling for a group of children between six and 18 years of age. The same study also uncovered evidence that girls and young women were worst affected, although it was limited to testing the effects attributable to the WASH-related health burden and differential time usage – in other words, the time adults and children spent fetching water – due to the cross-sectional nature of the data. In contrast, this study's findings were not consistent with the second study identified in the literature review, which evaluated the long-term human capital effects of latrine adoption in India and found no association with school enrolment measured in a single ten-year follow-up (Orgill-Meyer and Pattanayak, 2017). Possible explanations for this discrepancy include differences in the WASH exposure used in the two studies – this analysis relied on standard WHO/JMP definitions of unimproved sanitation – and the lack of available information regarding school enrolment and stoppage in the India study over time (results from the present study suggest individual gender, age and periods of school transition may all be important for understanding the long-term impact on education outcomes).

This study's findings also underscore the greater vulnerability of girls and young women to the adverse effects of unsanitary conditions in the home environment, which is supported by the evidence. Prior research has identified sex differences in the associations between infection, physical growth and cognitive development in infancy, as well as increased latrine adoption and/or usage among older girls (Bhalotra and Venkataramani, 2013; Dangour et al., 2013; Fuller et al., 2016; Sinha et al., 2017; Augsburg and Rodríguez-Lesmes, 2018). Several studies conclude girls benefit more from having access to improved sanitation compared to boys. Fuller and colleagues (2016) observed the main protective effect of increased sanitation coverage in girls' growth at 24 months in Ecuador, which they suggest may have been attributable to the earlier weaning and related pathogen burden among boys. A study from India also found significant associations between improvements in sanitation and child growth, measured in terms of standardised height-for age, were being driven by the effects on girls (Augsburg and Rodríguez-Lesmes, 2018). Further supportive research evidence comes from a study conducted in Mexico. Bhalotra and Venkataramani found positive effects of exposure to clean water (and the reduced incidence of infant diarrhoea) on girls' performance in cognitive assessments between nine and 14 years of age and in maths and reading tests at 15 years of age. The researchers also found that the earlier the exposure the greater the impact on cognitive development, observing that the large effects of exposure within the first 12 months were no longer present when exposure between two and five years was investigated. The present study's findings contribute to this evidence base by demonstrating how sex differences in the effects of unclean and unsanitary environments in early childhood may undermine later learning opportunities, illustrating the potential for targeted WASH interventions to improve education inequalities across gender groups.

This study had several limitations. It was not possible to account for household level water quality or quantity since neither were measured in the Cebu survey. There was also no available information regarding latrine usage among households who reported having access to one. This study concentrated on establishing the link between WASH exposure and schooling and did not test the proximate determinants of this relationship, including the influence of parental caregiving and early cognitive stimulation. Although it is not possible to say exactly what influence the inclusion of these variables might have had, they may have led to the attenuation of this study's findings. Future research focused on the empirical testing of mediating mechanisms would be valuable. Differences in school and teacher quality may also be important influences on school survival but were not measured in the CLHNS data.

Nonetheless, the present study improves our understanding of the impact of environmental risk factors in early childhood on education outcomes. In particular, it revealed the shorter school life of children exposed to unimproved sanitation and the potential vulnerability of girls and young women. These findings support the development of programmes aimed at improving early childhood development which include an explicit focus on conditions in the physical environment. Moreover, they serve to remind policymakers seeking to remedy persistent gender inequalities in education that some have their origins in early childhood.

Chapter 3.

The Effects of Neighbourhood Moves during Early to Middle Childhood on Cognitive Performance at 11 Years Old

3.1 Introduction

Residential moves are a potentially disruptive experience for children and may pose a risk to individuals during important periods of development. Moving is also a common event and has understandably attracted considerable research attention. Previous studies have linked housing instability during childhood to outcomes ranging from school performance and emotional and behavioural adjustment to physical and mental health (Simpson and Fowler, 1994; Pribesh and Downey, 1999; Dong et al., 2005; Jelleyman and Spencer, 2007; Rumbold et al., 2012; Coley, Leventhal, Lynch, & Kull, 2013; Cutuli et al., 2013; Anderson, Leventhal, and Dupéré, 2014; Mollborn, Lawrence, and Dowling Root, 2018). Researchers have also found associations between moving and child cognitive development (Roy, McCoy, and Raver, 2014; Ziol-Guest and McKenna, 2014; Fowler et al., 2015; Coley and Kull, 2016). However, existing research evidence has overwhelmingly come from high-income country (HIC) contexts, with little or no attention paid to the issue of mobility in resource-constrained settings. The present study addresses this gap and aims to improve our understanding of the impact of residential moves in low-income and lower-middle-income countries (LLMICs), investigating the effect of changes in neighbourhood during early to middle childhood on the cognitive performance of children approaching the end of primary school in a birth cohort from the Philippines.

3.1.1 Housing instability and child development

Housing instability is widely known to adversely affect individual development (Leventhal and Newman, 2010), but no clear pattern has emerged in the evidence to suggest mobility at one age is worse than at another. Empirical studies have found that higher frequency moves in the pre-school period are associated with poorer academic, socio-emotional and behavioural outcomes in young children (Anderson, Leventhal, and Dupéré, 2014; Ziol-Guest and McKenna, 2014; Schmitt and Lipscomb, 2016; Mollborn, Lawrence, and Dowling Root, 2018). Relatively little is known about the effects of moving in early childhood on developmental outcomes measured in older children. Behavioural problems related to pre-kindergarten mobility have been observed in children in the fifth grade (Roy, McCoy, and Raver, 2014); Rumbold and colleagues (2012) also found evidence of a connection between moves made in the first few years of life and internalizing behaviour scores when children were nine years old. These findings not only suggest support for the established view of the early years as a critical and/or sensitive period for development during which young children are particularly vulnerable, but also highlight the potential for some developmental deficits to manifest in older children.

Family or caregiver mobility during late childhood and adolescence has been linked to negative short- and medium-term effects on academic attainment, school dropout and completion rates (Haveman, Wolfe, and Spaulding, 1991; Simpson and Fowler, 1994; Voight, Shinn, and Nation, 2012; Coley, Leventhal, Lynch, & Kull, 2013; Cutuli et al., 2013). However, a large study conducted in the UK also found estimates of the dose-response relationship between residential moves and standardised school test scores were small and that frequent school moves were more influential (Hutchings et al., 2013). The combined effect of residential and school mobility on the educational outcomes of older children and adolescents is supported in the evidence (Pribesh and Downey, 1999; South, Haynie, and Bose, 2007). In general, researchers have not addressed differences in the effects of housing instability between male and female children, despite evidence suggesting boys are more likely than girls to develop emotional and behavioural problems exposed to stressful home environments (Pike et al., 2006).

Reviewing the published and grey literature showed a single study had previously examined residential and/or school mobility in a developing country context. In an urban cohort of children in South Africa no association was found between reported number of residential moves either before or after initial school enrolment and school grade failure, although a positive association was observed between number of moves following school entry and literacy and numeracy skills which were assessed when children were 16 years of age (Ginsburg, Richter, Fleisch, & Norris, 2011). No relationship was found between number of school changes and either outcome.

In summary, the evidence suggests the following: (1) residential mobility in early to middle childhood is disruptive, but moving at any age appears to undermine development; (2) longer-term or delayed effects are not well understood; (3) school moves potentially confound the association between residential mobility and cognitive development in school-age children; and (4) evidence of the associations between housing instability and child development obtained in HICs may not be applicable across different populations (e.g. those in resource-constrained settings in LLMICs). In addition, little consideration has been given to the role of gender by previous studies.

There is less research evidence on the effects of changes in residence on cognitive development. Nonetheless, several studies have explicitly examined the effects of exposure to mobility in childhood on cognitive performance. Results from a longitudinal study conducted in the US show that children who moved three or more times in the first five years of life achieved lower cognitive test scores when tested at five years old compared to children who did not move (Ziol-Guest and McKenna, 2014). In addition, the study authors found significantly higher levels of attention problems and internalising and externalising behaviours in the more residentially mobile group. Other studies found that moving even once in early to middle childhood was in low-income children related to lower levels of executive function – based on a formal assessment of underlying dimensions of cognition like working memory, inhibitory control and attention set shifting – as well as greater difficulty regulating attention, cognition and behaviour in the fifth grade according to teacher reports (Roy, McCoy, and Raver, 2014). There is also evidence suggesting the influence on cognitive skills might not be limited to early childhood and that duration of effects vary. In the US researchers linked housing mobility with deficits and/or delays in cognitive development among pre-school children, school-aged children and adolescents whose caregivers reported two or more moves in the previous 12 months (Fowler et al., 2015). However, the same study found children exposed to multiple moves partially made up the deficit over the three-year follow-up period. Coley and Krull (2016) also observed a relationship between moving in early to middle childhood and psychosocial functioning, although only moves made during middle childhood and early adolescence were negatively associated with cognitive skills and effects were short lived.

Possible mechanisms include move quality, which refers to the characteristics of the move and relative changes in neighbourhood environment. For example, moving to a resource-disadvantaged neighbourhood (compared to the neighbourhood of origin) may increase exposure to environmental risks at the same time as decrease the quality of education and healthcare services (Mollborn, Lawrence, and Dowling Root, 2018). Several studies conducted in the US have found move quality can have both positive and negative consequences for children (Sharkey and Sampson, 2010; Roy, McCoy, and Raver, 2014; Mollborn, Lawrence, and Dowling Root, 2018). The potential role played by quality of move is also consistent with theoretical explanations of the effects of housing instability on child development.

3.1.2 Theoretical perspectives on housing instability

The ecological model of human development emphasises the influence of biological and psychosocial factors on the reciprocal interactions between children and the people and objects in their immediate environment (known as proximal processes), while recognising the role played by more distal processes which connect children indirectly with wider social and cultural contexts (Bronfenbrenner, 1979). From this perspective, residential moves undermine the stability of the home environment, potentially disrupting the nurturing care essential to the acquisition of cognitive and socioemotional skills (Duncan and Brooks-Gunn, 1997; Shonkoff and Phillips, 2000). The model also incorporates the interplay between children and their neighbourhoods and schools, highlighting the influence that frequency and quality of changes in residence might have on development directly, as a result of transformed child-environment interactions, and indirectly via the impact moving may have on parents and other family members. Another element of the ecological model is the importance of developmental timing for the influence of exposure to risk factors (Bronfenbrenner and Morris, 2006). In other words, the impact of exposure to housing instability is expected to depend on whether it occurs during a developmentally important stage.

For example, young children rely more on their parents and spend more time at home than older children, suggesting pre-school-age children may be especially vulnerable to sources of instability that compromise proximal processes within the home environment (Shonkoff and Phillips, 2000). Equally, older children and adolescents may be more susceptible to the disruption of peer networks.

Ecobiodevelopmental models extend the ecological framework by describing the biological mechanisms underlying the association between adversity early in life and subsequent development (Shonkoff, 2010; Shonkoff et al., 2012). These models draw on empirical evidence indicating exposure to acute and/or chronic stress can cause disruption to the normal functioning of stress response systems including the hypothalamic-pituitary-adrenal (HPA) axis, which in turn affects cortisol response to stress (Repetti, Taylor, and Seeman, 2002). This physiological dysregulation is believed to compromise brain development leading to poorer cognitive, social and health outcomes in later life. Several studies examining the effect on cognitive skills found neighbourhood poverty, residential mobility, and parent unresponsiveness in early to middle childhood predicted variations in HPA activity, measured using basal cortisol levels, which were associated with lower levels of cognitive functioning (Blair et al., 2011; Suor et al., 2015).

The social capital explanation of the impact of housing instability on children is predicated on the weakening or loss of social connections (Coleman, 1988). From this perspective, residential mobility not only undermines the important bonds children make with their peers, but also the social, emotional and economic advantages their parents derive from social networks that enable the sharing of useful information or the help and support of nearby family members. However, children could also stand to benefit from mobility if social capital is gained from moving into a better-integrated community or a neighbourhood with more resources (Coleman, 1988; Pribesh and Downey, 1999).

The evidence shows residential mobility can disrupt children's lives with potentially serious consequences for their development. However, there exist important gaps in the evidence, particularly in our understanding of the effects on children living in LLMICs. The present study aimed to address some of these gaps. Using longitudinal data from a predominately urban cohort of children from the Philippines this study assessed the demographic, social and economic differences between more and less residentially mobile households. It examined whether moving between neighbourhoods in early to middle childhood was associated with child cognitive performance measured at 11 years of age, balancing observable differences between movers and non-movers using a combination of propensity score methods and inverse probability weighting. Based on the literature it also explored differential vulnerabilities between males and females, investigating the moderating effects of move quality (measured in terms of differences between origin and destination neighbourhoods) and subsequent school mobility (e.g. Hutching et al., 2013; Mollborn, Lawrence, and Dowling Root, 2018). The study tested four main hypotheses: (1) children from households who move neighbourhoods in early to middle childhood perform less well in cognitive assessments at 11 years of age; (2) male children are more vulnerable to the adverse effects of residential mobility than females; (3) the impact on cognitive performance is moderated by move quality; (4) school moves multiply the negative effects of residential mobility.

3.2 Data

The data used in this study were from the Cebu Longitudinal Health and Nutrition Survey (CLHNS), a population-based cohort study conducted in Metro Cebu. The study area includes the second largest city in the Philippines and its surroundings, households living in heavily populated urban areas, peri-urban areas, smaller rural towns and nearby mountain and island communities (Adair et al., 2010). A single stage cluster-sampling strategy was used to select 17 urban and 16 rural administrative units (barangays) at random. Households were then surveyed for pregnant women and those due to give birth between May 1983 and April 1984 were included in the sample, making it representative of births in Cebu for this period. Altogether 3,327 women were interviewed and those who had multiple births, a stillbirth or miscarriage, refused to participate further or out-migrated were excluded (Adair et al., 2010). The final sample was made up of 3,080 women who had given birth to a single live infant. Data were collected during the pregnancy, immediately after the birth and every two months for next 24 months. Follow up surveys were carried out every three to four years from 1991 onwards. The baseline survey gathered self-reported and interviewer observed information on the social, economic, and environmental conditions in each household, neighbourhood health services and infrastructure, in addition to the mothers' pregnancy and infant health. The number of residential moves were reported in first major follow-up survey in 1991 and cognitive tests were administered in 1991 and 1994. The richly detailed information in the CLHNS was well suited to this analysis. In addition, the longitudinal nature of the survey meant baseline covariates (pre-move individual, household and community characteristics collected in 1984) were unlikely to have been affected by the treatment (moving between 1986 and 1991, and collected in 1991); a necessary condition for estimating the propensity scores and the treatment effect, which can be problematic for studies using observational data collected concurrently.

The data used in the present study were from the baseline, birth, bimonthly, 1991 and 1994 surveys. Of the initial sample of 3,080 children, 2,252 underwent cognitive assessment in 1991 and 2,136 were tested in 1994 when the participants were between 11 and 12 years of age. The analytic sample was based on the availability of valid observations for cognitive test scores in 1991 and 1994, as well as individual and household information between 1984 and 1994, giving a final sample of 1,916 individuals (909 females and 1,007 males). Sample attrition occurring between baseline and the first follow-up in 1991 was primarily driven by infant death and out-migration (Adair et al., 2010). Missing data were assumed to be conditionally Missing at Random (MAR) – an assumption which requires that observed variables predictive of the missing data mechanism are included in the estimation model – since any systematic missingness was likely to be explained by covariates which were already the focus of this analysis. Specifically, socio-economic characteristics at baseline such as household income, assets and parents' education – which have been shown to predict attrition in similar longitudinal cohorts from LLMICs (Alderman et al., 2001) – early years infant growth, and residential location and mobility.

Cognitive performance (dependent variable). Child cognitive performance was measured at eight and 11 years of age using the Philippine Nonverbal Intelligence Test (Guthrie, Tayag, and Jacobs, 1977). Developed in the Philippines, the 100-item test was designed to assess the analytic and reasoning skills of Filipino children specifically. The test required participants to identify which one of five objects, numbers or designs on a picture-card did not belong in the same category as the other four (e.g. four arrows pointing up, and one pointing down). The 100 cards were arranged in order of difficulty, becoming increasing abstract, and the test was standardised by the developers at a large public elementary school in Manila using data from 100 boys and girls from each grade level. The test was deliberately developed for within-group comparisons only and is neither nationally representative nor can it be used to compare children across groups from different societies and cultures. However, it showed acceptable reliability and validity during test development and has since been used by multiple studies to investigate the influence of risk and protective factors on cognitive development in children from the Philippines (e.g. Daniels and Adair, 2004, Daniels and Adair, 2005).[5] Mean test scores were 51 points at eight years of age and 69 points at 11 years, while variability decreased with age (a standard deviation equalled 12.5 points in 1991 and 11.5 points in 1994).

Residential mobility (independent variable). Household mobility was based on whether mother and child were living in the same barangay in 1991 that were living in when surveyed six years earlier. If not, the mother was asked for the number of different barangays they had lived in before moving to their current location. The response gave the total number of moves between barangays over a six-year period, beginning when children were two years old and ending when they were eight. In this sample 89 percent of families reported not moving between 24 months and eight years

[5] KR-20 estimates of reliability ranged between 0.71 and 0.95. Validity coefficients (correlations between Philippine Nonverbal Intelligence Test scores and school grades obtained in Language, Maths and Social studies) were the same as those obtained using American tests given to American children. These coefficients were fairly low, but this was expected given the influence of classroom factors unrelated to ability (see Guthrie, Tayag, and Jacobs, 1977).

of age (n=1716); eight percent reported moving barangay once (n=147); and three percent had moved twice or more (n=49). Due to the small number of families moving more than once the final indicator of residential mobility was a binary variable which took one if a mother and child moved once or more and zero otherwise.

Other covariates. Variables measuring parental and household characteristics at baseline including maternal and paternal education, mothers' age and height, parity (whether the index child had been the woman's first pregnancy), neighbourhood urban development, household wealth, household income and size were used to adjust for potential confounding. A multicounty study analysing LLMIC cohort data including the CLHNS found breastfeeding duration was attenuated by mothers' age and parity; first-time mothers were less likely to breastfeed compared to multiparous women possibly as a result of biological or behavioural maturity (Fall et al., 2015). Neighbourhood urban development was measured using an index created by members of the CLHNS team for the purpose (Dahly and Adair, 2007). This combined neighbourhood population, population density, communications (cell phones, the internet, cable TV etc.), transportation (the density of paved roads, availability of public transportation), educational facilities, health services, and markets (including grocery stores and gas stations), each of which was assigned values from zero to 10 resulting in a scale ranging from zero to 70. Household wealth was based on reported ownership of an electric fan, furniture, a car, jeepney (local taxi vehicle widely used in the Philippines), colour television, cassette recorder, and refrigerator. Information on the quality of materials used to construct the house, access to running water, and whether the house had a flushing toilet was also added. Polychoric correlations between these items were estimated and principal components analysis conducted on the correlation matrix (Kolenikov and Angeles, 2004). The first principal component was then treated as a wealth index. Child characteristics measured at baseline included sex, age, standardised birth weight and length, whether they were breastfed, and height-for-age z-scores (HAZ) assessed at 24 months, which were calculated using World Health Organisation reference data.

3.3 Data analysis

The analysis had three main stages: (1) generating descriptive statistics for individuals and households from the moving and non-moving groups at baseline, as well as pre- and post-move neighbourhood characteristics; (2) calculating the propensity score for moving based on observed baseline covariates and ensuring a similar balance of covariates between movers and non-movers; and (3) using multilevel linear regression models with inverse probability of treatment weights (IPTWs) to estimate the effect of changes in neighbourhood on cognitive performance, and the moderating effects of move quality and school mobility. All analysis was carried out using Stata 14 software (StataCorp., 2015).

Propensity score methods like inverse probability of treatment weighting are an alternative to regression adjustment for confounding when using observational data (Rosenbaum and Rubin, 1983). The propensity score is defined as the probability of receiving treatment conditional on observed pre-treatment covariates (Austin and Stewart, 2015). The present study estimated propensity scores by regressing a binary treatment variable (which took the values one if families moved barangay once or more during the period between 24 months and eight years old and zero if not) on a set of covariates observed at the baseline using a logistic regression model. Covariates included demographic characteristics (children's age and sex, mothers' age), well-known developmental risk factors (birthweight and birth length, standardised height-for-age at 24 months, mothers' height, whether the child was breastfed, parity), measures of family socioeconomic position (mothers' and fathers' education, household income, household wealth) and environmental conditions (household size, access to adequate water and sanitation, and neighbourhood urban development).

Variable selection was based on the results of a review of the published and grey literature on risk factors affecting the cognitive development of children in LLMICs (e.g. Daniels and Adair 2004). Treatment assignment and outcome were expected to be affected by cluster-level characteristics, so the propensity score model also included neighbourhood dummy variables (Arpino and Mealli, 2008).

IPTWs that use propensity scores are formally defined as $w = \frac{T}{e} + \frac{1-T}{1-e}$ where T is the binary treatment variable and e is their propensity score (Stuart, 2010). However, treated individuals with very low propensity scores (i.e. close to zero), or individuals in the control group with very high scores, can lead to extremely large weights. To address the issue and improve the precision of the estimates, the weights were stabilised: $w = \frac{T\,\Pr(T=1)}{e} + \frac{(1-T)\,\Pr(T=0)}{1-e}$ (Austin and Stewart, 2015). This involved multiplying the inverse probability of receiving treatment by the marginal probability of treatment $\Pr(T = 1)$ and control $\Pr(T = 0)$. Each individual was therefore weighted by the inverse of the probability of moving between barangays, creating a weighted sample in which residential mobility was independent of observed confounding. The distribution of baseline covariates between movers and non-movers in this weighted sample was assessed by comparing the means and proportions of continuous and categorical variables respectively. Standardised difference in means, which gives the difference in means in units of pooled standard deviations was also assessed (Austin and Stewart, 2015).

In order to account for clustering in the data, which consisted of children and households nested within neighbourhoods, multilevel linear regression models were used. Failure to incorporate the hierarchical structure of data into models can lead to incorrect inferences (Austin, Goel, and van Walraven, 2001). Since there was a single index child in each household and therefore no variation within households the lowest level in the data referred to children in families and included both child and family covariates. The group level referred to the household's barangay or neighbourhood. Linear random effects models, which incorporated a barangay-specific random effect allowing the intercepts to vary across clusters, were specified as follows:

$$y_{ij} = \beta_0 + \beta_1 x_{1ij} + \beta_2 x_{2ij} + \cdots + \beta_k x_{kij} + u_j + e_{ij} \qquad (1)$$

where y_{ij} was the cognitive test score of child i in barangay j, β_1 was the average effect of residential mobility x_{1ij}, variables $(x_{2ij}, \ldots, x_{kij})$ included fixed individual, household and neighbourhood characteristics, and subject-specific random effects e_{ij} and cluster-specific random effects u_j were assumed to have independent normal distributions. Equation (1) was also estimated without covariates (widely known as a variance components model) to confirm the extent of the clustering. The final regression models included stabilised IPTWs and the full set of baseline covariates, since the combination of regression and weighting by propensity score has been shown to be 'doubly robust' to model misspecification (Rubin, 1979; Rubin and Thomas, 2000).

3.4 Results

At the baseline 73 percent of the sample were living in barangays classified as urban with the remaining 27 percent living in rural barangays or villages (movers and non-movers were similarly distributed). The proportion of urban-dwelling households remained constant over time, with 73 percent of non-movers and 72 percent of movers still living in an urban neighbourhood in 1991. 70 percent of families with multiple residences in this period (i.e. movers) reported last living in an urban barangay in Metro Cebu. This suggests moves occurred within the urban-rural strata.

However, it does not imply neighbourhoods of origin and destination were homogeneous for movers. Figure 3.1 displays quartiles of the distribution of mean neighbourhood development between 1986 and 1991.[6] Negative values indicated the barangay a family lived in 1991 was less developed than the area they were living in five to six years earlier; values around zero revealed there had been little or no change in neighbourhood development over the period; while positive values showed families living in more developed areas in 1991 than in 1986.

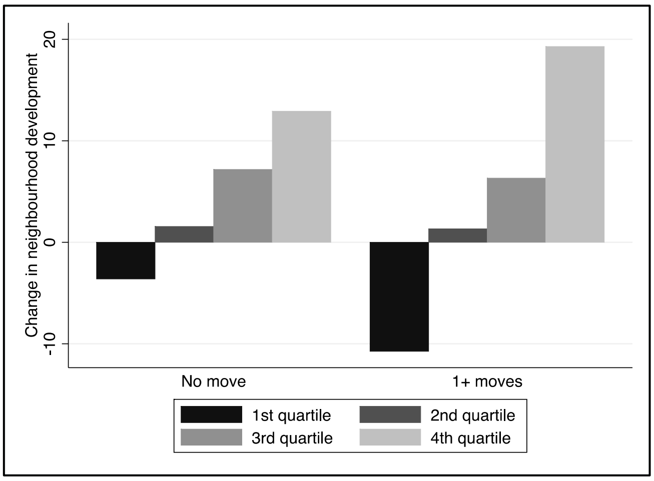


Figure 3.1 Mean change in neighbourhood urban development between 1986 and 1991.

[6] A multidimensional index based on barangay population size and density, communication, education, transport, health and market characteristics was used to measure urban neighbourhood development (see Dahly and Adair, 2010).

Metropolitan Cebu is a rapidly urbanising area and the sample mean change in neighbourhood development between 1986 and 1991 was positive (m=3.8, *SD*=7.6) In other words, changes in barangay development were evident regardless of move status. Yet the largest changes were observed among residentially mobile families. This suggests children who experienced residential relocation were exposed to changes in the quality of their environment that were more extreme than children who did not relocate.

Figure 3.2 shows the changes in neighbourhood quality were fairly evenly spread across the non-mover group: 27 percent of non-movers were living in barangays in 1991 that were less developed than they had been in 1986; 27 percent in barangays in which there had been little or no change in conditions; and 25 percent in barangays that were more developed. Among movers, changes were less evenly distributed: 32 percent of movers were resident in an area that was less resource-advantaged in 1991 than the neighbourhood they had lived in 1986. This shows approximately one-third of children who moved were exposed to a deterioration in environmental quality. Movers were also different from non-movers in terms of some of their social, economic and demographic characteristics.

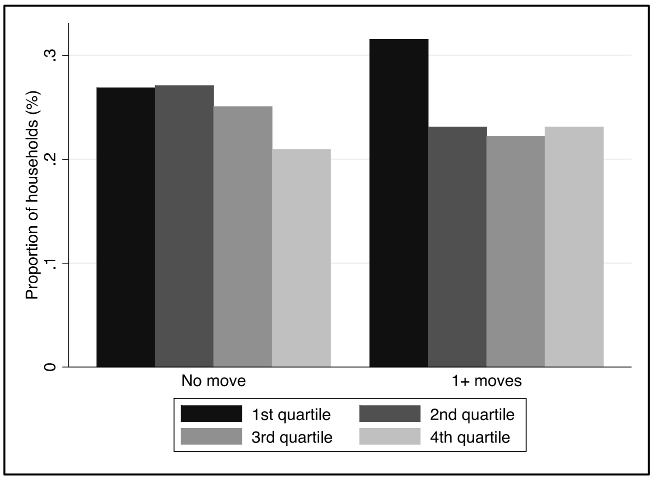


Figure 3.2 Distribution of households by change in neighbourhood development and move status.

Table 3.1 indicates the children who experienced a move were a couple of months older on average and marginally fewer were breastfed at birth. In addition, more women in the mover group were having children for the first-time compared to women in the non-mover group (27% and 19%, respectively). However, mothers from mobile families were significantly better educated (74% had completed primary school and 35% had finished secondary or high school) as were fathers (41% had finished secondary school). Families in the mover group also had a significantly higher level of household wealth than non-mover families.

Table 3.1 Baseline characteristics of non-mover and mover families in unweighted sample.[a]

	Mover: no (n=1720)	Mover: yes (n=196)	*p*-value	Standardised difference
Male	902 (52 %)	105 (54 %)	0.76	-0.02
Age, y	11.0 (± 0.5)	11.1 (± 0.5)	0.03	-0.17
Birth weight, g	3007.3 (± 426.3)	3024.9 (± 416.3)	0.58	-0.04
Length, cm	49.3 (± 2.1)	49.4 (± 2.2)	0.49	-0.05
Height-for-age (z-scores)	-2.28 (± 1.0)	-2.22 (± 1.1)	0.44	-0.06
Breastfed	1638 (95 %)	181 (92 %)	0.08	0.12
Mothers' age, y	26.0 (± 6.1)	25.7 (± 5.4)	0.34	0.07
Height, cm	150.7 (± 150.4)	150.5 (± 150.0)	0.58	0.04
Primiparous	335 (19 %)	52 (27 %)	0.02	-0.17
Mother primary school	1169 (68 %)	146 (74 %)	0.06	-0.14
Mother secondary	485 (28 %)	69 (35 %)	0.04	-0.15
Father primary school	1205 (70 %)	144 (73 %)	0.32	-0.08
Father secondary	603 (35 %)	81 (41 %)	0.08	-0.13
Household wealth	-0.05 (± 1.1)	0.13 (± 1.2)	0.03	-0.16
Income (log)	5.16 (± 1.0)	5.26 (± 1.0)	0.19	-0.10
Size	5.8 (± 2.8)	5.8 (± 3.0)	0.90	-0.00
Neighbourhood development	29.3 (± 12.8)	30.1 (± 13.6)	0.39	-0.06

[a] Continuous variables are expressed as mean ± standard deviation, and dichotomous variables are *n* (%).

The mean stabilised IPTW was 1.00 (SD=0.29) and the minimum and maximum values were equal to 0.45 and 1.87 respectively. Table 3.1 shows the standardised differences in the unweighted sample were 0.1 or greater for eight of the 17 covariates, where a standardised difference of 0.1 (10%) or more is believed to be indicative of a meaningful imbalance in baseline covariates (Austin, 2009). In other words, about 50 percent of the observed covariates were unbalanced prior to weighting. The largest standardized difference in the weighted sample was 0.03 (birthweight and birth length), which indicates that in the sample created using stabilised inverse probability of treatment weights observed pre-move characteristics were only modestly different

Table 3.2 Baseline characteristics of non-mover and mover families in weighted sample.[a]

	Mover: no (n=1711)	Mover: yes (n=205)	Standardised difference
Male	907 (53 %)	107 (52 %)	-0.00
Age, y	11.0 (± 0.5)	11.0 (± 0.5)	0.00
Birth weight, g	3009.1 (± 428.3)	3004.0 (± 417.1)	-0.01
Length, cm	49.3 (± 2.1)	49.2 (± 2.1)	-0.03
Height-for-age, (z-scores)	-2.28 (± 1.0)	-2.31 (± 1.1)	-0.03
Breastfed	1625 (95 %)	196 (96 %)	0.03
Mothers' age, y	26.1 (± 6.1)	26.2 (± 5.4)	0.01
Height, cm	150.6 (± 5.1)	150.6 (± 4.6)	-0.00
Primiparous	342 (20 %)	41 (20 %)	-0.00
Mother primary school	1181 (69 %)	139 (68 %)	-0.01
Mother secondary	479 (28 %)	57 (28 %)	-0.01
Father primary school	1198 (70 %)	144 (70 %)	-0.01
Father secondary	616 (36 %)	74 (36 %)	0.00
Household wealth	-0.04 (± 1.1)	-0.03 (± 1.1)	-0.01
Income (log)	5.17 (± 1.0)	5.16 (± 1.0)	-0.02
Size	5.8 (± 2.8)	5.8 (± 3.0)	-0.01
Neighbourhood development	29.4 (± 12.8)	29.3 (± 12.8)	-0.01

[a] Continuous variables are expressed as mean ± standard deviation, and dichotomous variables are *n* (%).

between mover and non-mover families (Table 3.2). The remaining imbalance after weighting was minimal (see propensity score kernel densities before and after weighting in Appendix 3A) and was addressed by the inclusion of the full set of baseline covariates in the final regression model.

Table 3.3 Association between residential mobility during early to middle childhood and cognitive performance assessed at 11 years of age (n=1916).

	Whole sample			**Female**			**Male**		
	Coef.	95% CI		Coef.	95% CI		Coef.	95% CI	
Moved neighbourhood	-1.560*	-3.147	0.027	-2.411**	-4.331	-0.491	-0.686	-2.563	1.191
Male	-0.365	-1.185	0.455	-	-	-	-	-	-
Age	1.482***	0.642	2.322	1.676**	0.436	2.917	1.302**	0.229	2.375
Birth weight	-0.001	-0.002	0.001	-0.001	-0.003	0.001	-0.001	-0.002	0.001
Birth length	0.122	-0.230	0.474	0.191	-0.263	0.645	0.028	-0.391	0.448
Height-for-age	0.414*	-0.070	0.899	0.800**	0.188	1.411	0.081	-0.484	0.646
Breastfed	-0.443	-3.129	2.243	-0.389	-3.515	2.738	-0.273	-3.361	2.814
Mothers' age	0.038	-0.048	0.123	0.073	-0.063	0.209	0.006	-0.098	0.109
Mothers' height	0.005	-0.072	0.082	0.069	-0.021	0.158	-0.042	-0.152	0.068
Primiparous	-0.500	-1.625	0.625	-0.751	-2.193	0.691	-0.198	-1.875	1.479
Mother primary school	1.415**	0.287	2.544	1.361*	-0.191	2.913	1.317*	-0.141	2.776
Mother secondary	1.401**	0.181	2.621	1.807***	0.742	2.871	1.273	-0.817	3.362
Father primary school	0.831	-0.392	2.054	0.576	-0.912	2.064	1.036	-0.676	2.749
Father secondary	0.570	-0.602	1.743	0.198	-1.382	1.778	0.739	-0.745	2.223
Household assets	0.597**	0.096	1.097	0.527	-0.133	1.187	0.723**	0.117	1.329
Household income	0.422	-0.170	1.014	0.894**	0.300	1.488	-0.033	-0.816	0.751
Household size	-0.015	-0.202	0.172	-0.138	-0.467	0.192	0.107	-0.151	0.364
Urbanicity	-0.029	-0.082	0.024	-0.046	-0.105	0.014	-0.026	-0.087	0.035
Cognitive test score	0.495***	0.446	0.543	0.472***	0.405	0.540	0.515***	0.443	0.587
Between barangay	2.613	1.187	5.755	2.487	0.563	10.981	2.191	0.696	6.897

*** $p < 0.001$, ** $p < 0.05$, * $p < 0.1$.

In the null model the intraclass correlation (ICC) – the correlation between the cognitive test performance of two children from the same neighbourhood – was calculated as 0.11. That is, without conditioning on individual, family, or community characteristics, 11 percent of the variance in the outcome was explained by differences between neighbourhoods. This was evidence of the within-cluster homogeneity in the data requiring a multilevel model. *Hypothesis 1* (children from households who move neighbourhoods in early to middle childhood perform worse in cognitive assessments at 11 years of age): The results from the main regression analysis reported in Table 3.3 reveal a negative association between residential mobility and cognitive performance. Specifically, children who moved between neighbourhoods in early to middle childhood had lower average cognitive test scores at 11 years of age, after adjusting for potential individual, household and community level confounding. Results from multilevel regressions including IPTWs and the full set of covariates showed that children who experienced a neighbourhood move had test scores that were on average 1.6 points lower than children whose families did not move (95% CI -3.14, 0.03, p=0.053). In the full model the estimated ICC was 0.03, suggesting that the within-barangay correlation or level of clustering after accounting for the fixed part of the linear random effects model was small.

Hypothesis 2 (male children are more vulnerable to the adverse effects of residential mobility than females): Subgroup analysis revealed that the cognitive deficit associated with moving was greater in female children, who scored 2.4 fewer points in testing than their female peers who did not experience a move (95% CI -4.33, -0.49, p<0.001). In contrast, the estimated effect of moving on male cognitive test scores was just 0.7 points and was not statistically significant. Appendix 3B presents the results from a two-way interaction model (household moves by gender) which also suggest that compared to male children the effects of moving on cognitive development were stronger for females (-1.93, 95% CI -4.30, 0.43, *p*=0.109). In addition, the impact was worse on girls living in poorer barangays compared to boys in similar circumstances (see Appendix 3C). However, some estimates lack precision and should be interpreted as evidence supportive of sex differences and/or as reason to reject the study hypothesis with caution.

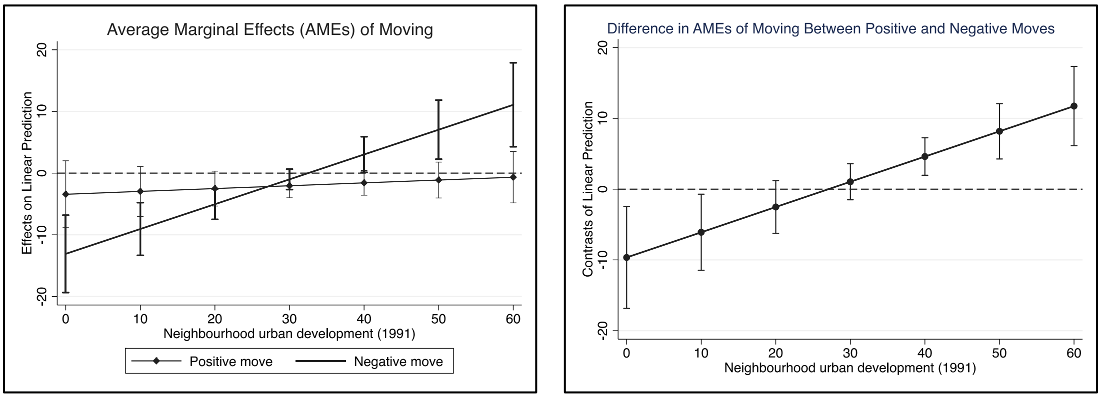


Figure 3.4 Average marginal effects and differences between the effects of positive and negative moves by destination neighbourhood development.

Hypothesis 3 (the impact on cognitive performance is moderated by move quality): A three-way interaction model (household moves by move quality by quality of destination neighbourhood) was specified to assess the extent to which the effect of residential mobility depended on the quality of move and quality of destination. Figure 3.4 shows the average marginal effect of a ‘negative move’ was significantly stronger when moves were to more or less developed neighbourhoods.[7] In other words, the effect on children of moving to a barangay with poorer infrastructure and services compared to the barangay they had been living in changed according to the quality of destination. So, children who experienced a negative move to a resource-

[7] Negative moves were defined as moves to a destination neighbourhood or barangay that was less developed in terms of infrastructure and services than the barangay of origin. A positive move was the opposite (i.e. the destination neighbourhood was more developed than the area the family had moved from).

disadvantaged barangay performed significantly worse in subsequent cognitive tests, although Figure 3.4 shows the adverse effects of a negative move decreased as resources in the destination neighbourhood increased.

Hypothesis 4 (school moves multiply the negative effects of residential mobility): The results in Appendix 3D show that the impact of residential mobility on cognitive test scores was unaffected by subsequent changes in school. That is, the children exposed to moves between neighbourhoods who then changed school did not perform worse than the children who experienced changes in neighbourhood but not in school.

3.5 Discussion

The present study examined the influence of residential mobility in early to middle childhood on cognitive test scores in a lower-middle-income country context using longitudinal data from a Philippine birth cohort. Descriptive statistics showed that move quality was variable, with roughly one-third of movers experiencing a deterioration in neighbourhood environment. However, the average mover was also better educated and more materially affluent than their non-mover counterpart. The main regression analysis indicated a significant negative association between exposure to a move and cognitive performance at 11 years of age. This association was obtained after balancing mother and child characteristics, well-known developmental risk factors, household socio-economic position and environmental conditions between mover and non-mover groups. There also appeared to be significant differences in the effect of changes in neighbourhood between male and female children, with females worst affected. The results from interaction models used to test the moderating effects of move quality, quality of destination neighbourhood and subsequent school mobility, revealed the impact on cognitive test scores of experiencing a 'negative' move (i.e. moves ending in relatively less developed areas) depended on the quality of destination neighbourhood, and that the adverse effects of moving were unaffected by subsequent changes in school.

These findings contribute to our understanding of the relationship between mobility in early life and child development in several ways. First, they confirm the period between two and eight years of age is developmentally sensitive and that exposure to housing instability (moving between neighbourhoods or barangays) during this time may have adverse effects on child development. This is broadly consistent with previous research which showed that mobility in middle childhood – the time roughly between four and 11 years of age – was linked to cognitive skills measured in the fifth grade, in contrast with mobility during the first few years of life (Coley and Krull, 2016). Second, these findings reveal that the detrimental effects of moving between neighbourhoods during early to middle childhood are not necessarily short-term. This has qualified support in the literature. Rumbold and colleagues (2012) found that young children exposed to two or more residential moves were more likely to have poorer mental health at nine years of age. Equally, Fowler and colleagues (2015) found evidence of persistent cognitive deficits in children between four and 14 years of age over the study's three-year follow-period despite partial catch-up, although the study only measured household moves in the 12-months immediately prior to each round of cognitive testing. Third, the results suggest there may be sex differences in the effects on cognitive performance, contrary to the study hypothesis that the most vulnerable group is male. This is inconsistent with the evidence suggesting boys are more sensitive to environmental risks than girls (Pike et al., 2006) but is supported by prior research from HICs showing the quality of home environments (including some parenting practices) is often lower for females than for males (Garcia, Heckman, and Ziff, 2017). Future investigations into the influence of social, cultural and economic factors more prevalent in low-income and lower-middle-income countries may help explain this finding.

Unlike this study, the only longitudinal research identified in a review of the literature to examine the impact of residential mobility in an LLMIC setting found moving had a positive effect on the numeracy and literacy skills of a cohort of children from South Africa (Ginsburg, Richter, Fleisch, & Norris, 2010). One possible explanation for this discrepancy is that the two studies took a different approach to measuring mobility. The South Africa analysis assessed the total number of changes in residence (the house a person lives in), whereas this study estimated the effects associated with residential changes in neighbourhood. Mobility as measured in this study might therefore represent a more substantial source of stress and/or disruption in children's lives.

The results from the present study suggest that residential moves may influence subsequent cognitive performance by exposing children to changes in environmental quality. From an ecobiodevelopmental perspective, moving to a neighbourhood with lower levels of infrastructure and services could alter children's exposure to stress (Shonkoff, 2010; Shonkoff et al., 2012). This is in line with prior research linking environmental conditions with neurocognitive development (Roy, McCoy, and Raver, 2014; Fowler et al., 2015). Anderson, Levanthal, Newman, and Dupéré (2014) also found that children moving between neighbourhoods during middle childhood were especially sensitive to environmental quality. The authors attributed this finding to physical, cognitive and socio-emotional growth leading to greater child independence and the increasing relevance of extrafamilial contexts, while highlighting changes in the role of institutional resources including social, health and educational services as children get older.

Unexpectedly, this study found no evidence that school mobility affected residential movers, which is inconsistent with the research evidence from studies in HICs which shows changes in school can have a disruptive effect equal to or greater than changes in residence (e.g. Pribesh and Downey, 1999). This suggests that the consequences of moving between schools within the same neighbourhood may be negligible among older children in LLMICs. However, this study was only able to measure sequential mobility (residential moves followed by school moves) and different combinations of residential and school moves are known to have different effects (Pribesh and Downey, 1999; Swanson and Schneider, 1999). Future investigators might also assess the role played by other important aspects of school mobility, such as school quality and the individual and/or parental motivations that lie behind changes in school, which were not measured in the Cebu data.

Every observational study contends with the issue of selection. In this case, the influence of non-random patterns of residential mobility on the estimation results. Interpreting these results therefore involves considering how factors potentially driving selection are related to the outcome. In studies conducted in high-income countries which observe a negative relationship between residential mobility and child development, socio-economic vulnerabilities associated with selection into the moving group are also often predictive of compromised development. In other words, children from more socially and economically vulnerable families are both more likely to be movers and to have developmental deficits regardless of move status. So, the influence of selection could bias estimates upwards making the effects of moving appear stronger than they actually are. Yet residentially mobile families in this study had significantly better socio-economic attributes compared to families who did not move; families reporting a change in neighbourhood had higher levels of education and greater household wealth, in common with more mobile families in other longitudinal surveys from LLMICs (Alderman et al., 2001). Therefore, although the potential for confounding remains a concern, estimates in this study may be likely to be biased by the kind of selection on observables that threatens to distort results from similar studies in HICs.

This study had several limitations, beyond the issue of selection. First, only a single definition of mobility was available in the data (whether a family moved when children were between two and eight years of age), which ignored within-neighbourhood changes in residence and failed to account for families' motives for moving. Second, the study was unable to identify the exact timing of reported moves, making it was impossible to determine move duration or the precise age of children at the time of moving. Third, the impact of exposure to moves during different developmental periods (e.g. moving in the first 24 months) could not be tested. Finally, complete data for residential mobility and cognitive test scores at eight and 11 years of age were only available for roughly 60 percent of the cohort. Since this means the analytic sample was selective towards less mobile families – over the course of the survey sample attrition was primarily driven by out-migration (Adair et al., 2010) – these results potentially represent a lower-bound estimate of the impact of moving on children in the population.

Despite these limitations, the results of this study emphasise the need for better understanding of the underlying mechanisms that link residential moves and child development outcomes in LLMICs, with an emphasis on the role played by risk factors in the neighbourhood environment. In view of this, these findings support policies and programme interventions designed to mitigate the impact of environmental adversity as well as efforts to reduce housing instability.

Chapter 4.

Social Marginalisation and Adolescent Mental Health: Disengaged Youth neither in Education, Employment nor Training

4.1 Introduction

The transition from school to work is a challenging time for adolescents and young adults worldwide, particularly those living in low- and lower-middle-income countries (LLMICs) with inadequate education systems, high levels of poverty and youth unemployment, and comparatively few social safety nets. One indication of the scale of the challenge in LLMICs is the prevalence of young people neither in education, employment nor training or ‘NEETs’. In 2015 the average percentage of young people between 15 and 19 years of age who were NEET in Organisation of Economic Cooperation and Development (OECD) countries was estimated at six percent. In OECD countries like Mexico and Turkey this figure was more than twice as high (OECD, 2016). The transitional period between school and work also overlaps with a substantial age-related rise in the prevalence of disorders like anxiety and depression (Patel, Fischer, Hetrick, & McGorry, 2007; Thapar, Collishaw, Pine, & Thapar, 2012). Adolescents and young adults in high-income country (HIC) settings who have become disengaged from important socialising institutions like school and the workplace have been shown to be especially vulnerable to mental health problems (Scott et al., 2013; O’Dea et al., 2014; Baggio et al., 2015; Cornaglia, Crivellaro, and McNally, 2015; Veldman et al, 2015; Goldman-Mellor et al., 2016; Henderson et al., 2017; Rodwell et al., 2018). Yet little consideration has been given to the mental health of disengaged youth in LLMICs. This represents an important evidence gap considering the majority of the global population of adolescents live in LLMICs and the diverse social and economic conditions observed in the developing world, which are direct determinants of the prevalence and severity of mental disorders. The present study intends to address this gap by examining the associations between adolescent mental health and disengagement from education and employment in a birth cohort of children from the Philippines.

4.1.1 Mental health and disengagement from school and work

The linkages between mental ill health and the domains of adolescent education and employment are complex. On the one hand, teenage NEETs are often from socio-economically disadvantaged backgrounds and are more likely to have experienced mental health problems as children (Goldman-Mellor et al., 2016). Evidence from several longitudinal studies shows young people with early behavioural and mental health problems are more vulnerable to becoming NEET during the subsequent move from school to work (Baggio et al., 2015; Veldman et al., 2015; Rodwell et al., 2018). There is also mixed evidence suggesting the onset of mental ill health in childhood adversely affects later school outcomes (Kessler et al., 1995; Fergusson and Woodward, 2002; Cornaglia, Crivellaro, and McNally, 2015). On the other hand, unemployment is well known to compromise mental health and wellbeing (Paul and Klaus, 2009). Moreover, though many are committed to finding work, NEETs are more likely to feel pessimistic about their future and struggle with loss of social connections and support (Goldman-Mellor et al., 2016; Huegaerts, Spruyt, and Vanroelen, 2018). A prospective cohort study conducted in New Zealand found unemployed young people had mental health problems, substance use disorders and suicidal behaviours between 1.4 and 8.4 times higher than their employed peers, even after adjusting for a wide range of confounders including individual history of psychiatric disorder and previous suicidal behaviour (Fergusson, Horwood, and Woodward, 2001).

It is also accepted that relations between risk factors in the social environment and mental health are bidirectional (Power et al., 2015). In other words, mental health problems that develop early in life increase the probability of adolescents and young adults becoming NEET. But NEETs are then exposed to independent risks that can cause mental illness and/or exacerbate pre-existing mental health vulnerabilities, which in turn increases the duration and severity of unemployment. This bidirectionality is consistent with the results of longitudinal research indicating that the social determinants of depressive disorders may interact and reinforce each other over time (Lund and Cois, 2017).

Yet existing research evidence from HICs may not apply to LLMIC populations in view of the different social and economic conditions observed in countries like the Philippines. One issue is the lack of any international standard underpinning the definition of NEET status (Elder, 2015). Equally important is the interpretation given to the term NEET when applied in LLMIC contexts. For example, individuals who have become disengaged from school but have not found work are not necessarily jobless. In other words, although unemployed according to the international standard, some people are not available to work. The majority of this subcategory are young women engaged in unpaid domestic work (Elder, 2015). Relatively little research attention has been paid to this group in HIC studies, although researchers in Mexico found NEET ‘homemakers’ – mostly married young women, with family and/or caring responsibilities – who had chosen to leave school and were not seeking wage work had fewer substance use disorders and suicidal behaviours compared to non-NEET women (Gutiérrez-Garcia et al., 2018). In addition, although the percentage of young people not completing secondary school is higher in LLMICs than in HICs, the risk of failing to finish school as a result of mental health problems is lower in developing countries suggesting family, community, sociocultural attitudes and/or characteristics of the education systems in LLMICs mitigate against some of the school effects associated with mental ill health in high-income country settings (Lee et al., 2009). Strong social support networks often based around extended, multi-generational family groups are common in many less developed countries and may also have protective effects on mental health. For example, research in South Asia found a grandmother living at home was protective against maternal depression (Maselko et al., 2015). In contrast, although poverty is consistently associated with increased prevalence of psychiatric disorders in low-income, middle-income and high-income countries, some poverty-related risks like food insecurity are potentially more acute in LLMICs. For this reason, the impacts of youth unemployment on mental health outcomes in less developed country contexts may be particularly acute (Patel, 2007; Paul and Moser, 2007; Lund et al., 2010). Finally, the extent to which the greater mental health vulnerabilities associated with female gender differ between LLMICs and HICs has not been established. However, there are grounds for thinking gendered social institutions including schools, marriage, and the workplace in LLMICs may cause different epidemiological patterns of common mental disorders among women (Maselko, 2017).

The impact of disengaging from education and employment on adolescent mental health in developing countries is not well understood. However, the limited research evidence available suggests it could be considerable. Researchers in Mexico found in an urban cohort of children between 12 and 17 years of age NEETs were at greater risk from mood, anxiety and behavioural disorders, substance abuse and suicidal feelings than the average adolescent (Benjet et al., 2012). The study also found children who reported exclusively attending school were the only group with better than average mental health, and that children who worked or combined school with working part-time were as vulnerable to mental ill health as NEETs. Follow-up of the same cohort eight years later not only showed that becoming NEET in adolescence was associated with adult mental health problems, but also indicated that individuals who had worked in their teens were no better off in terms of their subsequent mental health than young NEETs (Gutiérrez-Garcia et al., 2017). However, this evidence is for secondary-school-aged teens and the causes and consequences of disengagement from school and work may differ in older adolescents and/or in the social and economic context of the Philippines.

4.1.2 Education, employment and gender in the Philippines in the early 2000s

Prior to the Enhanced Basic Education Act of 2013 (Republic Act 10533) the Philippines was one of only three countries in the world to have a ten-year basic education cycle, alongside Angola and Djibouti (UNESCO, 2015, p.6). Compulsory schooling covered primary level education only – the first six years or until children were 11 years of age – though the following four years of lower secondary were free in public schools (Asis and Ruiz-Marave, 2013). Since the 1990s the private sector has also emerged as an important provider of education in the Philippines, with approximately one third of students attending private institutions between 2000 and 2005 (Jimenez and Sawada, 2001).

Although education is highly valued by parents in the Philippines, there was a decline in many education outcome indicators in the 2000s as a result of government underinvestment (Maligalig and Albert, 2008; Maligalig et al., 2010). However, secondary school enrolment held constant at around 60 percent between 2000 and 2005, while lower secondary school completion rates – measured as the percentage enrolled in the last year of basic education – were up to 79 percent in 2005 from 67 percent in 2000 (World Bank, 2018). Unexpectedly, completion rates were better among females than males (85% and 69%, respectively). This kind of gender gap is fairly unusual in developing countries and may have been the result of local efforts to encourage girls' education, a conservative culture that promoted better study habits among adolescent girls and/or greater employment opportunities for teenage boys (Tan et al., 2011). In 2005, labour force participation figures from the International Labour Organisation (ILO) indicate that of the 44 percent of young Filipinos between 15 and 24 years of age engaged in the labour market 53 percent were male and 34 percent were female (ILO, 2015).

This shows adolescent girls and young women in the Philippines studied for longer than their male counterparts, and presumably achieved higher qualifications. However, comparatively few appeared to find or seek wage work even after successfully completing school. ILO statistics confirm that 25 percent of Filipinos under 24 years of age were neither in education, employment nor training in 2006, but NEET prevalence was an estimated 32 percent for young women and 18 percent for young men (ILO, 2015). This pattern of low labour market engagement among young women is consistent with a subcategory of NEETs observed in many LLMICs – unemployed non-students, unavailable for work – possible explanations for which include the social and cultural pressure to undertake family and/or other unpaid domestic duties, as well as direct gender discrimination and institutional barriers to employment. Nonetheless, the potential mental health implications for girls and young women in the Philippines of high educational achievement combined with exclusion from the labour market are unknown.

Therefore, the present study examined not only the associations between mental health and exclusion or disengagement from education and employment in a birth cohort from the Philippines, but also differences across gender groups. In addition, it aimed to build on the evidence collected by Benjet and colleagues (2010), investigating the protective effects or health benefits of being in school and/or work compared to becoming NEET in a population of older adolescents. This study tested three main hypotheses: (1) disengagement from school and work during the transition to adulthood is negatively associated with adolescent mental health; (2) disengaged young women are more vulnerable to developing mental health problems than young men; and (3) the largest protective effects are associated with education.

4.2 Data

The cohort data used in this study were from the Cebu Longitudinal Health and Nutrition Survey (CLHNS). The study area comprised a major port and the second largest city in the Philippines (Cebu City), two adjoining cities (Mandaue and Lapu-Lapu) and the surroundings, including households living in peri-urban areas as well as smaller rural towns and nearby mountain and island communities (Adair et al., 2010). Using single stage cluster sampling 17 urban and 16 rural areas were selected at random from a total of 270 *barangays* (in the Philippines a barangay is a village or neighbourhood). Those randomly selected were surveyed for pregnant women and women due to give birth between May 1983 and April 1984 were recruited to the study, making it representative of births in Cebu for this period. In total 3,327 women were interviewed at baseline. Women who had multiple births, a stillbirth or miscarriage, refused to participate further or out-migrated were then excluded (Adair et al., 2010). The final sample of 3,080 women were interviewed during their pregnancy, immediately after the birth, and every two months for next 24 months. Follow up surveys were then carried out every three to four years from 1991 onwards.

This study was primarily concerned with data from the 2002 follow-up when children were between 17 and 19 years of age. In common with earlier rounds, the survey collected information on individual education, employment, and physical health, detailed demographic, social, and economic household level data, as well as barangay infrastructure and services. Unlike previous rounds the 2002 survey included a module on individual mental health for the first time. Data collected during the 1984 baseline, 1991, 1994 and 1998 rounds were also used. The CLHNS has relatively low levels of attrition compared to other LLMIC surveys of similar length (Alderman et al., 2001;

Outes-Leon and Dercon, 2008). However, there are still missing data. In this study there were 1,799 children (866 female and 936 male) who had complete data for each survey round between 1984 and 2002. Appendix 4A presents evidence of some systematic differences between the missing and complete data groups – children included in this study had slightly less well-educated parents and lived in more resource-disadvantaged neighbourhoods compared to excluded children – although no significant differences in household wealth or child development were observed.

Mental health (outcome variable). Adolescent mental health was assessed in 2002 using questions adapted from the Centre for Epidemiologic Studies-Depression Scale (CES-D; Radloff, 1977). The CLHNS research team in Cebu considered the cultural relevance and appropriateness of the original CES-D, excluding the following items 'Felt that I could not shake off the blues even with the help from my family or friends,' 'Felt depressed,' 'Thought my life had been a failure,' 'Talked less than usual,' 'Had crying spells,' 'Felt sad' and 'Could not 'get going''. The team added 'Had headaches,' 'Had poor digestion,' and 'Had the idea of taking my own life' (Hock et al., 2018). These were then translated from English to Cebuano, the local language in Cebu, and back to English by CLHNS researchers. Individuals responded to a total of 16 ordinal items, stating how often in the last four weeks they had experienced symptoms ranging from difficulty falling asleep to loneliness and suicidal ideation. Positive items included 'Felt happy,' Felt hopeful about the future' and 'Enjoyed normal daily activities'. Responses were given on a three-point scale from 'none of the time' (0), 'occasionally' (1) to 'most or all of the time' (2).

Although the measure has been used in previous studies (e.g. Hindin and Gultiano, 2006; McDade, Borja, Adair, & Kuzawa, 2012; Hock et al., 2018), the validity of some items adapted from the CES-D has been questioned. For example, the CLHNS contains four reverse-worded items which in the original CES-D were designed to assess positive affect and break up patterns of response (Radloff, 1977). Yet psychopharmacological research suggests positive and negative affect may be two distinct, uncorrelated dimensions of depression and should therefore be measured separately (Shelton and Tomarken, 2001; Nutt et al., 2007). There is also evidence showing positive worded items are only weakly related to psychological distress (Schroevers, Sanderman, van Sonderen, & Ranchor, 2000; Stansbury, Ried, and Velozo, 2006). Additionally, a validation study conducted in Hong Kong found positive emotions did not relate well to the construct of depression among adolescents in that culture (Lee et al., 2008), leading the study authors to conclude that a lack of positive affect did not appear to contraindicate depressive symptoms in non-Western samples. Researchers have also queried the validity of items measuring social and interpersonal problems (‘Felt that people disliked me’ and ‘Felt people were unfriendly’), cautioning they may be more symptomatic of anxiety than depressive disorders (Carleton et al., 2013). However, difficulties with interpersonal relations have been found to be more prevalent among depressed individuals from non-Western cultures compared to non-depressed individuals from the same culture, and more prevalent compared to depressed European Americans (Kanazawa, White, and Hampson, 2007). Moreover, several studies have found evidence supportive of the importance of social relations for the mental health and wellbeing of adult men in Manila, Philippines, as well as Filipino-American adolescents in Hawaii (Edman et al., 1999; Fernandez, Seyle, and Simon, 2018). To address these concerns this analysis used non-parametric item-response theory to test the psychometric properties of the 16-item measure (see Methods and Data Analysis sections).

Education and employment status (explanatory variables). Adolescent engagement in education and employment was measured using data on self-reported work and school status in 2002. Binary indicator variables were created for four separate subgroups: (1) young people who were neither working, studying at school or college, nor engaged in training (NEETs); (2) individuals who were only working; (3) individuals who were working and studying, either on the same or different days, or during the summer recess; and (4) young people who were only enrolled in education. The NEET group comprised teenagers who were not enrolled in school or vocational training but had not found a job, and young people actively looking for work as well as those who were not looking. Slightly less than half the group in education and part-time employment worked on the same day as they went to school (45%), one quarter worked on a different day (25%), and the remainder reported working only in the summer holiday or a combination of term time (on unspecified days) and during the school break. Adolescents who were exclusively enrolled at school were included in the education only group even if they had fallen behind their peers as a result of temporary dropout or grade repetition, meaning the group may have been heterogenous according to academic attainment.

Other covariates. To mitigate confounding, variables controlling for children's sex and age, birthweight, cognitive function, general health, completed schooling and marital status, in addition to parental education, level of neighbourhood urban development, household wealth and income and household crowding were included in the main structural models. Neighbourhood urban development was measured using a tailor-made index, which was based on results from an earlier factor analysis of data from the CLHNS (McDade and Adair, 2001; Dahly and Adair, 2007). This combined neighbourhood population, population density, communications (cell phones, the internet, cable TV etc.), transportation (the density of paved roads, availability of public transportation), educational facilities, health services, and markets (including grocery stores and gas stations), each of which was assigned values from zero to 10 resulting in a scale ranging from zero to 70. Household wealth was based on reported ownership of an electric fan, furniture, a car, jeepney (a small van), colour television, cassette recorder, and refrigerator at baseline. It also included information on the quality of materials used to construct the house, access to running water, and whether the house had a flushing toilet. Polychoric correlations between these items were estimated and principal components analysis conducted on the correlation matrix (Kolenikov and Angeles, 2004). The first principal component was used as an index of household wealth, with lower values indicating lower levels of wealth and higher values corresponding to higher levels. Crowding was calculated by dividing the of number of people currently resident in the household by the number of rooms.

4.3 Methods

Since the mental health module in the CLHNS had not been previously validated (i.e. it had no known factor structure) and included items adapted from the original CES-D that have had their measurement properties called into question, this study applied exploratory Mokken scale analysis to ascertain the scalability of the measure.

Mokken scale analysis (MSA: Mokken, 1971; Sijtsma and Molenaar, 2002) is from the item-response theory family of models that includes the Rasch model and has been widely used in psychology, psychiatry and public health research to assess relations between items and latent traits (Watson, Deary, and Shipley, 2008; Stewart et al., 2010; Stochl, Jones, and Croudace, 2012). However, non-parametric MSA relaxes the strict assumptions the Rasch model makes regarding the shape of the item response probabilities, making it more appropriate when the number of items is low (van Shuur, 2003). MSA is a probabilistic version of Guttman scaling (Guttman, 1944), which assumes the items in a unidimensional scale have a hierarchical structure. In other words, unidimensional scales are comprised of items ordered in terms of item difficulty – the level of the latent trait at which there is a median probability of an individual endorsing the item or category of item response – so that on a perfectly unidimensional scale individuals responding positively to extreme items will also respond positively to all other less extreme items. MSA therefore involves identifying items that selected to form a scale (i.e. a hierarchical set of items) reflect a single underlying or latent trait.[8]

[8] Although aggregated measures were not used in this study's analysis the sum of scales meeting the Mokken scale criteria is generally considered sufficient to identify an individual's position on the latent trait (Sijtsma and Molenaar, 2002).

The simplest Mokken scale must satisfy the three assumptions that define the model of monotone homogeneity (Mokken, 1971). The first is unidimensionality, meaning all items measure the same underlying trait. The second is the assumption of local independence. This refers to the statistical independence of each scale item after conditioning on a single (unidimensional) latent trait, so that an individual's response to one item is not affected by their response to another similar item. The third assumption is monotonicity, which refers to the probability that endorsing an item or response category is a non-decreasing function of the latent trait. In other words, the higher the value of the latent trait the greater the probability the respondent will answer items measuring the trait correctly.

Mokken scaling works by identifying unidimensional sets of items according to Loevinger's definition of homogeneity and his coefficient H (Loevinger, 1948). This is based on the extent to which items violate the Guttman model (van Shuur, 2003). On a perfectly unidimensional Guttman scale of mental health individuals are expected to respond negatively to items that measure more severe symptoms relative to their position on the latent trait and positively to all items that measure less severe symptoms. Item responses that do not follow the expected sequence are known as Guttman errors. The scalability coefficient H (Loevinger, 1948; Molenaar, 1997) is defined as a function of these errors and calculated to measure the homogeneity of individual items H_i, item-pairs H_{ij} and the overall scale H^s (Hemker, Sijtsma, and Molenaar, 1995). The H value equals $1 - \frac{observed}{expected}$, where observed errors are the number of times item responses do not follow item order and expected values are the probability items are chosen by chance. The closer H is to 1 the better the dimensionality or scalability of the items in the proposed scale; conversely, the closer

H is to zero the worse the scale's properties. Overall scale coefficients between 0.3 and 0.4 are generally interpreted as weak unidimensional scales, 0.4 to 0.5 as moderate and 0.5 or more as strong (Mokken, 1971).

4.4 Data analysis

The analysis was conducted in four main stages: (1) Mokken scaling techniques and confirmatory factor analysis (CFA) were used to investigate the dimensionality of the mental health measure; (2) classical test theory analyses were conducted to establish other psychometric properties of the scale(s) identified using MSA, including data quality, the targeting of items, scaling assumptions and scale reliability and validity; (3) the main structural equation models (SEM) estimated the associations between adolescent mental health, education and employment, adjusting for a set of individual, household and community level covariates; and (4) SEMs tested whether the parameters differed across male and female groups, conditional on establishing measurement invariance. Descriptive sample statistics were also generated for important individual and household characteristics.

In the first stage, a Mokken scale analysis was conducted to identify a scalable set of items using the user-written Mokken Scale Procedure (MSP) module in Stata 14 software (Hardouin, 2011). This involved setting a cut-off for values of coefficient H and allowing the programme to automatically select items to unidimensional scales. Following the suggested procedure by Hemker and colleagues (1995) cut-off c was initially set at $c \geq 0.30$. The programme then chooses the item-pair with highest H_{ij} above this threshold to form the foundation of the scale; the individual item with the next highest coefficient H_i is added to the scale if (1) it has a positive covariance with each of the items in the original pair, (2) the coefficient H_{ij} exceeds c, and (3) the item maximises the overall scale coefficient H^s (Hemker, Sijtsma, and Molenaar, 1995). The automatic MSP repeats the process until there are no more items in the set that meet these conditions, when it either begins constructing a new scale from the unselected items or stops. This procedure was re-run several times, increasing the cut-off for the scalability criteria c from 0.30 to 0.50 in 0.05 increments at a time to explore the dimensionality of the proposed scales. The overall scale coefficient H^s was evaluated for the final scale using confirmatory Mokken analysis (re-entering the group of items that the exploratory MSP found had good scale properties).

CFA was also used to test the fit of the proposed scales. Since the indicators were ordinal this was conducted using the weighted least squares mean variance (WLSMV) estimator and cluster robust standard errors in the R package *lavaan* (Rosseel, 2012). Fit was evaluated using χ^2, the comparative fit index (CFI), the Tucker-Lewis index (TLI), the root mean square error of approximation (RMSEA), and the standardised root mean squared residual (SRMR). Models with insignificant χ^2, CFI and TLI values above 0.95, RMSEA values equal to or less than 0.06 and SRMR values of less than 0.06 are generally considered to fit the data well (Hu and Bentler, 1999).[9] Local independence was assessed by examining the residual correlation matrix and where correlations were greater than 0.2 items were flagged for possible local dependencies (Reeve et al., 2007). The monotonicity of the scale was evaluated with the user-written *loevh* command in Stata 14 software (Hardouin, 2010). Violations of the assumption were checked by looking at the item traces – for ordinal items the traces are the proportion of responses to each category as a function of the latent score – and *crit* scores (Stochl, Jones, and Croudace, 2012). Following the heuristic suggested in Molenaar and Sijtsma (2000) items with *crit* values of 40 or less were not considered to be serious violations of monotonicity.

9 Some model fit indices are known to be sensitive to sample size (e.g. χ^2, CFI) and RMSEA is strongly affected by small degrees of freedom (Kenny, Kaniskan, and McCoach, 2015).

In the second stage the percentage of item-level missing data and the summed score distributions, in addition to the percentage of respondents scoring the minimum and maximum scores, were calculated to assess data quality and evidence of floor and/or ceiling effects.[10] Item-total correlations, the correlation between each item and the total score for all other items, were used to evaluate the scale assumption of a common underlying construct and were interpreted according to the recommended heuristic (Hobart and Cano, 2009). Internal reliability was measured using Cronbach's alpha as well as inter-item correlations. Convergent construct validity – the extent to which the scale(s) measured what it was being used to measure – was also assessed by estimating the association between MSA scale and self-reported rating of general health in a linear regression model adjusted for age and gender.

In the third stage of analysis, associations between mental health, education and employment were estimated using SEMs. A structural equation modelling approach was appropriate because it allowed the estimation of the latent mental health construct, which was indicated by multiple items, and accounted for the presence of measurement error in the indicators (Bollen, 1989). The main model examined whether being in school exclusively, studying and working part-time, working exclusively or being NEET during the transition from school to work predicted young peoples' mental health. Two different models were specified: the first compared the mental health of teenagers in each of the four education and employment groups to

[10] Floor effects were defined as the lowest possible score (and ceiling effects the highest possible score) regardless of the direction of the scale.

the average cohort teen, and the second contrasted the health of teens in school and/or work to NEET mental health (with NEETs as the reference category).

Measurement invariance was evaluated in the fourth and final stage to determine whether the mental health scale could be used to make group comparisons by gender. In other words, that the construct measured the same thing in the same way for young women and young men. This involved testing a series of increasingly restrictive models, following the process described in van de Schoot, Lugtig and Hox (2012). Firstly, the measurement model was fitted for both gender groups independently (configural invariance). Then the model was run for each group simultaneously but with the factor loadings constrained to be equal (metric invariance). Next, the model was rerun but factor loadings were allowed to vary across groups and the intercepts were constrained to be equal. Finally, the model was run with factor loadings and intercepts both constrained across groups (scalar invariance). The first SEM specification – comparing the mental health of teenagers in each of the education and employment groups to the average cohort teen – was then estimated as a two-group model and differences according to gender group were assessed using Wald tests, conditional on measurement invariance being supported.

4.5 Results

Table 4.1 outlines characteristics of the study sample according to work and school status. In the Cebu cohort 23 percent were NEET at 18 years of age, which was very similar to the official population NEET prevalence of 24 percent of youth between 15 and 24 years of age (ILO, 2018). Adolescent NEETs were not significantly different in terms of primary schooling, general health and the urban development of their neighbourhood compared to the average sample teen. Yet NEETs did meaningfully differ by secondary schooling, marital status, parents' level of education, household assets and household income. A significantly larger proportion of the NEET group in the sample was also female. Young people working exclusively made up the single largest group (36% of the total sample). They had the lowest levels of individual and parental education, were poorer in terms of household assets than the rest of the cohort and lived in less developed areas. A significantly larger number of individuals in employment were male (55%) and less than four percent of this group described their general health as poor. The household characteristics of the work and study and study only groups were broadly similar in terms of assets, income and neighbourhood urban development, which were each significantly higher than the sample mean averages. Adolescent and parental levels of education were highest among those who were studying exclusively, although roughly one-tenth of this group had not completed secondary school at 18 years of age. This suggests that a small proportion of the group had fallen behind their peers, which may have been due to delayed enrolment, grade repetition or a pattern of dropout and re-enrolment.

Table 4.1 Selected Sample Characteristics in 2002.

	NEET (n=411)	Work only (n=647)	Work and study (n=265)	Study only (n=476)
Adolescent characteristics		n (%)		
Male	195 (47.6)	357 (55.3)	134 (51.2)	247 (51.9)
Primary school	382 (93.2)	551 (85.3)	263 (100.0)	475 (99.8)
Secondary school	281 (68.5)	390 (60.4)	233 (88.9)	431 (90.6)
Married	72 (17.6)	68 (10.5)	5 (1.9)	3 (0.6)
General health				
Poor	27 (6.6)	26 (4.0)	16 (6.1)	27 (5.7)
Good	310 (75.6)	500 (77.4)	200 (76.3)	352 (74.0)
Excellent	73 (17.8)	120 (18.6)	47 (18.6)	97 (20.4)
Parent/household characteristics				
Mother primary school	152 (37.1)	183 (28.3)	136 (52.0)	297 (62.4)
Mother secondary school	60 (14.6)	78 (12.1)	83 (31.7)	186 (39.1)
Father primary school	180 (43.9)	22 (34.4)	151 (57.6)	314 (66.0)
Father secondary school	94 (22.9)	118 (18.3)	92 (35.1)	220 (46.2)
		means ± SD		
Household assets	-1.72 (0.9)	-1.78 (0.8)	-1.23 (0.9)	-1.06 (1.0)
Income	4.72 (3.8)	4.85 (3.7)	6.09 (4.4)	6.83 (5.8)
Urbanicity	41.34 (13.6)	38.04 (15.2)	44.48 (13.1)	44.68 (12.1)

[1] Values of p correspond to one-way ANOVA F-test statistics (continuous variables) or are based on Pearson's chi-square test (categorical variables).

The scales constructed in the exploratory Mokken scale analysis – which involved increasing the cutoff or minimum values of scalability coefficient H in 0.05 increments from 0.30 through 0.50 – are presented in Appendix 4B. At 0.30 three scales were formed and one item was excluded for failing to achieve the lowerbound value of H ('Felt happy'). Increasing the cutoff from 0.30 to 0.35 resulted in four scalable (unidimensional) item sets and two items were excluded ('Felt happy', 'Felt hopeful about the future'). However, the proposed scales at 0.35 were hard to interpret from a theoretical/clinical perspective. At 0.40 four unidimensional scales were identified and five items failed to reach or exceed the cutoff value ('Felt happy', 'Felt hopeful about the future', 'Felt couldn't overcome difficulties', 'Had difficulty falling asleep', 'Felt lonely').

Table 4.2 Descriptive item statistics and results of Mokken Scale Analysis (c = 0.40).

Item content	Item mean (SD)[a]	Item H (H_i)	Monot. (#vi)
Scale 1			
Thought of themselves as worthless	0.31 (0.51)	0.62	0
Felt life wasn't worth living	0.21 (0.44)	0.61	0
Wished they were dead	0.16 (0.39)	0.61	0
Had the idea of taking own life	0.10 (0.32)	0.60	0
Scale 2			
Had headaches	0.52 (0.54)	0.48	0
Had poor digestion	0.18 (0.39)	0.48	0
Scale 3			
Felt worried	0.76 (0.58)	0.40	0
Felt people were unfriendly	0.49 (0.58)	0.45	0
Felt people disliked them	0.46 (0.56)	0.44	0
Scale 4			
Felt able to face problems	0.75 (0.61)	0.40	0
Enjoyed normal daily activities	0.40 (0.55)	0.40	0

[a] Mean item response (0=None of the time; 1=Occasionally; 2=Most or all of the time).

Since the largest four-item scale (see Table 4.2) was unchanged from 0.40 through 0.50 it was decided there was no new information regarding the dimensionality of the CLHNS measure to be obtained by increasing the cutoff further. CFA showed a four-factor solution adequately fit the data (χ^2 = 219.82, df = 38, $p < 0.001$; CFI = 0.98; TLI = 0.97; RMSEA = 0.04; SRMR = 0.05) despite the high and significant value of chi-square, which was probably a consequence of the relatively large sample size. This factor structure was also consistent with empirical evidence from previous studies suggesting mood disorders like depression have multiple subdomains including positive affect, interpersonal and somatic symptoms (Ellard et al., 2010; Betancourt et al., 2014). Although it is difficult to conclusively establish local item independence none of the item pairs in the four scalable item sets had high residual correlation (greater than 0.2) suggesting the local independence of items within the proposed scales. The results presented in Table 4.2 also indicate there were no violations of monotonicity for any items according to #vi values which was supported by positive item traces. However, due to the minimum three-indicator per latent factor rule the main analysis was carried out using the four-item subscale only (Kline, 2016).

Descriptive statistics for the four-item depression subscale indicate no item-level missing data although there was evidence of substantial floor effects, with most respondents scoring zero (Appendix 4C). It is not uncommon for measures using three-point Likert-type scale indicators to exhibit floor effects (Deighton et al., 2014), but their presence suggests that while the scale measured severe symptoms of depression it was unable to discriminate between respondents with lower levels of symptom severity. The summed score distribution displayed in Appendix 4D highlights the extent of the issue. Internal reliability of the scale was acceptable (Cronbach's alpha was between 0.70 and 0.95) (Streiner, 2003). In addition, the alternative measure of item homogeneity (mean inter-item correlation) was also within the recommended range of 0.40 to 0.50 (Clark and Watson, 1995). Scale validity was underscored by the significant positive association between higher (worse) mental health scores and poorer self-reported rating of general health (Appendix 4E).

Prior to testing for gender differences in the relationship between mental health, education and employment, measurement invariance had to be established. Results indicate that the mental health construct was only partially invariant (Appendix 4F). That is, that scalar invariance could only be established after allowing one of the constrained factor loadings to freely vary. Nevertheless, partial measurement invariance is generally accepted as a sufficient condition for valid inferences to be made regarding between-group differences in latent factor means (Byrne, Shavelson, and Muthén, 1989; Dimitrov, 2010).

The main SEMs fit the data adequately (Appendix 4G). Standardised and unstandardised estimates of the model parameters are presented in Tables 4.3 and 4.4. *Hypothesis 1* (disengagement from school and work during the transition to early adulthood is negatively associated with adolescent mental health): Adolescents in the Cebu cohort who were neither in education nor employment had significantly poorer mental health compared to their peers (positive coefficients signified an increased prevalence of depressive symptoms). Specifically, the mean mental health score for a teenager in the NEET group was 0.151 higher or 0.07 standard deviations above (i.e. worse than) than the average sample teen. In contrast, young people who were in work or a combination of school and work had better than average mental health, indicated by the negative sign of the coefficient, although these scores were not statistically significant. The average mental health score among the study only group of teenagers was for practical purposes indistinguishable from the sample mean, with the interval estimate distributed evenly either side of zero.

Hypothesis 2 (disengaged young women are more vulnerable to developing mental health problems than young men): The results from the group comparison, which investigated sex differences in the relationship between adolescent mental health, education and employment, are also shown in Table 4.3. After running SEMs separately for each gender group both male and female NEETs appeared to be equally vulnerable, although depressive symptoms among working young men in the sample were noticeably less prevalent than in non-working males. A similar disparity was observed between young women in school and part-time work and the average female. Nonetheless, results from the formal testing of these differences lacked precision, so the study hypothesis was not supported.

Table 4.3 Standardised and unstandardised estimates of the association between adolescent depression symptoms, education and employment.[a]

	Whole sample (n=1799)		**Female** (n=865)		**Male** (n=934)	
	Coef. (95% CI)	Stand.	Coef. (95% CI)	Stand.	Coef. (95% CI)	Stand.
NEET	0.151** (0.02, 0.29)	0.068	0.232 (-0.07, 0.53)	0.075	0.172* (-0.03, 0.37)	0.077
Work only	-0.082 (-0.21, 0.04)	-0.042	-0.013 (-0.28, 0.26)	-0.040	-0.165* (-0.35, 0.02)	-0.088
Work and study	-0.102 (-0.27, 0.06)	-0.039	-0.223 (-0.60, 0.16)	-0.058	-0.039 (-0.28, 0.20)	-0.015
Study only	0.020 (-0.12, 0.16)	0.010	-0.054 (-0.33, 0.22)	-0.018	0.045 (-0.16, 0.25)	0.022

*** $p < 0.001$, ** $p < 0.05$, * $p < 0.1$.

[a] SEMs adjusted for child sex and age, whether the child was born underweight (less than 2500 grams), IQ score measured at 11 years of age, general health, whether primary and secondary school was completed, child marital status, household assets and income, level of neighbourhood urban development, household crowding, and parents' level of education.

Hypothesis 3 (the largest protective effects are associated with education): The results presented in Table 4 show that being in education and/or employment had a protective influence on young peoples' mental health in the Philippines. The coefficients in the unadjusted model were negative, which suggests engaging in any combination of school or work was better for adolescent mental health than being NEET (the negative sign is interpreted as a reduction in depressive symptoms or as a protective effect). The estimated average decrease in mental ill health was greatest in the combined work and study group, followed by the group that studied exclusively and the work only group (the coefficient for which was not statistically significant). Model 1 (Table 4.4) included individual covariates like sex, age and general health to adjust for potential confounding. There was no change in coefficient signs although there was a difference in magnitude compared to the unadjusted model.

Table 4.4 Standardised and unstandardised estimates of the association between adolescent depression symptoms, education and employment (reference category=NEET).

	Unadjusted		**Model 1**[a]		**Model 2**[b]	
	Coef. (95% CI)	Stand.	Coef. (95% CI)	Stand.	Coef. (95% CI)	Stand.
Work only vs NEET	-0.111 (-0.25, 0.03)	-0.060	-0.150* (-0.30, 0.00)	-0.077	-0.158** (-0.31, -0.01)	-0.081
Study only vs NEET	-0.165** (-0.32, -0.01)	-0.082	-0.10 (-0.26, 0.06)	-0.047	-0.099 (-0.27, 0.07)	-0.047
Work and study vs NEET	-0.250** (-0.43, -0.07)	-0.100	-0.201** (-0.39, -0.01)	-0.076	-0.196** (-0.39, -0.00)	-0.074

*** $p < 0.001$, ** $p < 0.05$, * $p < 0.1$.

[a] SEMs adjusted for child sex and age, whether the child was born underweight (less than 2500 grams), IQ score measured at 11 years of age, general health, whether primary and secondary school was completed and child marital status.

[b] SEMs adjusted for household assets and income, level of neighbourhood urban development, household crowding and parents' level of education (in addition to individual covariates in Model 1).

The protective effects of studying and working and studying only were reduced (while the coefficient for the study only group lost its significance). In contrast, after accounting for key individual characteristics the benefit of being exclusively in work increased in strength and was borderline statistically significant.

Results from the fully adjusted SEM (Table 4.4, Model 2) reveal that compared to being NEET young people who were working and studying had a mean latent mental health score that was -0.196 (0.07 SD) lower/better after adjusting for individual, household and community characteristics. The group of adolescents who were working only also had significantly lower mean scores than the NEET group. The mental health of teenagers who were exclusively studying was better than the mean score among NEETs, but contrary to the study hypothesis the protective effect of schooling was relatively weak for older adolescents (and not statistically significant). In general, these findings show that becoming NEET in adolescence is associated with poorer mental health in LLMICs, but that a combination of education and/or employment may mitigate the effects.

4.6 Discussion

This study examined the associations between adolescent mental health, education and employment in a group of 17 to 19-year-olds from Cebu, Philippines. The analysis revealed that young people who were neither in education nor employment (NEETs) had significantly poorer mental health than their age-group peers even after adjusting for characteristics like level of completed schooling, cognitive function, general health and family socio-economic position. However, the association was not significantly different for young women and men. This study also assessed the protective influence on adolescent mental health of having a job and/or being in school compared to being NEET and found that young people benefited from working full-time or combining part-time work with study. In contrast, differences between the mental health of teenagers exclusively enrolled in school and their NEET peers were fairly modest, potentially reflecting a shift in the relative value of employment over education among older adolescents. These findings show that becoming disengaged from school and the workplace during late adolescence may have important implications for individual mental health and wellbeing in LLMICs, underscoring the challenges facing young people during the transition from school to work and early adulthood.

The association between disengaged young Filipinos and mental ill health supports a number of different causal explanations. On the one hand, the findings suggest adolescents no longer in school but not in work are at greater risk of developing mood disorders like depression. This is in line with prior research which showed the connections between becoming NEET and adolescent mental health were independent of historical vulnerability to mental health problems (Goldman-Mellor et al., 2016; Gutiérrez-García et al., 2017). On the other hand, the reverse cannot be ruled out since mental health and being NEET were measured simultaneously, preventing this study from discounting the influence of pre-existing mental health vulnerabilities. So, the findings are also not inconsistent with longitudinal evidence suggesting an individual's history of mental health predicts their subsequent NEET status (Baggio et al., 2014; Veldman et al., 2015). In addition, the mechanisms underlying the association could be bidirectional (Goldman-Mellor et al., 2016). From this perspective, being NEET increases the risk of developing mental health problems, through loss of social support or exposure to stressful financial pressures, but prior mental ill health heightens the risk of becoming NEET, causing young people to leave school early or without the skills to succeed in the labour market. Pre-existing vulnerabilities and the exposure to new risks are both expected to negatively affect job search.

Although establishing causality was beyond the scope of this study, these findings demonstrate that disengaging from school and work in adolescence should be considered a mental health risk factor in the Philippines as well as potentially other LLMICs. They highlight the importance of mental health awareness in youth unemployment programmes in developing countries as well as the need to promptly identify and support at-risk youth in school. Moreover, the results of this study suggest more comparative and country-specific analyses conducted in LLMICs would be worthwhile, since mental disorders are so strongly socially determined and the economic, social and cultural conditions across developing countries are so diverse (Lund et al., 2018).

The epidemiological evidence indicates the overall burden of anxiety and mood disorders is higher in women than men, which is widely held to have both biological and social underpinnings (Patel and Kleinman, 2003; Seedat et al., 2009). Yet formal tests to determine whether the mental health effects of adolescent education and employment differed according to sex were inconclusive, leaving the potential role played by education, employment and gender in female mental health in the developing world to future investigations.

The protective influence of studying and working part-time found in this analysis is consistent with the inverse association between education and common mental disorders observed in many developing countries and emphasises the benefits of young people reducing their exposure to mental health risk factors (Araya, Lewis, Rojas, & Fritsch, 2002; Patel, 2007). However, after adjusting for individual, family and community characteristics the effects were noticeably smaller and only borderline significant. The reduced effect size may also capture the complex relationship between education and family socio-economic position in many LLMICs.

Compared to the NEET group there appeared to be modest health benefits to being exclusively enrolled in school, although estimates were not statistically significant. This contrasts with results from an earlier study conducted in Mexico that showed exclusive enrolment in school was the best (and only) buffer against adverse mental health outcomes (Benjet et al., 2012). Possible explanations for the discrepancy include differences in sample age – adolescents in the Cebu cohort were three to four years older than the teens in the Mexico study – suggesting the protective effects of education may diminish among older adolescents confronting the labour market's risks and rewards. Nonetheless, future research examining the influence of independent risk factors associated with post-compulsory schooling in low-income and lower-middle-income countries, including academic pressure, demanding workloads and the financial strain on families would be valuable.

Significantly improved mental health was observed among employed teens in Cebu after accounting for individual differences in cognitive functioning, education, and family socio-economic position. This suggests that at 18 years of age adolescent health may be less vulnerable to exposure to employment-related risk factors in developing countries than to the stress of needing or wanting to work and being unable to find a job (Chopra, 2009; Kortum, Leka, and Cox, 2011). Unemployment is major dimension of poverty in low-income and lower-middle-income countries and is linked with many of the underlying mechanisms connecting poverty and mental ill health, including food and housing insecurity and feelings of hopelessness (Patel and Kleinman, 2003). Adult unemployment has been found to be associated with increased mood disorders like depression as well as suicidal ideation (Vorcaro, Lima-Costa, Barreto, & Uchoa, 2001; Lund et al., 2010; Iemmi et al., 2016). However, no previous studies are known to have examined whether the positive mental health benefits of being employed extend to adolescents, so these findings represent an novel contribution to the evidence on mental health and youth employment in LLMICs and support policies and programmes aimed at reducing youth unemployment in the Philippines and elsewhere.

The present study has several limitations that must be considered along with these findings. First, there may have been unobserved confounding of the relationship between adolescent mental health, education and employment. For example, individual confounders such as maternal history of depression were not measured in the data. Second, the analysis was restricted to young men and women around 18 years of age, excluding younger and older NEET youth. Third, this study lacked the information to account for the duration and frequency of being NEET, which may affect disease severity. How and why individuals end up NEET is also known to modify the impact it has on mental health outcomes, although the data were not collected in the CLHNS (Gutiérrez-Garcia et al., 2017). Lack of data meant the study could not examine the role played by the gendered reasons behind the disengagement from education and employment of young women and men, which are expected to influence the transition to adulthood for young women in particular. Fourth, although scales like the CES-D measuring self-reported symptoms have been recognised as an important first step towards the preliminary identification of cases of depression, the information is not equivalent to a diagnosis and cannot replace a clinical interview (Breslau, 1985; Caracciolo and Giaquinto, 2002; Dardas et al., 2019). Finally, the scale showed considerable floor effects suggesting it was insensitive to less severe symptomology.

Notwithstanding these limitations, the implications of this study's findings are important. Adolescents who become disengaged from school and work have poorer mental health than their peers and could struggle to complete a successful transition from school to work. Strategies designed to increase employment and vocational training opportunities may be more effective than efforts focused solely on keeping older teens in education. However, these findings also support integrated policies aimed at tackling early school attrition among at-risk young people and mitigating the adverse impact of youth unemployment through increased mental health awareness.

Disengaged youth in low-income and lower-middle-income countries have received relatively little research attention, yet evidence from this study shows they are a vulnerable group who need support if they are to avoid becoming socially excluded as adults and developing countries are to take full advantage of the demographic dividend their young people represent.

Chapter 5.

Conclusions and Recommendations

5.1 Conclusions

Whether or not the global population of children and young people currently living in low-income and lower-middle-income countries (LLMICs) achieve their development potential has far-reaching consequences for them and for their children. It is also central to achieving the plan for sustainable human development set out in the United Nation's Sustainable Development Goals (SDGs) and agreed by the international community (United Nations General Assembly, 2015). The centrality of child and youth development to the 2030 Agenda for Sustainable Development is evident from the continuing commitment in the SDGs to the health, education and nutrition of young people first charted in the Millennium Development Goals (MDGs). Yet compared to the earlier goals, the SDGs adopt an unambiguously holistic approach to child and adolescent development.

For example, the goal for education (SDG4) builds on the original MDG of universal primary school enrolment, including ambitious targets with respect to enrolment in secondary education and a new emphasis on learning outcomes. SDG4 also links education to the developmental progression of skills across the first 8,000 days, recognising key inputs from early childhood development programmes and pre-school education, and outcomes such as leaving school with the skills to successfully make the transition out of education and into the workplace. Equally, the goal on health (SDG3) maintains the invaluable MDG-era commitment to reducing maternal and infant mortality but also targets the global burden of disease attributable to mental health disorders, acknowledging the serious risk to adolescent and young adult potential posed by mental health problems.

This shift has been driven by research into the problem of loss of child and adolescent development potential in LLMICs (e.g. Grantham-McGregor et al., 2007; Walker et al., 2007, 2011; Patton et al., 2016; Black et al., 2017). For example, the 2007 *Lancet Series* on child development in developing countries established the major poverty-related biological and psychosocial risk factors that cause inequalities of development among children under five years of age, as well as describing the social and economic costs of wasting the human potential of more than 200 million children worldwide. The second *Series* in 2011 reviewed the evidence on risks, such as inadequate cognitive stimulation and linear growth retardation, and protective factors, like breastfeeding and maternal education. It also summarised the evidence on the underlying links between exposure to adversity and brain development in young children. More recent research not only confirmed the life-course consequences of risks in the early years for the acquisition of cognitive and socioemotional skills during middle-childhood and adolescence, but also identified an urgent need to protect and promote development as children get older (Black et al., 2017). For example, the emotional and cognitive competencies necessary for adolescents to finish school, find work and engage in family and community life have been shown to be as vulnerable to exposure to risk factors as sensory-motor and cognitive-language development in early childhood (Patton et al., 2016).

These research efforts have also served to highlight the importance of ecological perspectives for our understanding of child and adolescent development (e.g. Bronfenbrenner and Morris 2006), which emphasise the direct and indirect influences on development of complex reciprocal interactions between individuals and families, neighbourhoods and the larger social contexts across the life course, as well as the role played by risk factors in the physical and social environment.

This investigation into the influence of risks on child and adolescent development in the Philippines further underscores the value of ecological frameworks like Bronfenbrenner's. First, its findings support the bioeocological view that exposure to one risk factor is likely to lead to the exposure to many, particularly in LLMICs where risks are expected to cluster together. For example, this research not only found that housing instability was associated with developmental deficits or delays among pre-school-age children, but its results also suggest that negative changes in the physical and/or social conditions in neighbourhoods involve additional exposures to risks. The results presented here are also consistent with the potential accumulation of multiple exposures to risks over time. For instance, the inability to stay in education and/or acquire the necessary skills for employment which this research has linked to adolescent mental health is likely to reflect the cumulative influence of infant malnutrition, compromised cognitive and socioemotional development, lack of school readiness and inadequate schooling.

Second, the results of this work draw attention to the dual role played by the biological and psychosocial dimensions of the *same* risk central to the bioecological model. For example, the evidence indicates the effects of exposure to excreta on infant linear growth appear to mediate the associations between inadequate WASH conditions in early life and subsequent schooling. Nonetheless, the consequences of malnutrition are not limited to physical development and may include reduced interactions between infants and their environment or with parents and caregivers which undermines the development of essential cognitive and socioemotional competencies.

Third, evidence from this research is consistent with ecological theory that states the primary mechanism underlying compromised child development is the disruption of key developmental processes, and the closer the risk factor to these processes the stronger the effects. However, it also supports the conclusion that these proximate risks are highly correlated with contextual factors. For example, this work found having a flushing toilet is less likely to benefit a family if community sanitation coverage is poor, while lacking improved sanitation facilities is not expected to pose the same risk if other household or neighbourhood vulnerabilities are low. Stressful neighbourhood environments may exert a greater influence on child development outcomes compared with within-family disruptions. Equally, individual adolescent health will reflect adverse conditions in the wider social and economic environment such as high levels of youth unemployment.

The current work responded to the evidence gaps, contributing to the literature on child and adolescent development in low-income and lower-middle-income countries, and shedding light on the influence of risk factors in the physical and social environment on health, education and development outcomes on which achieving the SDGs depends. Its specific findings are summarised as follows:

- Education inequalities are linked to the conditions in the physical home environment in early life. Specifically, the results presented in Chapter 2 indicate infant exposure to faecal-contaminated environments shortens the overall duration of schooling in a birth cohort from Cebu, Philippines. Subgroup analysis suggests malnourished girls are worst affected.

- Pre-primary-school age children living in predominately urban areas may be vulnerable to changes in their neighbourhood environment resulting from residential mobility, which is associated with developmental deficits or delays. The evidence from Chapter 3 shows the adverse cognitive effects observed in tests conducted at 11 years of age were moderated by move quality or the relative change in neighbourhood infrastructure and services. The effects of moving appeared to be stronger for female children than for males.

- Socially marginalised adolescents have an increased risk of developing mental health problems. Specifically, Chapter 4 results suggest young people in the Philippines who become disengaged from the important socialising institutions of school and the workplace may have a higher risk from mental illness. For older teens the protective effects associated with being employed (compared to being NEET) are greater than those linked to being in full-time education.

- The effects of exposure to risks in the physical and social environment at different developmental stages of childhood vary according to gender, with female children more vulnerable than males.

Several recommendations for policymakers and development practitioners working towards the sustainable development agenda in 2030 can be made on the basis of these findings:

1. Invest in children during middle childhood and adolescence. In the estimated 18 to 21 years it takes for a child to reach adulthood there are multiple developmentally important periods. Of these the period covering the first two or three years of life is rightly regarded as a priority. Yet there are several subsequent phases corresponding with age-specific biological changes and the adoption of distinctive social roles in new contexts which are also believed to be critical to achieving overall development potential (Patton et al., 2018). The evidence from this research underscores this important message and its findings highlight some of the possible consequences of underinvestment in older age-groups. For example, the investigation into mental health and social marginalisation revealed not only the serious health burden facing disengaged young people in LLMICs, but also how adolescent health and wellbeing underpins the transition from school to work and early adulthood and therefore guarantees earlier investments in nurturing care and quality education.

2. Address low-income and lower-middle-income country evidence gaps. Applying evidence collected in high-income country settings to LLMIC populations is potentially problematic, since the kind and/or prevalence of risk factors in Europe and North America may dramatically differ to those in the countries of the Global South. For example, the social determinants of health and development are heterogenous according to demographic, political, cultural and economic context, so the effects of risks in the social environment are likely to vary between Western, high-income countries and LLMICs by poverty and gender. However, findings from this work also highlight variation in the effects associated with exposure to risks between developing countries. Researchers in Mexico found that social marginalisation increased the risk of 15-year olds developing mental health problems, and similar results (presented in Chapter 4) were observed in a cohort of 18-year olds from the Philippines. Yet the Mexico study found the only form of social inclusion that improved teen mental health was being in school, whereas in the older Philippine group the strongest protective effect was seen in adolescents with jobs. Although differences in study design could account for the discrepancy, these findings suggest the need for flexible policies that are sensitive to local contexts, underscore the current limitations of the evidence from LLMICs and show the value of future comparative and country-specific analyses.

3. Consider multisectoral collaboration in response to cross-cutting issues in child development, health and education. The linkages between health and education are long established – *mens sana in corpore sano* – but systematic responses to the issues in LLMICs are limited, despite growing evidence of the ways a wide variety of sector-specific goals are interdependent. For example, the social determinants of common mental disorders include the demographic (age and gender), economic (income inequality and unemployment), neighbourhood (poor housing structures, overcrowding and inadequate recreation), environmental (trauma and distress caused by conflict or forced migration) and social and cultural (social exclusion) domains (Lund et al., 2018). New findings in this work also indicate the interconnectedness of water and sanitation environments with early childhood development and education, gender and housing, as well as the influence of education and youth employment on adolescent health. A lack of coordination in the application of evidence to the design and implementation of interventions is therefore expected to undermine efforts to achieve the ambitious sustainable development agenda.

A holistic ecological perspective is essential for understanding how children can be helped to realise their potential and low-income and lower-middle-income countries can make progress towards sustainable development. This research addresses important gaps in the evidence and makes several recommendations on the basis of its findings to policymakers and development practitioners. Making evidence-informed decisions can positively influence child development outcomes and improve real lives, but in LLMICs the evidence is often scarce. This work hopes to build the evidence base.

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7. Appendix

Table 1A Comparison of means between attriting and non-attriting families at baseline.

Characteristic	Non-attriting (n=1888)	Attriting (n=1189)	*p*-value[1]
	means ± SD		
Birth weight, g	3009.2 (438.8)	2969.1 (457.0)	0.016
Height-for-age	-1.68 (1.1)	-1.73 (1.2)	0.238
Weight-for-height	-0.95 (1.0)	-1.0 (1.1)	0.181
Mother height, cm	150.6 (5.1)	150.6 (5.1)	0.945
Mother age, y	26.1 (6.0)	26.0 (6.0)	0.595
Household income, pesos	278.2 (467.0)	284.5 (592.3)	0.744
Household wealth	-0.03 (1.2)	0.05 (1.3)	0.080
Household size	5.7 (2.8)	5.6 (2.9)	0.258
Household crowding	0.35 (0.2)	0.33 (0.2)	0.020
Neighbourhood urban development	29.5 (12.9)	32.3 (12.0)	<0.001
	n (%)		
Male	994 (53)	630 (54)	0.568
Mother primary school	1315 (70)	841 (72)	0.228
Mother secondary school	441 (23)	307 (26)	0.078
Father primary school	1350 (72)	883 (75)	0.022
Father secondary school	575 (30)	441 (38)	<0.001

[1] Values of *p* correspond to one-way ANOVA *F*-test statistics (continuous variables) or are based on Pearson's chi-square test (categorical variables).

Table 1B Attrition tests; determinants of school enrolment (probit regression)

	Full sample		Non-attriting sample	
	Coef.	*p*-value	Coef.	*p*-value
Male	-0.135	0.280	-0.204	0.162
Age	-0.228	0.229	-0.298	0.136
Birth weight	-0.000	0.837	-0.000	0.690
Height-for-age	0.117	0.081	0.144	0.066
Weight-for-height	0.080	0.204	0.070	0.343
Mother height	-0.000	0.981	-0.007	0.611
Mother age	-0.004	0.713	0.003	0.777
Mother years of schooling	0.060	0.015	0.070	0.016
Father years of schooling	0.004	0.839	0.006	0.810
Household income	-0.025	0.411	-0.004	0.923
Household wealth	0.114	0.239	0.201	0.126
Household size	0.064	0.034	0.056	0.119
Household crowding	1.584	0.002	1.605	0.006
Neighbourhood urbanicity	0.007	0.195	0.007	0.213
Constant	1.135	0.576	2.230	0.330
Chi-square test (prob>chi2)	-	-	18.55	0.183

Standard errors clustered at barangay level.

Table 1C Attrition tests; determinants of cognitive test performance (OLS regression).

	Full sample		Non-attriting sample	
	Coef.	*p*-value	Coef.	*p*-value
Male	-1.317	0.011	-1.756	0.002
Age	-0.106	0.925	0.050	0.968
Birth weight	0.000	0.466	0.000	0.892
Height-for-age	0.806	0.004	0.849	0.006
Weight-for-height	1.003	<0.001	0.897	0.003
Breastfed	0.585	0.630	1.033	0.456
Mother height	-0.025	0.643	-0.034	0.557
Mother age	-0.005	0.919	-0.009	0.866
Parity	1.174	0.118	0.979	0.235
Mother years of schooling	0.539	<0.001	0.531	<0.001
Father years of schooling	0.609	<0.001	0.636	<0.001
Household income	0.003	0.981	-0.066	0.612
Household wealth	0.396	0.187	0.224	0.495
Household size	0.007	0.952	0.012	0.933
Household crowding	4.154	0.011	4.289	0.016
Neighbourhood urbanicity	0.040	0.072	0.044	0.066
Constant	44.725	<0.001	47.890	<0.001
Chi-square test (prob>chi2)	-	-	21.28	0.170

Standard errors clustered at barangay level.

Table 2A Results from a comparison of sanitation risks.

	Model 1[a]			Model 2[b]			Model 3[c]		
	OR	95% CI		OR	95% CI		OR	95% CI	
Exposure risk 1									
Sanitation	1.638***	1.389	1.932	1.298**	1.094	1.539	1.181*	0.993	1.405
Exposure risk 2									
Sanitation/materials	1.738***	1.433	2.109	1.347**	1.102	1.645	1.195*	0.975	1.464
Exposure risk 3									
Sanitation/drainage	1.826***	1.478	2.256	1.405**	1.125	1.756	1.254**	1.001	1.570

*** $p < 0.001$, ** $p < 0.05$, * $p < 0.1$.

[a] Multilevel logistic regression model adjusted for child sex, age, birthweight.

[b] Model adjusted for child sex, age, birthweight, breastfeeding, mothers' age, height, parents' level of education, household size and neighbourhood urban development.

[c] Model adjusted for child sex, age, birthweight, breastfeeding, mothers' age, height, parents' level of education, household wealth and income, household size and neighbourhood urban development.

Table 2B Two- and three-way interaction models (sanitation risk by gender; gender by stunting; sanitation by risk by stunting).

	Female*sanitation			Female*stunted			Female*stunted*sanitation		
	OR	95% CI		OR	95% CI		OR	95% CI	
Unimproved sanitation	1.123	0.905	1.394	1.190	1.000	1.416	1.123	0.905	1.394
Female	0.744	0.629	0.881	0.737	0.638	0.851	0.754	0.632	0.900
Female*sanitation	**1.116**	**0.853**	**1.462**	-	-	-	0.939	0.703	1.254
Stunted	-	-	-	0.785	0.605	1.019	1.147	0.809	1.625
Female*stunted	-	-	-	**1.641****	**1.065**	**2.529**	1.002	0.547	1.835
Sanitation*stunted	-	-	-	-	-	-	0.440	0.264	0.735
Female*stunted*sanitation	-	-	-	-	-	-	**2.827****	**1.182**	**6.761**
Age	1.383	1.197	1.598	1.378	1.193	1.591	1.384	1.198	1.600
Breastfed	0.842	0.616	1.152	0.836	0.611	1.144	0.831	0.607	1.138
Birth weight	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
Mother’s age	0.983	0.971	0.994	0.983	0.972	0.994	0.982	0.971	0.993
Mother primary school	0.890	0.754	1.051	0.888	0.752	1.049	0.883	0.748	1.044
Mother secondary school	0.659	0.545	0.797	0.659	0.545	0.797	0.663	0.548	0.802
Father primary school	0.890	0.750	1.057	0.891	0.750	1.058	0.893	0.752	1.060
Father secondary school	0.783	0.658	0.932	0.783	0.658	0.932	0.783	0.658	0.932
Urbanicity	0.997	0.990	1.004	0.997	0.990	1.004	0.998	0.991	1.005
Household size	1.033	1.007	1.060	1.032	1.006	1.059	1.031	1.005	1.058
Household assets	0.800	0.742	0.862	0.802	0.745	0.864	0.802	0.745	0.865
Household income	0.986	0.970	1.003	0.986	0.969	1.003	0.987	0.970	1.004
Between Barangay	0.060	0.025	0.143	0.061	0.026	0.145	0.061	0.025	0.144

*** $p < 0.001$, ** $p < 0.05$, * $p < 0.1$.

Table 2C Multinomial logistic regression of competing risks (odds ratio).[a]

	Stopping school			Dropping out		
	OR	95% CI		OR	95% CI	
Unimproved sanitation	1.244***	1.062	1.458	1.009	0.671	1.515
Male	1.438***	1.148	1.802	2.351***	1.557	3.551
Age	1.426***	1.217	1.670	1.387***	1.146	1.678
Breastfed	0.921	0.707	1.200	1.206	0.611	2.380
Birth weight	1.000	1.000	1.000	1.000	1.000	1.000
Mother's age	0.983*	0.971	0.995	0.999	0.971	1.027
Mother primary school	0.802**	0.654	0.984	0.599***	0.427	0.840
Mother secondary school	0.627***	0.490	0.802	0.695	0.442	1.093
Father primary school	0.914	0.733	1.139	1.310	0.838	2.048
Father secondary school	0.758***	0.634	0.905	0.748	0.520	1.076
Urbanicity	0.996	0.989	1.003	0.995	0.987	1.003
Household size	1.037**	1.003	1.073	1.034*	0.981	1.090
Household assets	0.765***	0.706	0.829	0.719***	0.637	0.812
Household income	0.987**	0.971	1.002	0.998	0.953	1.046

*** $p < 0.001$, ** $p < 0.05$, * $p < 0.1$.

[a] Single level multinomial logistic regression model with standard errors clustered at barangay level.

Figure 3A Kernel densities of mover and non-mover propensity scores before and after matching.

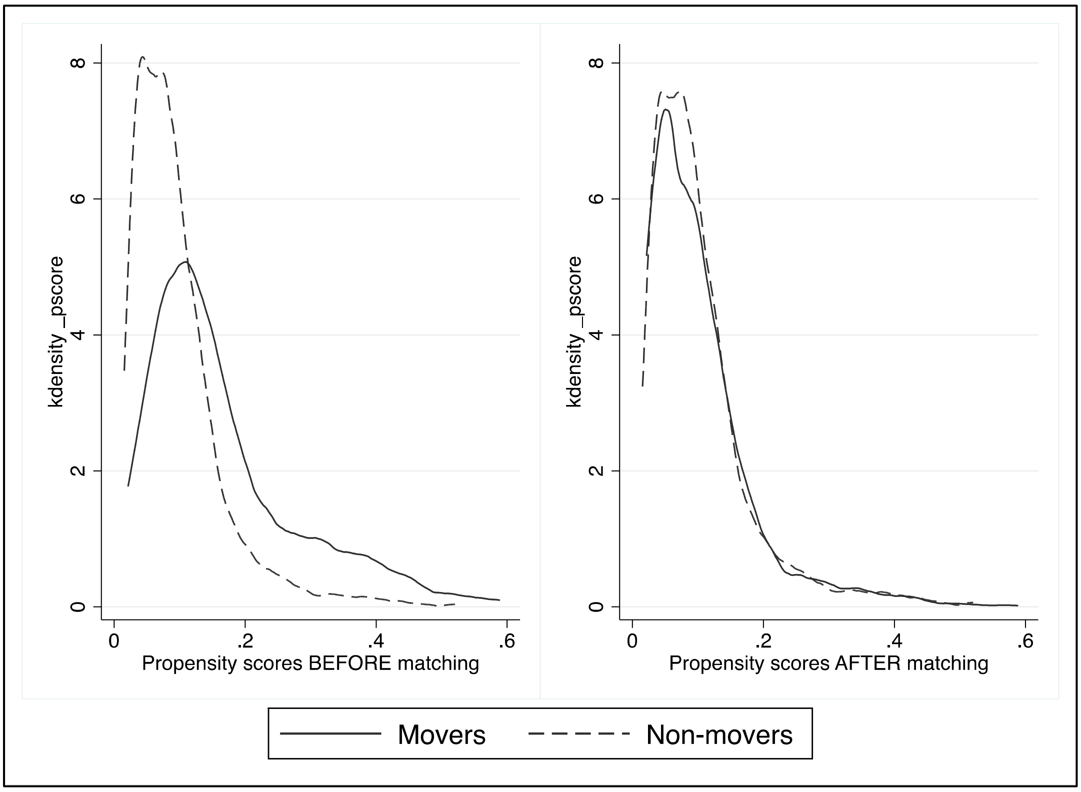

Table 3B Moderating effects of gender on associations between residential mobility and cognitive performance.

	Neighbourhood move [a]		
	Coef.	95% CI	
Male	-0.784	-2.375	0.806
Female	-2.613**	-4.519	-0.706
Female vs male [b]	-1.932	-4.302	0.436

*** $p < 0.001$, ** $p < 0.05$, * $p < 0.1$.

[a] Multilevel regression model with interaction term adjusted for cognitive test scores in 1991 and a complete set of baseline covariates (child sex, age, birthweight and length, HAZ, breastfeeding, parity, mothers' age, height, both parents' level of education, household wealth and income, household size, and neighbourhood urban development).

[b] χ^2 (1) = 2.56

Figure 3C Average marginal effects and contrasts between male and female children by level of neighbourhood development.

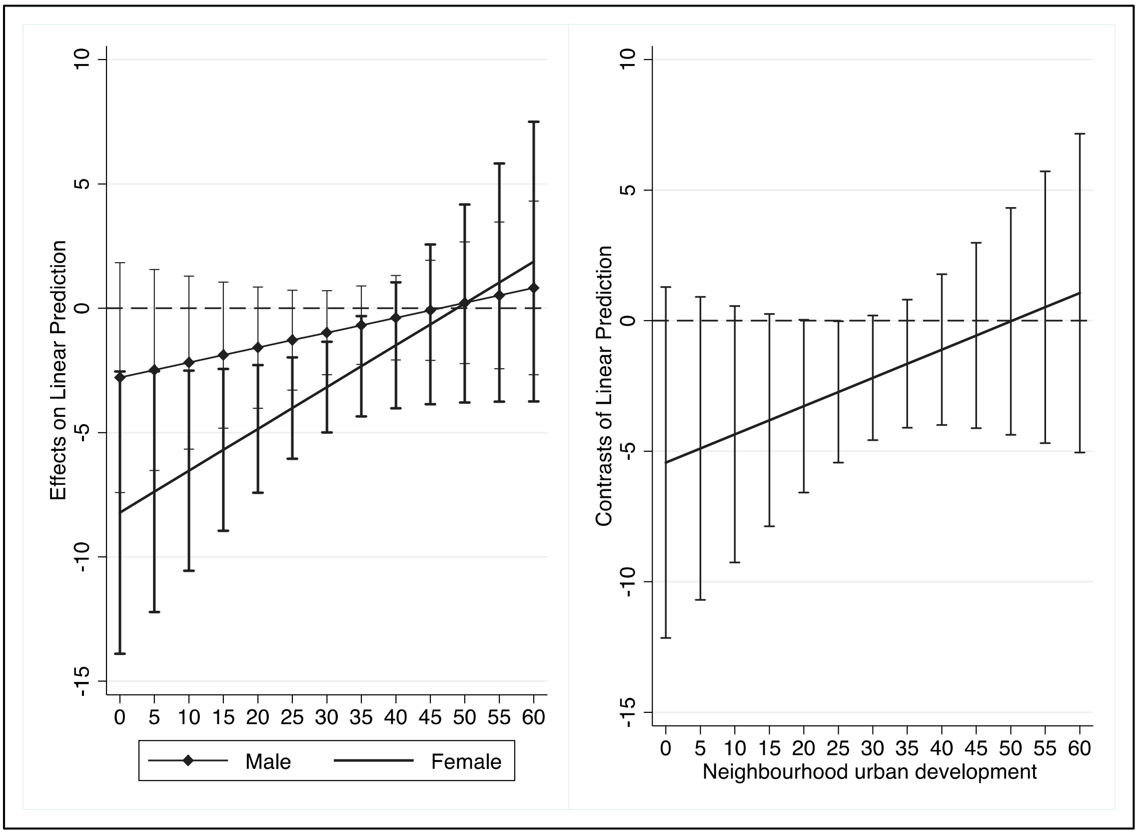

Table 3D Moderating effects of school moves on associations between residential mobility and cognitive performance.

	Neighbourhood move [a]		
	Coef.	95% CI	
School move			
No	-1.408	-3.598	0.783
Yes	-2.255*	-4.773	0.263
School move vs no school move [b]	-0.847	-4.520	2.825

*** $p < 0.001$, ** $p < 0.05$, * $p < 0.1$.

[1] Multilevel regression model with interaction term adjusted for cognitive test scores in 1991 and a complete set of baseline covariates (child sex, age, birthweight and length, HAZ, breastfeeding, parity, mothers' age, height, both parents' level of education, household wealth and income, household size, and neighbourhood urban development).

[b] χ^2 (1) = 0.20

Table 4A Characteristics of children and families in CLHNS at baseline.

Characteristic	Missing (n=1261)	Complete (n=1799)	*p*-value[1]
		means ± SD	
Birth weight, g	2976.0 (459.8)	3006.4 (435.5)	0.065
Birth length, cm	49.2 (2.2)	49.3 (2.1)	0.339
HAZ, 24 months	-2.3 (1.1)	-2.3 (1.1)	0.318
IQ score, 8yrs	51.1 (13.5)	51.5 (12.2)	0.508
Mother's height, cm	150.6 (5.1)	150.7 (5.1)	0.632
Mother's age, y	25.9 (6.0)	26.2 (6.0)	0.273
Mother's education, y	7.8 (3.8)	7.4 (3.7)	<0.001
Father's education, y	8.2 (4.0)	7.5 (3.8)	<0.001
Household income, pesos	297.0 (613.0)	269.1 (440.6)	0.144
Household assets	2.53 (2.0)	2.49 (1.9)	0.554
Household size	5.7 (2.9)	5.7 (2.8)	0.961
Neighbourhood urban development	32.4 (12.1)	29.3 (12.8)	<0.001
		n (%)	
Male	1261 (55)	1799 (52)	0.133
Flush toilet	1261 (7)	1799 (5)	0.031
Piped water supply	1261 (32)	1799 (22)	<0.001

[1] Values of *p* correspond to one-way ANOVA *F*-test statistics (continuous variables) or are based on Pearson's chi-square test (categorical variables).

Table 4B Partitioning of items into Mokken scales with increasing cutoff values of H.

Item content	Item mean (SD)	H= 0.30	H= 0.35	H= 0.40	H= 0.45	H= 0.50
Felt happy	0.72 (0.54)	DNS	DNS	DNS	DNS	DNS
Had headaches	0.52 (0.54)	Scale 2	Scale 2	Scale 2	Scale 2	DNS
Had poor digestion	0.18 (0.39)	Scale 2	Scale 2	Scale 2	Scale 2	DNS
Had difficulty falling asleep	0.40 (0.55)	Scale 1	Scale 4	DNS	DNS	DNS
Felt lonely	0.63 (0.63)	Scale 1	Scale 4	DNS	DNS	DNS
Felt hopeful about the future	0.71 (0.70)	Scale 3	DNS	DNS	DNS	DNS
Felt people were unfriendly	0.49 (0.58)	Scale 1	Scale 1	Scale 3	Scale 3	DNS
Felt worried	0.76 (0.58)	Scale 1	Scale 1	Scale 3	DNS	DNS
Felt couldn’t overcome difficulties	0.69 (0.60)	Scale 1	Scale 1	DNS	DNS	DNS
Felt able to face problems	0.75 (0.61)	Scale 3	Scale 3	Scale 4	DNS	DNS
Felt people disliked them	0.49 (0.56)	Scale 1	Scale 1	Scale 3	Scale 3	DNS
Enjoyed normal daily activities	0.40 (0.55)	Scale 3	Scale 3	Scale 4	DNS	DNS
Thought themselves as worthless	0.31 (0.51)	Scale 1	Scale 1	Scale 1	Scale 1	Scale 1
Felt life wasn’t worth living	0.21 (0.44)	Scale 1	Scale 1	Scale 1	Scale 1	Scale 1
Wished they were dead	0.16 (0.39)	Scale 1	Scale 1	Scale 1	Scale 1	Scale 1
Had the idea of taking own life	0.10 (0.32)	Scale 1	Scale 1	Scale 1	Scale 1	Scale 1

DNS = did not scale.

Table 4C Data quality, targeting, scale assumptions and reliability.

	CLHNS 4-item subscale
Data quality	
Item-level missing data (%)	0.0
Scaling Assumptions	
Item mean score (range)	0.10 - 0.31
Item SD (range)	0.32 - 0.51
Std. item-total correlations (range)	0.51 - 0.61
Targeting	
Mean score (SD) [95% CI]	0.77 (1.29) [0.72, 0.83]
Potential score (range)[a]	0 - 8
Actual score (range)	0 - 8
Floor/ceiling effect (%)[b]	64.0 / 0.10
Reliability	
Cronbach's alpha	0.77
Mean interitem correlation	0.46

[a] Higher scores correspond with poorer mental health.

[b] Floor effects = % of respondents scoring lowest possible score (best mental health); ceiling effects = % of respondents scoring highest possible score (worst mental health).

Figure 4D Distribution of summed scores.

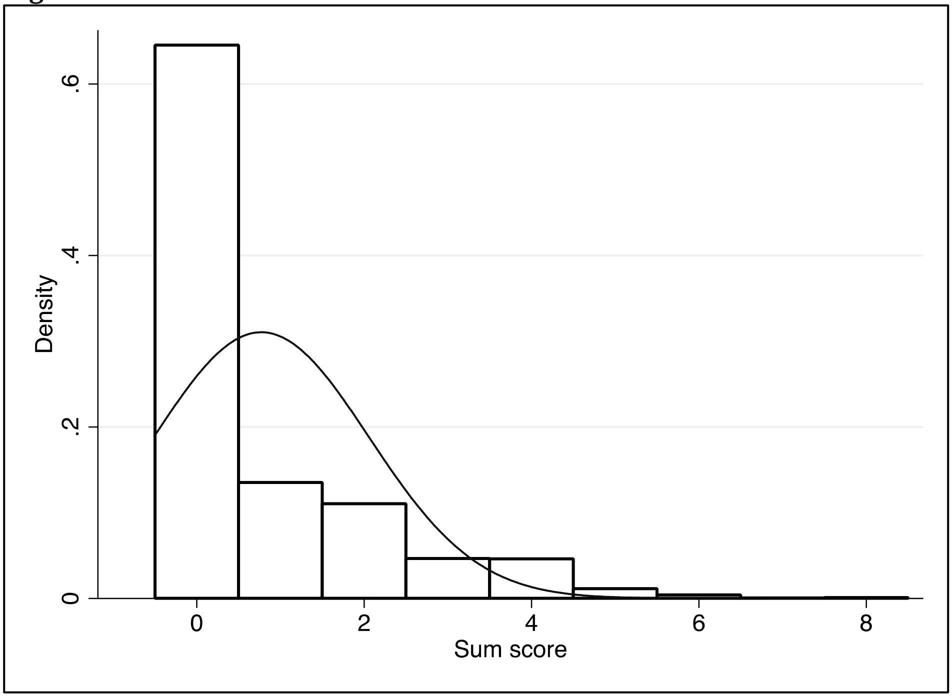

Table 4E Results of regression model testing convergent construct validity.[a]

	Coef.	95 % CI		*p*-value [a]
Excellent general health (reference category)	-	-	-	-
Good	0.243	0.125	0.361	<0.001
Poor	1.009	0.672	1.347	<0.001

Dependent variable=sum score.

[a] Model estimated with cluster robust standard errors and adjusted for adolescent sex and age.

Table 4F Test of measurement invariance by gender (n=1799).

							Chi-square difference test		
	Chi-square	df	*p*-value	CFI	TLI	RMSEA	Chi-square	df	*p*-value
Configural	60.47	4	<0.001	98.1	96.4	0.086	-	-	-
Metric	69.52	7	<0.001	99.0	97.3	0.079	11.78[b]	3	0.008
Metric (partial)[a]	63.45	6	<0.001	99.1	98.2	0.081	4.27[b]	2	0.118
Scalar (partial)[a]	65.84	10	<0.001	99.1	99.0	0.061	7.05[b]	6	0.316
							3.56[c]	4	0.468

Configural = invariant loadings and thresholds.
Metric = variant loadings and invariant thresholds.
Scalar = variant loadings and variant thresholds.
[a] Loading allowed to vary between gender groups for item ‘Thought of themselves as worthless’.
[b] Chi-square difference test compared to configural model.
[c] Chi-square difference test compared to partial metric invariant model.

Table 4G Model fit for structural equation models estimating the association between adolescent mental health, education, and employment.

	Chi-square	df	*p*-value	CFI	TLI	RMSEA	SRMR
NEET	85.56	62	0.025	0.983	0.998	0.026	0.062
Work only	83.57	62	0.035	0.984	0.998	0.025	0.062
Work and study	82.98	62	0.039	0.982	0.998	0.025	0.062
Study only	85.08	62	0.028	0.983	0.998	0.026	0.062
Study/work vs NEET	93.38	62	0.022	0.983	0.998	0.025	0.061