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'Mommy, I miss daddy'. The effect of family structure on children's health in Brazil

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ABSTRACT

This paper studies the relationship between single motherhood and children's height-for-age *z*-scores in Brazil. In order to isolate the causal effect between family structure and children's condition, we estimate an econometric model that uses male preference for firstborn sons and local sex ratios to instrument the probability of a woman becoming a single mother. Our results have a local average treatment effect interpretation (LATE). We find that children being raised by a single mother (whose marital status is affected by a firstborn girl and a low sex ratio) have a height-for-age *z*-score that is lower than that of children of similar characteristics that cohabit with both progenitors. We claim that the increasing trend of single motherhood in Brazil should be of concern in health policy design.

1. Introduction

The health of Brazilian children has constantly and significantly improved in recent decades. Data from the 2010 Census recorded that infant mortality in the first year of life had nearly halved in 10 years. In 2000, for every thousand live births, 29.7 babies died before their first birthday, while the figure in 2010 was 15.6.[2] Regarding anthropometric indices, stunting decreased during the period between 2003 and 2009 from 15.7% to 9.7% among white children and from 20.5% to 11.7% among blacks (Reis, 2012). This is an important achievement for the advancement of human capital investment as it is well known that early life conditions have persistent and profound impacts on an individual's later life in terms of chronic diseases, education and wages (see Barker, 1992; Almond and Mazumder, 2005, 2011; Almond and Currie, 2011; Currie and Vogl, 2012; Lin and Liu, 2014; among others).

The progress in maternal and child health that Brazil is achieving is the result of important changes over the last three decades in several domains. Brazil has experienced major economic growth with an increase in Gross National Income per capita of 47.2% only between 2004 and 2012 (according to World Bank data).[3] The financial gap between rich and poor has become less pronounced and cash transfer schemes (particularly *Bolsa Família*) have increased family income among the poorest groups (Soares et al., 2009). From an educational point of view, the major investment in elementary schooling during the 1990s has increased educational attainment among Brazilian

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[2] See Alves and Belluzzo (2004) for the socio-economic determinants that explain the major decline in infant mortality rates since the 1970s.

[3] In a recent study, de Oliveira and Quintana-Domeque (2014) show a robust correlation between GDP per capita and population stature in Brazil.

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mothers and the country has experienced an important decline in its fertility rate (Madalozzo, 2012). Brazil has also seen great improvements in terms of urbanization (access to water and sanitation) (Victora et al., 2011).

Specifically in terms of health policy, in the late 1980s, a three-tiered healthcare system of private security, social security and charitable institutions was replaced by a universal, tax-funded, national health system. Primary health care became the cornerstone of the system, and the geographical targeting of care led to the setting up of family health teams in the neediest areas of the country. Moreover, maternal-child health policies in Brazil underwent profound change: a vertical health program that started in the 1980s, PAISM, for the promotion of breastfeeding, oral rehydration and immunizations, as well as the implementation of many national and statewide programs have improved children's health and nutrition (Paim et al., 2011; Barros et al., 2010).

Despite all the progress, the prevalence of stunting (1 out of 10 children) is much higher than in well-nourished populations, indicating that there is still room for improvement. Moreover, even if the Brazilian case can be thought of as a successful example, the speed of the improvement has not been uniform and important differences still remain. Although convergence to a national standard level, the prevalence of stunting has been historically much higher in the poorest (North and Northeast) regions than in the wealthier Southeast. Also, although the racial gap in nutritional indicators has clearly decreased, white children still receive better nutrition than blacks (Reis, 2012).

While acknowledging regional and racial disparities, the main focus of this paper is to study the importance of family structure as a source of inequalities in children's health outcomes. Our hypothesis is that children raised by a single mother are more likely to have lower height-for-age *z*-scores after controlling for other socio-economic and demographic characteristics than children growing up with both parents. Several mechanisms may be at play. Single mothers may suffer higher levels of stress because of the difficulties dealing with the role of sole carer and primary breadwinner (Cairney et al., 2003; Mistry et al., 2007; Misra et al., 2010; Osborne et al., 2012). Women who raise their offspring alone have been found to have a higher risk of depression which, in turn, has been related with developmental delays, psychiatric disorders and behavioural problems in children (Schwarz et al., 2012; Surkan et al., 2011, 2012; Lieb et al., 2002; Sohr-Preston and Scaramella, 2006). At the same time, single mothers cannot count on the help and monitoring of a cohabiting partner. Similarly, they may have more difficulties obtaining care for their children because of a smaller extended family. All these mechanisms may explain the poorer health outcomes for children raised by a single mother. As far as we know, this is the first time that family structure has been taken explicitly into account when analysing children's health inequalities in the case of Brazil.

If our hypothesis is confirmed, our results have important implications both for social and health policy. According to data from Brazilian Censuses, the number of children under the age of 14 that did not live with their father multiplied almost fourfold from 1980 to 2010. Indeed, in Brazil, nearly 1 out of 5 children between 0 and 13 years of age were raised by a single mother in 2010. Historically, Brazil has had a high level of informal unions (which account for 36% of all unions in the most recent data). Also, it is common to have a child in each new union to fulfil the idea of having a *proper* family – this being particularly true for men who leave their children from previous relationships with the mothers (Greene, 1992). To this situation, we can add the violent context of high male homicide and incarceration rates (Murray et al., 2013; Reichenheim et al., 2011), which increase the probability of women rearing their children alone.

The isolation of a causal effect between family structure and children's health requires an econometric strategy that accounts for selection into marital status. We estimate a model that uses male preference for firstborn sons (as opposed to firstborn daughters) and local sex ratios to instrument the probability of a woman becoming a single mother. Family economics literature has well established that firstborn children are more likely to cohabit with a father figure if they are boys than if they are girls (see Dahl and Moretti, 2008; Ananat and Michaels, 2008; Ayllón, 2015). Local sex ratios are also used: the probability of a woman being a single mother is likely to be affected by the heterogeneity of her local marriage market (number of potential male partners and female competitors). Moreover, both instruments are complementary as fathers' preferences for cohabiting with their male children are likely to be mediated by the context where they take place. Firstborn gender has been used before as an instrument variable, for example, in Ananat and Michaels (2008) and Bennedsen et al. (2007), and sex ratios in Angrist (2002) and Charles and Luoh (2010). To the best of our knowledge, this is the first time that both instruments have been used in the context of child health analysis.

Our results indicate that children raised by a single mother in Brazil have a height-for-age *z*-score that is lower than that of children of similar characteristics that cohabit with both parents. Because of the use of instrumental variables, our results have a local average treatment effect (LATE) interpretation: they estimate the effect of single motherhood on children of mothers whose marital status probability is affected by the sex of their firstborn and the (strata- and age-specific) local sex ratios. Because of the use of two instrumental variables, our estimates are a weighted average of causal effects relative to each instrument compliant subpopulation. These findings are robust to different specifications that control for socio-economic and demographic characteristics and are likely to be conservative as we do not consider (biologically immature) teenage mothers. Moreover, the number of children in the household below the age of 10 is associated with lower height-for-age *z*-scores. In turn, mother's height, body mass index and employment status are positively related with children's height. Our results confirm that family structure should not be overlooked when designing health policy in Brazil: single motherhood may become a challenge for future maternal-child health programs (much more than racial or regional disparities).

After this introduction, the following section describes the acute consequences of malnutrition in early life.

Section 3 discusses the relationship between children's malnutrition and family structure according to previous literature. Section 4 describes the dataset used and presents some descriptives while Section 5 details the econometric strategy. Section 6 contains our main results and Section 7 concludes.

2. Malnutrition in early life

Malnutrition is typically caused by a combination of inadequate food intake and infections that impair the body's ability to absorb or assimilate food. It encompasses stunting, wasting, and deficiencies of essential vitamins and minerals, with obesity or over-consumption of specific nutrients as another form. From an economic perspective, infant nutritional status can be seen as the output of a health production function, where nutrient intake is a very important input. However, in addition to that and to the particular genetic variation, there are other relevant variables: appropriate care, household conditions (including food insecurity), parental education, access to quality health care and living environment.

There are two commonly used anthropometric indicators for children's nutritional status: *wasting* and *stunting*, which distinguish between short-term and long-term physiological processes (WHO, 1986). The first indicates a low weight-for-height and reflects current nutritional problems (diarrhoea, insufficient dietary intake and other childhood diseases). The weight loss associated with *wasting* can be restored quickly under favourable conditions and is generally seen as a short-lived problem.[4] The *stunting* index is measured by height-for-age and shows children's cumulative linear growth or a chronic restriction on their potential growth. It thus reflects the child's present and past inadequate nutrition and/or frequent illnesses, expressing a long-run health condition and is not usually reversible. The WHO recommends it as a reliable measure of overall social deprivation (WHO, 1986).[5] These indicators are measured in terms of a standard deviation *z*-score which accounts for the difference between the value for an individual and the median value of the reference population (same age and sex) divided by the same reference population's standard deviation.

Extensive epidemiological literature has focused on the early childhood environment, nutrition in particular, and its relationship with health outcomes in adulthood.[6] It emphasizes that linear growth failure is largely confined to the intrauterine period and the first years of life, and is caused by inadequate diet and frequent infections (Black et al., 2008; Grantham-McGregor et al., 2007). Malnutrition in the early years of life may also cause later deficiencies in cognitive development (Mendez and Adair, 1999; Crookston et al., 2011; Behrman et al., 2014). There is some disagreement as to whether the effects of malnutrition on cognitive development are more acute when the child is a baby (0–6 months), as Dobbing (1976) argues, or a toddler (6–36 months), but there seems to be a consensus that the first 2 years are very important (Waber et al., 1981; Pollitt et al., 1995; Martorell, 1995; Glewwe and King, 2001).

Early growth failure will lead to reduced adult stature unless there is compensatory growth in childhood (so-called *catch-up growth*), which is partly dependent on the extent of maturational delay that lengthens the growth period. Because maturational delays in low-income and middle-income countries are usually shorter than 2 years, only a small part of growth failure can be compensated. Individuals who remain in the setting where they developed childhood malnutrition tend to become short adults (Martorell et al., 1994, 2005).[7]

The epidemiological literature reveals that the nutritional status of a woman before and during pregnancy is important for a healthy pregnancy outcome (Kramer, 1987; Kramer and Victora, 2001). Short maternal stature is a risk factor for caesarean delivery and low maternal body mass index is associated with intrauterine growth restriction (WHO, 1995; Fishman et al., 2004; Ronsmans et al., 2006). In the short term, the results of maternal/child malnutrition are mortality, morbidity and disability. In the long term, the consequences are adult size, intellectual ability, economic productivity, reproductive performance, and metabolic and cardiovascular diseases (Black et al., 2008).

Regarding academic achievement, particularly in developing countries, the economic literature has found that childhood malnutrition is a significant determinant of delayed enrolment and low test scores (Glewwe and Jacoby, 1995; Glewwe et al., 2001; Alderman et al., 2001; Wisniewski, 2010; Monk and Kingdon, 2010). Malnutrition in childhood can also undermine labour outcomes in adulthood because of the probability of low school attainment being greater for unhealthy children than for those who were well-nourished (Deaton, 2008).

In the Brazilian case, Machado (2008) showed that for children between 7 and 14 years old and from the Northeast and Southeast regions, low height-for-age status increases the chances of late entry into school. The study by Gomes-Neto et al. (1997) uses data from a major education intervention program, EDURURAL (conducted in three Northeastern states in the 1980s), which measured students' health. This program was designed to reduce low achievement and high drop-out rates in these rural areas. The authors studied the complementarities between

[4] Glewwe et al. (2001) argue that weight-for-height can be useful for evaluating intervention programs because it quickly responds to changes in nutritional status.

[5] Weight-for-age is a third index which allows the measurement of *undernutrition*, but it must be interpreted cautiously because of its inability to distinguish between *stunting* and *wasting*, and consequently between acute and chronic malnutrition. Because it is a combination of the other two indices, it does not distinguish between small but well-fed children and tall but thin ones. It also changes rapidly.

[6] This is known as the *fetal and infant origins hypothesis* and was developed by Barker (1992) who has argued that inadequate nutrition in uterus "programs" the fetus to have metabolic characteristics that can lead to future diseases. See Sotomayor (2013) for a recent example of evidence that confirms the fetal origins hypothesis and Doyle et al. (2009) for arguments on the importance of health investments during the prenatal period.

[7] Improvements in living conditions, such as those brought on through adoption, can trigger catch-up growth but do so more effectively in very young children.

health and educational attainment/cognitive proficiency of children and found that changes in nutritional status play an important role in explaining cognitive differences among children (but do not have an impact on the final degree of schooling).

More recently, Gigante et al. (2006) used a large cohort of data from all children born in the city of Pelotas (on the country's Southern region) in 1982, with follow-ups in adulthood (21–23 years old) to analyse various later health and economic outcomes. This study confirms that height-for-age predicts school or cognitive test performance in later life, and stunting between 12 and 36 months of age predicted poorer cognitive performance and/or lower school grades attained in middle childhood. Similarly, Curi and Menezes-Filho (2009) calculated height-education and height-wage elasticities for the Brazilian adult population (21–65 years old). Their results show that height has a positive and significant effect on the completion of school cycles (higher for males than females) and on the income of individuals.

In short, the previous literature has reached a strong consensus on the acute consequences of malnutrition during the first years of life across the whole life cycle.

3. Family structure and infant health: a review

Studies that have analysed the impact of marriage on health systematically find a positive association between the two. In general, the literature indicates that married people live longer, have fewer problems related to alcohol, engage in fewer risky behaviours and have better mental health (see Lillard and Panis, 1996; Espinosa and Evans, 2008; Rendall et al., 2011; Curran et al., 1998; Duncan et al., 2006; Umberson, 1992; among others).

The literature has established that the marriage–health relationship can be explained by selection, causation or both.[8] According to the selection framework, healthier people are more prone to marriage because they are more desirable in the marriage market (e.g. physical attraction, mental well-being, higher income, greater self-sufficiency, likely longevity, etc.). In this case, it is not possible to attest that marriage itself provides a better health condition. So, the advantages observed in the health status of married people (compared to those without a partner) only reflect selection (Goldman, 1993).

On the other hand, the causation alternative advocates that marriage does cause benefits to the health of the people involved (adults and children). This is known as the *marriage protection hypothesis*, which sees marriage as a mechanism that protects people's health through: (i) social channels as integration, (ii) accomplishment of some social roles, (iii) support, (iv) financial resources and economies of scale, and (v) joint and full monitoring between spouses. The most common view is that marriage leads to emotional support: it is seen as an institution that gives a sense of meaning (or purpose) of responsibility/commitment to others, as well as closeness.[9] It also works as a sign of fulfilment of an adult social role, which together with parenthood reduces the chance of health-harmful and risk-taking attitudes (Ross et al., 1990; Waite, 1995; Hibbard and Pope, 1993; Rendall et al., 2011). Other factors include family income/wealth and economies of scale that may increase access to medical care and the purchase of better nourishment and housing (Becker, 1991; Hahn, 1993).

Despite the difficulties disentangling the influence of selection and protection, studies that have dealt with the issue are consistent regarding a marriage premium. They generally find that the association between marriage and health is a combination of both (Lillard and Panis, 1996; Murray, 2000; Rendall et al., 2011), so marriage affects and is affected by its members' health status. Consequently, the positive results of marriage on health are also extended to the children living in two-parent families.

The monitoring mechanism is especially relevant when there are children in the household because they are treated as a collective good by both parents (Weiss and Willis, 1985).[10] In this case, any opportunistic behaviour by one of the parents is avoided by the monitoring between them, which may encourage both to invest in the common good. Theoretical models provide intuition for the empirical studies that have evaluated the effect of marriage on children's health through specific inputs, for example, quality/time of prenatal care, nutrition and/or abstention from tobacco, alcohol or illegal substances during pregnancy (Currie and Gruber, 1996; Joyce, 1999; Evans and Lien, 2005; Abrevaya and Dahl, 2008; Almond and Mazumder, 2011; Evans and Ringel, 1999). The findings indicate that the quality of care during the prenatal period, including the significance of the mother's nutrition, is very important for ensuring that the infant is born in healthy conditions, thus affecting the child's subsequent ability to accumulate human capital.

Moreover, despite the increased diversity of family arrangements, women are still disproportionately the primary caregivers for children in Brazil. As a result, women are more likely than men to be faced with the dual role when unmarried: that of sole caregiver and primary breadwinner. This situation affects access to resources and psychological well-being and must have an impact on the quality of care provided to children (Osborne et al., 2012). A mother's stress (or depression) has a negative impact on her child's health, not just in key periods such as pregnancy and breastfeeding but also during early childhood (Chung et al., 2004).[11]

Regarding the father's involvement, some research suggests that it plays a role in linking family structure to children' health outcomes. Married and cohabiting fathers are more likely to be involved with their children by virtue

[8] See Wood et al. (2007) and Ribar (2004) for extensive reviews of studies that relate marriage to adult's health and mortality as well as the impact of marriage on children's development outcomes (intergenerational health effects of marriage).

[9] This approach started with the seminal work by Durkheim (1951) on suicide.

[10] In empirical studies, children's quality may be represented either by their health condition or by some educational outcome.

[11] See Amato (2005), Fomby and Cherlin (2007) and Wu (1996) for further studies that have documented the adverse associations of family instability and child outcomes.

of proximity, while men who are not married to or living with their child's mother find it harder to be involved, and hence participate less (Furstenberg and Cherlin, 1991). Besides, when a union dissolves, the father's involvement drastically declines (Tach et al., 2010), especially if the relationship ends when children are very young. In general, the literature suggests that fathers may have an impact on infant health by encouraging (or not) women to seek prenatal care and refrain from unhealthy behaviours, such as smoking and consuming alcohol while pregnant (Teitler, 2001). Misra et al. (2010) have suggested that the father's involvement may improve birth and infant health outcomes by reducing maternal stress.

The relevance of each aforesaid mechanism is difficult to evaluate since they depend on the interaction between family members.[12] Therefore, no empirical study of the impact of family structure on children's development has succeeded in isolating each mechanism.

A significant number of empirical studies have examined the potential effects of family structure on children's development, and these generally confirm the disadvantage of children from households led solely by women.[13] Most of the literature focuses on cognitive and educational factors, as well as factors that are relevant during youth (early pregnancy, smoking, anxiety and economic inactivity). Studies that specifically analyse the impact of family structure on children's health indicate that unmarried mothers invest less in prenatal care and, consequently, their children have poorer indicators at birth. Thus, their children are at a disadvantage to others in terms of prematurity, birth weight and mortality (Bennett, 1992; Bennett et al., 1994; Bird et al., 2000; Peacock et al., 1995).[14]

However, few of these studies have addressed a causal relationship. The challenge for disentangling the effect of family structure on children's outcomes resides in the difficulty finding exogenous variations in the mothers' marital status. Ribar (2004) argues that in most studies the identification strategies used to address the issue of selectivity into marriage are unconvincing.[15] One of the most common methods used is that of instrumental variables, and it is usually found that exclusion restrictions reduce the association between family structure and well-being (which is consistent with selectivity) but do not eliminate the association (which is consistent with causality). Such is the case with Buckles and Price (2013), who measure the effect of marriage on four child indicators: low birth weight (<2500 g), prematurity (<37 weeks), mortality in the first year of life and the Apgar test. They use a sample of siblings to show that accounting for selection with regard to observable and time-invariant characteristics significantly reduces estimates of the marriage premium, but the positive causal effect remains.

This paper uses the instrumental variable method to analyse the causality between family structure and children's malnutrition (measured by height-for-age *z*-scores) by proposing a new exclusion restriction that has not been previously used for a similar research question.

4. Data, definitions and descriptives

Data is from the Brazilian Household Budget Survey (*Pesquisa de Orçamentos Familiares*, POF) relative to the 2008–2009 period. The survey contains 190,159 individuals from 55,970 households and is representative of the population living in private households in Brazil.[16] Our working sample is restricted to all (single-birth) children aged 6–60 months that are sons or daughters of the head of household (or spouse), whose mothers are between 20 and 44 years of age (both inclusive) and did not have their child as teenagers.[17] This last restriction is imposed because we do not want to confound the effect of teenage biological immaturity (which may have consequences on children's health outcomes) with that related to single motherhood.[18] Moreover, in the POF, family ties are defined in relation to the head of household: (i) spouse/partner, (ii) child (son/daughter), (iii) another relative, or (iv) non-relative. Such a survey structure allows us to link 88.4% of the infants in the sample to their mothers. Children classed as 'another relative' (11.4%) or 'non-relative' (0.2%) cannot be linked to their mothers even if they are in the household.[19] As explained above, the 60 month threshold seems most appropriate when seeking to capture early-life health. In total, 9255 records have been used.

The POF data contains anthropometric measures for all individuals (weight, height and length) as well as information on such characteristics as age, race, education, position in the family, availability of a private health insurance plan for each person in the family, household income, governmental cash transfers, living conditions (sanitation, street paving, public lighting, health facility in the community, quality of public services, violent area, etc.) and geographic variables.

Despite its importance, the raw anthropometric data collected by the POF is of limited value as an indicator of

[12] Note, for example, that single mothers may also be exposed to some degree of monitoring by other family members (parents, siblings, etc.).

[13] For a literature review see Ribar (2004) and Sigle-Rushton and McLanahan (2004).

[14] In an extreme case, it has also been found that parental orphans are shorter than children with a living father in Tanzania (see Alderman et al., 2006).

[15] It seems that the literature that analyses the impact of family structure on child outcomes (such as education or early adult results such as teenage pregnancy, tobacco or drug consumption) address causality more often. Some examples are Ermisch and Francesconi (2001), Finlay and Neumark (2010) and Craigie (2008).

[16] More detailed information about the Brazilian Household Budget Survey can be found in Reis (2012), de Oliveira and Quintana-Domeque (2014), and on the website of the Brazilian Institute of Geography and Statistics: http://www.ibge.gov.br/english/xml/pof_2008_2009.shtm

[17] Children aged 0–5 months are not included in the sample because length has been imputed in nearly 20% of the cases. This sample restriction is often used in the literature – see, for example, Reis (2012) for the same dataset, or Duflo (2000, 2003).

[18] Alves and Belluzzo (2004) show that children of mothers younger than 16 have a lower height-for-age than the rest of children in Brazil.

[19] Strictly speaking, this means that our sample is not representative of all children aged 6–60 months in Brazil. However, it is also true that about only 5% of children under the age of 6 do not live with their mothers (according to the 2010 Census data) so, we are not missing as many mother–child links as it may seem. Also note that the restriction of mother's age is also related to survey structure limitations. By only using the sample of women between 20 and 44 we disregard the possibility of considering grandmothers that are rearing their teenage daughter's child.

malnutrition in its own right, partly because weight and length/height depend on both age and gender. So, in order to assess the adequacy of a child's growth it is necessary to compare these indicators with their distribution in a "healthy" reference group and identify "extreme" or "abnormal" departures from the reference values.[20] To do this, the *Anthro* software created by the World Health Organization (WHO) was used, which requires input of each infant's sex, age in months, weight and length/height data.[21] From this data, z-scores for height-for-age have been calculated. Table 1 shows that the average height-for-age z-score is nearly zero with a standard deviation of 1.67 and a minimum and maximum value of −5.99 and 5.92, respectively. Indeed, descriptives indicate that 11.1% of children in the sample suffer chronic malnutrition (*stunting*). On the other hand, 9.9% are obese.

Regarding family structure, a child's mother is considered to be a single mother if none of the registered members of the household is her partner. This means that the single mother may be divorced, separated, never married or widowed. As a matter of fact, the POF data does not make it possible to distinguish the mother's marital status.[22] In total, 10.0% of the children in the sample cohabit with a single mother.

As in many other parts of the world, single motherhood has become an increasing phenomenon in Brazil as shown in Fig. 1, which uses data from the Censuses from 1970 to 2010 to create a sample with the same characteristics as those explained above.[23] A thorough analysis of this trend goes beyond the scope of this paper but two explanations stand out in the case of Brazil. Firstly, the analysis by Greene and Rao (1995) indicates a squeeze in the Brazilian marriage market (a shortage of men) driven by a decrease in mortality rates. As, on average, men marry younger women, the simple reduction of mortality in younger cohorts increases the supply of women in relation to men. The result of this compression would be an increase in cohabitation, with men going through various unions during their life cycle.[24] Secondly, it is important to take into account how the Brazilian context of violence also increases the probability of a woman having to raise her children alone. The male homicides rate (per 100,000 inhabitants) increased from 21.3 to 49.0 between 1980 and 2009, but for the young male population (15–29) it grew 189% during the same period (from 36.1 to 104.4) according to data from the Brazilian Ministry of Health's Mortality Information System database.[25] The same is true for incarceration rates.

The other rows in Table 1 show summary statistics for the groups of variables that will be used as controls when modelling the effect of single motherhood on children's height-for-age z-scores: other demographic factors, children's and mothers' characteristics, household economic status, and regional/urbanization descriptives.

As for other demographic characteristics, summary statistics indicate that there are between 1 and 4 children in each household under 60 months of age (average 1.31) and that 41% of the children in the sample have at least a brother or sister between the ages of 5 and 10 and 24% between the ages of 11 and 15.

Regarding children's characteristics, 49.2% of the sample are girls; 51.4% are black, coloured or indigenous and, on average, they are 34.7 months old.[26] As for mothers' characteristics, mean age is 29.7 and mean age at birth is 27.3. Mothers' average height is 158 cm and their body mass index is 24.8. Moreover, 4% of the mothers have never been to school and nearly 18% of them did not complete primary education.

At least two of the mothers' characteristics are particularly important for the children's health: height and age at birth. Regarding mother's height, extensive literature highlights its correlation with some reproductive outcomes. Taller mothers tend to have easier births (Liljestrand et al., 1985), heavier babies at birth (Kirchengast et al., 1998), fewer stillbirths (Pollet and Nettle, 2008) and higher survival rates among their children. Thus, the mother's height carries direct information on genetic inheritance and indirect information about her own past situations of malnutrition and poverty.

The information on the mother's age at birth is a measure of her biological maturity, which is another relevant factor for the child's health outcomes. Finlay et al. (2011) analyse the effect of the mother's age at first birth on infant mortality, its anthropometric failure (stunting and underweight), diarrhoea and anaemia. Using data from Demographic and Health Surveys (conducted from 1990 to 2008) for 55 low and middle-income countries, these authors found that the risk for the child's health is lowest for women who gave birth for the first time between the ages of 27 and 29. The authors conclude that firstborn children not only of adolescent mothers but also of those in their early twenties are the most vulnerable to mortality and poor health outcomes.

The variables referring to household economic conditions are poverty status and being a *Bolsa Família*

[20] A total of 68 observations were considered extreme cases and were deleted from the final analysis.

[21] *Anthro* can be consulted at the web page http://www.who.int/childgrowth/software/en. This software also provides the WHO Child Growth Standard measures for weight-for-age (WAZ), weight-for-height (WHZ) and body mass index for age (BMI). Since 2006, the WHO has used new curves to assess the growth and development of children around the world. This new standard was obtained by combining longitudinal follow-ups from birth to 24 months and a cross-sectional survey of children aged 18–71 months. The sample has information from 8440 healthy breastfed infants and young children from a wide variety of ethnic backgrounds and cultural settings (Brazil, Ghana, India, Norway, Oman and the USA).

[22] Throughout the paper, we use single mother to refer to women that are raising their children without a partner in the household, regardless of marital status.

[23] The graph includes infants aged 0–5 months because Census data does not allow identifying age in months.

[24] A qualitative analysis by Greene (1992) showed that in Brazil it is common to have a child in each new union to fulfil the idea of having a proper family. A husband and wife may wish to have a child together regardless of the number of children either of them had before. This is particularly true for men who leave their children from previous relationships with the mothers.

[25] This implies a loss of over 45,000 male lives each year at the end of the decade.

[26] Coloured refers to children than come from black and white parents.

Table 1
Summary statistics.

Variable	Mean	Std. Dev.	Min.	Max.
Height-for-age *z*-score (*haz*)	−0.037	1.666	−5.99	5.92
% in malnutrition ($haz < -2$)	0.111	0.313	0	1
% being obese ($haz > +2$)	0.099	0.298	0	1
Family structure				
Single mother	0.100	0.300	0	1
No. of children up to 60 months	1.311	0.527	1	4
No. of children 5–10 years old	0.531	0.731	0	5
No. of children 11–15 years old	0.321	0.641	0	5
Child's characteristics				
Girl	0.492	0.499	0	1
Black, coloured or indigenous	0.514	0.499	0	1
Age (in months)	34.73	15.81	6	60
Mother's characteristics				
Age	29.73	5.980	20	44
Age at birth	27.32	6.007	17	44
Height (in cm)	158.91	7.000	130	188
Body mass index (BMI)	24.85	4.656	13.86	54.67
Never at school	0.039	0.196	0	1
Primary school incomplete	0.178	0.382	0	1
Primary school completed	0.294	0.455	0	1
Secondary school completed	0.363	0.481	0	1
University degree	0.118	0.323	0	1
Household economic status				
In poverty	0.165	0.371	0	1
Receiving *Bolsa Familia*	0.236	0.425	0	1
Log(income per capita)	5.758	1.017	1.244	9.572
Urbanization/regional characteristics				
Piped water	0.877	0.328	0	1
Paved street	0.593	0.491	0	1
North	0.099	0.299	0	1
Northeast	0.302	0.459	0	1
Southeast	0.379	0.485	0	1
South	0.142	0.349	0	1
Central West	0.078	0.268	0	1
Urban	0.801	0.399	0	1
Instrumental variables				
Firstborn girl	0.489	0.499	0	1
Low sex ratio	0.323	0.467	0	1
N			9255	

Source: *Pesquisa de Orçamentos Familiares*, 2008–2009. Author's computation. Weighted results. *Note*: Children age 6–60 months.

beneficiary. Poverty is defined using the same concept applied by the government cash-transfer program, *Bolsa Famîlia*, which has been operating since 2003. This study uses the values that were established from June 2008 onwards: families with an income per capita up to 30 USD (monthly) are considered ''extremely poor'' and families with an income per capita above 30 USD but below 60 USD are considered ''poor''. Here the sum of these two groups defines a poor family. In the sample, 16.5% of the children live in a poor family and 23.6% receive the cash transfer.

It is important to note that the *Bolsa Famîlia* is a secondary indicator of household poverty but also works as a health input control. Its inclusion was necessary for two reasons. First, because of its conditionality regarding the health of children up to 7 years old, which requires beneficiary families to monitor the vaccination card and make frequent visits to the doctor to follow-up the children's growth and development. This condition exposes all children whose families are beneficiaries (partially) to the same health input: access to medical services (prescriptions and monitoring).[27] Second, the *Bolsa Famîlia* and poverty status are both included (despite being based on the same criteria) because not every poor family receives this cash transfer (and vice versa). Household income per capita in quartiles has also been included in our analysis.

In order to control for the level of urbanisation, we also consider the variables ''piped water'' and ''paved street''. The first is used as an indicator of sanitation conditions

[27] Women aged 14–44 years should also follow up and, if pregnant or nursing (lactating), should carry out pre-natal tests and monitor their health and the baby's health.

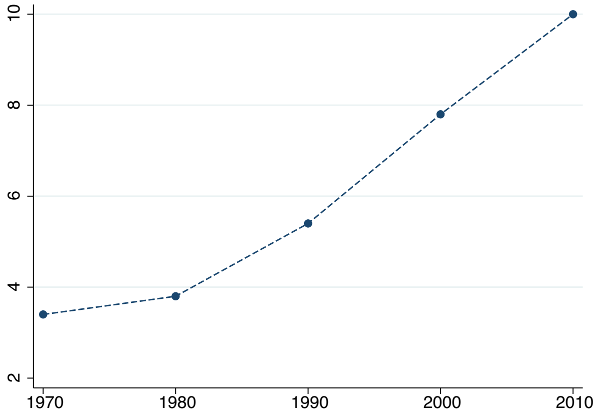

Fig. 1. Evolution of the percentage of children under 60 months of age cohabiting with their mother (20–44 years) with an absent father in Brazil, 1970–2010. Source: Own calculations on Brazilian Census Data, 1970–2010. Weighted results.

and the second as a *proxy* for the public services that come with urbanization. In the sample, 87.7% of the children live in an area with piped water and 59.3% with a paved street. As for regional characteristics, the Brazilian territory is divided into five areas (North, Northeast, Southeast, South and Central West). Most children live in the Northeast region of the country, followed by the Southeast and South. Also, 80.1% of children live in an urban area. Finally, and as for the instrumental variables (see the econometric section below), 49% of the children live in a household with a firstborn girl and 32% in a low sex ratio area.

5. Econometric strategy

The simplest modelling strategy to capture the relationship between family structure and child malnutrition would be an OLS regression accounting for changes in the outcome of interest according to a set of observable characteristics. Formally,

$$haz_i = \beta_0 + \beta_1 singlemother_i + \beta_2' X_i + \epsilon_i \qquad (1)$$

being haz_i the height-for-age z-score. As already explained above, among nutritionists, height-for-age is a widely recognised index to identify malnutrition. Among economists, it is the most useful indicator of a child's health and welfare in developing countries and a good predictor of many later-life outcomes. X_i is the vector of explanatory variables, and finally, ϵ_i is the error term.

An OLS regression is always a powerful description device – it provides the best linear approximation to the conditional expectation function. However, the OLS estimate of β_1 will capture a causal effect of the mother's marital status on child malnutrition as long as ϵ_i has mean zero conditional on the vector of explanatory variables X_i. This conditional mean independence assumption will be violated if factors that are related to both child malnutrition and mother's marital status are omitted from the regression equation (for example, if they are unobservable). If that is the case, the OLS estimator will be biased.

To account for this possibility, we use an instrumental variables (IV) strategy and apply a two-stage least squares (2SLS). The first-stage equation is defined by:

$$singlemother_i = \alpha_0 + \alpha_1' X_i + \alpha_2' IV_i + u_i \qquad (2)$$

where IV_i is the vector of instrumental variables.

In the second stage, we estimate:

$$haz_i = \beta_0 + \beta_1^{IV} \widehat{singlemother}_i + \beta_2' X_i + \nu_i \qquad (3)$$

Ideally, IV allows us to use the variation of mother's marital status that is uncorrelated with any omitted variable that affects both child's height-for-age and mother's marital status to identify a causal effect of the mother's marital status on child's height-for-age.

In this paper, we use two instrumental variables: "firstborn girl" and the local sex ratios. As for the first one, family economics literature has shown the existence of a relationship between the probability of divorce or single motherhood and the sex of the firstborn in a family. Dahl and Moretti (2008), Ananat and Michaels (2008), Bedard and Deschênes (2005) and Ayllón (2015) have shown for the United States that firstborns are less likely to cohabit with a father figure if they are girls. In other words, firstborn girls (as opposed to firstborn boys) increase the probability of divorce or separation among natural parents and decrease the probability of their mother finding a new partner after dissolution of a partnership.[28]

Several hypotheses have been used to explain the association between partnership stability and firstborn males as summarised in Dahl and Moretti (2008). First, it could be that men are (simply) *gender biased* and have a greater preference for raising boys than girls, which would explain why they are more likely to cohabit with their male offspring (Dahl and Moretti, 2008). Second, the *role model hypothesis* argues that fathers may believe that their sons are more in need of a male role model (than daughters) and that is why they are more likely to stay (Biblarz and Raftery, 1999). Third, the *technological reasons* or *differential costs hypotheses* assume that men could be more efficient at raising boys than girls so, since it is easier for them, they are more likely to cohabit with their male children (Lundberg and Rose, 2003). This is similar to the idea that girls could be more costly in terms of money and time (Ben-Porath and Welch, 1976). Finally, the *compensatory behaviour hypothesis* assumes that boys have more health problems and are harder to look after than girls, so fathers are more likely to stay in the household to compensate (Angrist and Lavy, 1996).

We hypothesize that, as in the United States, "firstborn girls" (and all their brothers and sisters) are more likely to live without their father in Brazil than children in families

[28] Note that the same idea can be applied to widows. The death of a partner is a natural event that does not affect all women with the same likelihood. But we claim that once a woman becomes a widow, her chances of having a new partner are conditional on the sex of her firstborn child.

where a boy is born first. Indeed, a simple regression probit with the dataset at hand confirmed that this is the case: firstborn girls are more likely than boys to live with a single mother (with a significance level of 5%). We have confirmed these results with data from the five Brazilian Censuses from 1970 to 2010 extracted from the International Integrated Public Use Microdata Series (IPUMS-International) (see Minnesota Population Center, 2013). The significance level for the gender effect was even higher, which is readily explained by the fact that the Census contains a huge number of observations compared to the POF and therefore the gender effect can be more easily captured.[29]

To reinforce our instrumental variables strategy, we have also included information relative to the local sex ratios for the first time when women in our sample were in the marriage market. It is reasonable to think that the probability of a woman being a single mother is likely to be affected by the heterogeneity of her local marriage market (Angrist, 2002). The variation of sex ratios across markets at a point in time and within markets over time allows us to identify the effect of interest.[30] Moreover, local sex ratios help us to put into context male bias towards cohabitation with their male children given that preferences are likely to be mediated by the context where they take place. Thus, males living in a low sex ratio area where the number of women (potential new partners) is greater than the number of men (potential competitors) may have more room to express their bias for cohabiting with boys and therefore are more likely to create another family (in the event of having had a girl with their previous partner) than otherwise. The mechanisms by which each instrument affects single motherhood are complementary of each other.[31]

The use of IV estimation implies that several assumptions need to be satisfied in order to separate a causal effect from confounding factors (Angrist and Pischke, 2009). First of all, the instruments must have a clear effect on the probability of single motherhood. Indeed, this is confirmed in the next section by using an *F*-statistic on the relevance of the instruments that is above the "rule-of-thumb" of 10 proposed by Stock et al. (2002).

Second, the instruments should be as good as randomly assigned. In the case of firstborn gender, we assume that child sex is random (women cannot induce the sex of their babies) and that there is no sex-selective abortion in Brazil as confirmed by Chiavegatto Filho and Kawachi (2013). In the case of local sex ratios, the POF data set unfortunately does not contain any kind of information on possible migration status of the mothers in our sample. However, descriptive statistics from the 2010 Census indicated that 94.6% of single mothers between the ages of 20 and 44 (either never married, divorced, separated or widowed) with children below the age of 5 remained in the same major and minor administrative unit for the last 5 years.[32] Only 2.8% would have changed minor unit (municipality) while remaining in the same major unit (state), 2.4% would have changed state and 0.2% would have changed country. The corresponding percentages among married mothers are 91.7%, 4.1%, 3.9% and 0.3%. In short, descriptive statistics for mobility do not indicate a major concern: women that are rearing a young child in Brazil do not seem to be very mobile – especially if doing so alone.[33]

Third, the exclusion restriction attests that the instruments must operate through a single known causal channel. Phrased differently, the instruments must be independent of the error term in the second-stage equation. There are different reasons why this assumption could be violated. In the case of firstborn gender, if there is some sex-biased preference that also affects the child's height-for-age, then the validity of the exclusion restriction would be threatened. To test this possibility, we used a sub-sample of the children up to 47 months (73.3% of the total sample) for whom we have information about breastfeeding. We believe that if there is any gender preference that may have an impact on the determinants of early nutrition, it is very well represented by breastfeeding practices. A probit regression for breastfeeding on the sex of the child and birth order showed that although the younger children in the household are more likely to be breastfed (which seems reasonable), this has no relation with the sex of the child, which supports the validity of the exclusion restriction for this instrument.[34]

In the case of local sex ratios, the exclusion restriction implies that sex ratios must not be correlated with child outcomes other than through family structure.

[29] Mothers's previously imposed age limitations help us to ensure that the child observed to be the eldest really is the mother's firstborn. In Brazil, mean age at leaving home is above 20, so it is unlikely for the firstborn in a family to have already left the parental home while the mother has a child that is younger than 6. Median spacing between children below the age of 20 of the same mother is 3 years (with an average of 3.4) according to Census data from 2010.

[30] In order to calculate local sex ratios, we have matched information from the Brazilian population data estimates provided by the Brazilian Institute of Geography and Statistics (IBGE) with the information on geographic strata available in POF. Each sex ratio refers to the age range between 15 and 29, when women in our sample were in the marriage market for the first time and had their first child – average age at marriage in Brazil is 22 according to the 2010 Census. For women younger than 29, the sex ratio has been calculated from when they were 15 up to their age in the survey year (2009). Finally, we have averaged sex ratios across mothers' ages in each geographic State-strata. The Brazilian marriage market is characterised by a high level of interracial marriage so sex ratios are strata- and age-specific (but not race-specific) (Ribeiro and Silva, 2009). Moreover, and in order to reduce the number of functional form assumptions, we have dichotomised the information and we refer to a "low sex ratio" when the number of women exceeds the number of men (males/females < 1).

[31] Moreover, the inclusion of two instruments gives us the possibility to work with an overidentified equation and thus we can take advantage of running a series of tests that confirm the validity of our strategy. As explained in Baum (2006), "although overidentification might sound like a nuisance to be avoided, it is actually preferable to working with an exactly identified equation. Overidentifying restrictions produce more efficient estimates in large samples" (Baum, 2006: 191).

[32] We use the variable `mgrate5` within IPUMS.

[33] Brazil had a large flow of migration from rural areas towards cities from the 1950s to the late 1990s, but since then it has stabilised (Camarano and Beltrão, 2000).

[34] These results are available from the authors upon request. Our results differ from evidence found in other countries with a well-documented social preference for boys. See, for example, Jayachandran and Kuziemko (2011) for the case of India.

Noticeably, we use the sex ratio that refers to the period of time when the mother was in the marriage market for the first time. This strategy allows us to avoid problems with reverse causality between the sex ratios and measures of economic and social conditions. Nonetheless, we include variables related to ''locality'' that strengthen control: a distinction between urban and rural areas, and whether the family lives in an area with piped water and paved streets.

One could also believe that sex ratios might affect children's outcomes through the mother's labour market and fertility decisions (Finlay and Neumark, 2010). For example, women in low sex ratio areas could have poorer labour market prospects because they have to compete with more females for certain jobs. To avoid correlation between the instrument and the error term through this channel, we include controls for mother's labour market status and educational level. Similarly, if women with poorer prospects in the marriage market (due to an excess of female competitors) have fewer children and the number of children in the household affects height-for-age of all siblings, failure to control for such a characteristic could lead to correlation. Thus, we include controls for number of children by age group.

Finally, the effect of the instruments on single motherhood must be monotone, that is, ''while the instrument may have no effect on some people, all those affected are affected in the same way'' (Angrist and Pischke, 2009: 154).[35] This means that some men may not care about the sex of their firstborn, but for those for whom such a characteristic is important and, should they find a favourable marriage market (low sex ratio), their probability of searching for a new partner is always higher. Using the literature's terminology, there are no *defiers*. Importantly, a consequence of monotonicity is that our estimates should be interpreted as a local average treatment effect (LATE). This means that we estimate an effect on a subset of children (called *compliers*) for whom the probability of living with a single mother is affected by the sex of the eldest child in the household and the local sex ratio. So, our results cannot be read as the estimated effect of single motherhood on the height-for-age *z*-score for all children, but is restricted to those affected by the instruments. We discuss the external validity of our results below.

6. Empirical results

The first rows in Table 2 show the results relative to the instrument variables used in the first-stage estimation.[36] Coefficients indicate that controlling for other demographic and socio-economic characteristics, the probability of living with a single mother in Brazil differs depending on the gender of the firstborn and the local sex ratio. Indeed, both coefficients are positive and statistically significant at 95% and 99% confidence level, respectively.

As explained above, the use of an overidentified estimation allows us to run a series of tests that support the instrument variable strategy used. In order for ''firstborn girl'' and ''low sex ratio'' to be valid instruments for single motherhood, the first stage in the instrumental variable regression estimation must be sufficiently strong. This is

Table 2
Coefficients of the IV estimation of height-for-age *z*-scores in Brazil (2SLS), children aged 6–60 months.

	Coefficient	Std. error	*P*-value
	Single motherhood (first-stage)		
Instrument variables			
Firstborn girl	0.017	0.007	0.033
Low sex ratio	0.042	0.008	0.000
	Height-for-age *z*-score (second-stage)		
Family structure			
Single mother	−2.654	1.098	0.016
No. of children up to 60 months	−0.140	0.068	0.041
No. of children 5–10 years old	−0.160	0.057	0.005
No. of children 11–15 years old	−0.008	0.070	0.909
Child's characteristics			
Girl	0.158	0.040	0.000
Black, coloured or indigenous	−0.005	0.069	0.937
Age (in months)	−0.040	0.010	0.000
Age (squared) (in months)	0.000	0.000	0.001
Mother's characteristics			
Height	0.036	0.004	0.000
Body mass index (BMI)	0.016	0.005	0.002
Age at birth	0.591	0.206	0.004
Age at birth (squared)	−0.020	0.007	0.006
Age at birth (cubic)	0.000	0.000	0.008
Ref. Never at school			
Primary school incomplete	0.177	0.112	0.114
Completed primary	0.135	0.110	0.224
Completed secondary	0.164	0.155	0.292
University degree	0.217	0.227	0.339
Employed	0.475	0.164	0.004
Household economic status			
Ref. 'Not poor, not receiving Bolsa Familia'			
Not poor, receiving Bolsa Familia	0.067	0.087	0.442
Poor, not receiving Bolsa Familia	0.113	0.182	0.534
Poor and receiving Bolsa Familia	0.107	0.172	0.535
Household income per capita – first quartile			
Second quartile	−0.081	0.100	0.419
Third quartile	−0.245	0.169	0.148
Fourth quartile	−0.379	0.199	0.060
Urbanization/regional characteristics			
Piped water	0.109	0.096	0.254
Paved street	0.218	0.057	0.000
Ref. Southeast			
Northeast	−0.039	0.068	0.568
North	−0.204	0.095	0.031
South	0.089	0.051	0.081
Central West	0.051	0.091	0.575
Urban	0.081	0.094	0.388
Constant	−11.33	1.699	0.000
N		9255	

Source: Pesquisa de Orçamentos Familiares (POF), 2008–2009. Weighted results. Heteroskedasticity-robust standard errors, clustered at State-strata level.

[35] As explained by Angrist and Pischke (2009), ''without monotonicity, IV estimators are not guaranteed to estimate a weighted average of the underlying individual causal effects'' (Angrist and Pischke, 2009: 154). See also Imbens and Angrist (1994).

[36] To save space, we do not show the full set of coefficients from the first-stage equation but they are available from the authors upon request.

confirmed with an *F*-statistic that takes the value of 14.12. Similarly, the results from the Sargan and Basmann tests for overidentifying restrictions assert that both instruments are uncorrelated with ν_i. Also, the Hansen's J and the difference-in-Sargan C tests indicated that both instruments satisfy the orthogonality condition (see Baum, 2006).[37,38] In short, from the tests, we gain confidence that we are not drawing spurious conclusions that are based on weak instruments.

After confirming the validity of our estimation strategy, lets move on to our main findings. The second panel in Table 2 clearly indicates that, controlling for other observable factors, being raised by a single mother is associated with lower height-for-age *z*-scores or, phrased differently, with a higher probability of suffering malnutrition in Brazil compared to children in two parents households. The coefficient for single mothers is −2.65 and is statistically significant at 2%. Importantly, remember that this result refers to the subgroup of children who have a single mother because the family's firstborn is a girl and they live in a low sex ratio area but who would not otherwise be living in this type of family structure. As a matter of fact, the use of two instrumental variables implies that our results are a weighted average of two LATEs (each specific to the instrument compliant subpopulation) with the weight defined by the strength of each instrument in the first stage.[39] In the case of "firstborn girl", compliers make up 8.3% of the treated group, and in the case of "low sex ratios", 13.1%. Less than 1.5% of the non-treated group are compliers in both cases.[40]

Arguably, our results refer to a small compliant population but, at the same time, this marginal group of children should be of primary interest in health policy intervention. Our *compliers* are likely to be children whose fathers take less responsibility and are less committed and whose mothers may have greater difficulties to find a new father figure (who could benefit the children, for example, in terms of economic security). More generally, low sex ratio contexts are associated with decreased female bargaining power in the marriage market which tends to shift resources in a way that normally does not favour women and their children (Angrist, 2002). It is true that we cannot guarantee the external validity of our results with the dataset at hand, but it is our understanding that our findings should be taken as a warning sign for health policy design.

In comparison, results from a simple OLS would clearly underestimate the influence of single motherhood on the height-for-age *z*-score of children in Brazil. As a matter of fact, the coefficient for single mother from an OLS regression controlling for the same observed characteristics yields a coefficient of −0.31 with a *P*-value of 0.000. The high divergence of OLS and 2SLS results, in addition to selection bias, is due to the fact that our study captures the impact of a small complier group while the OLS shows the mean for the whole sample. In the OLS context, the extreme cases are diluted by the mean.[41] Similarly, the coefficients of "firstborn girl" and "low sex ratio" in the reduced form regression are statistically significant at standard levels but are small (−0.080 and −0.085, respectively). Again, this is explained because in the reduced form, which captures *intention-to-treat* (ITT), the instruments are only an imperfect proxy for the treatment received (single motherhood) thus, the reduced form understates the causal effect of treatment *per se* (Angrist, 2006). Non-compliance dilutes ITT effects.[42] Only the 2SLS estimates capture the causal effect of single motherhood for compliers (undiluted by non-compliance and unaffected by selection bias).

As for the other variables for family structure, the number of children in a household up to 10 years of age is associated with lower height-for-age *z*-scores in the studied sample, but not in the case of older siblings (11–15 years). It is reasonable to believe that malnutrition is more prevalent in families with a large number of children since food needs to be shared among more members.

Regarding the other explanatory variables, we focus first on children's characteristics. Being a female infant is positively related with higher height-for-age *z*-scores while being black, coloured or indigenous does not have any explanatory power. Our results indicate that the nutritional racial gap is closing in Brazil once controlling for the endogeneity of family structure. Children's age (in months) is negatively associated with height-for-age *z*-score, which indicates that younger children are (in relative terms) taller than older children.

The mothers' characteristics offer some of the factors with the greatest explanatory power. Mothers' height is associated with higher children's height-for-age *z*-scores and the same is true for body mass index. Both variables are keys in our specification as they control for the child's

[37] We have also tested the possibility that one of the instrument variables could be redundant (and thus, the efficiency of the estimation would not be improved by including them). The likelihood-ratio test of redundancy of "firstborn girl" was rejected at 90% confidence level and the one for "low sex ratio" at 99% which more clearly indicates that the variable is not redundant.

[38] We have used the option `gmm2` in our estimation to guarantee consistent and efficient estimates in the presence of errors that are not i.i.d. Additionally, the use of the option `liml` for Limited Information Maximum Likelihood estimation made no difference to our results which is reassuring.

[39] Within the potential outcomes framework of IV estimation, it is important to understand that different instrumental variables identify different causal parameters, each specific to the subgroup of compliers for that instrument. With multiple instruments, "(...) 2SLS is a weighted average of causal effects for instrument-specific compliant subpopulations" (Angrist and Pischke, 2009: 175). In our case, the relative strength of the instrument in the first stage indicates a weight of 0.86 for "low sex ratio" and of 0.14 for "firstborn girl". Moreover, the use of covariates produces an average of covariate-specific LATEs while increasing the precision of 2SLS estimates. Also, we run a regression with a mutually exclusive set of three dummies built from the two instruments and we reach the same conclusions.

[40] According to Angrist and Pischke (2009), "the proportion of treated who are compliers is given by the first stage, times the probability the instrument is switched on, divided by the proportion treated" (Angrist and Pischke, 2009: 168).

[41] Noticeably, the average height-for-age *z*-score of children living in a single mother household in a low sex ratio area and with a female eldest sister is as low as −0.65. Note, however, that compliers cannot be individually identified in the dataset.

[42] Non-compliers are, for example, firstborn boys that live with a single mother.

genetic code. As for the mother's age at birth: the older the mother was when she gave birth to the child, the higher her children's scores. This association slightly reverses at older ages, following the usual inverted-U shape.[43] A mother's status in the labour market is another key variable for understanding children's height-for-age z-scores in Brazil: children with an employed mother have higher than average height-for-age z-scores, with a coefficient that is statistically significant at 1%. Differently, the mother's level of schooling does not have any explanatory power.

Household economic status was controlled by building a categorical variable that combined poverty status and receipt of *Bolsa Família* (BF). Similarly, we included household income (per capita) quartiles. However, none of these variables have a statistical effect that is different from zero at least at 95% confidence level.[44] In terms of the level of urbanization, the results show that living in an area with paved streets is positively related with higher height-for-age z-scores while piped water does not have any explanatory power. As for regional characteristics, children being raised in the North have a significantly lower height-for-age z-score than those in the (relatively richer) Southeast.

Other variables were included in the specification to test their importance for height-for-age z-scores but they did not yield statistically significant results and were left out of the final specification. These include whether the child has a private insurance plan, sewage disposal in the household, mail delivery service, dangerous location in the sense of being close to a dump, sewer, polluted river or a hill subject to sliding, garbage service and street-lighting. Note that the inclusion of none of these variables (or groups of them) made any difference to the main results.

7. Conclusions

This paper has studied the causal relationship between family structure and children's health conditions in Brazil. In particular, we assessed whether single motherhood has any influence on the height-for-age (z-score) of children aged 6–60 months. We use data from the *Pesquisa de Orçamentos Familiares* collected between 2008 and 2009 and our results are the outcome of an instrumental variable econometric strategy that has been used for the first time in the context of child health analysis. More precisely, we use male preference for firstborn sons and local sex ratios to instrument the probability of a woman becoming a single mother.

We find that children being raised by a single mother have a height-for-age z-score that is lower than that of children of similar characteristics that cohabit with both progenitors. Importantly, and given our use of IV estimation, our results are a weighted average of two local average treatment effects (LATEs): estimates refer to a small subgroup of children who have a single mother because the family's firstborn is a girl and they live in a low sex ratio area but who would not otherwise be raised in this type of family structure.[45] Our results do not have implications for the effect of single motherhood on children in households whose family structure is not affected by these characteristics. However, we have argued that our subgroup of *compliers* is likely to be children in the greatest need of policy intervention because they possibly have a less committed father and a mother with little bargaining power in the marriage market. Even if our results cannot be referred to all children, we believe that they should be of concern in policy design and understood as a warning sign. As in other parts of the world, each year, more and more children are being raised by single mothers in Brazil. Health policy designers should not only focus on the (shrinking) inequalities by region or race but should also shift their attention to the (increasing) inequalities by family structure.

The mechanisms that help to understand the positive relationship between stunting and single motherhood in Brazil, while controlling for other socio-economic and demographic factors, are difficult to disentangle. The previous literature has indicated that single mothers probably suffer higher levels of stress (and depression) given the need to deal with the dual role of sole carer and primary breadwinner.[46] Stress may affect women's capacity to care for children and therefore child development may be jeopardised. A single mother also lacks monitoring by a cohabiting partner that could, for example, dissuade her from health-harmful and risk-taking attitudes with regard to the child. Finally, women that raise their children on their own may have a smaller extended family and also fewer social relations (due to the amount of time and energy that needs to be devoted to childcare) and, as a result, may find it harder to seek or obtain help (in all domains).

Future research should take up some of the limitations of our paper. As explained, we were unable to distinguish between divorced, separated, never married or widowed mothers but each path into single motherhood may have different consequences for children's health outcomes.[47]

[43] Mother's age was not included in order to avoid multicollinearity problems.

[44] A possible explanation to these results is that, in POF 2008–2009, the reference period for income information is 12 months prior to completing the questionnaire. Given the heterogeneity of income sources (cash and seasonal non-cash) plus the large reference period, the variable of income per capita in this survey may not quantify very well the purchasing power of the household. We tried substituting income quartiles with log income but observed no difference in the results.

[45] Our results are probably conservative as we have intentionally not considered teenage mothers in order not to confound biological immaturity with single motherhood (Finlay et al., 2011). Linnemayr et al. (2008) indicate the possibility that not only the biology of adolescence in utero can explain poorer children's health outcomes among teenage mothers but also their mother's inexperience at child rearing or a weaker bargaining position within the household regarding scarce resources.

[46] For example, Paxson and Waldfogel (2002) found for the United States that higher levels of absent fathers, especially absent fathers and working mothers, were associated with higher rates of child maltreatment.

[47] Corak (2001) finds that children's outcomes are more weakly associated with parental death than divorce.

It may also be important to account for the amount of time that a mother has acted as the sole carer in order to find out whether changes in family structure are short or long-term triggers of health problems. Additionally, the interpretation of our results within the potential outcomes framework highlights the fact that other identification strategies are necessary in order to confirm our findings.

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