Social Science & Medicine 211 (2018) 9–15

Contents lists available at ScienceDirect

Social Science & Medicine

journal homepage: www.elsevier.com/locate/socscimed

Effect of a conditional cash transfer program on length-for-age and weight-for-age in Brazilian infants at 24 months using doubly-robust, targeted estimation

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A R T I C L E I N F O

Keywords:
Brazil
Conditional cash transfer
Child growth
Health policy
Targeted estimation
Targeted
Doubly robust estimation
Sensitivity analysis

A B S T R A C T

Objective: Conditional cash transfer programs are popular internationally and represent a large investment in child health. Evidence of their impact on child nutrition status remains weak and inconsistent, particularly for Bolsa Família, the Brazilian conditional cash transfer program and one of the world's largest. Our objective was to estimate the effect of the Brazilian conditional cash transfer program, Bolsa Família (BF), on child nutritional status as measured by length-for-age z-score (LAZ) and weight-for-age z-score (WAZ) at 24 months.
Methods: We analyzed the 1703 children eligible for BF from the 2004 Pelotas Birth Cohort. Children were divided into three exposure groups by total amount of money their household received from BF in 24 months: no BF, low BF (≤R$1000) and high BF (> R$1000). Using a doubly robust semiparametric estimation method we estimated the effect of receiving low and high levels of BF on LAZ and WAZ at 24 months.
Results: After adjustment for measured confounders, the expected difference in LAZ between children that received low or high levels of BF compared to no BF was −0.14 [95% confidence interval (CI): −0.27, −0.02] and −0.20 (95% CI: −0.33, −0.08) respectively. For WAZ the estimated differences were −0.04 (95% CI: −0.17, 0.08) for low levels versus no BF and −0.18 (95% CI: −0.30, −0.05) for high levels versus no BF. The expected difference in population LAZ had all eligible households received it and population LAZ under no BF was −0.15 (95% CI: −0.26, −0.04). Sensitivity analyses suggested only a strong confounder could explain away these results.
Conclusions: Among participants of the 2004 Pelotas Birth Cohort, BF was associated with a reduction in LAZ and WAZ in 24 month old children.

1. Introduction

Conditional cash transfers (CCTs) are government programs that give cash payments to poor families provided they meet specific requirements (Rawlings, 2005). Most CCTs have conditions that relate to improving child health and education such as requiring health checkups, vaccinations, perinatal care for the mother, school enrolment or a specific level of school attendance (Lagarde et al., 2009). CCTs have been implemented by a large number of developing countries, including nearly all Latin American countries, in an attempt to reduce intragenerational inequalities and break the intergeneration poverty cycle (Fiszbein et al., 2009).

Bolsa Família (BF), the Brazilian CCT program, began in 2004 and by 2015, served over 27 million households (Fiszbein et al., 2009; Ministério de Desenvolvimento Social e Combate à Fome, 2016) and paid out a total of R$28.5 billion in 2016 alone (9.0 billion USD) (Controladoria-Geral da União Brasileira, 2017). BF provides families with monthly payments conditional on their children receiving yearly health checkups, getting their scheduled vaccinations and maintaining 85% school attendance. The program also includes an unconditional

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https://doi.org/10.1016/j.socscimed.2018.05.040
Received 25 October 2017; Received in revised form 16 April 2018; Accepted 22 May 2018
Available online 24 May 2018
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component for the poorest families. In 2004, families could receive between R$15–95 monthly depending on their household per capita income and number of children (see appendix, Table A1) (Soares, 2012).

There have been a number of studies on the impact of CCTs on child anthropometric measures but we will focus on the evidence from Latin America because BF is a Brazilian CCT. A randomized control trial of a Mexican CCT found a positive effect on child height (Rivera et al., 2004) but had potential sources of bias such as contamination between arms (Rivera et al., 2004), substantial attrition (Parker and Teruel, 2005) and possible differential loss to follow-up (Stecklov et al., 2005). The literature is summarized by two systematic reviews, which found a high risk of bias and demonstrate a wide range of effect estimates (Lagarde et al., 2009; Manley et al., 2013). Even within Brazil, results are mixed. Of four studies of the effect of BF (or BF's pre-cursor Bolsa Alimentação) on the height of children under 5 years old, one found decreased odds of stunting (Paes-Sousa et al., 2011), two found null results (Oliveira et al., 2011; Saldiva et al., 2010), and one found a decrease in height for age (Morris et al., 2004). All four of these studies used cross-sectional data and controlled for a very limited number of confounders. Larger studies and alternative methods may lead to better estimates of the health impacts of CCTs.

Although a variety of anthropometric measures are used to evaluate nutritional status, height-for-age (HAZ) or length-for-age (LAZ) is of particular interest for two reasons. First, HAZ and LAZ are good proxies for cumulative nutritional status allowing us to study the long-term nutritional consequences of BF (Lejarraga, 2012). Second, poor growth and stunting among children, particularly in the first 24 months, is an important predictor of poor future health and poor cognitive outcomes (Grantham-McGregor et al., 2007; Sudfeld et al., 2015; Victora et al., 2008). Therefore, any improvement in HAZ may be due to an improvement in past nutritional status and may improve health and cognitive outcomes later in life. Weight-for-age (WAZ) is also used to evaluate nutritional status and can indicate shorter term changes in nutritional status (Lejarraga, 2012).

Our study aims to estimate the effect of BF on LAZ and WAZ at 24 months using a high quality birth cohort linked to a government database. We chose the period from birth to 24 months because that is when growth faltering is most likely to take place (Shrimpton et al., 2001).

2. Methods

2.1. Study population

We used data from the 2004 Pelotas Birth Cohort in southern Brazil, which recruited 99% of hospital births in all five hospitals within the city limits of Pelotas in 2004. Questionnaires on prenatal, parental and socioeconomic variables were administered shortly after delivery. Measurements of the newborns were taken within 24 h of birth. Follow-up of the cohort occurred at 3, 12, 24, 48, 72 months and 11 years (Santos et al., 2011). Only data from birth, 12 and 24 months were used in this analysis. Trained field workers conducted house visits to measure the infants and administer questionnaires to the parents. Non-singleton births and infants who died before 24 months were excluded from the analysis. We limited our analyses to families reporting a household per capita income of R$100 or below because households with a larger per capita income were not eligible to receive BF. More information on the 2004 Pelotas Birth Cohort is available elsewhere (Santos et al., 2011, 2014).

2.2. Measures

Our primary outcomes were LAZ and WAZ at 24 months. Length was measured with the child barefoot and in the recumbent position using an infantometer. Weight was measured by subtracting the weight of the mother from the weight of the mother holding the child. LAZ and WAZ were calculated using the 2006 World Health Organization growth standards (WHO Multicentre Growth Reference Study Group, 2006).

Data on the amount of money received from BF were obtained by linking the cohort with the Portal de Transparência (http://www.portaltransparencia.gov.br), a publicly available database that contains information on monthly payments to BF recipients (Schmidt et al., 2017). The linkage was done probabilistically using the child's name and birth date as well as the parent's name and age. Of the 1796 matches, 83.2% were perfect matches and 12.2% only differed in the parent's name by at most, four letters. All non-perfect matches were reviewed by hand.

The crude relationship between money received from BF and LAZ at 24 months was found to be non-linear using a scatterplot and local linear regression (see appendix, Fig. A1). Therefore, the BF was categorized into three groups by total amount of BF received in the first 24 months: no BF (received R$0), low BF (received R$1000 or less) and high BF (received more than R$1000). The R$1000 cutoff was chosen based on examination of the crude relationship between money received from BF and LAZ and WAZ. In secondary analyses with the outcome measured at 12 months, the cutoff between low and high BF groups was R$500.

We only adjusted for baseline variables. A full list of covariates can be found in Table 1.

2.3. Statistical analysis

Our target parameters consisted of the expected difference in population LAZ and WAZ under five contrasts where we contrast the population average outcome difference when everyone in the population is set to each level:

1) Low BF versus no BF
2) High BF versus no BF
3) High BF versus low BF
4) The observed BF level versus no BF
5) The level of BF they were eligible to receive versus no BF

In contrast 4 observed BF refers to the level of BF the household was observed to have received. In contrast 5 we set BF to the level a person was eligible for given their household composition and income. We also the estimated the same five parameters using the outcomes measured at 12 months.

These parameters were estimated using targeted maximum likelihood estimation (TMLE). TMLE is a doubly robust, semiparametric, efficient substitution estimator of the statistical parameters, equaling the population average treatment effects under the necessary causal assumptions (van der Laan and Rose, 2011). TMLE relies on estimation of the outcome regression given BF and covariates, and estimation of the exposure mechanism given the covariates. If either the outcome regression or exposure mechanism is estimated consistently, then TMLE will yield consistent estimates of the parameter of interest. In other words, TMLE provides two chances to control for measured confounding. We viewed this as an important advantage in our context because the exposure and outcome models are social and biological mechanisms, respectively. To minimize the risk bias due to regression model misspecification, we used a data-adaptive method (van der Laan et al., 2007) to estimate the complex mechanisms that can describe child growth or selection into BF particularly with the large number of covariates we adjust for. TMLE and data-adaptive estimation make our data more robust to residual confounding than previous methods used to study the effects of CCTs. Cross-validation provided the optimal combination of estimates from a wide range of algorithms (see appendix, Table A2 for a list of algorithms used).

Standards errors were estimated using the estimated influence curve

Table 1
Anthropometric, perinatal, parental and socioeconomic characteristics of the households from the 2004 Pelotas Birth Cohort who reported earning a household per capita income below R$100.

Variables	No BF	Low BF	High BF
Outcomes			
LAZ at 12 months	−0.28 (1.28)	−0.43 (1.29)	−0.60 (1.24)
LAZ at 24 months	−0.34 (1.14)	−0.55 (1.19)	−0.69 (1.14)
WAZ at 12 months	0.19 (1.12)	0.11 (1.25)	−0.03 (1.23)
WAZ at 24 months	0.12 (1.13)	0.02 (1.19)	−0.19 (1.11)
Birth			
Length at birth (cm)	48.0 (2.6)	48.2 (2.5)	48.3 (2.7)
Weight at birth (g)	3112 (536)	3161 (522)	3188 (560)
Female	0.48	0.54	0.46
Caesarean	0.38	0.29	0.27
Gestational age (weeks)	38.5 (2.5)	38.7 (2.1)	38.7 (2.4)
Health problem at birth	0.13	0.10	0.09
Prenatal			
Weight gain during pregnancy (kg)	11.3 (6.3)	10.8 (5.9)	10.4 (6.1)
Smoked during pregnancy	0.37	0.40	0.41
Drank alcohol during pregnancy	0.04	0.05	0.03
Mother worked during pregnancy	0.27	0.26	0.27
Number of prenatal visits	7.1 (3.2)	6.8 (2.9)	6.7 (3.1)
Parity	2.5 (1.8)	3.1 (2.0)	3.9 (1.9)
Mother hospitalized during pregnancy	0.12	0.12	0.10
Parental			
Mother's age	24.2 (7.0)	25.9 (7.2)	27.1 (6.3)
Father's age	28.3 (8.6)	30.1 (8.5)	30.6 (8.5)
Mother's education (years)	6.6 (3.0)	6.0 (2.6)	5.4 (2.5)
Father's education (years)	6.5 (3.2)	5.6 (2.9)	5.3 (2.9)
Mother's skin color*			
White	0.66	0.57	0.58
Black	0.27	0.36	0.32
Father's skin color*			
White	0.57	0.46	0.53
Black	0.18	0.24	0.21
Mother's height (cm)	158 (6)	157 (7)	157 (6)
Mother was born premature	0.16	0.19	0.20
Socioeconomic			
Per capita income (R$)	42 (34)	42 (32)	45 (28)
Number of children living in the household	2.1 (1.4)	2.6 (1.6)	3.3 (1.6)
Number of people in the household	5.3 (2.1)	5.5 (1.9)	5.9 (1.9)
Other family members living in household	1.5 (2.2)	1.1 (1.8)	0.7 (1.7)
Other people living in household	0.1 (0.4)	0.1 (0.7)	0.0 (0.2)
Own black and white television	0.24	0.21	0.30
Number of color televisions sets	1.1 (0.8)	0.9 (0.7)	0.8 (0.6)
Own a car	0.15	0.09	0.08
Own fridge	0.82	0.77	0.75
Own freezer	0.17	0.10	0.05
Own vhs player	0.20	0.11	0.08
Own washing machine	0.42	0.34	0.32
Own microwave	0.07	0.02	0.01
Own air conditioner	0.01	0.00	0.00
Own computer	0.05	0.00	0.00
Own landline	0.37	0.30	0.27

and the delta method (Moore and van der Laan, 2009). We performed all analyses in R 3.1.2 (R Core Team) using the ltmle package (Schwab et al., 2016). Computations were run on the supercomputer [redacted].

There was missing data both due to loss to follow-up and item non-response. We used multiple imputation to impute missing values using the mi package in R (Su et al., 2011). We ran analyses on five imputed datasets and pooled the results using Rubin's Rules (Rubin, 2004).

We used a recently-developed sensitivity analysis to determine the characteristics of an unmeasured, binary confounder that would make the estimates presented in our analyses null (Ding and VanderWeele, 2016). Using this method, we produced graphs showing the maximum difference in the outcome by level of unmeasured confounder required to make our estimates null for a given ratio of BF prevalence between levels of the unmeasured confounder. The same graphs can be interpreted for a categorical unmeasured confounder. In that case, the x-axis is the maximum ratio of BF prevalence within levels of the unmeasured confounder and the y-axis is the largest difference in the outcome between any two levels of the unmeasured confounder within levels of BF. These graphs were presented using centimeters and kilograms as the outcome rather than z-scores because the method requires a positive continuous outcome. As a second sensitivity analysis, we fit the models without parental education variables and socioeconomic variables in Table 1 to determine how much adjustment for these variables changed our estimates.

Lastly, we verified the positivity assumption by checking the overlap in propensity scores within each level of treatment.

The data collection and analyses were submitted to and approved by the Research Ethics Committee [redacted]. The Research Ethics Office of [redacted] also approved this study.

3. Results

3.1. Sample characteristics

Of 4231 infants enrolled at birth in the cohort, 84 twins were excluded plus another 88 children who had died before 24 months. An additional 2356 families were above the R$100 household per capita income threshold and were excluded from the analysis (Fig. 1). Of the 1703 families in the analysis, 110 were in the low BF group and 319 were in the high BF group at 12 months and 291 were in the low BF group and 355 in the high BF group at 24 months. On average, participating households received BF for 16 (SD = 8) of the 24 months after the birth of their infant. Of those who initiated BF, 96% were still receiving it when the outcomes in this study were measured 24 months after their child was born. The variation in the number of months households received BF is due to households not receiving BF immediately after birth, not receiving some payments because they violated a condition of BF or because they did not withdraw the payment. Loss to follow-up at 24 months was 7.5%. A descriptive table of those lost to follow-up can be found in the appendix (Table A3). BF participation was lower among those lost to follow-up (22%) than in the rest of the participants (39%). This is likely because the mechanisms behind being lost to follow-up are similar to those that prevent participation in BF (e.g. migration, lack of a fixed address). Item non-response was under 5% for all covariates except weight gain during pregnancy (12%), father's education (12%) and whether the mother was born premature (36%).

Mothers receiving BF were older, shorter, less educated, attended fewer prenatal visits, gained less weight during pregnancy, were more likely to smoke and less likely to work during the pregnancy relative to mothers who did not receive BF (Table 1). Birth length, birth weight, gestational age and reporting health problems at birth were all similar by BF level. All socioeconomic variables indicated that exposed households were notably poorer and had larger households. In the entire 2004 Pelotas Birth Cohort there is a low prevalence of LAZ growth deficit (5.0%) and WAZ growth deficit (2.4%) by WHO standards (WHO Multicentre Growth Reference Study Group, 2006) but the same prevalences are nearly double among those eligible for BF.

3.2. BF association with LAZ and WAZ

Results estimated using multiply imputed data differed from complete case analyses (see appendix, Table A4) therefore we will only present the former. We compared the estimated propensity scores for

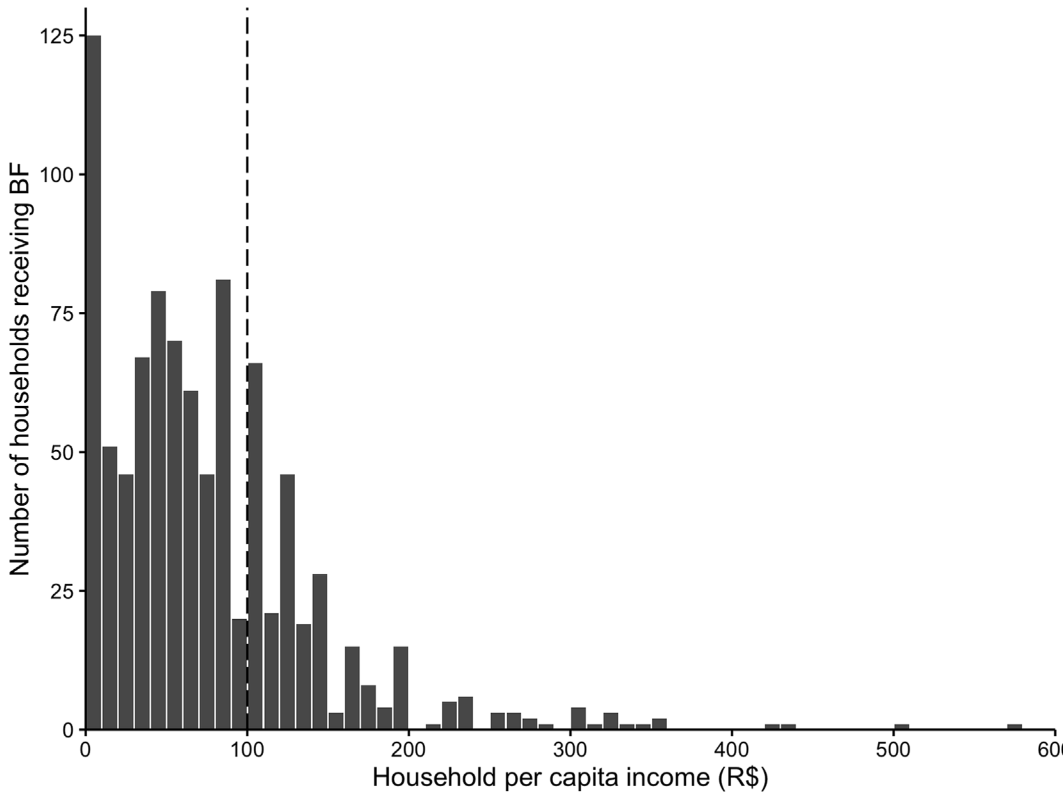

Fig. 1. Household per capita income of Bolsa Família participants: Histogram of the number of households from the 2004 Pelotas Birth Cohort who received at least one Bolsa Família payment within two years of the birth of their infant by household per capita income. The dashed line represents the upper limit of household per capita income for Bolsa Família eligibility.

Table 2
Estimated population average outcomes (95% CI) had all households received each level of BF and estimated differences between each level (95% CI).

Outcome	Age (months)	Estimated population average			Contrast between levels of BF		
		No BF	Low BF	High BF	Low vs no BF	High vs no BF	High vs low BF
LAZ	12	−0.33 (−0.41, −0.25)	−0.40 (−0.60, −0.20)	−0.52 (−0.62, −0.42)	−0.07 (−0.28, 0.14)	−0.19 (−0.30, −0.08)	−0.12 (−0.34, 0.09)
	24	−0.39 (−0.49, −0.29)	−0.54 (−0.65, −0.42)	−0.59 (−0.69, −0.50)	−0.14 (−0.27, −0.02)	−0.20 (−0.33, −0.08)	−0.06 (−0.20, 0.09)
WAZ	12	0.15 (0.08, 0.22)	0.15 (−0.03, 0.33)	0.08 (−0.02, 0.19)	0.00 (−0.18, 0.18)	−0.07 (−0.19, 0.05)	−0.07 (−0.27, 0.14)
	24	0.08 (−0.01, 0.18)	0.04 (−0.08, 0.16)	−0.09 (−0.19, 0.01)	−0.04 (−0.17, 0.08)	−0.18 (−0.30, −0.05)	−0.13 (−0.29, 0.02)

each level of treatment and concluded there was sufficient support in the data for our analyses (see appendix, Fig. A2). We found a negative association between BF and LAZ and WAZ at nearly all BF levels (Table 2). The association was highest for receiving high BF in LAZ and WAZ at 24 month with a reduction around 0.19 z-scores in both cases. Comparisons between low and high BF were all negative but had confidence intervals also consistent with positive values. Estimates were smaller at 12 months than at 24 months. Results from all estimation methods and in the full dataset can be found in the appendix (Table A4).

Due to the low BF participation rate, the observed BF distribution changed the population average outcome only slightly and the confidence intervals are consistent with both a positive and negative difference (Table 3). Had these households received the full amount of BF they were eligible for, the average population LAZ at 12 and 24 months would have decreased by about 0.15 z-scores whereas the average population WAZ at 12 and 24 months would have decreased by around 0.06.

To demonstrate the robustness of our findings to unmeasured confounding, we characterized a categorical unmeasured confounder that would render the observed estimates null (Fig. 2). In each graph, the line represents the largest difference in the outcome between any two levels of the unmeasured confounder within levels of BF that would be required to make the estimate null given a relative prevalence of the unmeasured confounder in those exposed and unexposed to a specific level of BF. For example, an unmeasured confounder that is twice as common in the high BF group as the unexposed (x-axis), a difference in height in infants with and without the unmeasured confounder would need to be at least 1.47 cm (y-axis) for the high BF estimate to be null. For low BF, the difference in height in infants with and without the unmeasured confounder would need to be 0.70 cm to make the estimate null. In comparison, the *crude* association of smoking with LAZ at 24 months is −0.42 cm (−0.50, −0.33) and smoking is only 13% more prevalent in the high BF group relative to the low BF group. Therefore, any potential U would need more than twice as strong as the crude association of smoking with LAZ and be even more prevalent in the high BF group relative to the unexposed group.

Lastly, we found that the largest difference between the models

Table 3
Estimated population average outcome (95% CI) under three BF distributions: no BF, the observed BF distribution and the distribution had all families received the full amount of BF they were eligible to receive (full compliance). Also presented are the estimated differences in population outcome between the observed BF distribution and no BF and the full compliance distribution of BF and no BF.

Outcome	Age (months)	Estimated population average			Contrasts between levels of BF	
		No BF	Observed BF	Full compliance BF	Observed vs no BF	Full compliance vs no BF
LAZ	12	−0.33 (−0.40, −0.25)	−0.37 (−0.59, −0.15)	−0.48 (−0.61, −0.36)	−0.04 (−0.25, 0.17)	−0.15 (−0.28, −0.02)
LAZ	24	−0.39 (−0.49, −0.30)	−0.46 (−0.62, −0.30)	−0.55 (−0.65, −0.44)	−0.07 (−0.20, 0.07)	−0.15 (−0.26, −0.04)
WAZ	12	0.15 (0.08, 0.22)	0.14 (−0.06, 0.34)	0.10 (−0.01, 0.22)	−0.01 (−0.20, 0.17)	−0.05 (−0.17, 0.08)
WAZ	24	0.08 (0.00, 0.17)	0.04 (−0.12, 0.20)	0.01 (−0.10, 0.12)	−0.04 (−0.18, 0.09)	−0.07 (−0.18, 0.04)

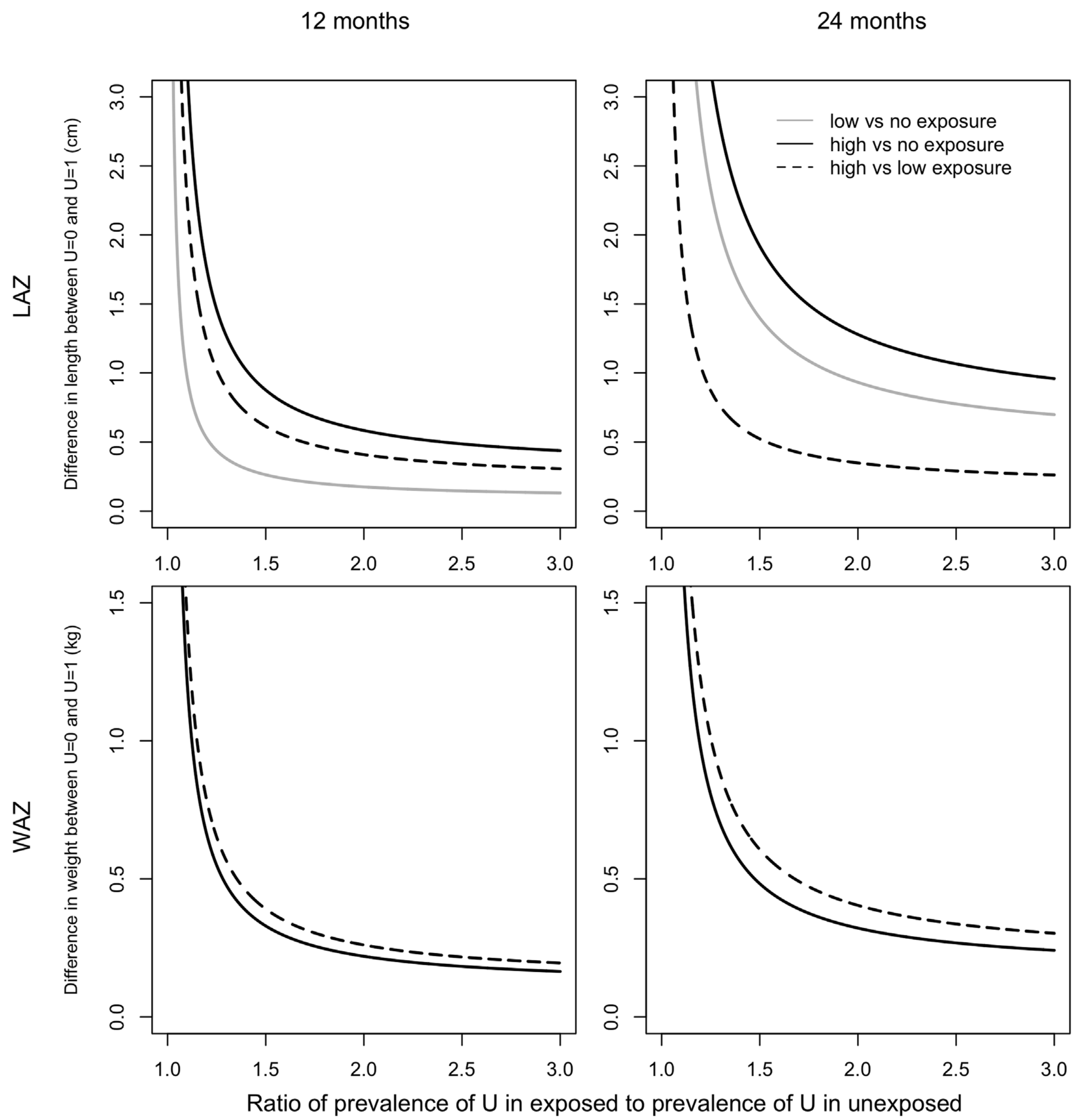

Fig. 2. Sensitivity to unmeasured confounding: The areas above each line represents combinations of associations between BF and an unmeasured confounder (U) and U-outcome associations that would be required to explain away the estimated effect sizes.

including or excluding socioeconomic variables was 0.04 (see appendix, Table A5). This suggests that any potential residual confounding due to unmeasured socioeconomic variables would have to be stronger than the confounding adjusted for with the socioeconomic variables already included in the analysis to render the estimates null.

4. Discussion

Our results suggest that in the participants of the 2004 Pelotas Birth Cohort, receiving BF within the first two years of life was negatively associated with LAZ and WAZ. At 24 months, this negative association is equivalent to about three weeks of delayed growth in LAZ and over five weeks of delayed growth in WAZ (21).

Although many studies have found a negative crude association between BF and HAZ or LAZ, we are only aware of one other that found a negative adjusted estimate. Morris et al. (2004) found that Bolsa Alimentação, a BF pre-cursor, reduced HAZ by 0.11 (95% CI: −0.23, 0.01) and WAZ by 0.13 z-scores (95% CI: 0.01, 0.25) in children between 0 and 84 months using administrative errors as a quasi-experiment. Overall, the literature on the association between CCTs and HAZ/LAZ and WAZ has produced very different estimates though mostly null or positive (Ranganathan and Lagarde, 2012; Manley et al., 2013), even within Brazil (Morris et al., 2004; Paes-Sousa et al., 2011). One national study in Brazil found that BF coverage on the municipal level was associated with a strong decrease in under-five mortality due to malnutrition (Rasella et al., 2013).

One explanation of the wide range of effect estimates could be uncontrolled confounding. For example, it is possible given the baseline characteristics that the poorest families were the first targeted to receive BF. If the discrepancy in estimates is due to unmeasured confounding, our study seems the least likely to suffer from this as we have controlled for more covariates than any other study to our knowledge and we are the only ones to employ semi-parametric and doubly-robust estimation to have two chances to correctly control for measured confounding. We acknowledge that despite our best efforts, the possibility of residual confounding cannot be eliminated. Using a recently proposed sensitivity analysis we did find, however, that a strong confounder would be required to explain away our result for LAZ at 24 months. Measurement error in household income may also bias our estimates but would likely do so toward the null if non-differential or upward if people with higher incomes are misreporting their income in order to be eligible for BF.

Another explanation for the wide variety of estimate in the literature is that the effects of CCTs are truly heterogeneous depending on the design of the program, how the program is targeted, when the program is implemented and the population within which it is implemented. In Brazil, municipalities are tasked with targeting populations and verifying that some conditions are met (de Janvry et al., 2005; Lindert et al., 2007; Soares, 2012). If one municipality targets the households most in need of social assistance while another targets those most motivated to participate, it is possible that the effects in these two municipalities could be qualitatively different. Unfortunately, we were not able to locate records of the BF targeting practices in Pelotas during the period of this study.

If the results are not due to residual confounding, there are possible causal mechanisms that would allow BF to have a negative effect on child LAZ and WAZ. Some studies have suggested that receiving BF may be associated with buying less nutritious food (de Bem Lignani et al., 2011). A study in North Eastern Brazil found that families who received BF had over three times higher odds of consuming junk food than families who did not receive BF (Saldiva et al., 2010). A poorer diet may lead to increased infections (Dewey and Begum, 2011) or micronutrient deficiencies which may, in turn, reduce LAZ (Black et al., 2013). There also exist, however, studies demonstrating a positive relationship between BF and diet and nutrition (Martins et al., 2013; Martins and Monteiro, 2016).

We have already highlighted our efforts to minimize residual confounding by using TMLE and data-adaptive estimation as well as to minimize selection bias using multiple imputation. Another important strength of our study is the data quality from 2004 Pelotas Birth Cohort for all outcomes and covariates. Also, to our knowledge, we are the first study of a CCT program to obtain a monthly record of payments allowing us to calculate the precise amount of money received during the study period.

As with all observational studies, our effect estimates require the unverifiable assumption of no unmeasured confounding be met. It is also possible that losses to follow-up or missingness in the data are not missing at random which would bias our estimates because our multiple imputation models would lack variables predictive of missingness. There is also the possibility of bias in the process of matching the cohort participants with BF records. If households with children who experienced lower growth were more likely to be matched in the BF payment database, this would bias our estimates downward. We believe that any mismatching that occurred was likely non-differential because the matching was done solely with names and birth dates. Lastly, low BF participation rates were observed because the cohort coincided with the beginning of BF. BF participation rates have been increasing (Schmidt et al., 2017) and it now covers a much larger proportion of the population now which may limit the generalizability of this study.

In conclusion, we found a negative between BF and LAZ and WAZ. Despite this finding, the LAZ and WAZ population averages were only slightly reduced because BF participation was low. It should be noted that though we found a negative association, the prevalence of stunting has been steadily decreasing in Brazil particularly among the poor, who are more likely to receive BF, though these improvements began well before BF and little improvement has been observed in the south of Brazil since the 1990s (Victora et al., 2011). Given that our study took place when BF was started additional studies with more recent populations that are nationally representative would be required to determine whether this negative association holds in the present.

Funding

JAL was supported a Banting and Best Doctoral Award from the Canadian Institutes of Health Researhc. JSK is a Canada Research Chair in Health Disparities. ES is supported by a Chercheur boursier Junior 2 from the Fonds de recherche du Québec – Santé. AM, ISS and AJDB are supported by the CNPq. This article is based on data from the 2004 Pelotas Birth Cohort, conducted by the Postgraduate Program in Epidemiology at Universidade Federal de Pelotas, with the collaboration of the Brazilian Public Health Association (ABRASCO). The World Health Organization, National Support Program for Centers of Excellence (PRONEX), Brazilian National Research Council (CNPq), Brazilian Ministry of Health, and Children's Pastorate supported the study in its two first years.

Acknowledgements

We would like to thank Dr. Mark van der Laan for answering our questions about specific aspects of TMLE. The operation of guillimin, the supercomputer used in this study, is funded by the Canada Foundation for Innovation (CFI), ministère de l'Économie, de la Science et de l'Innovation du Québec (MESI) and the Fonds de Recherche du Québec - Nature et Technologies (FRQ-NT).

Appendix A. Supplementary data

Supplementary data related to this article can be found at http://dx.doi.org/10.1016/j.socscimed.2018.05.040.

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