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The role of microcredit in older children's nutrition: Quasi-experimental evidence from rural China

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ABSTRACT

This article evaluates the causal impact of rural households' borrowing, through formal microcredit, on child nutrition in poor northwest China. The analysis exploits the panel data in rural Gansu between 2000 and 2004. Unobserved differences between borrowers and non-borrowers are controlled for in a dynamic fuzzy regression-discontinuity design creating a quasi-experimental environment for causal inference. Both anthropometric and micronutrient measures of child nutrition are investigated. Borrowing formal microcredit improves parent-reported health status and weight, and alleviates anemia and zinc deficiency. All effects nevertheless appear to exist in the short-term only.

Introduction

The impact of microfinance has been widely examined in many developing countries and a positive role for microfinance in reducing poverty measured by income or consumption has been documented (e.g., Imai et al., 2010 for India; Imai and Azam, 2012 for Bangladesh; Li et al., 2011 for rural China). In addition to improving economic well-being, it is also believed that access to microcredit may relax agents' credit constraints and provide a means of consumption-smoothing and therefore, improve social well-being by increasing expenditure on health and education (Armendáriz de Aghion and Morduch, 2005). However, evidence is not yet conclusive. Some literature finds positive correlation between access to credit institutions and health outcomes in Indonesia (Deloach and Lamanna, 2011) and between access to microcredit and expenditure on health in poor peri-urban Vietnam (Doan et al., 2011), but there is no significant impact of microfinance on borrowers' schooling and health in India (Banerjee et al., 2009). To my knowledge, in the context of rural China, however, there has been no empirical and causal examination of the effectiveness of microcredit *per se* on child nutritional status.

Children's nutrient intake in rural China has recently gained much attention (e.g., Luo et al., 2012; Shi et al., 2012). Recent reviews for developing countries (Currie and Vogl, 2012; Alderman, 2012) show that nutritional insults in childhood can pose long-term challenges for adult earning-potential and human capital formation. In China, childhood health also exhibits strong effects on adult health outcomes (Smith et al., 2012). However, a significant number of rural children do not have adequate access to micronutrients in rural China, despite a 12.6% annual increase in per capita net income and 10.8% annual growth in per capita consumption for Chinese rural households over the past two decades (1990–2010).[1] 12.8% of rural children (<15 years old) in 7 central and eastern provinces are anemic (Li et al., 2013). 65.3% of rural children aged 1–2 years in northern China have vitamin D-deficiency, leading to a rickets prevalence of 41.6% (Strand et al., 2009). In a poor western province, Shaanxi, where rural household per capita net income ranked within the bottom 5 out of 31 provinces between 1990 and 2010,[2] 38.3% of fourth-year primary school students suffer from iron deficiency (Luo et al., 2012).

It is suggested by the aforementioned studies that the government should take responsibility for micro-nutrition supplementation to rural children. Although supplements would be integral to child development, praxis may be deterred by the tight budgets of local governments, especially in poor areas. Health education

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[1] Author's calculation based on data from Rural Household Survey conducted and published by the National Bureau of Statistics in China Statistical Yearbook 2011. Monetary variables in this article are transformed to real terms at 1990 prices by the rural consumer price index from the same source.

[2] Author's calculation based on data from Rural Household Survey conducted and published by the National Bureau of Statistics in China Statistical Yearbook 2011.

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for parents informing them of their children's malnutrition could not alleviate the problem either (Shi et al., 2012), but would incur high costs (Ma et al., 2008). By conducting field experiments, Wang et al. (2009) find that food fortification in terms of nutrient fortified complementary food supplements can effectively prevent iron deficiency for children aged 6–12 months in poor rural Gansu province. Nevertheless, it would be difficult to extend such programs of direct therapeutic food distribution to older children until their adolescence in poor rural areas in China. For one thing, such large-scale interventions are less likely to be disbursed adequately and continuously by local (county) governments after decades of fiscal decentralization, especially in poor rural areas, in view of their limited fiscal capacity and strong incentives to achieve goals of economic growth compared to other aspects of human well-being. For another, at the household level, Chinese rural households confront various shocks and have limited capability to cope with them. 71% of rural households in 10 provinces are credit rationed, which brings about losses of 12–13% in income and 15–16% in consumption (Rui and Xi, 2010). These problems would undermine households' sustainable widespread use of food supplements for children and weaken demand for healthcare when the child is in ill health.

Given the above circumstances, microcredit might offer a solution. There are a number of ways borrowing microcredit can affect child nutrition and health status. In rural China microcredit enables households to earn more income by engaging in more profitable production and out-migration (Li et al., 2004) and raises consumption (Li et al., 2011). More income in turn increases food consumption as well as household demand for a higher quality diet (Yu and Abler, 2009).[3] Microcredit also serves as an effective means to smooth consumption for rural households in the presence of health shocks (Armendáriz de Aghion and Morduch, 2005; Gertler et al., 2009; Islam and Maitra, 2012). With more credit, households may be more able to smooth their food consumption against various risks and shocks and therefore, maintain children's nutritional and health status. To the extent that microcredit brings about higher income and enhances the household capacity for resisting risk, a potential role played by borrowing microcredit could be expected to increase care and investment in child health and hence to improve child nutritional status.

Different from the existing studies examining the benefits of access to microfinance (e.g., Deloach and Lamanna, 2011; Doan et al., 2011; Foster, 1995; Gertler et al., 2009), this article provides the first empirical attempt to test the causal impact of borrowing through formal microcredit *per se* on child nutritional outcomes. The present study also contains methodological advantages in dealing with endogeneity between borrowing and child nutritional outcomes. The analysis uses a fuzzy regression-discontinuity design (RDD) to reflect the unobservables between borrowers and non-borrowers. RDD is known to have milder assumptions and higher internal validity compared to other widely used 'natural experiment' strategies of impact evaluation, such as difference-in-difference, instrumental variables and propensity score matching, and can provide a quasi-experimental environment to extract more credible causal inferences (Lee and Lemieux, 2010). More importantly, the analysis takes dynamic impact of microcredit into account. As indicated by Islam (2011), it may take rural households a long time to build their capacity in order to meet the requirements of microcredit programs and the impact of borrowing behavior on their livelihood may also take time to be realized. Such possible progressive effects will be investigated further in a dynamic fuzzy RDD. The results of this article will also contribute to the on-going debate on the effectiveness of microfinance on raising borrowers' social well-being in the aspect of nutrition by providing new evidence from rural China and may inform policy intervention to tackle child malnutrition in poor regions.

In general, the empirical analysis reveals positive influence brought by formal microcredit uptake to rural children's nutrition outcomes. Borrowing formal microcredit appears to improve parent-reported health status and weight in the short-term, while this positive impact dissipates quickly over time. In regard to micronutrients, formal microcredit can increase zinc intake and alleviate child anemia temporarily for new borrowers.

The article proceeds as follows. The next section describes the dataset. 'Mehodology' section sets up the model. 'Results and discussion' section shows the appropriateness and credibility of the usage of fuzzy RDD and discusses the estimation results. Concluding remarks and possible implications for policy are summarized in 'Conclusion' section.

Data

The present study uses the Gansu Survey of Children and Families (GSCF) in 2004, which was initiated by the World Bank and recently supported by the United Kingdom Economic and Social Research Council/Department for International Development (ESRC/DfID) Joint Scheme for Research on International Poverty Reduction. The surveys were conducted locally by the National Bureau of Statistics Gansu Branch in collaboration with the Northwest Normal University and the Centre for Disease Control. Gansu has long lagged behind other provinces in China's progress to economic prosperity. The rural household per capita net income amounted to only 25.8% of that of Beijing and 54% of the national average in 2010 and had then been in the bottom 3 of the 31 provinces for a decade.[4]

The dataset includes 1918 older children aged between 12 and 20.[5] Each young adult was selected from one sample household spread across 100 villages in 20 counties. Some samples are dropped when estimating the assignment variable $\hat{p}_i$ and Eqs. (5), (6) because of missing values. The final sample size in regression reduces to 1128–1283, which will be shown in Tables 3 and 4 in the next section. The following descriptive statistics in this section therefore only focus on the upper sample size of 1283 in 95 villages out of 20 counties.[6]

543 sample households (42.3%) lived below the World Bank international poverty line of US$1.25/day, measured by their per capita consumption. The present study focuses on formal microcredit only, as China has stringent regulations for financial markets and restricts the entry of non-governmental financial institutions. Rural Credit Cooperatives (RCCs) are the largest formal microcredit provider in rural China (Li et al., 2011). RCCs services are available in almost every township and most villages in rural China (Li et al., 2011), which mitigates the non-random placement bias. Household borrowing from RCCs appears to be limited, although the government has been introducing financial reforms to facilitate lending to rural households since the early 2000s. Only 424 households (33.1%) arranged formal loans. 491 households (38.3%) felt it

[3] They estimated the quality-corrected income elasticities of demand for grain, vegetables and dairy products as 0.31, 0.35 and 0.18 in turn.

[4] Author's calculation based on data from Rural Household Survey conducted and published by the National Bureau of Statistics in China Statistical Yearbook 2011.

[5] 87% of them were 16-year old or younger. Only 3 observations were of 18- to 20-year olds.

[6] One may be concerned with biased estimators generated by the reduced sample size. I inspected this by comparing the descriptive statistics for key variables such as microcredit borrowing decisions, child health indicators and household wealth between remaining households in final regressions, those having been dropped during estimation procedures, and by testing for the difference of the mean of each variable. The t-statistics are all less than 2 (from 0.17 to 1.24), indicating that we cannot reject the null hypothesis of zero difference between two groups.

was more difficult to borrow than before.[7] Table 1 lists all variables in the estimation.

I employ two anthropometric indicators and two micronutrient indicators to reflect as comprehensively as possible children's nutritional status. The former class includes the parent-reported child health status, which has five categories from very poor to very good health, and the body-mass-index (BMI) calculated as the ratio of weight over squared height.[8] The latter class includes the hemoglobin level and the hair zinc indicating iron and zinc deficiency, respectively. 26.7% of sample children were reported as not experiencing higher than 'average health status'. 18.8% were underweight according to the BMI-for-age criteria for school-age children set by the World Health Organization (WHO).[9] 10.4% were anemic[10] and 0.9% suffered from zinc deficiency.[11] These figures testify to malnutrition, especially the aspects of perceived health status and underweight.

Methodology

We employ the RDD framework to evaluate the impact of RCCs.[12] It is notable that there is no single criterion for obtaining a loan from RCCs, but the manager evaluates the risk for the intended rural households based on their demographic characteristics, living and production situation and then issues them certificates stating the maximum loan permitted. The actual participation and borrowing decisions are made on a voluntary basis by households within their loan limit. Therefore, we choose the FRD model specification within the RDD framework.

Specifically, given these features of RCCs, we first construct an exogenous assignment variable determining the treatment receipt. Using a standard probit model specification, we regress households' observed borrowing decision on a set of variables including factors that are considered by the RCCs managers when granting the loans and which influence households' demand for credit.[13] The predicted probability of borrowing for the household i ($\hat{p}_i$) is then calculated from the estimates. It can also be understood as the household's ability to borrow, considering the components of our independent variables in the probit regression. A household is assigned to receive the treatment (b_i) when the predicted probability of borrowing surpasses 50%:

$$b_i = \mathbf{1}(\hat{p}_i \geqslant 0.5) \quad (1)$$

It should be noted that having higher ability than 0.5 does not necessarily mean borrowing RCCs. Those households can choose whether to actually receive the treatment. In view of voluntary participation of RCCs, I use a 'fuzzy' regression-discontinuity (FRD) design (Lee and Lemieux, 2010) to model the treatment.[14]

Although some of the unobservables may correlate with the household's treatment status in the presence of self-selection of treatment, households are unable to control precisely the assignment variable $\hat{p}_i$.[15] Consequently, every household close to the cut-off would have similar chance of having their $\hat{p}_i$ higher or lower than 0.5, as if a 'local' randomized experiment. Without restricting self-section behavior, the causal impact of borrowing RCCs can be obtained by comparing the sub-sample just above (the treatment group) and just below (the control group) the threshold. According to Hahn et al. (2001), this treatment effect is defined as the difference in nutritional outcomes at the treatment threshold divided by the fraction of households induced to take up the treatment:

$$\begin{aligned} E[\tau_{FRD}|\hat{p}_i = 0.5] &= \lim_{e\downarrow 0} E[\tau_{FRD}|b_i(0.5+e) - b_i(0.5-e) = 1] \\ &= \frac{\hat{\alpha}_{y_+} - \hat{\alpha}_{y_-}}{\hat{\alpha}_{b_+} - \hat{\alpha}_{b_-}} = \frac{\lim_{\hat{p}\downarrow 0.5} E[y_i|\hat{p}_i] - \lim_{\hat{p}\uparrow 0.5} E[y_i|\hat{p}_i]}{\lim_{\hat{p}\downarrow 0.5} E[b_i|\hat{p}_i] - \lim_{\hat{p}\uparrow 0.5} E[b_i|\hat{p}_i]} \end{aligned} \quad (2)$$

where y_i is the child nutritional outcome. I employ two kinds of non-parametric approaches to estimate Eq. (2). First, Hahn et al. (2001) develop a 'local Wald' estimator calculating the four limits as $\lim_{p\downarrow 0.5} E[y_i|\hat{p}_i] = \frac{\sum_{i\in\Psi} y_i w_i}{\sum_{i\in\Psi} w_i}$, $\lim_{p\uparrow 0.5} E[y_i|\hat{p}_i] = \frac{\sum_{i\in\Psi} y_i(1-w_i)}{\sum_{i\in\Psi}(1-w_i)}$, $\lim_{p\downarrow 0.5} E[b_i|\hat{p}_i] = \frac{\sum_{i\in\Psi} b_i w_i}{\sum_{i\in\Psi} w_i}$ and $\lim_{p\uparrow 0.5} E[b_i|\hat{p}_i] = \frac{\sum_{i\in\Psi} b_i(1-w_i)}{\sum_{i\in\Psi}(1-w_i)}$ where the indicator variable $w_i = I\{0.5 \leqslant \hat{p}_i < 0.5 + h\}$ defines whether the observation

7 If counting those feeling 'the same as before', the figure will amount to 61.1%.

8 Strauss and Thomas (1998) raise the concern of possible large measurement errors in self-assessed health status. In order to attain robustness, BMI is used as an objective indicator as well. However, admittedly, BMI is only an indicator for 'temporary' nutritional status. A more informative measure for 'chronic' nutritional outcomes is child height, as suggested by Strauss and Thomas (1998). Given that our samples include older children only (12–20 years old), BMI-for-age is preferred to height z-scores.

9 This includes 4.5% severe thinness (BMI-for-age z-score < -3) and 14.3% thinness ($-3 \leqslant$ BMI-for-age z-score < -2). The method to calculate BMI-for-age z-scores can be found in WHO (2009) and the criteria of malnutrition are available at http://www.who.int/growthref/who2007_bmi_for_age/en/index.html (accessed 10 October 2012).

10 Haemoglobin is less than 120 g/L for children over 12 (WHO, 2008).

11 Hair zinc is less than 70 μg/g (Qin et al., 2009).

12 The regression-discontinuity design is selected because of four considerations. First, RDD has milder assumptions than other quasi-experimental tools and in particular can deal with the unobservables accounting for selection. Specifically, DID assumes that any changes in RCCs receipt do not correlate with changes in any omitted variables that affect child health outcomes. PSM also has strong identifying assumptions that the observables can fully explain households' selection into borrowing and there is no remaining selection on the unobservables. By contrast, RDD allows both observables and unobservables to affect households' borrowing decision and derives a local average treatment effect which has higher internal validation than DID and PSM (global) estimators. Second, the characteristics of RCCs make DID and PSM less appropriate than RDD. There is no single and strictly exogenous selection criterion to obtain RCCs in China as in Bangladesh. The manager of RCCs collects annually the household's various information, including income, consumption, assets, demographics and so on, to evaluate the maximum credit available to this household to borrow and then issues it a certificate stating this maximum credit. On receiving this certificate, the household is free to choose whether, when and how much to borrow within the credit limit. In this context, the selection criterion is multi-dimensional and partly hinges on the household's ability to borrow which could be unobserved. The RDD framework with our constructed assignment variable (the index of household ability to borrow) can better accommodate this situation. Third, I experimented with the DID-PSM estimation and the treatment effect is 0.73 with standard errors being 0.77 (p-value = 0.345). However, the DID-PSM estimator is less valid than the RDD estimator as the balancing test cannot be passed. The mean of some variables, in particular the child health outcomes, differs between the control and treated groups at 10% significance level even after matching. This invalidates the DID-PSM estimation. Fourth, I also implemented a standard IV approach and re-estimated both static and dynamic treatment effects (Columns 1–2 and 4–5 of Tables 3 and 4). According to van der Klaauw's (2008), I used the indicator b_i as the excluded instrument of the household's actual selection into borrowing RCCs. Results are shown in Table A.2 in Appendix. It can be seen that both static and dynamic treatment effects have the same signs as before, so the conclusion of FRD holds. However, the estimators lost statistical significance and F-stats suggest that the regressions for two nutritional indicators also lost overall significance. This is consistent with van der Klaauw's (2008) finding that IV estimation results in larger standard errors in impact evaluation than RDD.

13 The independent variables include household size, the number of family members currently not living in the household, the number of employed members, average educational level and age of all household members, household perception of its total income in the previous year (i.e., whether its income was enough to support daily life), the quality of housing in terms of log monetary values of the house, the ratio of irrigated land over the total farmland owned by the household, the wealth status of the household in the village, whether borrowing RCCs to pay school fees, and village dummies controlling for the unobservables. Estimation results are reported in Table A.1 in Appendix.

14 The FRD design means that the treatment is determined *partially* by whether the household's ability to borrow surpasses 50%. This fits the situation that various observed and unobserved factors not included in our model may jointly affect take-up of RCCs but our regression of estimating $\hat{p}_i$ cannot include all of those factors. This leads to imperfect compliance among the treated group in FRD, while the standard sharp RDD assumes perfect compliance for those above 50%. 'Identification problems in RDD' section will use evidence from GSCF to justify the use of FRD and the threshold 50%.

15 This argument will be tested formally in Identification problems in RDD' section.

Table 1
Descriptive statistics.

Variable	Mean	SD
Child health status (parent-reported health status from 1 (very poor) to 5 (very good))	0.028	0.998
BMI (body-mass index)	18.093	2.374
Ln(Hemoglobin-count (g/L))	4.904	0.127
Ln(Hair zinc (μg/g))	5.031	0.310
Microcredit in 2004 (borrow = 1)	0.352	0.478
Microcredit in 2000 (borrow = 1)	0.457	0.498
Age	15.087	1.159
Gender (boy = 1)	0.531	0.499
Ethnic minority (yes = 1)	0.982	0.134
Siblings' average health (averaged parent-reported health status from 1 (very poor) to 5 (very good) for the sample child's siblings)	4.062	0.910
Health at birth in the mother's eye (categories from 1 (lowest) to 5 (highest))	3.658	0.942
Water (categories from 1 (natural water in the field) to 4 (tab water in room))	1.965	1.104
Severity of food insecurity (average frequency of eating meat, aquatic product, eggs, fresh vegetables and fruit and drinking dairy product; there're three categories from 1 (seldom) to 3 (once a week or more) indicating times of eating in last month for each food variety; the quality of food is the average of the frequency across these varieties.)	1.776	0.483
Father's edu. in yrs.	6.818	3.962
Mother's edu. in yrs.	4.200	3.629
Ln(hh per capita wealth)	7.397	1.084
Hh social capital (whether at least one of the parents participates in community social groups in the village, yes = 1)	0.233	0.423
Hh migration (1 if any of the household member migrated to work in other places)	0.327	0.469
Women's empowerment in issues with children (1 if the husband decides, 2 if both decide, 3 if moth decides)	1.911	0.580
Ln(village average per capita net income)	7.170	0.883
Share of pop. having telephones in the village	0.102	0.074
Ratio of No. of medical staff over No. of village clinic	3.577	4.320
Distance to the nearest township hospital (km)	4.714	4.124
Quality of the nearest township hospital	2.151	0.628

Table 2
Non-parametric estimation of treatment effects on child nutrition.

	Child health status		BMI		Ln(Hemoglobin-count)		Ln(Hair zinc)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\hat{\alpha}_{y_+} - \hat{\alpha}_{y_-}$	0.107* (0.064)	0.167** (0.083)	0.332 (0.363)	0.301 (0.358)	0.011 (0.012)	0.015 (0.013)	0.034 (0.069)	0.076 (0.097)
$\hat{\alpha}_{b_+} - \hat{\alpha}_{b_-}$	0.064 (0.070)	0.070 (0.076)	0.072* (0.045)	0.110*** (0.024)	0.081 (0.058)	0.273*** (0.040)	0.060 (0.071)	0.201*** (0.045)
$\hat{\tau}_{FRD}$	1.669 (2.889)	2.386 (5.856)	4.633 (6.758)	2.736 (1.710)	0.140 (0.176)	0.054 (0.049)	0.577 (1.912)	0.361 (0.463)
Covariates		Yes		Yes		Yes		Yes

Note: Columns 1, 3, 5 and 7 are Hahn et al.'s (2001) local Wald estimators. Columns 2, 4, 6 and 8 are estimated by the triangle kernel with covariates listed in Tables 3 and 4.
*** 1% significance level.
** 5% significance level.
* 10% significance level.

lies barely above the cut-off; $\Psi = \{i | i \in (0.5 - h \leqslant \hat{p}_i < 0.5 + h)\}$ confines the sub-sample in the vicinity of the cut-off used for the estimation. The optimal bandwidth h is selected by Imbens and Kalyanaraman's (2009) procedure. Second, to achieve robustness, I also estimate the left- and right-hand sides limits relative to 0.5 of conditional expectations in Eq. (2) by using a triangle kernel regression which proves to have better properties at boundaries (Ludwig and Miller, 2010) and controlling for other covariates in the right-hand side of four limits.

One may be concerned that the information drawn from the sub-sample around the threshold could not fully inform the treatment effect for the study population. As a complement to the non-parametric approaches, I also implement two-step semi-parametric control-function estimation proposed by van der Klaauw (2008) to utilize richer information from the full sample. The first step estimates the probability of treatment receipt by a standard probit model specification:

$$E[b_i|\hat{p}_i] = \Pr(b_i = 1|\hat{p}_i) = \gamma_0 + \gamma_1 \cdot \mathbf{1}(\hat{p}_i \geqslant 0.5) + g(\hat{p}_i) \quad (3)$$

where γ_1 measures the discontinuity at the cut-off; $g(\hat{p}_i)$ is a quadratic piecewise function and parameterized by:

$$g(\hat{p}_i) = \lambda_0 + \lambda_1 \hat{p}_i + \lambda_2 \hat{p}_i^2 + \{\lambda_3(\hat{p}_i - 0.5) + \lambda_3(\hat{p}_i - 0.5)^2\} \cdot \mathbf{1}(\hat{p}_i \geqslant 0.5) \quad (4)$$

In the second step, based on the estimate of $E[b_i|\hat{p}_i]$ in the control function Eq. (3), a reduced-form outcome equation is written as:

$$y_i = \beta_0 + \tau_0 E[b_i|\hat{p}_i] + X_i\beta_1 + X_h\beta_2 + X_v\beta_3 + k(\hat{p}_i) + \varepsilon_i \quad (5)$$

where ε_i denotes the disturbance following the *i.i.d.* normal distribution; X_i, X_h, and X_v control for characteristics of the sample child, his/her family and the village respectively; $k(\hat{p}_i)$ captures directly the effect of household ability to borrow and ought to be a smooth and continuous function to insure that, in the absence of the treatment, nutritional outcomes are a smooth function of the ability to borrow and hence, differential nutritional outcomes are the only source of discontinuity around the cut-off. Empirically, we follow van der Klaauw (2008) and let $k(\hat{p}_i)$ take a semi-parametric form, $k(\hat{p}_i) \approx \sum_{j=1}^{J} \eta_j \hat{p}_i^j$, to accommodate non-linearity and reduce misspecification, where the power J is left determined by generalized cross-validation of the data.$\hat{\tau}_0$ reflects the causal effect of borrowing on nutrition as defined in Eq. (2). Among control variables, X_i incorporates child demographic characteristics and in particular his/her siblings' average self-reported health status and his/her own health status at birth in the mother's eye. The former could not only reflect allocation of limited resources among children within the family, but also help identify the impact of parents' educational attainments in X_h, because an average level of health status for children could reflect parents' attitudes towards investment in child health and nutrition which usually correlate positively with their educational levels. The child's health at birth captures the 'initial health status' that might have persistent impact on the child's growth. This would help disentangle the impact of microcredit $\hat{\tau}_0$.The specification of Eq. (5) also justifies the theoretical issues. There are a number of ways that the presence of microfinance institutions affects child nutritional outcomes as recently summarized by DeLoach and Lamanna (2011). According to them, access to microfinance can improve many-faceted individual and community development which in turn would contribute to children's better nutrition. I control explicitly for these developments in order to disentangle the impact of borrowing *per se*. Specifically, X_h also includes the quality of water accessed by the household, the severity of food insecurity, household wealth, women's empowerment in terms of mother's power in decision-making on children's issues, household social capital measured as whether the household participates in organizations in the village and whether at least one of the family members is in out-migration. These factors are selected according to findings in the previous literature and can affect the manner in which inputs allowed by more credits to produce nutrition are used. Specifically, the quality of water and food insecurity in terms of skipped meals because of financial restrictions could affect household health and nutritional outcomes (Kosec, 2012; Reis, 2012). It is also necessary to control for women's empowerment and household social capital should we aim to estimate the impact of microcredit more precisely. Imai et al. (2012) find that women's higher relative bargaining power over men's alleviates underweight and stunting for Indian children under age 3 both temporarily and long-term. Obtaining microcredit could entail households' participation in community organizations that would facilitate development of social capital. This would in turn (1) make them better able to cope with adverse shocks and hence suffer less from child malnutrition (Carter and Maluccio, 2003); and (2) allow parents to circulate information about communicable diseases (Nobles and Frankenberg, 2009) and to increase maternal knowledge which is catalytic for augmenting the impact of income on child nutritional intake (Christaensen and Alderman, 2004). Last, household out-migration can also be important for child nutrition with increasingly large-scale migration of workers from rural to urban China, as indicated by studies by de Brauw and Mu (2011).At the village level X_v, I consider the local economic development in terms of village mean per capita income and infrastructure measured by the share of telephone holders. Given the fact that, in the dataset, about 67% of villages needing to consult doctors opted to visit village clinics as their first choice, followed by 15% choosing township hospitals, I also take health facilities into account by including the quality of village clinics, the distance from the village to the nearest township hospital and its perceived quality. Based on the existing literature (e.g., Imai et al., 2012), it is expected that a higher level of economic development and better health facilities would improve child health and nutritional status.It is notable that $\hat{\tau}_0$ is virtually an average treatment effect of RCCs, including both the immediate consequences of current borrowing behavior in 2004 and cumulative influence of the household's previous participation in RCCs. To distinguish between the two, as van der Klaauw (2008), I further include lagged borrowing behavior in Eq. (5):

$$y_i = \beta_0 + \tau_0 E[b_i|\hat{p}_i] + \tau_1 m_{i,2000} + X_i\beta_1 + X_h\beta_2 + X_v\beta_3 + k(\hat{p}_i) + \varepsilon_i \quad (6)$$

where $m_{i,2000} = 1$ if the household borrowed RCCs in 2000 and 0 otherwise. Therefore, $\hat{\tau}_1$ captures the long-term effect of one-time borrowing behavior; and $\hat{\tau}_0$ picks up the contemporaneous effect only.

Results and discussion

Identification problems in RDD

It is necessary to justify the use of RDD before proceeding to interpretation of the results, as suggested by Lee and Lemieux (2010). The crucial assumption of quasi-experimental environment for the sub-group households around the cut-off requires that households cannot 'manipulate' their ability to borrow ($\hat{p}_i$) by changing their characteristics to deliberately stay below or above the treatment threshold. That is, the assignment to treatment at the cut-off should be random for households. I inspected this by plotting the distribution of $\hat{p}_i$. Fig. 1 does not reveal a significant jump in the distribution of $\hat{p}_i$ at the cut-off: the McCrary's (2008) density test cannot reject the null hypothesis of continuity in distribution. The estimated log difference in height of the bins around the cut-off is 0.043 with the standard error of 0.183. According to him, this supports the assumption of a quasi-experiment in RDD and furnishes the causal inferences.

Moreover, I show the appropriateness of using an FRD rather than other forms of RDD to model the treatment. For one thing, Fig. 2 shows that the actual probability of borrowing increases with households' higher ability to borrow. Households' borrowing behavior changes significantly at the cut-off: the probability of actual borrowing jumps by about 10 percentage points once above the threshold with different levels of statistical significance (Row 2 of Table 2). Note that the probability of treatment receipt does not reach 100% as soon as the household crosses the cut-off, but only gradually converges to 100% in the treatment group. This fits the fact that the treatment recipiency is not determined solely by our constructed household ability, nor is it mandatory to borrow RCCs for the treatment group. FRD can match this imperfect compliance for the treatment group. The discontinuity in household borrowing behavior at the threshold also assures a non-zero denominator in Eq. (2).

For another, on observing the discontinuity in borrowing decisions at the cut-off, one should identify discontinuity in outcome variables as well; otherwise borrowing behavior would not bring about any difference to child nutrition. I inspected this in Fig. 3 where child nutrition is drawn against the assignment variable. There is a jump in children's four nutritional outcomes when the household receives RCCs, i.e., the ability to borrow crosses the cut-off. This implies that the treatment recipiency of RCCs could be a reason for changes in child nutrition, but the extent of causal impact and statistical significance may vary across different measures.

One may be concerned with the arbitrariness attached to the choice of the cut-off. The selection of 50% is inspired by the literature on vulnerability to poverty, which classifies a household as vulnerable if its predicted probability to be poor is higher than 50% (e.g., Christaensen and Subbarao, 2005). Moreover, comparing the estimated household vulnerability profile and actual observed poverty in rural China, Zhang and Wan (2009) demonstrate that setting the vulnerability line at 50% can best predict who are actually poor. Essentially, our assignment variable is a predicted probability of borrowing RCCs. Thus, it could be reasonable to assume that a household begins to borrow if its propensity to borrow is beyond 50%. However, the problem of arbitrariness does exist and may affect the estimated treatment effects. To check the robustness of our estimates, I also experimented with other levels of threshold determining treatment recipiency. By re-plotting Fig. 2 for various new thresholds, no discontinuity can be found with thresholds lower than 50% or higher than 75%. If using thresholds lying between 50% and 75%, the jump in actual borrowing probability at the cut-off, i.e., the denominator of Eq. (2), will first increase to about 20% at the threshold of 65% and then shrink gradually to zero after the threshold of 75%. This is predictable: more households would be assigned RCCs under higher thresholds, but less significant changes in households' borrowing behavior would occur at some very high thresholds, because many of them may have already borrowed RCCs. At the same time, the estimated impact of borrowing on child nutritional outcomes without considering partial compliance, i.e., the numerator of Eq. (2), also becomes larger for all four nutritional measures, possibly because those more able households not only could afford but also would like to devote more effort to improving children's nutrition. Altogether, the treatment effects of RCCs become larger at thresholds higher than 50%. I calculated the marginal threshold treatment effect (MTTE) according to Dong and Lewbel (2012). A marginal increase in the threshold 0.5, say from 0.5 to 0.51, would bring 0.221 increase in child health outcomes. In this sense, our estimators based on the cut-off at 50% may provide a lower bound of impact of microcredit on child nutrition.

Discussion of estimation results

Table 2 presents the non-parametric estimation. RCCs seem to enhance parent-reported child health status by 1.67 categories and increase BMI by 4.6 (Columns 1 and 3 of Table 2). Regarding micronutrients, borrowing RCCs could increase hemoglobin by 14% and zinc intake by 57.7% (Columns 5 and 7 of Table 2). Nevertheless, none of these effects shows statistical significance as non-parametric estimation tends to generate larger standard errors. Moreover, after controlling for other covariates listed in Tables 3 and 4 and drawing upon triangle kernel to estimate Eq. (2), there is no significant difference in non-parametric estimators by comparing Columns 1–2, 3–4 and 5–6 of Table 2. This not only adds robustness to non-parametric estimators, but also reaffirms that the source of discontinuity in child nutritional outcomes lies mainly in households' borrowing behavior rather than other factors.

I also check the possible sensitivity of estimation results to the choice of bandwidth. Although I used the 'optimal' bandwidth derived by Imbens and Kalyanaraman (2009), the smoothness of child nutrition as a function of household ability to borrow in Fig. 3 and thus the discontinuity of the outcome variables at the cut-off are contingent to the value of bandwidth. Fig. 4 depicts the local Wald estimators (without controlling for other covariates) for different bandwidths in the range of 50% and 150% of the optimal. It can be seen that, by and large, the above findings on the signs and magnitude of treatment effects of RCCs still hold. Again, large standard errors in non-parametric estimation lead to statistical insignificance.

Drawing upon the full sample, I further estimated Eqs. (5), (6) by the two-step control function with the power $J = 3$ determined by the dataset. As shown in Tables 3 and 4, the signs of the treatment effects remain consistent with non-parametric results, while statistical significance appears in association with parent-reported health status, BMI and zinc intake. Specifically, obtaining credit from RCCs substantially improves the parent-reported health status by 1.427 categories (Columns 1 of Table 3), which, however, is likely to be contemporaneous with borrowing in 2004. After including households' lagged borrowing behavior in 2000 (Column

Table 3
Semi-parametric estimation of treatment effects on child nutrition (anthropometry).

Independent variable	Child health status			BMI		
	(1)	(2)	(3)	(4)	(5)	(6)
Treatment effects						
$\hat{\tau}_0$	1.427** (0.733)	1.664* (0.963)	1.703* (0.985)	4.879* (2.736)	4.645 (4.326)	12.018** (4.688)
$\hat{\tau}_1$		−0.297 (0.202)			−0.848 (0.811)	
Child						
Age	−0.026 (0.046)	−0.031 (0.048)	−0.017 (0.040)	0.681*** (0.066)	0.686*** (0.064)	0.588*** (0.093)
Gender	0.047 (0.057)	0.040 (0.058)	0.040 (0.058)	−0.764*** (0.139)	−0.747*** (0.132)	−0.505*** (0.175)
Ethnic minority	−0.014 (0.337)	−0.025 (0.340)	−0.057 (0.352)	0.112 (0.739)	0.121 (0.745)	0.265 (0.722)
Siblings' average health	0.588*** (0.071)	0.606*** (0.077)	0.566*** (0.059)	0.082 (0.096)	0.092 (0.102)	0.028 (0.119)
Health at birth	0.046 (0.038)	0.043 (0.038)	0.091* (0.049)	0.198** (0.079)	0.194** (0.080)	0.136 (0.113)
Household						
Water	0.057* (0.033)	0.057* (0.033)	0.060* (0.032)	0.130* (0.076)	0.139* (0.073)	0.130 (0.091)
Severity of food insecurity	0.115 (0.102)	0.126 (0.105)	0.158 (0.119)	0.334 (0.240)	0.272 (0.198)	0.136 (0.192)
Father's edu.	0.004 (0.010)	0.006 (0.011)	0.004 (0.010)	0.054*** (0.018)	0.051 (0.019)	0.073*** (0.028)
Mother's edu.	0.015* (0.008)	0.014* (0.008)	0.006 (0.010)	0.010 (0.025)	0.007 (0.024)	0.047 (0.029)
Ln(hh per capita wealth)	0.069 (0.069)	0.070 (0.070)	0.055 (0.065)	0.230 (0.236)	0.216 (0.223)	0.649*** (0.241)
Hh social capital	0.012 (0.098)	0.042 (0.111)	0.043 (0.111)	0.276 (0.228)	0.291 (0.239)	0.647* (0.331)
Women's empowerment	0.002 (0.051)	−0.016 (0.055)	0.015 (0.049)	0.0001 (0.142)	−0.018 (0.154)	−0.117 (0.151)
Hh migration	−0.148** (0.076)	−0.142* (0.076)	−0.229** (0.104)	−0.317 (0.300)	−0.287 (0.275)	−0.815*** (0.267)
Village						
Ln(village average per capita net income)	0.197*** (0.063)	0.211*** (0.066)	0.152** (0.063)	0.014 (0.138)	−0.003 (0.145)	−0.132 (0.160)
Share of villagers using telephones	−1.015 (0.816)	−1.203 (0.819)	−1.399* (0.847)	3.608** (1.697)	3.281** (1.574)	6.435*** (2.090)
Ratio of no. of medical staff over no. of village clinics	0.011 (0.008)	0.015 (0.009)	0.003 (0.008)	0.001 (0.019)	0.003 (0.018)	−0.051 (0.032)
Distance to the nearest township hospital (km)	0.009 (0.010)	0.014 (0.012)	0.004 (0.009)	0.028 (0.021)	0.030 (0.022)	0.022 (0.028)
Town						
Quality of township hospitals	0.084 (0.058)	0.082 (0.058)	0.107* (0.060)	0.011 (0.130)	0.023 (0.134)	0.111 (0.170)
County dummies	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	1454	1454	643	1331	1331	765
R^2	0.421	0.421	0.421	0.208	0.208	0.242

Note: Standard errors in parentheses are clustered by the household ability to borrow in order to mitigate possible misspecification problems, as suggested by Lee and Card (2008).
*** 1% significance level in turn.
** 5% significance level in turn.
* 10% significance level in turn.

Table 4
Semi-parametric estimation of treatment effects on child nutrition (micronutrient).

Independent variable	Ln(Hemoglobin-count)			Ln(Hair zinc)		
	(1)	(2)	(3)	(4)	(5)	(6)
Treatment effects						
$\hat{\tau}_0$	0.141 (0.201)	0.135 (0.191)	0.639*** (0.229)	0.509 (0.655)	0.485 (0.590)	0.983* (0.600)
$\hat{\tau}_1$		−0.021 (0.036)			0.115 (0.115)	
Child						
Age	0.001 (0.003)	0.001 (0.003)	−0.007 (0.004)	0.023** (0.010)	0.023** (0.010)	0.011 (0.019)
Gender	0.004 (0.007)	0.005 (0.007)	−0.011 (0.011)	−0.030* (0.018)	−0.032* (0.018)	−0.069** (0.030)
Ethnic minority	0.049* (0.029)	0.049* (0.029)	−0.005 (0.035)	−0.127 (0.082)	−0.129 (0.083)	−0.210 (0.146)
Siblings' average health	−0.007 (0.005)	−0.006 (0.005)	−0.0001 (0.006)	−0.007 (0.012)	−0.008 (0.013)	0.040* (0.023)
Health at birth	−0.004 (0.004)	−0.004 (0.004)	−0.016** (0.005)	−0.004 (0.009)	−0.004 (0.010)	−0.012 (0.020)
Household						
Water	−0.0003 (0.004)	−0.0001 (0.004)	0.004 (0.005)	0.029** (0.014)	0.028** (0.013)	0.001 (0.018)
Severity of food insecurity	0.014 (0.011)	0.013 (0.010)	0.015 (0.014)	0.030 (0.032)	0.038 (0.027)	0.107* (0.056)
Father's edu.	0.001 (0.001)	0.001 (0.001)	−0.002 (0.001)	0.002 (0.003)	0.002 (0.003)	−0.004 (0.005)
Mother's edu.	0.001 (0.001)	0.001 (0.001)	0.005** (0.002)	0.0001 (0.003)	0.0003 (0.003)	0.0004 (0.005)
Ln(hh per capita wealth)	−0.006 (0.012)	−0.006 (0.012)	−0.024* (0.013)	0.022 (0.031)	0.021 (0.029)	−0.006 (0.029)
HH Social capital	0.002 (0.010)	0.002 (0.011)	0.027* (0.016)	0.010 (0.033)	0.007 (0.034)	0.104** (0.051)
Women's empowerment	−0.001 (0.009)	−0.001 (0.009)	0.003 (0.011)	0.013 (0.021)	0.015 (0.023)	−0.023 (0.026)
HH Migration	−0.013 (0.015)	−0.012 (0.014)	−0.045*** (0.016)	0.040 (0.044)	0.036 (0.041)	−0.021 (0.052)
Village						
Ln(village average per capita net income)	−0.007 (0.006)	−0.007 (0.006)	−0.014 (0.009)	−0.016 (0.017)	−0.014 (0.018)	−0.066** (0.033)
Share of villagers using telephones	−0.016 (0.074)	−0.025 (0.069)	0.129 (0.106)	0.442** (0.190)	0.484*** (0.175)	−0.015 (0.461)
Ratio of no. of medical staff over no. of village clinic	−0.002 (0.001)	−0.002* (0.001)	−0.005*** (0.002)	−0.001 (0.002)	−0.001 (0.002)	−0.001 (0.004)
Distance to the nearest township hospital (km)	−0.0004 (0.001)	−0.0003 (0.001)	−0.00004 (0.002)	0.001 (0.003)	0.001 (0.003)	0.007 (0.006)
Town						
Quality of township hospitals	0.011 (0.008)	0.012 (0.008)	0.019 (0.015)	−0.031 (0.020)	−0.033 (0.021)	0.060 (0.041)
County dummies	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	1322	1322	762	1275	1275	358
R^2	0.041	0.041	0.062	0.081	0.082	0.111

Note: Standard errors in parentheses are clustered by the household ability to borrow in order to mitigate possible misspecification problems, as suggested by Lee and Card (2008).
*** 1%, significance level in turn.
** 5% significance level in turn.
* 10% significance level in turn.

2 of Table 3), $\hat{\tau}_0$ is still statistically positive, but its magnitude (1.664) turns out to be 16.6% larger than that of Column 1. $\hat{\tau}_1$ is negative but statistically insignificant. These indicate that the benefit of RCCs on child nutrition is rather temporal and would dissipate quickly with elapsed borrowing over time. A similar finding also holds for BMI. As shown in Columns 4 and 5 of Table 3, children in borrowing families tend to experience 4.879 higher height-standardized weight than those living in non-borrowing families, whereas insignificant $\hat{\tau}_1$ points out no long-run accrued benefit. I further concentrate on the 'new borrowers' only by re-estimating Column 3 for sub-samples who did not borrow in 2000. We can see from Columns 3 and 6 of Table 3 that when the household borrowed RCCs in 2004, $\hat{\tau}_0$ rises dramatically in all cases compared with that of Columns 2 and 5 and is still statistically significant. This reaffirms the role of RCCs as short-term palliative in combating child malnutrition.

For the micronutrient status, Columns 1–2 and 4–5 of Table 4 suggest no warrant for raising the hemoglobin or zinc level regardless of time span. However, if focusing on new borrowers only, we find that RCCs increased hemoglobin by 64% and almost doubled the amount of hair zinc immediately after borrowing (Columns 3 and 6 of Table 4).[16]

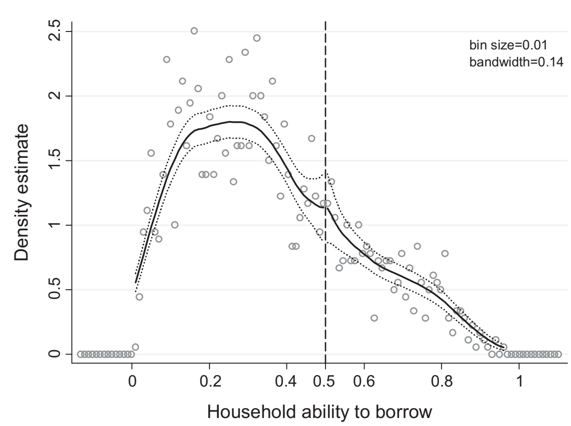

Fig. 1. Distribution of household ability to borrow in 2004. Note: The circle represents the average of household ability to borrow in each bin for the histogram. The optimal bin size and bandwidth are selected by McCrary's (2008) procedure. The solid line is the smoothed bin midpoints by local linear smoothing regressions.

One may be concerned with the possibility that the favorable impact of microcredit on older children's nutrition might have been overshadowed by the positive relationship between microcredit and household income in rural China as reviewed in 'Introduction' section. I inspected this by inserting logarithmic household per capita income as an additional regressor to all columns in Tables 3 and 4. The estimated coefficients of income are insignificant in all cases. The signs of $\hat{\tau}_0$ and $\hat{\tau}_1$ remain consistent as before and the magnitude varies only fractionally.[17]

[16] Re-estimating Columns 1–2 and 4–5 of Tables 3 and 4 by OLS without dealing with endogeneity yields 15–30% smaller estimated treatment effects. Statistical significance attached to anthropometric indicators disappears, while the contemporaneous treatment effects on two micronutrient indicators become significant. These indicate that endogeneity in borrowing not only biases the impact of RCCs downward, but could also affect qualitatively the main conclusion. Estimated coefficients of other covariates generally hold, while a few lose statistical significance.

[17] Results are not reported given limited space, but are available from the author.

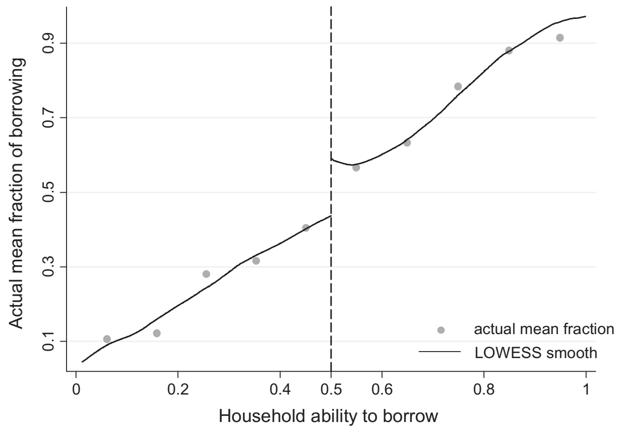

Fig. 2. Household actual borrowing behavior as a function of its ability to borrow. Note: Children are grouped into 10 bins left and right to the cut-off, respectively. The dot is the cell mean of the indicator for whether the household actually borrows RCCs, which reflects the actual probability of borrowing. The solid line represents predicted probability of borrowing from LOWESS smoothing of those actual probabilities.

It is worth investigating further why borrowing has no causal impact on child nutrition. As mentioned in Section 1, borrowing may improve child nutrition by increasing the level of food consumption and/or help smooth food consumption. To test for the former channel, I regressed log food consumption in 2004 on various factors such as borrowing credit from RCCs in 2004 and 2000 respectively, household and child characteristics, parents' education, women's bargaining power in children's issues, whether seeking advice from doctors when the child is ill, health facilities in the village, and village dummies. The estimated coefficient of borrowing in 2004 is 0.371 and statistically insignificant with p-value of t-statistics being 0.608. Neither is the estimated coefficient of borrowing in 2000 statistically significant. I also experimented with log health expenditure, the proportion of food consumption and health expenditure in the household's total consumption expenditure as the dependent variable. The estimated coefficients are all statistically insignificant. Overall, the evidence denies the first channel that borrowing increases the quantity of food consumption and in turn improves child nutrition and health. To test for the latter channel, I employed Islam and Maitra's (2012) model specification and regressed volatility of food consumption in terms of the log difference of household food consumption between 2000 and 2004 on control variables mentioned above. The impact of borrowing behavior in 2004 reduces substantially the volatility of food consumption with the estimated coefficient being −1.156 and significant at 10% level. However, previous borrowing in 2000 does not have a statistically significant impact on reducing volatility of food consumption. I also experimented with the volatility of health

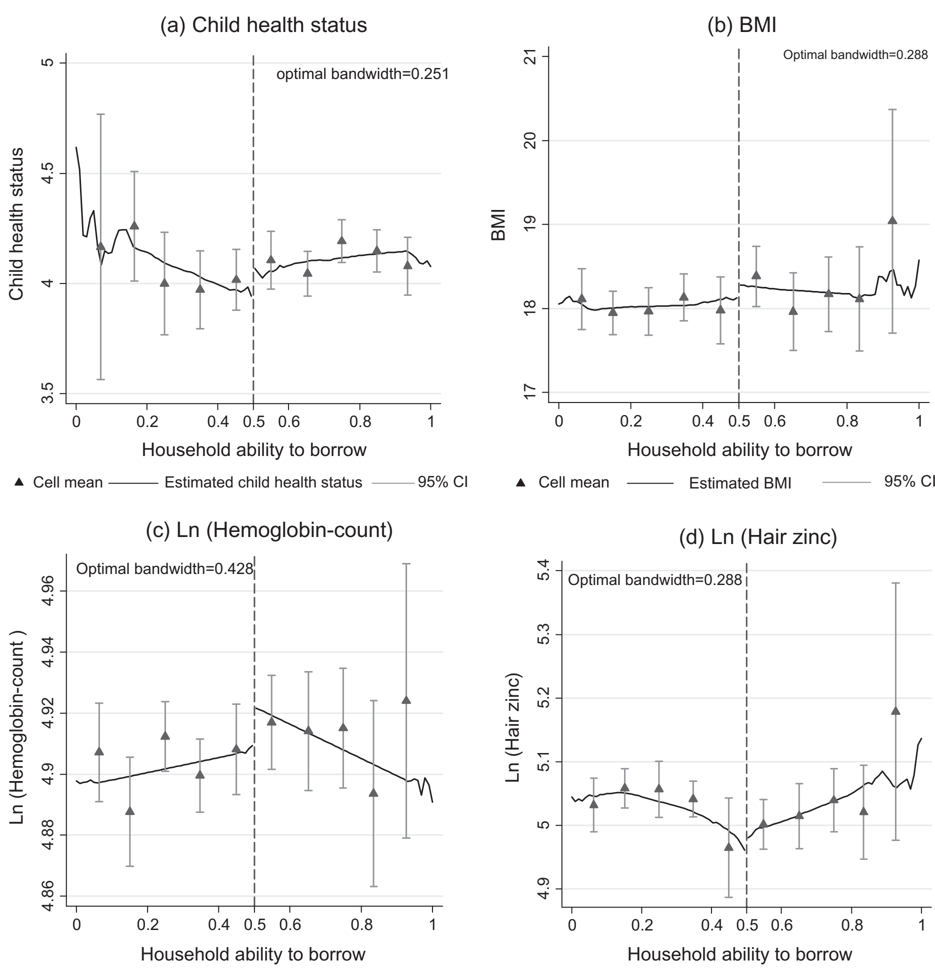

Fig. 3. Child nutrition as a function of household ability to borrow. Note: Households are grouped into five bins left and right to the cut-off, respectively. The triangle measures the average child nutrition indicators for those falling in the same bin. The solid line is the kernel-weighted linear regression of the cell averages.

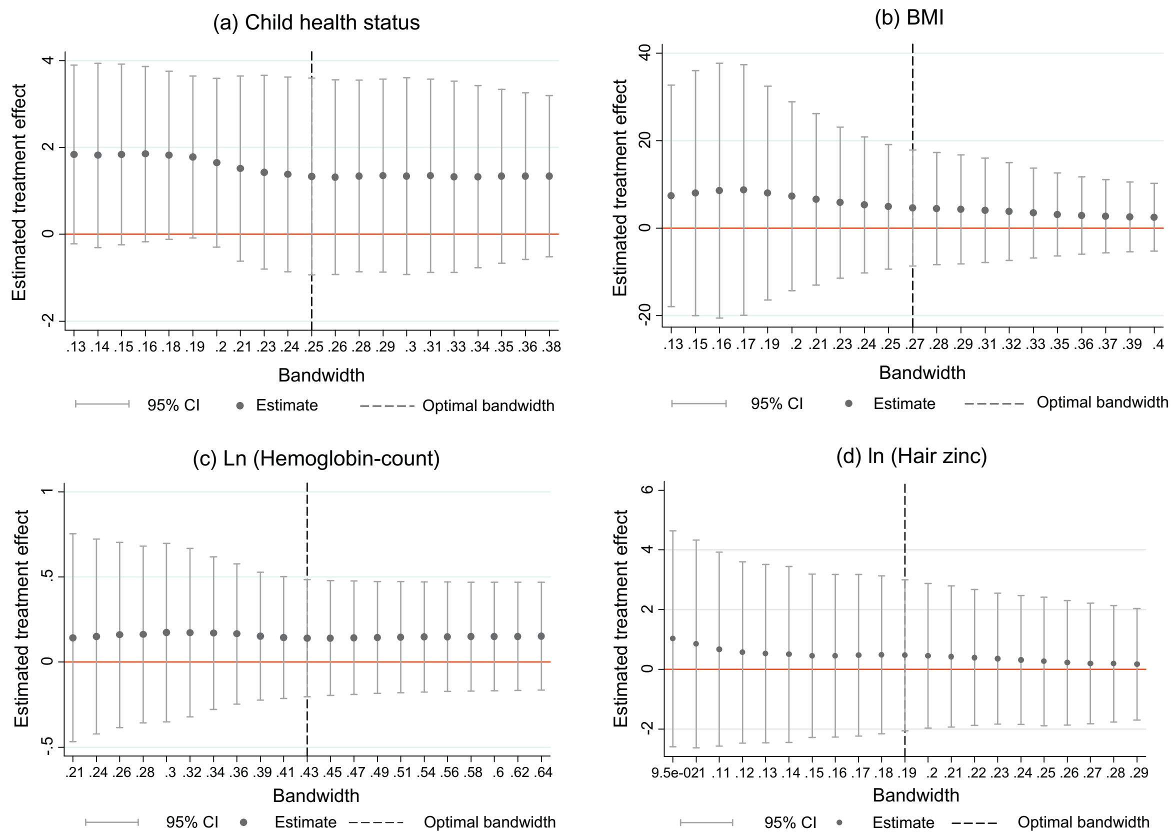

Fig. 4. Sensitivity of the estimates to bandwidth.

expenditure for the sample child calculated in the same way as food volatility. The estimated coefficient of borrowing in 2004 is −8.068 with 1% significance level, while the impact of borrowing in 2000 is insignificant. These findings support the second channel that borrowing improves child health by helping smooth food consumption, and are also consistent with the results of FRD in Tables 3 and 4 that microcredit has no prolonged causal impact (insignificant $\hat{\tau}_1$).

The comparison of static and dynamic causal impacts of RCCs prompts the question as to why microcredit exhibits distinct influence on child nutrition over time. A possible reason may be found in parents' behavior and attitudes towards children after borrowing RCCs. As the negative side of obtaining loans, borrowers would be involved more in expanded production which would require more child labor and less care for children. The data suggest that, although borrowing households in 2000 experienced 4.2 percentage points higher annual growth rate of household per capita income than initial non-borrowers over the period 2000–2004, 23.6% of initial borrowing parents said (in the 2004 survey) that they asked children to go to work as opposed to 18.8% for initial non-borrowing parents. Child labor in 2004 in terms of hours per week doing housework, agricultural production, household business and other forms of income generation activities was 16.3% higher in initial borrowing families than in non-borrowing families. Parents also reduced the time allocated to children after borrowing RCCs. The share of initial borrowers reporting 'no time' to take their ill children to hospitals in 2004 was 4.7 times as high as that of initial non-borrowers. It therefore could be inferred that RCCs might result in heavier work-load and less child care in at least the medium run, although helping smooth consumption at the same time. These two contradicting forces would have an ambiguous 'net' long-term consequence on child health and nutrition.

The remainder of this sub-section proceeds to discuss other attributes to child nutrition. Weight and zinc-intake increase as the child grows up, while other nutrition indicators are insensitive to age. There is an evident gender difference in weight and zinc: girls suffer from less underweight than boys and absorb more zinc. This might be at odds with conventional impressions of male advantage in many developing countries. However, using China Health and Nutrition Surveys (CHNS), Mangyo (2008) shows that, in rural China, nutritional intake is more elastic for males than for females and attributes this to faster decreases in males' marginal utility and production than in females as income and food resources available to households increase. Ethnic minorities show higher hemoglobin than subjects of Han nationality, which has also been found in CHNS (Li et al., 2013). Siblings' average perceived health conditions relate strongly and positively to the sample child's perceived health condition (Columns 1–3 of Table 3), but negatively to zinc intake. The former finding is predictable as siblings' better health in general could reflect parents' attitudes and care of children's health. The latter observation may imply competition among children for nutrition-rich food in poor rural areas. Children's health status at birth can persist into their adolescent weight (Columns 4–5 of Table 3), which has also been found in CHNS (Smith et al., 2012). It is nevertheless unlikely to influence perceived health status or acquisition of micronutrients.

At the household level, improved water access is positively associated with most subjective and objective indicators except the hemoglobin level. The direct impact of more nutrient-dense food is positive for all health and nutrition indicators, but is largely insignificant except for zinc in the short-term. Mothers' education raises child health status with the marginal effect of one more year in education being 0.014–0.015 (Columns 1–2 of Table 3). Moreover, although the magnitude of the impact of mothers' education turns out to be larger than that of fathers' education, the difference is not statistically significant – the *F*-statistics for equal estimated coefficients between mothers and fathers are 0.58–0.36 in Columns 1–2 with *p*-values being 0.447 and 0.550 respectively. In comparison, Columns 4–5 of Table 3 and Column 3 of Table 4 reveal some benefit of educating fathers on reducing underweight rather than mothers, with the marginal effect of one year in education being 0.05–0.07 higher BMI and it is significantly higher than that of mothers' education (*F*-statistic 3.35 with *p*-value of 0.067). Women's empowerment does not suggest consistent and significant influence on various child health indicators.[18]

We also observe negative impact of wealth on child nutrition in terms of hemoglobin-count and zinc and a 1% increase in wealth may even reduce hemoglobin by 0.024% in the short-term (Column 3 of Table 4). This echoes the recent review that income growth is not a sufficient condition for reduction of child malnutrition (Alderman, 2012), especially in developing countries. In the Chinese context, households' wealth has a detrimental effect on nutrition through worse quality of diet – it shifts away from high-carbohydrate foods toward high-fat, high-energy foods that cause diet-related non-communicable diseases (Du et al., 2004). Li et al. (2013) also find a negative link between wealth and hemoglobin in CHNS. As for anthropometry, a 1% increase in wealth would nevertheless raise the young adult's BMI by 0.645 in the short-term (Column 6 of Table 3). Similarly, 1% increase in household income raises height-for-age *z*-scores by 0.541 in CHNS over the period 1991–2006 (Chen et al., 2010). By using the same dataset GSCF, Hannum et al. (2012) also find that being in the wealthiest quintile in 2004 is associated with additional 0.271 height-for-age *z*-scores and 0.279 weight-for-age *z*-scores. It seems that wealth relates positively to anthropometric parameters for older children, but this cannot underlie an interpretation of strengthened nutritional status.

Adolescents from households with more social capital appear to be healthier. Participation in village organizations brings about 0.647 higher BMI and increases hemoglobin-count and hair zinc by 2.7% and 10.4%, respectively. However, these substantial benefits exist only in the very short-term for new borrowers.

In line with de Brauw and Mu (2011), the estimation documents a negative consequence of household members' migration on the left-behind older children's parent-perceived health and weight (Table 3), despite more income and consumption resulting from working in cities. The above authors attribute this to the fact that the remaining older children take on more household chores and this can also be found in our dataset. Older children living in households with migrating members spent 38.3% more time per week (10.97 h) doing household agricultural production, business and housework than those in families without outmigration (7.93 h per week). Migration is less likely to affect nutrition.

At the aggregated level, access to health facilities (village clinics and township hospitals which bear the brunt of 'primary care') do not contribute to improvement in child nutrition, except in the short-term for parent-reported child health. The distance between households and the nearest township hospital is irrelevant to child health and nutrition and a higher ratio of medical staff over the number of village clinics even relates negatively to hemoglobin. The insignificant or even negative role of health facilities might be caused by low efficiency in Chinese hospitals (Ng, 2011). Moreover, these insignificant estimators raise concern about the effectiveness of the government's recent campaign of enlarging investment in medical infrastructure announced in 2009 and efforts to enhance the role of primary care in providing health services for rural households.

A one percentage point increase in the village income raises parent-reported child health status by 0.152–0.211 categories. Interestingly, it does not affect the other three objective nutritional measures which relate to village basic infrastructure represented by the proportion of villagers using telephones. A one percentage point increase in telephone users in the village leads to 3.281–6.435 higher BMI (Columns 4–5 of Table 3) and 0.442–0.448% increase in hair zinc (Columns 4–5 of Table 4). This implies that the relationship between economic growth and better child nutritional outcomes which has been widely observed in the existing literature for developed countries should be interpreted with caution in the Chinese context. That is, local economic development in terms of income growth alone would be less likely to propel other aspects of human well-being such as child nutrition. This, in contrast, may need more comprehensive development like improved infrastructure.

Conclusion

Using rural household data for north-western China (which lags behind other areas), this article has discussed the promise of formal microcredit initiatives as a means of counteracting malnutrition in this poor rural part of China. Overall, results partially support the conjecture in 'Introduction' section. There are indications of substantial causal impact of RCCs on older children's nutrition. A causal relationship runs from borrowing RCCs to both improved anthropometry and micronutrient take-up in the short-term but ceases to persist into the longer-term.

Our findings carry important implications for policy makers. Current (local) interventions to prevent child under-nutrition in rural China include, for example, providing one egg a day per child in Shaanxi and Ningxia provinces since 2009/2010 by provincial governments. The local leaders began to recognize the importance of treating micronutrient deficiencies after the publication of a series of reports (e.g., Luo et al., 2012). However, Kleiman-Weiner et al. (2013) find that giving children chewable vitamins rather than 'one-egg-a-day' policy has significant impact on child nutrition but local governments are reluctant to hand out vitamins. Another attempt to reduce incidence of under-nutrition among infants and children is the Rural Primary Health Care Project (RPHC) initiated in 10 western provinces in 2001 by the Ministry of Education of China and co-funded by the United Nations Children's Fund (UNICEF). It consists of a three-level 'county-township-village' health-care network to provide nutritional consultation, monitor children under 3 years old and offer those with low birth weight feeding and dietary instruction through the above three-level health facilities. An evaluation of RPHC over the period 2001–2005 suggests that children's stunting rate reduced by 0.4% with statistical significance, but the heightened concentration index by 60% as poor families tended to live in under-served areas and were

[18] I implemented the Durbin–Wu–Hausman test to investigate the potential endogeneity problem in women's empowerment. Specifically, I followed Imai et al. (2012) and instrumented women's empowerment by the proportional difference of the father's age and the mother's, as they noted that older fathers may be more powerful in bargaining but may not affect directly children's nutritional status. The null hypothesis of exogeneity cannot be rejected at all three conventional statistical levels. The *p*-values range from 0.367 to 0.993 for different columns of Tables 3 and 4.

less likely to be covered than their wealthier counterparts (Pei et al., 2012).

It would be very difficult and time consuming to tackle the problem of child under-nutrition if one relied exclusively on persuading the government to adopt some supplement interventions that appear to be effective. The government may nevertheless decide to try other means for some unknown considerations as in the above 'one egg' case. Moreover, government-led policies like RPHC aiming at reducing under-nutrition rates may differ from those taking care of inequality between the poor and the wealthier, because they fail to deal with root causes of under-nutrition (Pei et al., 2012). It would be better for policy makers to put the problem in a broader context. Microcredit could be an effective approach, especially in an area where both poverty and inequality prevail: borrowing households become wealthier and less credit-constrained, so they are more able to afford a nutrient-rich diet and smooth consumption. As a result, young adults' under-nutrition would be alleviated.

It is noteworthy that the analysis in this article does not mean to contend that successfully addressing malnutrition for older children in rural China hinges solely on formal microcredit provision. On the contrary, direct nutritional supplementation schemes are not immaterial to those young adults. As argued by Alderman et al. (2006), such schemes could be complemented by other policy interventions in households' production and consumption behavior, such as formal microcredit found in this paper, to improve the youth's health and nutritional status.

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Appendix A

Table A.1
Correlates of the probability of borrowing RCCs.

Independent variable	Estimator (SD)
hh size	−0.062 (0.067)
No. of members not living in the hh	0.043 (0.078)
No. of employed hh members	−0.077 (0.214)
Average edu. of hh members	−0.124 (0.127)
Average age of hh members	0.008 (0.005)
HH wealth status within the village (from low = 1 to high = 4)	−0.063* (0.032)
Quality of housing	0.074* (0.038)
Share of irrigated land	−0.178 (0.208)
Whether borrowing RCCs for child edu. (yes = 1)	0.537*** (0.087)
Whether hh income was sufficient in the past yr. (yes = 1)	−0.237*** (0.065)
Village dummies	Yes
No. of obs.	1773
R^2	0.162

Heteroscedasticity-robust standard errors are in parentheses. Village dummies and the constants are not reported.
** 5% significance level.
*** 1% significance level.
* 10% significance level.

Table A.2
IV estimation of treatment effects on child nutrition.

Independent variable	Child health status		BMI		Ln(Hemoglobin-count)		Ln(Hair zinc)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Treatment effects								
$\hat{\tau}_0$	1.863 (1.350)	1.805 (1.473)	6.438 (9.757)	6.235 (9.157)	0.175 (0.349)	0.172 (0.336)	0.709 (2.495)	0.663 (2.266)
$\hat{\tau}_1$		−0.172 (0.567)		−1.201 (1.862)		−0.034 (0.069)		0.118 (0.459)
Child								
Age	0.020 (0.018)	0.020 (0.018)	0.882*** (0.249)	0.880*** (0.239)	−0.002 (0.009)	−0.001 (0.009)	−0.003 (0.070)	−0.002 (0.065)
Gender	−0.045 (0.040)	−0.045 (0.040)	−0.606** (0.249)	−0.613*** (0.236)	0.003 (0.009)	0.003 (0.009)	−0.053 (0.047)	−0.052 (0.043)
Ethnic minority	−0.013 (0.207)	0.012 (0.207)	−1.468 (1.685)	−1.546 (1.672)	0.063 (0.059)	0.067 (0.064)	−0.009 (0.285)	−0.006 (0.286)
Siblings' average health	0.472*** (0.030)	0.472*** (0.030)	−0.076 (0.180)	−0.107 (0.210)	−0.005 (0.007)	−0.004 (0.008)	0.010 (0.046)	0.012 (0.055)
Health at birth	0.129*** (0.025)	0.129*** (0.026)	0.278** (0.130)	0.296** (0.143)	−0.005 (0.004)	−0.005 (0.005)	−0.019 (0.030)	−0.020 (0.035)
Household								
Water	0.024 (0.021)	0.024 (0.021)	0.179 (0.142)	0.166 (0.130)	−0.001 (0.005)	−0.001 (0.005)	0.017 (0.026)	0.018 (0.022)
Severity of food insecurity	0.034 (0.055)	0.033 (0.055)	0.166 (0.286)	0.253 (0.250)	0.007 (0.011)	0.005 (0.011)	0.077 (0.073)	0.068 (0.044)
Father's edu.	0.006 (0.006)	0.006 (0.006)	−0.077 (0.050)	−0.079 (0.050)	0.002 (0.002)	0.002 (0.002)	0.005 (0.009)	0.005 (0.009)
Mother's edu.	0.003 (0.007)	0.003 (0.007)	−0.023 (0.047)	−0.023 (0.046)	0.002 (0.002)	0.002 (0.002)	0.002 (0.007)	0.002 (0.007)
Ln(hh per capita wealth)	0.046 (0.030)	0.045 (0.030)	0.324 (0.524)	0.302 (0.475)	−0.006 (0.020)	−0.006 (0.019)	−0.044 (0.146)	−0.040 (0.128)
HH social capital	0.024 (0.051)	0.024 (0.051)	−0.284 (0.623)	−0.325 (0.663)	0.003 (0.020)	0.004 (0.022)	0.067 (0.121)	0.070 (0.128)
Women's empowerment	0.093*** (0.035)	0.092*** (0.034)	0.154 (0.227)	0.187 (0.250)	0.002 (0.010)	0.001 (0.011)	−0.012 (0.048)	−0.015 (0.060)
HH migration	−0.063 (0.056)	−0.063 (0.056)	0.351 (0.671)	0.312 (0.597)	−0.011 (0.025)	−0.010 (0.023)	−0.039 (0.174)	−0.033 (0.149)
Village								
Ln(village average per capita net income)	0.136*** (0.037)	0.136*** (0.037)	0.173 (0.219)	0.180 (0.219)	−0.008 (0.009)	−0.008 (0.009)	−0.036 (0.053)	−0.036 (0.055)

(continued on next page)

Table A.2 (*continued*)

Independent variable	Child health status (1)	(2)	BMI (3)	(4)	Ln(Hemoglobin-count) (5)	(6)	Ln(Hair zinc) (7)	(8)
Share of villagers using telephones	−1.026** (0.515)	−1.031** (0.514)	1.563 (3.053)	2.046 (2.584)	0.007 (0.127)	0.007 (0.106)	0.728 (0.578)	0.675 (0.395)
Ratio of no. of medical staff over no. of village clinics	0.010** (0.005)	0.010** (0.005)	0.021 (0.027)	0.018 (0.026)	−0.001 (0.001)	−0.001 (0.001)	−0.004 (0.005)	−0.004 (0.004)
Distance to the nearest township hospital (km)	0.008 (0.006)	0.008 (0.006)	0.013 (0.034)	0.006 (0.032)	−0.001 (0.001)	−0.001 (0.001)	−0.001 (0.010)	−0.0003 (0.007)
Town								
Quality of township hospitals	0.021 (0.039)	0.021 (0.039)	−0.028 (0.245)	−0.026 (0.240)	0.009 (0.011)	0.009 (0.010)	−0.037 (0.050)	−0.036 (0.041)
County dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	1454	1454	1174	1174	1165	1165	1123	1123
F-stat (*p*-value)	22.02 (0.000)	1.60 (0.009)	3.17 (0.000)	3.29 (0.000)	0.75 (0.874)	0.77 (0.861)	1.01 (0.454)	1.06 (0.363)

Note: Heteroskedasticity-robust standard errors are in parentheses.
* 10% significance level.
*** 1% significance level.
** 5% significance level.

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