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Early-life Exposure to Land Reform and Children's Health: Evidence from Cuba

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Early-life exposure to land reform and children's health: Evidence from Cuba

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Abstract

Despite the potential for land reform to improve children's health through nutritional intake, there is little evidence of its effectiveness. Cuba implemented land reforms in 2012 to increase agricultural production, similar to land reforms in other transitional countries. Leveraging the variation in land reform diffusion in Cuba, we evaluate the impact of the land reform in 2012 on children's physical development. Our results show that prenatal exposure to land reform improved children's birthweight. We also find rural-specific effects of postnatal exposure on children's weight and height, and that postnatal exposure is a key driver of health improvements among children, particularly those in more vulnerable populations. Thus, despite the restricted land reforms in Cuba compared with reforms in other countries, they have positive impacts on children's health. Moreover, considering the limited distribution system in Cuba, the land reform demonstrates substantial positive impacts on children's health after birth, with these benefits being more pronounced in rural areas characterized by higher agricultural intensity.

Keywords: Land reform; Fetal origins hypothesis; Early-life health; Usufruct rights; Cuba

Subject classification codes: I15; P26; Q15

Early-life exposure to land reform and children's health: Evidence from Cuba

1. Introduction

Early-life exposure to shocks is a key determinant of human capital formation. Building on the Fetal Origins Hypothesis (FOH), a growing body of literature shows that exposure to shocks—disasters, pandemics, and anti-poverty programs—during the prenatal and postnatal periods affects human capital outcomes in both the short and long terms (Almond and Currie, 2011; Almond et al., 2018; Galiani and Schargrodsky, 2004; Heckman, 2007; Vogl, 2007). Identifying the effects of such shocks is essential to the design of policies to enhance children's human capital, particularly in developing countries, where investments in human capital are crucial for growth.

Early-life exposure to anti-poverty programs contributes to the improvement of children's human capital (Galiani and Schargrodsky, 2004; Vogl, 2007). Although land reforms are not explicitly designed to improve human capital, they may indirectly affect child health through changes in household income and food availability. Treating land reform as an early-life shock influencing children's health (Ding et al., 2025; Kosec and Shemyakina, 2024; Xu, 2021), the literature suggests that greater autonomy among farmers can improve children's human capital by increasing household income and food availability. Consider the Chinese land reform, which marked the transition from the production team system to the household responsibility system (i.e., HRS reform). The HRS reform transformed collective management system into one of individual farm management, giving farmers greater autonomy and incentives to increase their effort. Ding et al. (2025) demonstrate that prenatal exposure to the HRS reform contributed to an increase in birthweight by increasing household income, which in turn improved nutritional intake. Xu (2021) finds that exposure to the HRS reform during early life affected long-term outcomes in adulthood, including health, education attainment, and economic status. In the early 1990s, the government of the Kyrgyz Republic reallocated agricultural land from collective farms to individual households, including previously landless individuals. Kosec and Shemyakina (2024) find that longer prenatal exposure to this reform led to improvements in children's height and weight, and these positive effects were attributed to an increase in consumption of home-grown food.

Cuba is well known for providing free access and high-quality healthcare, and the prevalence of child health problems is as low as in developed countries. According to the World Health Organization (WHO) (2025), the stunting ratio among children in Cuba was about 7% in 2022 compared with a global average of around 22%. Despite a highly developed healthcare system, child nutrition remains another fundamental determinant of early childhood development in Cuba. Low agricultural production has led to limited food availability; in response, the government has implemented agricultural reforms, including land reforms, since 2007 (Nova González, 2012a). Because agricultural reforms directly increase food availability and improve household nutrition, the reforms presumably influence children's health through improved nutrition intake (Ding et al., 2025; Kosec and Shemyakina, 2024). Building on this literature, we analyze whether land reform in Cuba similarly improves children's health, as observed in other countries.

This study aims to identify the causal impact of early-life exposure to Cuba's land reforms on children's health outcomes. Under the land reform in Cuba, since 2008, state-owned idle land with land use rights (usufruct right) have been distributed. The farmers who gain land through this land reform, known as usufruct farmers, are considered more productive since they possess greater autonomy in production. After the reforms, the number of usufruct farmers increased, and the reforms contributed to an increase in production of several crops (Nova González and Alfonso, 2018). These increases in agricultural production are expected to contribute to improving children's health.

However, unlike the land reforms in China and the Kyrgyz Republic, Cuba's reform has distinct characteristics: (1) the distributed land consists of idle plots that are difficult to cultivate immediately after being granted, (2) usufruct farmers are new entrants to the agricultural units, and (3) the reform in Cuba imposes stricter limitations on usufruct farmers. No empirical study has examined the effects of such restricted land reforms, with the exception of Hashiguchi and Murakami (2026), who analyze the effects on agricultural productivity using crop-level panel data. Cuba thus provides new empirical evidence on whether land reform can improve children's health under more restrictive conditions.

Exploiting the variation in the diffusion of land reform across provinces, we employ the regional-level cohort difference-in-differences (DID) model. We focus on the reform in 2012 because provincial-level data from the National Office of Statistics and Information in Cuba (*Oficina Nacional de Estadísticas e Información*, ONEI) on the diffusion of land reform are only available from 2011 onward. Furthermore, micro-level data on children's health are sourced from the Multiple Indicator Cluster Surveys (MICS) provided by the United Nations International Children's Emergency Fund (UNICEF), which are available for only two survey round, 2014 and 2019. Thus, evaluating the impacts of the 2008 reform is not feasible under such data constraints. The laws to distribute idle land, Decree-Law 259 in 2008 and Decree-Law 300 in 2012, were exogenously enforced and diffused nationally with regional variations. Consequently, different birth cohorts were exposed to different intensities of the reforms. We expect that the diffusion of land reforms increases the food supply within a province, thereby improving maternal and child nutrition in the province. This allows us to hypothesize that children with higher exposure to the reforms will have better nutrition and improvements in their health status. To the best of our knowledge, this study is the first to empirically analyze the causal effects of land reform in Cuba using micro-level data, thereby making a novel contribution to the literature.

The remaining paper is organized as follows. Section 2 describes the post-2008 land reforms in Cuba. Section 3 discusses the identification strategies. Section 4 describes the data and section 5 presents the empirical specifications. Sections 6 and 7 present the estimation results and the robustness checks, respectively. Section 8 concludes.

2. Land reforms in Cuba

The agricultural sector in Cuba consists of a state sector and a non-state sector. Until 2008, the non-state sector was primarily constituted by agricultural cooperatives, specifically three types: Cooperatives of Credit and Services (*Cooperativas de Créditos y Servicios*, CSSs), Agricultural Production Cooperatives (*Cooperativas de Producción Agropecuaria*, CPAs), and Basic Units of Cooperative Production (*Unidades Básicas de Producción Cooperative*, UBPCs), which possessed greater autonomy in production than those in the state sector (Alvarez, 2004). Thus, their productivity was higher and played an essential role in total agricultural production. However, because agricultural production was insufficient to satisfy national demand, the government implemented a land reform in 2008 to create a new category of farmers called usufruct farmers—private producers with the highest level of autonomy in Cuba.

Land reforms, which began in 2008, were subsequently revised in 2012 and 2018. The initial reform in 2008 distributed state-owned idle land to individuals and entities through the provision of usufruct rights (Gaceta Oficial de Cuba, 2008). Individuals received land of up to 13.42 hectares for a contract period of 10 years. However, beneficiaries were prohibited from transferring their land to a third person, including through leasing and renting, and utilizing unplanned activities such as planting fruit trees on the land (Gaceta Oficial de Cuba, 2008). Four years later, the revised reform in 2012 allowed farmers to grow fruit crops and expand their land area up to 67.10 hectares, provided that they already held land and were associated with agricultural cooperatives or state farms (Gaceta Oficial de Cuba, 2012). Under the reforms, the usufruct farmers have greater residual claim of their output compared with other types of farmers including those in cooperatives (Nova González, 2012b). Nevertheless, the land reform still obligates farmers to return their land and its assets to the state at the expiration of usufruct rights, to use it rationally and sustainably, and to fulfill the commitments of the rationing system (Gaceta Oficial de Cuba, 2008; 2012). Given these restrictions, Cuban usufruct farmers face weaker incentives to invest than those affected by land reforms in other countries.

The reforms ultimately expanded usufruct farming, increased the production of several crops, and promoted more efficient land reallocation (González-Corzo, 2019; Mesa-Lago and González-Corzo, 2021; Nova González & Alfonso, 2018). Table 1 shows the number of agricultural producer units from 2011 to 2017. Under the 2008 reform, approximately 160,000 usufruct farmers were newly incorporated as agricultural producers. The number of usufruct farmers doubled following the enforcement of the Decree-Law 300 in 2012. Table 2, which presents the change in the distribution of farmland by types of farmers from 2007 to 2016, shows that the share of land held by the CCSs and private farmers, including the usufruct farmers, increased.

[Insert Tables 1 and 2 here]

Agricultural production also increased as a result of the reform. We regress the number of usufruct farmers per 1,000 people on each crop's output at provincial level. In Table 3, all crop categories are positively correlated with the number of usufruct farmers, leading us to hypothesize that the reforms also contributed to children's health improvement[1]. Therefore, we evaluate the impacts of the reform in 2012 on children's health.

[Insert Table 3 here]

3. Identification strategies

Health endowments established during early life, including the prenatal and postnatal periods, are key determinants of future human capital. This idea, known as the FOH, suggests that shocks affecting health in early life have lasting effects on health in later life (Almond and Currie, 2011; Almond et al., 2018; Heckman, 2007). Following the FOH, early-life nutrition shapes health outcomes for children at each

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[1] We do not use crop outputs in our main estimation, because these data are unavailable for the years before 2014 and have limited credibility owing to substantial missing values and outliers.

stage of development. We thus assume that Cuba's land reform increased food supply and access, thereby improving maternal nutrition during the prenatal period and children's nutrition during the postnatal period, which in turn leads to better children's health.

We focus on an underlying mechanism distinct from those discussed in previous studies, which propose two potential mechanisms linking land reform to children's health (Ding et al., 2025; Kosec and Shemyakina, 2024; Xu, 2021): the income channel, through which increased household income resulting from higher agricultural productivity enables households to allocate more resources, including food, to their children (Ding et al., 2025; Xu, 2021); and the consumption channel, through which increased household's access to land enables households to consume more of their own agricultural production rather than purchasing food in the market (Kosec and Shemyakina, 2024). Data constraints make it difficult to identify both the income and consumption channels directly[2]. We still hypothesize that greater regional food supply increases access to nutrition for both farming and non-farming households, resulting in improvements to children's health. Therefore, our analysis examines the impact of land reform on the health of all children living in these regions.

We exploit variation in the diffusion of Cuba's land reforms and across birth cohorts to estimate their effect on children's health. Despite the national scale of the post-2008 Cuban land reforms, the pace and extent of diffusion of the reform has varied across the country. Figure 1 presents the diffusion patterns in each province from 2011 to 2019. Almost all provinces rapidly extended the number of usufruct farmers after the 2012 modification; following that, the diffusion paths of the reform diverged across provinces. We expect that children exposed to more advanced reforms during early life exhibit better health outcomes compared with those exposed to less advanced reforms. This heterogeneity in both provincial diffusion and cohort exposure allows us to identify the effects of land reform. Note that although a never-treated province does not exist, we utilize the difference in diffusion across provinces to identify the causal effects.

[Insert Figure 1 here]

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[2] Even though one of the surveys we use, the 2019 MICS, contains agricultural information, it lacks data on income and consumption. As a cross-sectional survey, it does not allow us to identify the income and consumption channel using DID model.

We also hypothesize that the adoption of land reform does not exhibit spillover effects across provinces. The private farmers, including usufruct farmers, have faced a restriction of food transport in Cuba. For instance, private farmers need an acceptance to transport their output to other provinces or connection with some supplier (Thiemann and Spoor, 2019). However, there are *de facto* difficulties gaining permission and connection. Instead, farmers sell their outputs in their living province only. Under this unique market condition, our assumption necessitates us to focus our analysis solely on the direct, localized impacts within each region rather than effects spreading through inter-provincial channels.

4. Data

This study uses micro-level data for children's health status and provincial level data for land reform exposure. The micro-level data were collected by the UNICEF in 2014 (MICS2014) and 2019 (MICS2019). Both datasets include children's health status, specifically measured weight and height at the survey time, for children under 5 years of age. We exclude observations for children aged 2–5 years from the MICS2014 sample, because provincial level data on the corresponding exposure indicator are unavailable for these cohorts. Both datasets include children's health status, demographic characteristics, parent's educational attainment, and household infrastructure.

Outcome variables in our analysis include birthweight, the z-scores of weight-for-age, height-for-age and weight-for-height, and health risk indicators based on these z-scores: underweight, stunting, wasting, and overweight. To explore the influence of exposure during the prenatal period, we restrict the sample to children under 2 years of age because the MICS surveys only collected birthweight data for children within this age range. To evaluate the impacts of postnatal exposure, we construct z-scores by employing the WHO guidelines using measured weight and height data in each survey[3]. Following the WHO and UNICEF (2019), the health risk indicators—underweight, stunting, and wasting—are each defined as a dummy variable equal to 1 if the child's respective z-score (i.e., weight-for-age, height-for-age, and weight-for-height, respectively) is less than -2. Overweight is defined as a dummy variable equal to 1 if the child's z-score of weight-for-height is greater than 2. Figure 2 shows the mean weight-for-age and height-for-age z-scores by age cohort for each province. In almost all provinces, younger cohorts exhibit higher z-scores than older cohorts.

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[3] To construct the metrics, we employ WHO's child growth standards Stata macro package (https://github.com/unicef-drp/igrowup_update).

[Insert Figure 2 here]

Another dataset is provincial level agricultural data sourced from ONEI[4]. We utilize the annual number of usufruct farmers in each province from 2012 to 2019 to measure the degree of individual exposure to the land reform during the prenatal and postnatal period. Owing to the inaccessibility of more disaggregated data such as municipality[5] and monthly levels, we calculate the degree of exposure by weighing the annual level data based on the month and year of childbirth. In addition, we include cultivated area of land as control variable for a time-varying agricultural level, which might encourage the entry of new usufruct farmers. Both agricultural variables are primarily expressed in per capita metrics[6]. Table 4 lists the descriptive statistics of each variable. The surveys provide non-response adjusted expansion weights. Using the weights, our sample represents the country at national, provincial, and urban/rural levels. The weights are used for all estimations in this study.

[Insert Table 4 here]

5. Empirical specification

We utilize a regional-level birth cohort DID model to identify the impact of land reform, which is similar to the specification used by Xu (2021) to evaluate the impact of land reform on children's human capital in China. We estimate the following model to explore the impact of prenatal exposure on children's health using a selected sample of children under 2 years of age:

$$Y_{ipas} = \alpha + \beta WEprenatal_{pas} + \boldsymbol{X}'\boldsymbol{\gamma} + \mu_p + \delta_m + \theta_s + \epsilon_{ipas}, \tag{1}$$

where i, p, a, s, m, and t denote the index individual, province, age in month, survey year, birth-month, and birth-year, respectively. Y_{ipas}, the outcome variable, is the birthweight.

$$WEprenatal_{pas} = \left(\frac{m}{12} \times Exposure_{pt-1}\right) + \left(\frac{12-m}{12} \times Exposure_{pt-2}\right), \tag{2}$$

$$s \in (2014, 2019)$$

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[4] Data sourced from ONEI for this study can be found at https://www.onei.gob.cu/.

[5] Since the microdata only include the living region at the provincial level, we cannot use municipality level agricultural data, even if such data were accessible.

[6] We also use the logarithmic the number of usufruct farmers as an alternative measure of the degree of individual exposure.

$WEprenatal_{pas}$ is the rate of exposure to the land reform during the prenatal period and consists of the rate of exposures in two years in Equation (2)[7]. The rate of exposure, $Exposure_{pt-1}$, is the number of usufruct farmers per capita in each province. Owing to data limitations, the number of usufruct farmers is only available as annual data published in December. We thus calculate the rate of exposure using weights derived from the child's birth month. The variable $WEprenatal_{pas}$ is lagged by one year since the impact appears at least one year after a farmer newly gains usufruct rights under the reform[8]. For example, the rate of exposure for a child born in September 2018, whose prenatal period is from September 2017 to September 2018, is calculated based on the land reform status from September 2016 to September 2017.

$\boldsymbol{X'}$ is a vector of covariates, including the provincial agricultural level, mother's educational level, number of siblings, and gender (all variables are shown in Table 1)[9]. The agricultural level and diffusion patterns across provinces may be endogenous, thereby leading to a biased estimate of the key parameter of interest β. To control for this time-variant influence, we include the agricultural level lagged by one year in our regression.

The mother's years of schooling is a potential determinant of nutritional intake during pregnancy and child-feeding. Furthermore, the number of siblings and the child's gender are potential determinants of parental investment in nutritional supply to the child. We thus include these variables as additional control variables.

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[7] Following Kosec and Shemyakina (2024), we assume a 12-month prenatal period. However, we will conduct robustness checks by re-defining the prenatal period as nine months in section 7.

[8] Results using the unlagged variable $WEprenatal_{pas}$ are similar to our main findings and are available upon request.

[9] We do not include migration information across provinces, because the MICS2014 lacks these data. However, based on the migration information available in the MICS2019, migration across provinces is minimal within our sample.

Finally, μ_p, δ_m, and θ_s represent province fixed effects, birth-month fixed effects, and survey year fixed effects, respectively[10]. Note that health infrastructure is an important determinant of children's health. Since 2011, the Cuban government has undertaken a nationwide reorganization of the health system to enhance its sustainability (Morales Ojeda et al. 2018). Because Cuba had already established its health system before our study period and the health-system reforms were implemented uniformly across provinces, province fixed effects are expected to account for cross-provincial differences in the initial level of health infrastructure, while survey-year fixed effects absorb the common effects of the health-system reforms. Birth-month fixed effects account for seasonal variation in nutrient availability arising from the crop production cycle. The survey year fixed effects capture nationwide time-varying factors including macro-economic shocks and policy reforms, which affects all children observed in a survey year.

To evaluate the effects of exposure during the postnatal period on children's health status, we estimate the following model using the full sample (all age cohorts):

$$Y_{ipas} = \alpha + \beta WEpostnatal_{pas} + \boldsymbol{X'\gamma} + \mu_p + \delta_a + \theta_s + \epsilon_{ipas} \tag{3}$$

where n denotes survey-month. Y_{ipas} is the z-scores or health risk indicators in this model.

$$\begin{aligned} WEpostnatal_{pas} = {} & \frac{n}{12} \times Exposure_{ps-1} + \frac{12 \times 1[t-s<0] - m}{12} \times Exposure_{pt-1} \\ & + \sum_{T=s-4}^{s-1} Exposure_{pT-1} \times 1[T-t>0] \end{aligned} \tag{4}$$

$$s \in (2014, 2019)$$

The explanatory variable, $WEpostnatal_{pas}$, measures the rate of exposure to the land reform during the postnatal period. It consists of three components: exposure during the survey year, exposure in the birth year, and exposure accumulated between these two years. Specifically, $\frac{n}{12} \times Exposure_{ps-1}$ captures partial year exposure in the survey year; $\frac{12\times1[t-s<0]-m}{12} \times Exposure_{pt-1}$ indicates exposure in the birth year; and $\sum_{T=s-4}^{s-1} Exposure_{T-1} \times 1[T-t>0]$ measures exposure in years between the birth year and the survey year. Given that our analysis focuses on children aged 5 or younger, T, representing the years between the two periods, ranges from 4 years before the survey year ($s-4$) to 1 year before the

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[10] Given the limited number of observations for this specification, the birth cohort-level fixed effects are defined by birth month in this model.

survey year ($s-1$). Since $WEpostnatal_{pas}$ aggregates rates of exposure during the postnatal period, the key coefficient of interest, β, identifies the cumulative effect of postnatal exposure to the reform.

$\boldsymbol{X}'$ in this model additionally includes the rate of exposure to the land reform during the prenatal period, in addition to the vector of covariates in Equation (1). Finally, δ_a captures age in month fixed effects, instead of birth month fixed effects used as the birth cohort-level fixed effects in Equation (1).

6. Results

Table 5 presents the estimation results for the impact of prenatal exposure on birth weight using Equation (1). The coefficients of both explanatory variables, number of usufruct farmers per capita and log of the number of usufruct farmers, show positive effects on children's birth weight. On average, a one-standard deviation increases in the rate of exposure, corresponding to an increase of about 1.2 usufruct farmers per 100 people, increases birthweight by 0.137 kg[11]. Our estimated effect is larger than that of Ding et al. (2025), who report that prenatal exposure to the HRS reform increases birth weight by approximately 55 g. However, since Ding et al. (2025) define treatment as an indicator that takes the value of one if a child's county had already implemented the HRS reform during the prenatal period, the two effects are not directly comparable because of differences in empirical specifications.

[Insert Table 5 here]

Table 6 presents the estimation results of Equation (3). In all the models, the estimates show no significant cumulative effects of exposure to land reform during the prenatal period. However, we investigate the effects on children's health separately for rural and urban subsamples since land reform was primarily implemented in rural areas where agriculture is more intensive. Table 7 presents the estimation results obtained from the subsamples. Children's health in rural areas improved with higher cumulative postnatal exposure, even after controlling for postnatal exposure (see Models 1, 3, and 5). Thus, children living in rural areas benefit especially from the land reform because they have easier

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[11] The estimated effect is calculated as the coefficient β from Model 1 in Table 5 (11.5539) multiplied by one-standard deviation of $WEprenatal_{pas}$ from Table 4 (0.0119).

access to the increased food supply attributed to the reform[12]. By contrast, the positive effects of land reform do not appear in urban areas (see Models 2, 4, and 6). Possibly, limited market channels restrict the delivery of food produced by private farmers, including usufruct farmers, from rural to urban areas in Cuba (Thiemann and Spoor, 2019). The magnitude of positive effect on height-for-age is about 0.65 standard deviations in response to a one-standard deviation increase in the rate of exposure during the postnatal period[13]. Because an additional month of exposure to land reform in the Kyrgyz Republic increases height-for-age by 0.017 standard deviations, the effect estimated in our study is approximately equivalent to the effect of exposure to land reform in the Kyrgyz Republic for 38 months in early life on height-for-age (Kosec and Shemyakina, 2024).

[Insert Tables 6 and 7 here]

Table 8 presents the effects of postnatal exposure on health risks. Higher postnatal exposure to land reform reduces health risks, including underweight and stunting (see Models 1 and 2). Consistent with the results using the z-score outcomes, the positive impacts of reducing the likelihood of health risks are greater for children living in rural areas (see Table 9).

[Insert Tables 8 and 9 here]

7. Robustness Check

First, following Rosales-Rueda (2018), we estimate our main specification using a nine-month prenatal period to establish the robustness of the 12-month assumption. Table 10 presents the results based on the nine-month definition. All the results are consistent with those from the baseline specification.

[Insert Table 10 here]

Second, we investigate whether differences in the initial land reform intensity led to improvements in children's health. The land reform was extensively implemented in the initial year in 2013 (see Figure 1). This suggests that initial intensity may drive heterogeneity in children's health. Table 11 shows the effects of prenatal exposure on birth weight based on the DID using only the initial year

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[12] The estimation results using the log of the number of usufruct farmers as the treatment variable do not show statistically significant positive effects. However, using the number of usufruct farmers per capita is more appropriate because it captures population differences across provinces.

[13] The estimated effect is calculated as the coefficient β from Model 3 in Table 7 (13.2412) multiplied by one-standard deviation of $WEpostnatal_{pas}$ from Table 4 (0.0491).

variation[14]. The insignificant coefficient indicates that it is important to consider the variation in the reform's diffusion over multiple years rather than relying solely on the initial implementation.

[Insert Table 11 here]

Finally, we employ quantile regressions using the rural subsample to investigate whether the association between postnatal exposure and improvements in children's health differed across the health distribution, particularly among children in the lower quantiles. Figure 3 shows that the effect of postnatal exposure varied substantially across the distribution of z-scores in rural areas. The quantile regression estimates show that the positive impact is larger at lower quantiles and gradually diminishes toward higher quantiles. This pattern suggests that land reform benefits vulnerable children by improving their health status.

[Insert Figure 3 here]

8. Conclusion

Land reform can function as a potential shock, influencing human capital formation through nutritional intake (Ding et al., 2025; Kosec and Shemyakina, 2024; Xu, 2021). This study investigates the impact of land reform exposure in early life on children's health in Cuba by leveraging variations in the diffusion of reform across provinces. Since 2008, Cuba has implemented a series of land reforms to increase national food production. We show that despite the restrictive nature of Cuba's reform, it improved child health by increasing the regional food supply.

The results confirm a positive effect of prenatal exposure on birth weight. In addition, postnatal exposure improved the health outcomes of children, especially those living in rural areas. Considering the limited distribution system in Cuba, land reform demonstrates substantial positive impacts on children's health after birth, with these benefits being more pronounced in rural areas characterized by higher agricultural intensity. Moreover, the results of the effects on health risks indicate that postnatal exposure contributes to a decline in health risks for children. These findings imply that land reforms have a more effective impact on vulnerable children.

The findings provide evidence supporting the role of land reform as an anti-poverty policy. Existing literature suggests that land reforms can generate positive effects on subsequent generations, particularly those living in rural areas. Our study, which focuses on a more restrictive form of land

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[14] The details of DID are explained in Appendix.

reform, provides new evidence on the important role of land reform in improving the health of vulnerable children in the context of poverty.

While this study focuses only on the 2012 reform owing to data limitations, the initial reform in 2008 had a stronger impact on agricultural production and productivity (González-Corzo, 2019; Mesa-Lago and González-Corzo, 2021; Hashiguchi and Murakami, 2026). Evaluating the effects of the 2008 land reforms is an intriguing topic for future research. Moreover, our analysis does not directly identify the mechanisms underlying the causal effects of land reform on children's health. Since data on food price changes across provinces during the reform period are unavailable, we cannot empirically verify the hypothesized channel of an increase in regional food supply and cannot rule out other potential mechanisms. Therefore, identification of the underlying mechanisms remains an important topic for future research.

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Tables

Table 1

Units of agricultural producers.

		2011	2012	2013	2014	2015	2016	2017
Legal entities	Total	na	4856	5781	7584	6610	6702	8578
	State-owned agricultural enterprise	na	931	413	779	2133	1874	2565
	UBPCs	na	1872	1723	1603	1651	1629	1691
	CPAs	na	982	914	984	867	879	962
	CCSs	na	na	1249	1352	1223	1224	2306
	Other	na	1071	1482	2866	736	1096	1054
Natural person	Total	na	426622	454919	447715	412615	419973	408446
	Usufruct famer	157948[a]	300810	312752	312296	279021	287107	274635
	Other small farmers	na	125812	142167	135419	133594	132866	133811

Notes: CSSs: Cooperatives of Credit and Services (*Cooperativas de Créditos y Servicios*), CPAs: Agricultural Production Cooperatives (*Cooperativas de Producción Agropecuaria*, CPA), and UBPCs: Basic Units of Cooperative Production (*Unidades Básicas de Producción Cooperative*. [a] The number of usufruct farmers in 2011 represents the cumulative total of farmers established under Decreto-Ley 259 (during 2008–2011). Source: ONEI agricultural statistics at https://www.onei.gob.cu/agricultura (Accessed 21 March, 2024).

Table 2

Distribution of farmland by types of farmers, in 2007 and 2016.

		2007			2016		
		Agricultural	Cultivated	Not Cultivated	Agricultural	Cultivated	Not Cultivated
Area (1000 hectares)							
Total		6619	2988	3631	6226	2733	3493
State		2371	694	1677	1912	521	1391
Non-state	UBPCs	2448	1189	1258	1528	840	688
	CPAs	585	305	280	503	267	236
	CSSs and Private	1214	799	415	2283	1103	1180
Distribution (%)							
Total		100	100	100	100	100	100
State		35.8	23.2	46.2	30.7	19.1	39.8
Non-state	UBPCs	37.0	39.8	34.6	24.5	30.7	19.7
	CPAs	8.8	10.2	7.7	8.1	9.8	6.8
	CSSs and Private	18.3	26.7	11.4	36.7	40.4	33.8

Notes: CSSs: Cooperatives of Credit and Services (*Cooperativas de Créditos y Servicios*), CPAs: Agricultural Production Cooperatives (*Cooperativas de Producción Agropecuaria*, CPA), and UBPCs: Basic Units of Cooperative Production (*Unidades Básicas de Producción Cooperative*.

Source: Author's own calculations based on the data sourced from ONEI (2008; 2017).

Table 3

Estimation results from regressions of crop outputs on the number of usufruct farmers per 1,000 people.

	Tons per capita					
	Tuberous	Vegetables	Plantains	Cereals	Fruits	Beans
No. usufruct farmer per 1000 people	0.0110***	0.0070**	0.0027***	0.0076***	0.0026***	0.0006***
	(0.0024)	(0.0031)	(0.0006)	(0.0011)	(0.0006)	(0.0002)
Constant	-0.1754**	-0.0118	-0.0207	-0.1516***	-0.0205	-0.0075
	(0.0771)	(0.0982)	(0.0179)	(0.0335)	(0.0186)	(0.0055)
Observations	58	58	58	58	52	58
R-squared	0.2664	0.0822	0.2910	0.4793	0.2745	0.1876

Notes: Standard errors are in parentheses. ***, **, and * indicate significance at the 1 percent, 5 percent, and 10 percent levels, respectively. We utilize annual data on the number of usufruct farmers per 1,000 people and crop output in each province from 2014 to 2019 sourced from ONEI.

Table 4

Descriptive statistics.

	N	Mean	SD	Min	Max
Weight-for-age Z-score (z_wei)	5256	0.2408	1.1887	-5.43	4.99
Height-for-age Z-score (z_len)	5250	-0.1831	1.3814	-5.73	5.94
Weight-for-height Z-score (z_wfl)	5238	0.4895	1.2644	-4.64	4.99
Underweight (z_wei<-2)	5256	0.0282	0.1657	0	1
Stunting (z_len<-2)	5250	0.0789	0.2697	0	1
Wasting (z_wfl<-2)	5238	0.0211	0.1436	0	1
Overweight (z_wfl>2)	5338	0.1222	0.3276	0	1
Birthweight (kg)	1860	3.2874	0.5639	0.604	7.5
Weighted rate of exposure during the prenatal (units per capita)	5338	0.0268	0.0119	0.0019	0.0597
Weighted rate of exposure during the postnatal (units per capita)	5338	0.0678	0.0491	0.0005	0.2115
Weighted rate of exposure during the prenatal (log of units)	5338	9.7626	0.5127	7.3658	11.0352
Weighted rate of exposure during the postnatal (log of units)	5338	10.4455	0.9751	6.3075	12.0508
Agricultural level (cultivated area per capita)	5338	0.2787	0.1397	0.0086	0.8556
Household head's education	5338	11.4463	3.7824	0	18
Mother's education	5338	12.8032	2.9241	0	18

Gender	5338	0.5006	0.5000	0	1
Siblings	5338	0.7494	0.8745	0	10
Urban	5338	0.6617	0.4732	0	1
Pipe water	5338	0.8312	0.3746	0	1
Flush toilet	5338	0.7766	0.4165	0	1
Age in month	5339	30.5598	17.5130	0	60
Province	5338	8.1166	4.4052	1	16
Survey year	5338	2017.6260	2.2323	2014	2019

Source: Multiple Indicator Cluster Surveys 2014 and 2019 and the author’s own calculations based on data from ONEI agricultural and population statistics.

Table 5

Effect of prenatal exposure on birthweight.

	Rate of exposure = usufruct farmers / population	Rate of exposure = log (usufruct farmers)
	Model 1	Model 2
WE during the prenatal period	11.5539**	0.2518**
	(4.5646)	(0.0900)
Agricultural level	0.0872	-0.0377
	(0.3951)	(0.3439)
Other controls	Yes	Yes
Province FE	Yes	Yes
Birth month FE	Yes	Yes
Survey-year FE	Yes	Yes
Observations	1,860	1,860
R-squared	0.0967	0.0974

Notes: Robust standard errors clustered at the provincial level are shown in parentheses. ***, **, and * indicate significance at the 1 percent, 5 percent, and 10 percent levels, respectively. FE: Fixed effects.

Table 6

Effect of early-life exposure on z-scores.

	Exposure rate = usufruct farmers / population			Exposure rate = log (usufruct farmers)		
Dependent variable	Weight-for-age	Height-for-age	Weight-for-height	Weight-for-age	Height-for-age	Weight-for-height
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
WE during the postnatal period	-1.3462	-4.2849	1.9914	-0.3611	-0.4800	-0.2277
	(1.2574)	(3.1144)	(2.8973)	(0.3494)	(0.9297)	(0.3469)
WE during the prenatal period	-6.5503	-1.4644	-9.3062	-0.1957	-0.2491	-0.0888
	(4.6662)	(8.7217)	(8.3788)	(0.1547)	(0.3412)	(0.3570)
Agricultural level	1.8874**	2.1590	1.3525	2.2644***	2.8873*	1.2818
	(0.6941)	(1.6699)	(1.4262)	(0.6466)	(1.4047)	(1.2211)
Other controls	Yes	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	Yes	Yes	Yes	Yes
Age in month FE	Yes	Yes	Yes	Yes	Yes	Yes
Survey-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	5,256	5,250	5,238	5,256	5,250	5,238
R-squared	0.1204	0.1886	0.0891	0.1209	0.1881	0.0882

Notes: Robust standard errors clustered at the provincial level are shown in parentheses. ***, **, and * indicate significance at the 1 percent, 5 percent, and 10 percent levels, respectively. The number of observations varies across models because each health indicator is collected separately, and some values are missing or excluded as outliers. For instance, a child might have a z-score for weight-for-age but lack valid z-scores for height-for-age and weight-for-height. FE: Fixed effects.

Table 7

Effect of early-life exposure on z-scores for rural and urban samples.

	Rate of exposure = usufruct farmers / population					
Dependent variable	Weight-for-age		Height-for-age		Weight-for-height	
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
WE during the postnatal period	17.4730***	-3.2615**	13.2412**	-4.8606	12.9858***	-0.1602
	(5.3160)	(1.4114)	(5.7373)	(2.8455)	(3.1446)	(1.9870)
WE during the prenatal period	-13.5465	-3.9523	1.6534	3.1270	-19.1868***	-10.9575
	(9.1024)	(5.2255)	(14.2408)	(8.3963)	(5.0292)	(6.6692)
Agricultural level	2.2724	1.3802*	2.5134	2.3976	0.6890	0.5156
	(1.6172)	(0.7854)	(2.0725)	(1.5066)	(1.2998)	(1.4734)
Sub-sample	Rural	Urban	Rural	Urban	Rural	Urban
Other controls	Yes	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	Yes	Yes	Yes	Yes
Age in month FE	Yes	Yes	Yes	Yes	Yes	Yes
Survey-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1,490	3,766	1,487	3,763	1,487	3,751
R-squared	0.2057	0.1483	0.2194	0.2602	0.1712	0.1148

	Rate of exposure = log (usufruct farmers)					
Dependent variable	Weight-for-age		Height-for-age		Weight-for-height	
	Model 7	Model 8	Model 9	Model 10	Model 11	Model 12
WE during the postnatal period	0.9406	-0.8008**	-0.4049	-0.4376	1.0076	-0.6302
	(0.8529)	(0.2775)	(1.7523)	(0.6887)	(0.8709)	(0.3605)
WE during the prenatal period	-0.6465**	0.0182	-0.3158	0.0394	-0.6220**	-0.0951
	(0.2912)	(0.2043)	(0.5307)	(0.2103)	(0.2596)	(0.2512)
Agricultural level	2.1725	1.6851**	3.7846	2.8715**	0.2053	0.6393
	(1.7459)	(0.5977)	(2.5687)	(1.3364)	(1.7503)	(1.2481)
Sub-sample	Rural	Urban	Rural	Urban	Rural	Urban
Other controls	Yes	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	Yes	Yes	Yes	Yes
Age in month FE	Yes	Yes	Yes	Yes	Yes	Yes
Survey-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1,490	3,766	1,487	3,763	1,487	3,751
R-squared	0.1968	0.1490	0.2150	0.2579	0.1659	0.1156

Notes: Robust standard errors clustered at the provincial level are shown in parentheses. ***, **, and * indicate significance at the 1 percent, 5 percent, and 10 percent levels, respectively. FE: Fixed effects.

Table 8

Effect of early-life exposure on health risks.

	Rate of exposure = usufruct farmers / population				Rate of exposure = log (usufruct farmers)			
Dependent variable	Underweight	Stunting	Wasting	Overweight	Underweight	Stunting	Wasting	Overweight
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
WE during the postnatal period	-0.3851***	-1.0612**	-0.0890	-0.2559	-0.0577***	-0.0536	0.0268	-0.0427
	(0.1241)	(0.4968)	(0.1747)	(0.4290)	(0.0208)	(0.0511)	(0.0326)	(0.0622)
WE during the prenatal period	0.3355	1.7527	-0.2735	0.6032	0.0028	0.0822	-0.0265	0.0192
	(0.9720)	(1.6819)	(0.9364)	(1.3933)	(0.0316)	(0.0521)	(0.0357)	(0.0481)
Agricultural level	0.1129	-0.1470	0.0285	-0.1682	0.1654	-0.1115	0.0686	-0.1656
	(0.1013)	(0.1985)	(0.0673)	(0.1889)	(0.1028)	(0.1400)	(0.0587)	(0.1547)
Other controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age in month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Survey-year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	4,575	5,233	4,683	5,321	4,575	5,233	4,683	5,321

Notes: Coefficients are the marginal effects of the probit estimation. Robust standard errors clustered at the provincial level are shown in parentheses. ***, **, and * indicate significance at the 1 percent, 5 percent, and 10 percent levels, respectively. The number of observations varies across the models for two reasons. First, each health indicator is

collected separately and some values are missing or excluded as outliers, as described in the note to Table 6. Second, in the probit estimations, we exclude regional birth cohort cells with no within-cell variation in the health-risk dummy (i.e., cells in which the dummy takes a value of 1 or 0 for all children). FE: Fixed effects.

Table 9

Effect of early-life exposure on health risks for rural and urban samples.

	Rate of exposure = usufruct farmers / population							
Dependent variable	Underweight		Stunting		Wasting		Overweight	
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
WE during postnatal period	-3.5717***	-0.5016***	-2.2986***	-1.2358**	-1.7262*	-0.1745	1.8067**	-0.3445
	(0.8897)	(0.1818)	(0.8231)	(0.5134)	(0.9787)	(0.1387)	(0.8447)	(0.3815)
WE during prenatal period	0.7800	0.7917	-0.8879	2.4662	-3.3235	0.8125	-1.5719	2.7518
	(1.1953)	(1.0447)	(2.0789)	(2.0357)	(4.2543)	(0.7659)	(1.3732)	(1.7772)
Agricultural level	0.4604	0.0388	-0.0432	-0.1182	0.1751	-0.0029	-0.1414	-0.2538
	(0.2878)	(0.1344)	(0.2789)	(0.1952)	(0.2068)	(0.0874)	(0.3165)	(0.2067)
Sub-sample	Rural	Urban	Rural	Urban	Rural	Urban	Rural	Urban
Other controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age in month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Survey-year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	668	2,806	1,380	3,743	665	2,938	1,418	3,821

	Rate of exposure = log (usufruct farmers)							
Dependent variable	Underweight		Stunting		Wasting		Overweight	
	Model 9	Model 10	Model 11	Model 12	Model 13	Model 14	Model 15	Model 16
WE during postnatal period	-0.3859***	-0.0491	-0.0392	-0.0933*	0.1226	0.0236	0.0692	-0.0941
	(0.0655)	(0.0480)	(0.0979)	(0.0557)	(0.1603)	(0.0438)	(0.1912)	(0.0686)
WE during prenatal period	0.0368	0.0013	-0.0207	0.1095*	-0.1663	-0.0071	-0.0388	0.0748
	(0.0392)	(0.0429)	(0.0766)	(0.0595)	(0.1056)	(0.0396)	(0.0541)	(0.0643)
Agricultural level	0.7606**	0.0961	-0.0071	-0.0981	0.1045	0.0215	-0.1841	-0.3475*
	(0.3053)	(0.1361)	(0.3100)	(0.1526)	(0.2496)	(0.0825)	(0.3987)	(0.1944)
Sub-sample	Rural	Urban	Rural	Urban	Rural	Urban	Rural	Urban
Other controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age in month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Survey-year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	668	2,806	1,380	3,743	665	2,938	1,418	3,821

Notes: Coefficients are the marginal effects of the probit estimation. Robust standard errors clustered at the provincial level are shown in parentheses. ***, **, and * indicate significance at the 1 percent, 5 percent, and 10 percent levels, respectively. The number of observations varies across the models for two reasons. First, each health indicator is

collected separately and some values are missing or excluded as outliers, as described in the note to Table 6. Second, in the probit estimations, we exclude regional birth cohort cells with no within-cell variation in the health-risk dummy (i.e., cells in which the dummy takes a value of 1 or 0 for all children). FE: Fixed effects.

Table 10

Estimation results with nine months prenatal exposure specification.

(a) Effects on birth weight.

	Rate of exposure = usufruct farmers / population Model 1
WE during the prenatal period	14.4687***
	(4.4730)
Agricultural level	0.0104
	(0.4944)
Other controls	Yes
Province FE	Yes
Age in month FE	Yes
Survey-year FE	Yes
Observations	1,860
R-squared	0.0987

Notes: Robust standard errors clustered at the provincial level are shown in parentheses. ***, **, and * indicate significance at the 1 percent, 5 percent, and 10 percent levels, respectively. FE: Fixed effects.

(b) Effect on z-scores for rural and urban samples.

	Rate of exposure = usufruct farmers / population					
Dependent variable	Weight-for-age		Height-for-age		Weight-for-height	
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
WE during the postnatal period	15.9128***	-3.9830***	12.1353**	-5.3366**	11.4703***	-0.8202
	(5.0528)	(1.0872)	(5.6356)	(2.4394)	(2.9937)	(1.8429)
WE during the prenatal period	-15.7676	-4.3221	-2.4419	7.5898	-21.3913**	-16.7139**
	(9.4552)	(7.0173)	(15.7191)	(9.4795)	(7.5717)	(7.2243)
Agricultural level	0.6884	-0.5489	-0.4193	1.4414	0.9995	-0.9309
	(1.2218)	(2.0852)	(1.3762)	(2.4920)	(1.4302)	(1.9740)
Sub-sample	Rural	Urban	Rural	Urban	Rural	Urban
Other controls	Yes	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	Yes	Yes	Yes	Yes
Age in month FE	Yes	Yes	Yes	Yes	Yes	Yes
Survey-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1,490	3,766	1,487	3,763	1,487	3,751
R-squared	0.2026	0.1468	0.2159	0.2581	0.1705	0.1167

Notes: Robust standard errors clustered at the provincial level are shown in parentheses. ***, **, and * indicate significance at the 1 percent, 5 percent, and 10 percent levels, respectively. FE: Fixed effects.

(c) Health risks in rural and urban areas.

	Rate of exposure = usufruct farmers / population							
Dependent variable	Underweight		Stunting		Wasting		Overweight	
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
WE during the postnatal period	-3.5769***	-0.4215***	-2.5568***	-1.1317***	-1.8446**	-0.1652	1.8463**	-0.3326
	(0.8909)	(0.0776)	(0.8716)	(0.4330)	(0.7749)	(0.1432)	(0.8536)	(0.3207)
WE during the prenatal period	1.7866	0.9094	-0.0679	1.5712	-2.0356	0.3936	-1.5416	-0.0487
	(1.4197)	(1.1153)	(1.8092)	(1.6149)	(3.8222)	(0.9792)	(1.8975)	(2.0680)
Agricultural level	1.1294***	0.2935	0.5655***	-0.0296	0.3713	-0.0151	0.0158	-0.9633**
	(0.2424)	(0.3715)	(0.2156)	(0.2057)	(0.3373)	(0.2144)	(0.4162)	(0.3793)
Sub-sample	Rural	Urban	Rural	Urban	Rural	Urban	Rural	Urban
Other controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Province FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age in month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Survey-year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	668	2,806	1,380	3,743	665	2,938	1,418	3,821

Notes: Coefficients are the marginal effects of the probit estimation. Robust standard errors clustered at the provincial level are shown in parentheses. ***, **, and * indicate significance at the 1 percent, 5 percent, and 10 percent levels, respectively. The number of observations varies across the models for two reasons. First, each health indicator is

collected separately and some values are missing or excluded as outliers, as described in the note to Table 6. Second, in the probit estimations, we exclude regional birth cohort cells with no within-cell variation in the health-risk dummy (i.e., cells in which the dummy takes a value of 1 or 0 for all children). FE: Fixed effects.

Table 11

DID estimate of initial exposure intensity effect on birthweight.

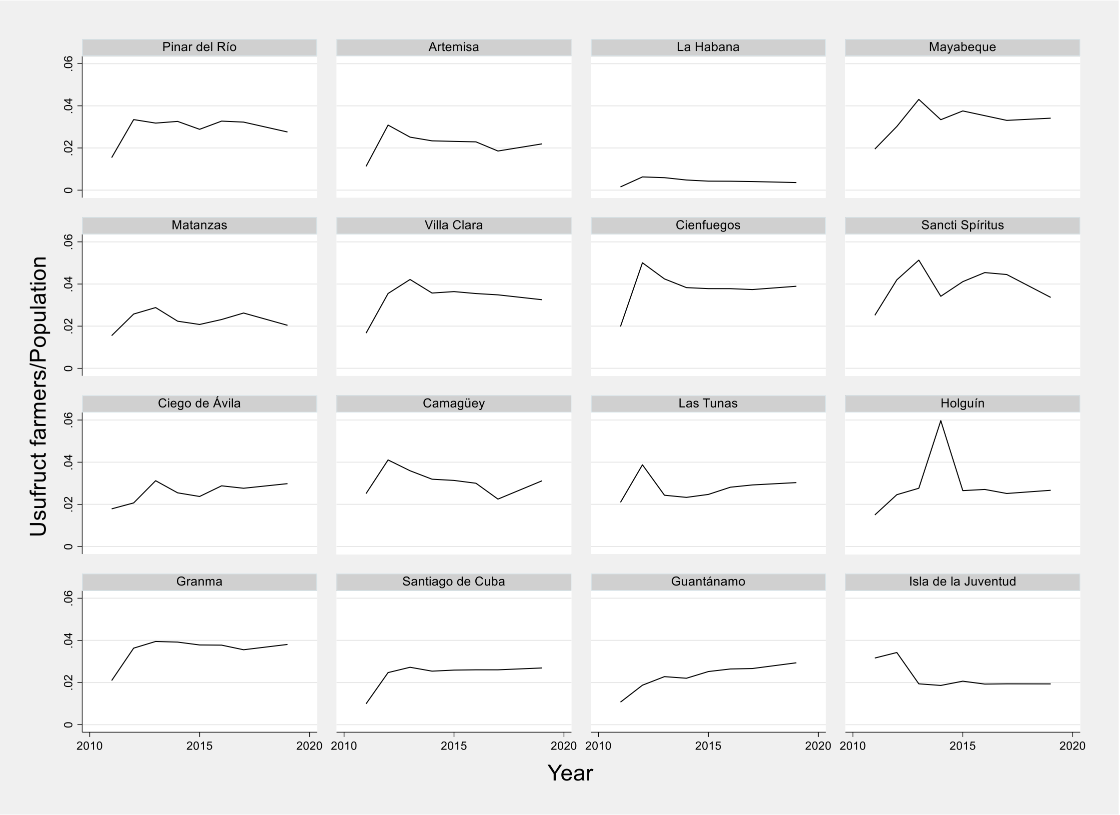
	Birthweight
$Time_{as} \times IC_p$	0.0836
	(0.1500)
Other controls	Yes
Province FE	Yes
Age in month FE	Yes
Survey-year FE	Yes
Observations	1,860
R-squared	0.1349

Notes: Robust standard errors clustered at the provincial level are shown in parentheses. ***, **, and * indicate significance at the 1 percent, 5 percent, and 10 percent levels, respectively. FE: Fixed effects.

Figures

Figure 1

Provincial diffusion of the land reform since 2011.

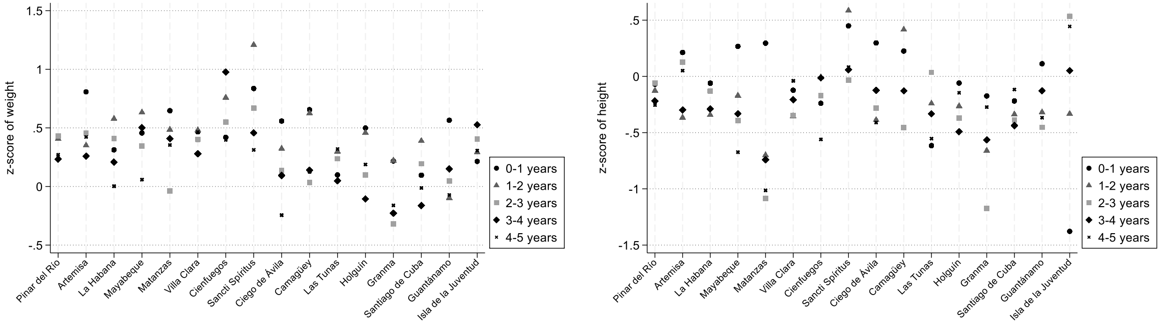


Notes: This figure demonstrates variations in the diffusion of land reforms across Cuban provinces. Holguín's 2014 figure for usufruct farmers is unusually large, suggesting potential measurement errors. Nevertheless, we retain this value because we could not verify misreporting.

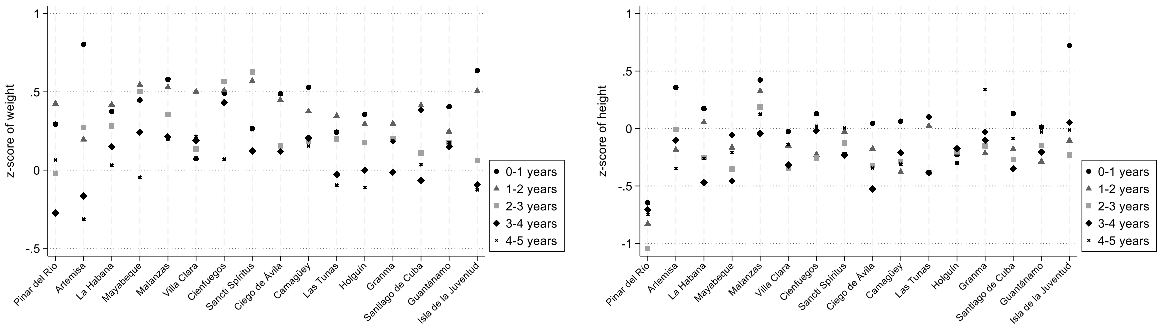
Figure 2

Mean z-scores by age cohort and province.

(a) Mean z-scores of weight-for-age and height-for-age in MICS 2014.



(b) Mean z-scores of weight-for-age and height-for-age in MICS 2019.

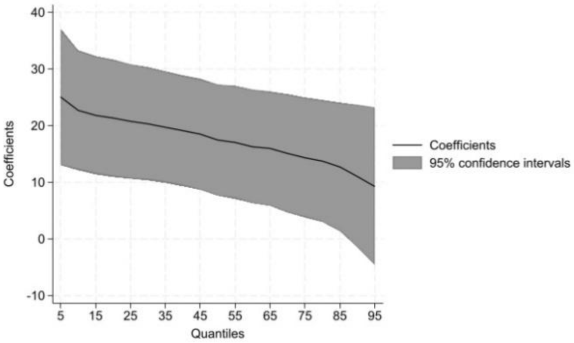


Notes: The figures plot mean z-scores by age cohort separately for each province. The vertical axis shows the corresponding z-score. All statistics are calculated by the author using the MICS.

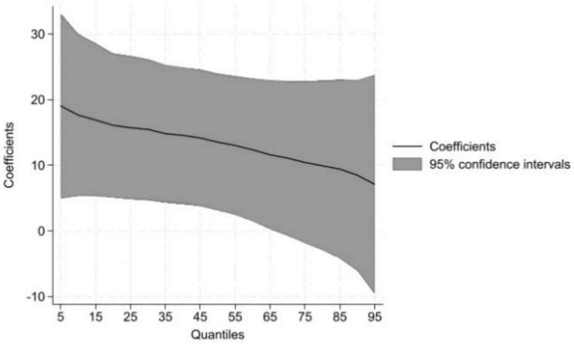
Figure 3

Quantile regression results for rural children.

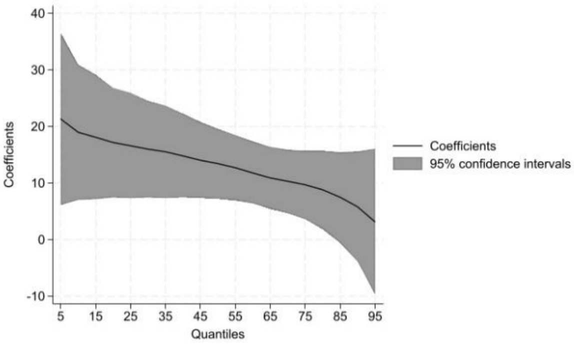
Weight-for-age



Height-for-age



Weight-for-height



Notes: The solid line plots the estimated quantile regression coefficients and the shaded area represents the 95% confidence intervals based on robust standard errors clustered by province.

Appendix

We estimate the effects of land reform by treating the initial changes in 2012 as paramount for determining the subsequent spread of the reform. Based on this assumption, our DID specification is

$$Y_{ipas} = \alpha + \beta Time_{as} IC_p + \boldsymbol{X}'\boldsymbol{\gamma} + \mu_p + \delta_m + \theta_s + \epsilon_{ipas}, \tag{A1}$$

where $Time_{as}$ is a dummy variable indicating the exposure to a reform shock during the prenatal period. Given that the reform was implemented on September 20, 2012, children born from 2014 onward are considered to have been exposed to reform shock during their fetal development.

Consistent with our main estimation, where lagged variables are considered, $Time_{as}$ takes the value for children born from November 2014 onward. Furthermore, leveraging the availability of exact birth month data and the reform enforcement date, we include a continuous dummy variable that increases by 0.083 (1/12) for each month. This continuous variable captures the intensity of births immediately after the reform shock. IC_p indicates the initial change in the number of usufruct farmers per capita and is a proxy for the initial land reform intensity in each province.