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REGULAR ARTICLE

The long-term impact of education on dietary diversity among women in Zimbabwe

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Abstract

Education is perceived to have a positive impact on a variety of health outcomes. However, it is unclear how causal this association is or what could account for the observed relationship, especially in low-income countries. This study examined the educational gradient in dietary diversity among young women using individual-level survey data from Zimbabwe. A parametric fuzzy regression discontinuity design was used in the empirical analysis, with school reform exposure serving as an instrumental variable for educational attainment. The results show that increased schooling improves dietary diversity among women and that this effect is large and statistically significant. An examination of the potential mechanisms by which education improves dietary diversity revealed that women with more education are more likely to engage with print media by reading newspapers or magazines, to be literate, to access prenatal care when pregnant, to be wealthier, to have fewer children, and to live in metropolitan areas. These findings suggest that expanding educational opportunities, particularly for young girls in developing countries like Zimbabwe, could be a useful policy strategy to promote healthy eating among young women and, as a result, could enhance population health and nutrition outcomes.

KEYWORDS

causal mediation analysis, dietary diversity, education, quasi-experiment, regression discontinuity design, young women, Zimbabwe

JEL CLASSIFICATION

C36, D13, I26

1 | INTRODUCTION

The physiological requirements for pregnancy and breastfeeding expose women of reproductive age to nutrition-related problems (Black et al., 2013). Pregnant or breastfeeding women have more significant nutritional needs than other adults. Inadequate micronutrient intake before and during pregnancy can negatively affect women and their infant's development, especially during the critical first two years of life (Cusick & Georgieff, 2016). A diverse and high-quality diet for the mother is strongly associated with improved health outcomes for her children, including a lower risk of low birth weight, preterm delivery, and developmental delays in later life (Huang et al., 2018; Madzorera et al., 2020; Rammohan, Goli, Singh, Ganguly, & Singh, 2019; Zhao et al., 2021). However, in many low-income countries, like Zimbabwe, women of reproductive age eat bland, starchy staple foods deficient in essential micronutrients (Arimond et al., 2010; Martin-Prevel et al., 2017). Improving access to education, particularly for girls, has been at the forefront of many governments' developmental agendas in low-income countries due to its well-documented positive effects on several health outcomes (Grossman, 2006). Given the disproportionate share of the burden of diet-related non-communicable diseases in low-income countries (Global Nutrition Report, 2022), there is a greater need to understand better what role education might play in improving nutrition-related outcomes today.

While previous research has found strong, positive correlations between education and various outcomes, evidence on the causal effects of education, including the mechanisms through which it influences outcomes, is limited and primarily concentrated in developed countries. Only a few studies have been conducted in low-income countries, particularly in sub-Saharan Africa (SSA) (Mensch, Chuang, Melnikas, & Psaki, 2019). The evidence in SSA has been limited partly due to data unavailability and the difficulty associated with establishing a causal interpretation of schooling (Agüero & Bharadwaj, 2014; Makate & Makate, 2018a). Health outcomes in the SSA region remain unsatisfactory, with several countries less likely to meet their Sustainable Development Goals (SDGs) obligations (Mejía-Guevara, Zuo, Bendavid, Li, & Tuljapurkar, 2019). For instance, several countries in the continent continue to experience higher levels of undernutrition among children under age 5, with others less likely to meet the second goal of the SDG, which calls for a global end to hunger through improved food security and better feeding practices (World Health Organisation, 2018). Expanding education opportunities, especially for young girls in low-income countries, remains among several policy options available to address salient health inequalities in the region.

This paper studies how education affects dietary diversity among women in Zimbabwe. To do so, we exploit a powerful quasi-experiment that increased schooling opportunities for Zimbabwean children following the country's independence from Britain in 1980. Compared to their relatively older counterparts who were 14 years and older, children who were within the secondary school starting age of 13 years or younger had greater exposure to this reform. We assume, as in previous related studies, that other post-independence policy changes did not affect children differently based on their age at independence (Agüero & Bharadwaj, 2014; Makate & Makate, 2018a). A clear understanding of the education gradient in dietary diversity among young women is essential for health policy and planning, particularly given the SSA's growing burden of noncommunicable diseases (World Health Organization, 2018). This paper is one of many that are looking into the causal effect of education in SSA using school reforms as natural experiments for identification.

The findings show that the 1980 school reform, which increased educational opportunities for children in Zimbabwe, had a positive and appreciably significant impact on young women's dietary diversity several years later. These large effects could reflect the combined effect of the school reform and other post-independence initiatives aimed at promoting nutrition in the country. While our empirical strategy could not differentiate between these effects due to a lack of data, including region fixed effects may mitigate some of the resulting bias. Furthermore, we identified potential explanations for the impact of education that can be connected to improved access to information, literacy, use of prenatal care, asset wealth, residence location, and intra-household competition for food resources.

The remainder of this paper is structured as follows. Section 2 reviews the empirical literature and briefly overviews the 1980 school reform in Zimbabwe. The data source, key variables, and econometric strategy are introduced and presented in Section 3. Section 4 presents the empirical results. Section 5 discusses the empirical findings. Finally, Section 6 concludes the study.

2 | BACKGROUND AND LITERATURE

2.1 | Literature review

The nonmarket effects of education have long been studied and empirically contested. Recent evidence has moved away from identifying simple associations and toward investigating causal relationships using quasi-experimental techniques (Mensch et al., 2019). These studies have mainly relied on exogenous variation in education brought on by universal schooling policies to identify and determine the causal impact of education. More research is required to fully understand the health effects of education, especially in low-income countries, according to a recent systematic review of the literature and meta-analysis, which found that the evidence currently available is far from conclusive (Hamad, Elser, Tran, Rehkopf, & Goodman, 2018; Mensch et al., 2019). In addition, a significant portion of this literature examines the relationship between education and outcomes for children, with very few studies investigating the causal relationship between education and outcomes for mothers or women.

According to the health economics literature, educational gains can lead to improved outcomes directly through allocative or productive efficiency (Grossman, 2006) or indirectly through educational benefits, such as increased wealth and earnings. Behrman (2015b) conducted a study using data from Uganda and Malawi that support the latter mechanism (Behrman, 2015b). Allocative efficiency occurs when more educated people choose an optimal mix of inputs (e.g., time and money) to maximize health gains compared to their lowly educated counterparts. In contrast, productive efficiency occurs when more educated people "obtain a larger output from a given level of health inputs than the less educated" (Grossman & Kaestner, 1997). Education enables individuals to process better health information that may be accessible through books, television, radio, and the internet (Bundorf, Wagner, Singer, & Baker, 2006; Wagner, Hu, & Hibbard, 2001). The latter mechanisms have also been supported in studies conducted in SSA (Agüero & Bharadwaj, 2014; Makate & Makate, 2016, 2018a, 2018b).

While there is growing evidence that education improves health outcomes in SSA, the evidence is mixed and still limited. Also, most of these studies focus on the impact on child health outcomes, with relatively few studies focusing on health outcomes for women (Mensch et al., 2019). Moreover, research has yet to be conducted to investigate the impact of education on dietary diversity among young women. For example, studies in several SSA countries discovered that having more education reduced the probability of adolescent sexual activity, marriage, and fertility (Alsan & Cutler, 2013; Baird, Chirwa, McIntosh, & Özler, 2010; Behrman, 2015a; Duflo, Dupas, & Kremer, 2015; Makate & Makate, 2018b; Moussa & Omoeva, 2020; Osili & Long, 2008; Ozier, 2018). Other studies have found that having more education improves one's knowledge of HIV (Agüero & Bharadwaj, 2014) and lowers the likelihood of testing positive for HIV (Bago, Ouédraogo, & Lompo, 2021; Behrman, 2015b; De Neve, Fink, Subramanian, Moyo, & Bor, 2015). In a different study, Behrman, Peterman, and Palermo (2017) show that an additional year of schooling reduces the risk of sexual violence by an estimated 9 percentage points among Ugandan women, but no such effect was found for Malawian women (Behrman et al., 2017). However, none of these studies have examined the causal impact of education on women's dietary diversity. Of the few studies that have explored factors associated with female dietary diversity in selected countries, none of them has examined the causal effect of education on dietary diversity (see, for example, Delil, Zinab, Mosa, Ahmed, & Hassen, 2021; Gitagia et al., 2019; Kiboi, Kimiywe, & Chege, 2017). This paper studies how education affects dietary diversity among women in Zimbabwe. To do so, we exploit a powerful quasi-experiment that increased schooling opportunities for Zimbabwean children following the country's independence from Britain in 1980 as a source of exogenous variation in education.

2.2 | The 1980 education reform in Zimbabwe

Before 1980, the Zimbabwean education system was characterized by inequalities in access to learning and grade progression for black children. Children for whites had access to free and compulsory education, whereas black children had neither free nor compulsory education (Dorsey, 1989). The colonial government made insufficient investments in black education, ensuring that black children faced numerous difficulties with grade progression (Nhundu, 1992). Educational reforms, according to the new government, were the only panacea to unlocking the envisioned endless economic opportunities for every citizen and included free primary and secondary education for all, the removal of age-related restrictions that allowed older children to re-enroll, the incorporation of vocational subjects such as food and nutrition into school curricula, the construction of new schools in disadvantaged areas, and the training of more teachers (Nherera, 1994). Consequently, enrolment at the primary and secondary school levels increased rapidly (see Figures 1 and 2), with the most notable increases witnessed at the secondary school level (Ansell, 2002; Mackenzie, 1988). Figures 1 and 2 summarize the changes in primary and secondary school enrolment rates. Primary school enrolment increased from about 800,000 in 1980 to slightly >2 million in 1987 and then to about 2.5 million in 1997.

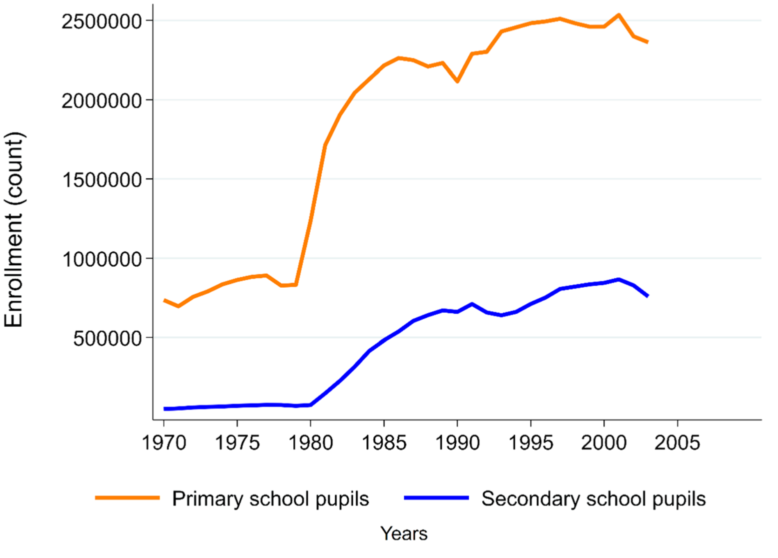

FIGURE 1 Primary and secondary school enrolments in Zimbabwe, 1970–2003. The author's elaboration is based on data from the World Development Indicators. [Colour figure can be viewed at wileyonlinelibrary.com]

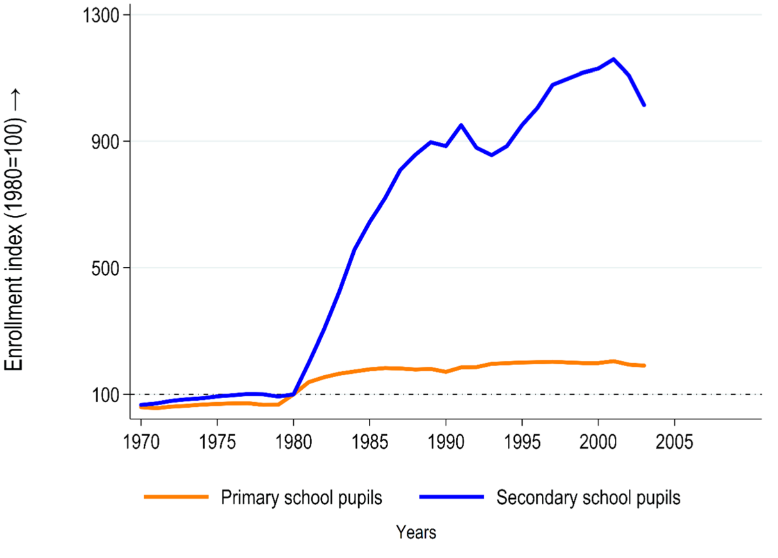

FIGURE 2 Primary and secondary school enrolments indexes in Zimbabwe, 1970–2003 (1980 = 100). The authors' elaboration is based on data from the World Development Indicators. The figure shows the evolution of enrolment rates at both primary and secondary school levels relative to 1980 (i.e., 1980 as the base or index year). [Colour figure can be viewed at wileyonlinelibrary.com]

At the secondary school level, enrolment increased from around 74,000 in 1980 to around 700,000 in 1991. The construction of new learning facilities and hiring more teachers accompanied enrolment increases.

Figure 2 depicts the distribution of the changes in enrolment rates at the primary and secondary school levels since 1970. In this case, 1980 serves as the base or index year against which we compare changes in enrolment over time. We show that the most significant changes or increases in enrolment occurred at the secondary school level. For instance, we observe a rapid increase in enrolment relative to 1980, from 100% in 1980 to *c*. nine times (900%) more significant by 1990. While the data at the primary level show significant improvements relative to 1980, the changes at the secondary level were much more pronounced. This observation may be explained by the fact that, prior to independence, numerous bottlenecks made it virtually impossible for black children to progress to secondary school. Children of blacks, for example, were often required to pass additional tests or exams before they could transition from primary to secondary school. Despite high levels of racial discrimination, there is no evidence that the education system before independence was gender biased.

Children in Zimbabwe attend school at age six and complete seven years of primary school before moving on to a four-year secondary education. After completing secondary school, students can enrol in advanced-level studies for two years before moving on to university if they meet the requirements.

3 | METHODS

3.1 | Data source

This study uses data from the 2005/2006 Demographic and Health Survey (DHS) for Zimbabwe drawn from Integrate Public Use Microdata Series (IPUMS) platform (Zimbabwe Central Statistcal Office and Macro International Inc, 2006). The DHS is a nationwide survey conducted by the Zimbabwe National Statistics Agency in collaboration with Inner City Fund (ICF) International. This survey collects health information from women aged 15–49 and their children. It employs a stratified two-stage cluster approach in which enumeration areas (EAs) are sampled first, followed by a list of households in EAs but excluding households in institutional arrangements. The Zimbabwe population census is used as a sampling frame for the survey. Several kinds of data, including employment, household assets, education, and maternal health indicators, are collected. The 2005/06 survey collected data on specific food items consumed by each respondent in the 24 hours preceding the survey (Office & Inc, 2007). We focus on young women for which information on dietary diversity was collected ($N = 2{,}706$). However, our analysis sample was determined by the data-driven optimal bandwidth selection procedure for the empirical analysis.

3.2 | Outcome variables

The primary outcome variable of this study is dietary diversity among young women. In particular, we rely on three different dietary diversity measures for the analysis. The first outcome measure is based on a continuous composite score reflecting the consumption of goods in the past 24 hours before the survey. The other two measures are binary indicator variables measuring the consumption of foods rich in vitamin A or iron. The first outcome measure is a composite score derived from the number and types of foods consumed by the respondent 24 hours before the survey. A complete description of these outcome measures is provided in Table 1. Consuming a food item from any food group was coded as one and zero otherwise. The food groups are (1) grains, white roots and tubers, and plantains; (2) pulses (beans, peas, and lentils);

(3) nuts and seeds; (4) milk, cheese, or other food made from dairy; (5) meat, poultry, fish, and other organ meats; (6) eggs; (7) any dark green, leafy vegetables; (8) vitamin A–rich yellow fruits and vegetables; (9) any other fruit or vegetables; and (10) fruit juice or sugary drink. We also created indicators measuring the consumption of foods rich in vitamin A, iron, and minimum dietary diversity—a measure of micronutrient adequacy and an essential marker for diet quality (FAO, 2016). Vitamin A is an essential micronutrient to the immune system, is instrumental in maintaining epithelial tissues in the body, and is associated with a lower risk of maternal mortality (WHO, 2009). Iron deficiency is a known risk factor for anemia—a known risk factor for mortality among pregnant mothers and their children.

3.3 | Explanatory variables

Education is measured using the number of years of completed schooling and as a dummy variable for secondary school completion as observed at the survey date. Our models include controls for weight-to-height ratio and region or province fixed effects. We excluded several variables that are potentially endogenous in the dietary diversity equation, including marital status, urban residence, partner's education and employment status, prenatal care utilization, and the number of children who are five years and younger (Chou, Liu, Grossman, & Joyce, 2010; Osili & Long, 2008). Education could impact dietary diversity through potential mechanisms, including access to information, empowerment, literacy, marital status, prenatal care utilization, partner's education, assets, number of children, and place of residence. Table 1 provides a complete description of these variables.

3.4 | Empirical strategy

This study examines the effect of education on dietary diversity using a fuzzy regression discontinuity (FRD) design. In an FRD design, exposure to the intervention is probabilistic rather than deterministic (Lee & Lemieux, 2010). The causal effect from an FRD design can be estimated through a parametric or nonparametric approach (Lee & Lemieux, 2010). Given the relatively small sample size and the absence of a critical mass of observations around the cut point, we first estimate the causal effect of education using the parametric approach and provide the nonparametric estimates as a sensitivity check (Jacob, Zhu, Somers, & Bloom, 2012). Following Imbens and Lemieux (2008), we define the FRD estimator as the ratio of the jump in dietary diversity and the discontinuous change in the probability of exposure to the 1980 school reform at the limit as we approach the cut point from either below or above (Lee & Lemieux, 2010):

$$\tau_{FRD} = \frac{\lim\limits_{c\to 13-} E[DDW|a80=c] - \lim\limits_{c\to 13+} E[DDW|a80=c]}{\lim\limits_{c\to 13-} P[reform=1|a80=c] - \lim\limits_{c\to 13+} P[reform=1|a80=c]}$$

where *reform* equals one if the woman was 13 years or younger in 1980 (the year of reform) and zero otherwise. The causal effect of the treatment can be identified under the assumption that there are no discontinuous shifts in observed and unobserved characteristics at the time of the policy except for the jump in the running variable. Under certain assumptions, Hahn, Todd, and Van Der Klaauw (2001) show that the FRD estimator is equivalent to an instrumental

TABLE 1 Summary statistics for selected variables used in the analysis

Variables	Variable descriptions	Aged 13 years or younger in 1980 ($N = 2{,}555$)	Aged 14 years or older in 1980 ($N = 151$)	*t*-Test (*p*-value)	Overall sample ($N = 2{,}706$)
Years of education	Completed years of schooling as at survey date	7.997	3.029	0.000	7.817
Secondary school or higher	=1 if completed secondary school or higher and zero otherwise	0.626	0.078	0.000	0.606
Age at survey	Age (years) at the time of the survey	26.030	43.275	0.000	26.654
Age in 1980	Age (years) in 1980	1.150	18.412	0.000	1.774
Weight/height	Weight divided by height	0.028	0.028	0.586	0.028
Employed	=1 if working at time of survey and zero otherwise	0.338	0.269	0.193	0.336
Regularly reads newspapers	=1 if respondent reads newspapers frequently and zero otherwise	0.372	0.130	0.000	0.363
Literate	=1 if able to read and write and zero otherwise	0.817	0.316	0.000	0.799
Married	=1 if the respondent was married and zero otherwise	0.840	0.853	0.680	0.841
Asset quintile 1 (poorest)	=1 if poorest and zero otherwise	0.237	0.364	0.014	0.242
Asset quintile 2	=1 if poorer and zero otherwise	0.216	0.293	0.109	0.219
Asset quintile 3	=1 if average wealth and zero otherwise	0.175	0.231	0.115	0.177
Asset quintile 4	=1 if rich and zero otherwise	0.216	0.076	0.002	0.211
Asset quintile 5 (richest)	=1 if richest and zero otherwise	0.156	0.036	0.001	0.152
Maternal dietary diversity score	An index representing food consumption from any of the 10 food groups in the 24 hours leading to the survey	2.974	1.942	0.000	2.936
Minimum dietary diversity	=1 if the respondent consumed food items from any of five food group and zero otherwise	0.196	0.086	0.008	0.192
Vitamin A–rich foods	=1 if the respondent consumed at least one food item that was rich in vitamin A in the 24 hours preceding the survey and zero otherwise	0.610	0.370	0.000	0.601

TABLE 1 (Continued)

Variables	Variable descriptions	Aged 13 years or younger in 1980 ($N = 2{,}555$)	Aged 14 years or older in 1980 ($N = 151$)	*t*-Test (*p*-value)	Overall sample ($N = 2{,}706$)
Iron-rich foods	=1 if the respondent had consumed at least one food item that was rich in iron in the 24 hours leading to the survey and zero otherwise	0.471	0.281	0.003	0.464
Urban resident	=1 if respondent lives in urban area and zero otherwise	0.291	0.065	0.000	0.283
Daily household purchases	=1 if the respondent has final say on household purchases for daily needs and zero otherwise	0.740	0.754	0.252	0.740
Large household purchases	=1 if the respondent has final say on making large household purchases and zero otherwise	0.768	0.760	0.109	0.767
Prenatal care visits	Number of prenatal care visits for the most recent pregnancy that occurred in the last 3 or 5 years before the survey	4.727	4.228	0.053	4.709
Uneducated husband/ partner	=1 if the respondent's husband or partner had never attended school and zero otherwise	0.024	0.207	0.000	0.030
Child under age 5	Number of children who are 5 years and younger in a household	1.742	1.959	0.040	1.750
Region 1	Manicaland	0.120	0.200	0.002	0.123
Region 2	Mashonaland central	0.116	0.102	0.962	0.115
Region 3	Mashonaland east	0.084	0.045	0.237	0.083
Region 4	Mashonaland west	0.096	0.079	0.830	0.095
Region 5	Matabeleland north	0.063	0.120	0.008	0.065
Region 6	Matabeleland south	0.048	0.059	0.309	0.048
Region 7	Midlands	0.148	0.116	0.450	0.146
Region 8	Masvingo	0.155	0.214	0.222	0.157
Region 9	Harare	0.119	0.057	0.048	0.117
Region 10	Bulawayo	0.051	0.008	0.035	0.050

Note: Estimates are weighted to be representative of women aged 15–49 years who had given birth to at least one surviving child under the age of three years.

variables (IV) estimator and estimable through two-stage least squares (2SLS) with the binary indicator for exposure to the reform serving as the excluded instrument. Children in Zimbabwe commence secondary learning around the age of 13 years. As such, we use the woman's age in 1980 as the running variable. In this instance, exposure to the school reform is measured using a dummy indicator that equals one if an individual was 13 years or younger in 1980 and zero otherwise. Following Lee and Lemieux (2010) and the related literature in low-income countries, the first stage regression takes the following form:

$$\text{educat}_i = a_0 + a_1\text{reform}_i + X'a_1 + f(a80_i - 13) + p_i + (p_i \times a80_i) + u_i \tag{1}$$

The second stage regression is formulated as follows:

$$DDW_i = \beta_0 + \beta_1\widehat{educat}_i + X'\delta_2 + f(a80_i - 13) + p_i + (p_i \times a80_i) + \varepsilon_i \tag{2}$$

where DDW_i is dietary diversity for the individual woman i; $educat_i$ represents her level of education and $\widehat{educat}_i$ is the predicted level of education from the first-stage regression; X is a vector of exogenous controls believed to be associated with dietary diversity; p_i are region fixed effects; $f(a80_i - 13)$ is a smooth function of age in 1980; $(a80_i - 13)$ is the distance between age in 1980 and the secondary school starting age; and e_i and u_i are disturbance terms. The term $f(a80_i - 13)$ incorporates the component of the running variable so that potential shifts observed at the cut point can be construed as changes in the intercept. This function $f(.)$ also includes an interaction term $[(a80_i - 13) \times \text{reform}_i]$ to account for potential differences in the gradients on either side of the cut point (Jacob et al., 2012). The distance from the discontinuity point viewed from below or above the cut point is paramount. We use a data-driven approach to bandwidth selection as suggested by Calonico, Cattaneo, Farrell, and Titiunik (2017). The optimal bandwidth chosen using the mean squared error-optimal bandwidth selection procedure was $(+/-4)$ years on either side of the discontinuity. However, given that there is always a trade-off between bias and precision of the estimates, we also considered a smaller $(+/-3)$ and wider bandwidths of sizes $(+/-5)$ and $(+/-6)$. As the bandwidth gets smaller, the estimates become imprecise but less likely to be biased. On the contrary, the bigger the bandwidth, the more precise the estimates and the higher the probability of getting biased estimates (Calonico et al., 2017). Following the recommendations of Gelman and Imbens (2019), we include only first-order polynomials and avoid the inclusion of higher-order polynomials, given the possibility of generating noisy estimates.[1] It is conceivable that the school reform in Zimbabwe was likely to have differential effects across the 10 regions. Thus, we include the interaction between the region and the individual's age in 1980, $p_i \times a80_i$ as a proxy measure for the intensity of the school reform.

Our interest lies in β_1—a measure of the effect of female education on dietary diversity among women and is identified as the local average treatment effect (LATE). In the instance when the system is exactly identified, the coefficient β_1 is comparable to the coefficient τ_{FRD} and can be interpreted as the weighted LATE with the weights representing the relative ex-ante probability that an individual lies within the neighborhood of the cut point (Lee & Lemieux, 2010). We use linear models to estimate Equations 1 and 2 and test for the functional form of these equations through the inclusion of low-order polynomials, as well as interactions of these polynomials with the treatment group dummy as is common practice in the regression discontinuity (RD) literature (Gelman & Imbens, 2019). A central tenet of RD designs is that no one can manipulate the running variable (in our case, age at the time of school policy in 1980) in relation to the cutoff age. To formally verify this, we visualize the density of the running variable using a histogram, as suggested by McCrary and Royer (2011). Figure A1 in the Supplementary Appendix S1 shows no evidence of bunching around the age 13 cutoff.

In addition, we provide the reduced form or intent-to-treat (ITT) estimates where we express dietary diversity among women as a function of the education reform dummy, the polynomial function $f(a80_i - 13)$ together with other explanatory variables as mentioned earlier. The reduced form equation we estimate takes the following form:

$$DDW_i = \gamma_0 + \gamma_1 reform_i + X'\gamma_2 + f(a80_i - 13) + p_i + (p_i \times a80_i) + \epsilon_i \tag{3}$$

where the coefficient γ_1 measures the ITT effect, which also has a causal interpretation of the offer of treatment (Angrist, 2009). All our models correct for potential specification error in the polynomial by calculating robust standard errors that are clustered at two dimensions, namely, age cohort in 1980 and region of residence. Given the smaller number of clusters involved, we calculate bootstrapped standard errors with 500 replications to enhance the precision of our estimates. Our preferred estimates are those calculated using the optimal bandwidth of size +/−4, which was selected using a data-driven procedure as suggested in Calonico, Cattaneo, and Titiunik (2014). We present the parametric estimates as the main empirical results but present their nonparametric counterparts as supplementary material and as a robustness check. We also conduct a formal causal mediation analysis to identify the potential mechanisms through which female education might influence dietary diversity (see Appendix S1). All analysis was conducted using Stata version 15.1 (StataCorp., 2017).

4 | RESULTS

4.1 | Descriptive statistics

The sample-weighted descriptive statistics presented in Table 1 show that the average years of completed schooling for women aged 13 years or younger in 1980 was 7.997 compared to 3.029 among those aged 14 years and older in 1980 and less likely to have benefited from the school reform. This difference in years of education is statistically significant at the 1% level. Secondary school completion is significantly higher in the cohort of women younger than 13 years in 1980 (the exposed cohort) compared to the comparison cohort (62.6 vs 7.8%) and statistically significant at the 1% level. The *t*-test results indicate significant differences between the treatment and comparison groups on some observable characteristics except employment, household asset quintiles 1 and 3, and marital status. Also, the urban residence is much higher among the younger cohort of women than those who were relatively older at the time of the reform (29.1 vs 6.5%).

4.2 | The 1980 school reform and female education in Zimbabwe

Figure 3 shows that the 1980 school reform in Zimbabwe increased both the average number of years of completed schooling (top graph) and the probability of completing secondary school

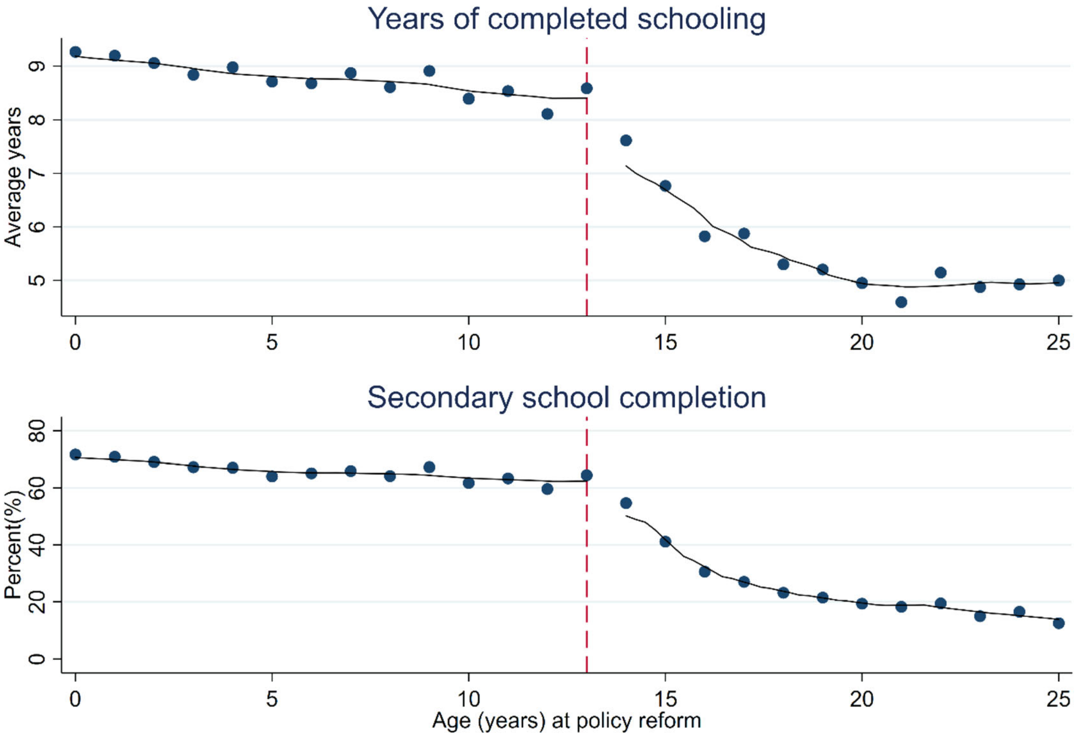

FIGURE 3 Distribution of average years of schooling and secondary school completion by age at the time of school reform in 1980. Author's elaboration based on data from the 2005–06 Zimbabwe Demographic and Health Survey (ZDHS). [Colour figure can be viewed at wileyonlinelibrary.com]

(bottom graph). The top graph in Figure 3 shows the effect of the school reform on average years of completed schooling with the sample divided into treatment (the likely beneficiaries who were aged 13 years and younger in 1980), left panel, and comparison (those who were less likely to benefit from who were 14 years and older in 1980), right panel. In Zimbabwe, a typical child officially starts her first year of secondary school at the age of 13 years with some starting even a year earlier or just few years older. Thus, data points to the left of the vertical line correspond to women born in 1967 or later, while the data points to the right of the vertical line comprise women born in 1966 or earlier. The discontinuity or disconnect in average years of completed schooling and secondary school completion appears to be more pronounced at the age 13 cut point. Each dot in the graph represents the average years of completed schooling or the percentage of females who completed secondary school in each age cohort with age referring to the age at the time of the reform in 1980.

Figure 4 shows the impact of the school reform on educational attainment. Note that this graph was constructed using all the 2005-06 Zimbabwe Demographic and Health Survey (ZDHS) data and incorporates sample probability weights to ensure representativeness of the data. Women who were aged 13 years and younger in 1980 had the highest secondary school completion rate (about 59%) compared to their counterparts who were aged 14 years and older (17%). In our analysis sample, about 25% of individuals who were older than 13 years in 1980 never attended school compared to about 2% among individuals who were aged 13 years and younger in 1980. Most individuals (57%) in the relatively older cohort completed primary school

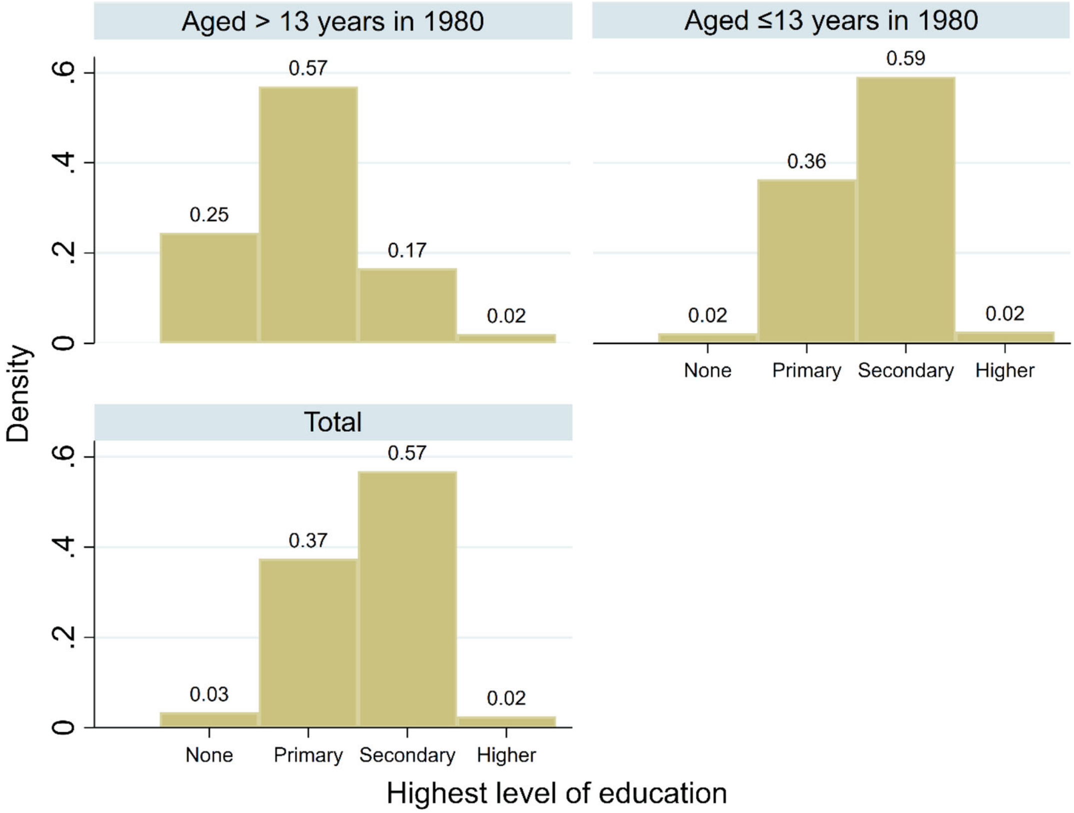

FIGURE 4 The effect of the 1980 school reform on educational attainment of young women in Zimbabwe. Author's elaboration based on data from the 2005–06 Zimbabwe Demographic and Health Survey (ZDHS). [Colour figure can be viewed at wileyonlinelibrary.com]

education compared to 37% in the cohort aged 13 years and younger. This graph shows the general differences in educational attainment for individuals who were young enough to have benefited from the reform compared to those who were relatively older and less likely to have benefited from the school reform.

4.3 | First-stage regression estimates

The first-stage regression estimates of the effect of the 1980 education reform on years of completed schooling and the probability of finishing secondary school are presented in Table 2. The results show that exposure to the school reform was associated with increases in years of schooling ranging from 3.498 to 4.111 for the +/3 years bandwidth and +/−5 years bandwidth, respectively, and all statistically significant at the 1% level. Similarly, the probability of completing secondary school ranged from 0.215 to 0.326 for the +/−4 and +/−6 years bandwidths, respectively, and all were statistically significant at the 1% level.

The standard checks for the validity of the IV (binary variable for exposure to the 1980 school reform) used in the first-stage regression and presented in Table 3 show that the IV is strong and valid. An IV is considered to be strong if the first-stage *F*-test statistic exceeds

10 (Staiger & Stock, 1997; Stock & Yogo, 2005; Wooldridge, 2010). Given the results of the first-stage regression showing a strong and significant influence of the reform on schooling together with the F-test statistics reported in Table 3 that range from 20.079 ($p < 0.024$) to 84.846 ($p < 0.017$), our IV is indeed strong and relevant.

4.4 | Female education and maternal dietary diversity in Zimbabwe

To assess the extent to which maternal dietary diversity outcomes vary with the level of education, Figure 5 plots the raw and observed relationship generated using local polynomial regressions in which dietary diversity is regressed on the number of years of schooling. The results show a linear and increasing association between dietary diversity and years of completed schooling. However, these results are primarily descriptive and not suggestive of any significant associations.

Figure 6 shows the impact of the school reform on the dietary diversity composite score (top graph) and minimum maternal dietary diversity (bottom graph). Both graphs have a vertical line dividing data points for women in the treatment group (data points to the left) and women in the comparison group (data points to the right). The results show that average dietary diversity scores are consistently higher for women in the treatment group than in the comparison group. The discontinuity in schooling coincides with the gap in maternal dietary diversity scores. The bottom section of Figure 6 also shows that women in the treatment group are more likely to have a minimum maternal dietary diversity than their counterparts in the comparison group.

Figure 7 compares the average maternal dietary diversity scores regarding the consumption of food items considered rich in vitamin A (top panel) and iron for the treatment and control groups. The results show that, on average, women aged 13 years and younger in 1980 had relatively higher scores for vitamin A–rich foods than their relatively older counterparts. A similar pattern can be observed for scores related to iron-rich foods.

4.5 | The effect of female education on dietary diversity

Table 3 presents the main findings (parametric estimates) for the relationship between the dietary diversity of young women and increased schooling. The ordinary least squares (OLS) estimates, which do not consider the possibility that there are observed and unobserved factors that could affect a person's level of education and dietary diversity, are shown in the topmost section of Table 3. The results show that increased female schooling is positively and significantly associated with dietary diversity, starting with our preferred estimates in Column (2). More specifically, a 1-year increase in education is associated with an approximate 0.14 increase in dietary diversity score and is statistically significant at the 1% level. The findings in Columns (1), (3), and (4) demonstrate a positive correlation between female education and dietary diversity, serving as a robustness check.

The second-stage results, in which we account for the potential endogeneity of female education, are presented in the bottom section of Table 3. In this case, the binary indicator for exposure to the 1980 school reform is an IV for years of completed schooling. The results indicate that a 1-year increase in schooling increases dietary diversity by about 0.474 for individuals in the +/−4 age cohort and is statistically significant at the 1% level. Given that the average

TABLE 2 First-stage results of the effect of the 1980 education reform on the schooling of young women in Zimbabwe

Specifications	Size of bandwidths for above (+) and below (−) the cut point: +/−3 years (1)		+/−4 years (2)		+/−5 years (3)		+/−6 years (4)	
Panel 1: Years of completed schooling								
Exposure to the 1980 school reform	3.498*	(0.335)	4.061*	(0.828)	4.111*	(0.539)	3.661*	(0.436)
Average outcome variable—Treatment	7.015		7.238		7.368		7.403	
Average outcome variable—comparison	3.507		3.594		3.658		3.648	
Observations	200		280		367		468	
Panel II: Probability of completing secondary school or higher								
Exposure to the 1980 school reform	0.256*	(0.044)	0.215*	(0.040)	0.270*	(0.043)	0.326*	(0.043)
Average outcome variable—Treatment	0.474		0.500		0.523		0.526	
Average outcome variable—comparison	0.104		0.094		0.097		0.102	
Observations	200		280		367		468	

Notes: Robust standard errors in parentheses and clustered at the region of residence and age in 1980. In all specifications, we included linear age polynomials (above and below the cut point), individual's weight-to-height ratio and region fixed effects.
*significant at 1% level.

dietary diversity scores among women in the +/−4 age cohort were about 2.722 (see Table 1), the 0.474 increase following a 1-year rise in schooling translates to an estimated 17.41% $\left[\left(\frac{0.474}{2.722}\right) \times 100\right]$ increase in overall dietary diversity. The results also show that our estimates are robust, considering a smaller (+/−3) bandwidth and wider age bandwidths of +/−5 and +/−6, respectively. For example, a one-year increase in female schooling increases dietary diversity by about 0.373 among women in the +/−3 age bandwidth and is statistically significant at the 1% level. As the age bandwidth increases beyond the +/−4 years, the effect of education decreases.

In addition, Table 3 shows that the OLS estimates are much smaller than their 2SLS equivalents. The primary reason for the difference in these estimates is endogeneity because the OLS and IV estimates have identical signs and statistical significance. This endogeneity may arise due to the possibility that education is measured with error, especially in cross-sectional survey data sets such as the DHS we use here (Becker, 2016). The latter is conceivable, given that the DHS measures education as observed at the survey date. Some females would not have completed schooling then, resulting in an underestimation of education. Given that the IV estimator is not affected by the measurement error in the treatment variable, it tends to be larger than its OLS equivalent (Becker, 2016).

Table 4 shows the 2SLS estimates of the impact of female education on the probability of consuming food items rich in vitamin A (Panel I of Table 4) and iron (Panel II in Table 4). Starting with the estimates in Panel I, the results from Column (2) (our preferred estimates) show that a 1-year increase in female education increases the probability of consuming at least one food item that is rich in vitamin A by about 11.4 percentage points (pp) and statistically

TABLE 3 Two-stage least squares (2SLS) estimates of the effect of education on dietary diversity among young women in Zimbabwe

Specifications	Size of bandwidths for above (+) and below (−) the cut point							
	+/−3 years (1)		+/−4 years (2)		+/−5 years (3)		+/−6 years (4)	
OLS estimates								
Years of completed schooling	0.129**	(0.017)	0.140**	(0.016)	0.140**	(0.012)	0.135**	(0.012)
Observations	200		280		367		468	
IV/2SLS estimates								
Years of completed schooling	0.373**	(0.050)	0.474**	(0.048)	0.296**	(0.073)	0.212*	(0.080)
Average maternal dietary diversity score —Treatment	3.000		2.869		2.944		2.934	
Average maternal dietary diversity score—Comparison	2.119		2.189		2.183		2.148	
First-stage F-statistic (*p*-value)	84.846	(0.017)	20.079	(0.024)	50.555	(0.030)	62.723	(0.026)
Observations	200		280		367		468	

Notes: Robust standard errors are shown in parentheses and clustered at the region of residence and age in 1980. Additional control variables included as in Table 2.

Abbreviations: IV, instrumental variable; OLS, ordinary least squares.

*significant at 5% level, and.

**significant at 1% level.

significant at the 1% level. When we consider variations in the bandwidth, the results show a positive and significant impact of female education on the probability of consuming at least one food item rich in vitamin A in the 24 hours preceding the survey. The latter finding suggests that the main estimates presented in Column (2) are robust to alternative bandwidth sizes. Panel II of Table 4 presents the model results estimating the probability of consuming at least one food item rich in iron in the 24 hours preceding the survey. The estimates in Column (2) show that a 1-year increase in female schooling increases the probability of consuming at least one food item rich in iron by an estimated 9.6 pp and is statistically significant at the 1% level. When considering tighter and wider age bandwidths, the results suggest that our Column (2) estimates are robust to these changes. We also report the results from reduced form estimates of the impact of schooling reform on dietary diversity (see Table A6 in Appendix S1). The results show that the school reform significantly impacted the dietary diversity of young women.

4.6 | Pathways through which education might influence dietary diversity

Several variables were considered potential explanations for the observed education gradient in dietary diversity among young women in Zimbabwe. These variables are described in Table 1.

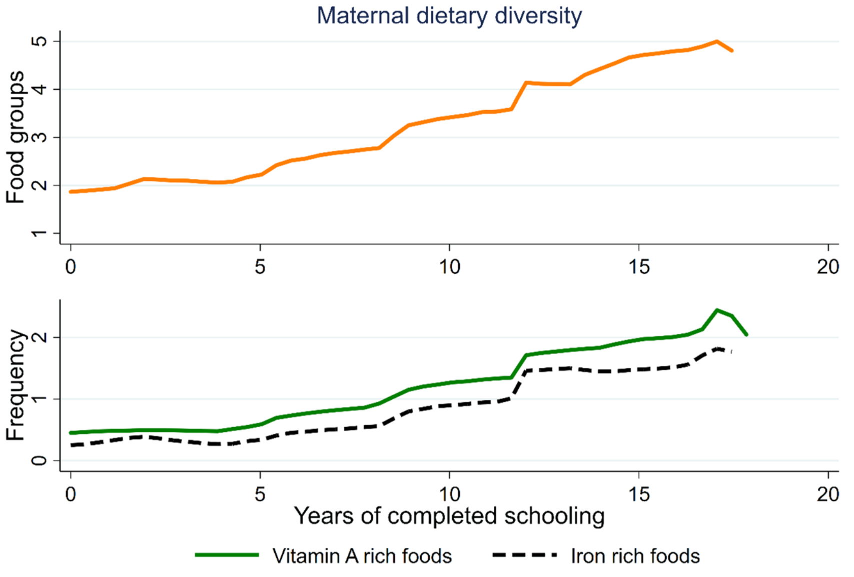

FIGURE 5 The distribution of the dietary diversity of young women by years of schooling. The figure plots local polynomial regressions of dietary diversity on years of completed schooling. The dependent variable is either the dietary diversity score or the number of vitamin A–rich foods or the number of iron-rich foods, and the explanatory variable is years of completed schooling of young women. [Colour figure can be viewed at wileyonlinelibrary.com]

According to Table 5, exposure to school reform was associated with a higher likelihood of reading newspapers, which was statistically significant at the 1% level. Reading newspapers mediated an estimated 16.6% of the total effect of the intervention on dietary diversity. The results also show that exposure to education reform was linked to increased literacy. Literacy mediated *c*. 50.7% of the total effect of the intervention and was statistically significant. The frequency of prenatal care utilization mediated an estimated 9.3% of school reform's effect on young women's dietary diversity. Furthermore, women exposed to the school reform were less likely to come from households with fewer children under five. In other words, the number of children under age five within a household mediated about 7.4% of the effect of the school reform on dietary diversity. The findings also show that exposure to the reform was associated with a higher likelihood of living in an urban area. We discovered that living in an urban area mediated *c*. 35.6% of the total effect of the school reform on dietary diversity. In addition, exposure to school reform was associated with a lower probability of coming from a low-income family-low household wealth mediated an estimated 21.98% of the total effect of the school reform on dietary diversity. Overall, the top mediating factors identified were access to information through reading newspapers, improved literacy, increased frequency of prenatal care use, decreased fertility, decreased probability of having low wealth, and increased probability of living in an urban area. A series of sensitivity checks on these findings ensured the robustness of our estimates (see Supplementary Appendix S1).

5 | DISCUSSION

Using representative survey data, we examined the education gradient in dietary diversity among young women in Zimbabwe. The 1980 school reform expanded secondary school

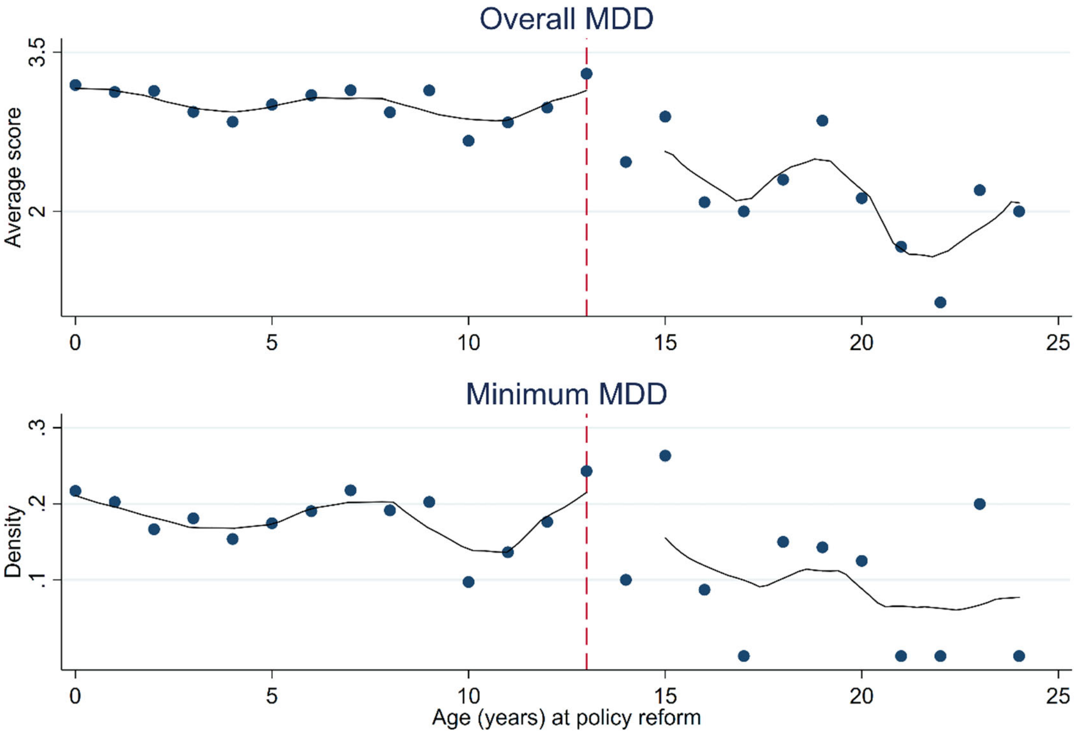

FIGURE 6 The effect of the 1980 school reform on dietary diversity and minimum dietary diversity scores of young women in Zimbabwe. Author's elaboration based on data from the 2005–06 Zimbabwe Demographic and Health Survey. [Colour figure can be viewed at wileyonlinelibrary.com]

opportunities for children throughout the country. Using the reform as an exogenous instrument for schooling, we found a strong and significant education gradient in dietary diversity. Increased female schooling was also associated with consuming vitamin A or iron-rich foods and adhering to the minimum recommended diet. We also found evidence supporting mechanisms related to frequent engagement with print media by reading newspapers or magazines, improved literacy, access to prenatal care when pregnant, improved wealth, reduced fertility, and residence in metropolitan areas. These results are robust to several sensitivity checks and consistent with previous evidence (For example, Keats, 2018; Makate & Makate, 2018c; Murendo, Nhau, Mazvimavi, Khanye, & Gwara, 2018; Zanin, Radice, & Marra, 2015).

Increased schooling promotes the dietary diversity of young women. Several factors could account for this outcome. Increased literacy and health knowledge assist women in comprehending the benefits of a well-balanced diet. Increased use of prenatal care could also explain the education gradient in dietary diversity. Given that the analysis sample consists of young women between the ages of 15 and 49, maternal healthcare utilization may explain the education gradient due to its link to dietary diversity through health education, advice, and counselling provided to women during pregnancy (Lincetto, Mothebesoane-Anoh, Gomez, & Munjanja, 2006). Healthy eating is crucial for pregnant mothers and their unborn children (Lincetto et al., 2006). Pregnant women may learn about the value of consuming foods high in vitamin A and iron during prenatal care appointments (World Health Organization, 2013). Compared to uneducated women, educated women are more likely to internalize and process

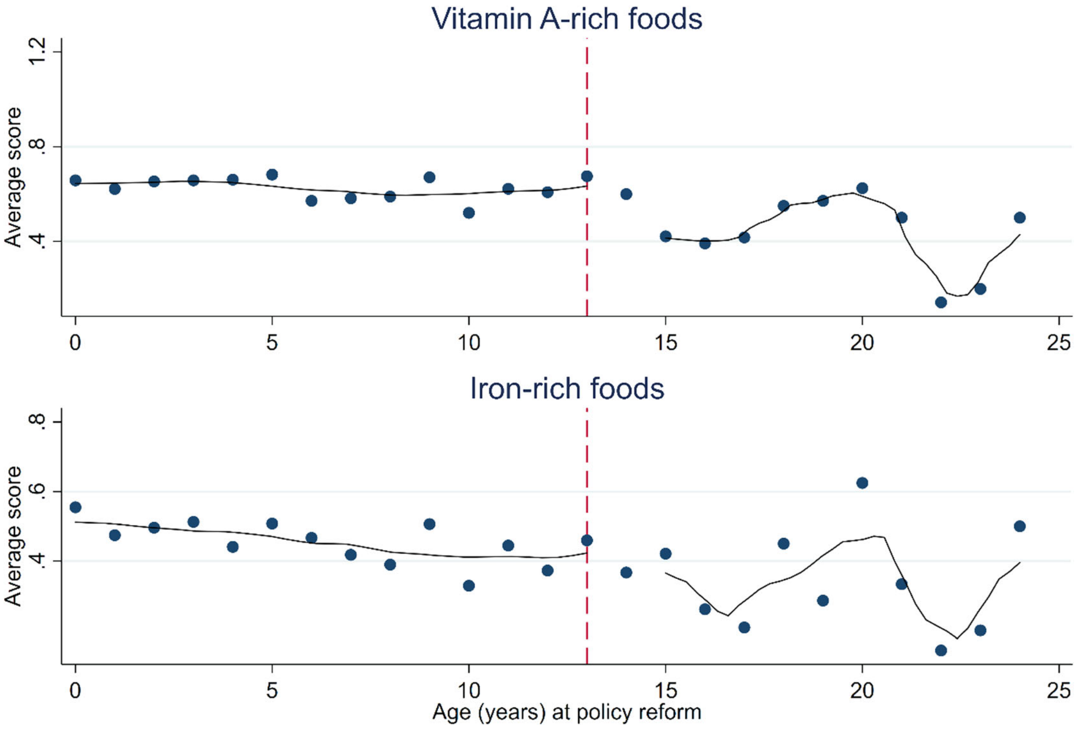

FIGURE 7 The effect of the 1980 school reform on the average number of vitamin A- and iron-rich foods consumed in Zimbabwe. Author's elaboration based on data from the 2005–06 Zimbabwe Demographic and Health Survey. [Colour figure can be viewed at wileyonlinelibrary.com]

nutritional advice better (Makate & Makate, 2018a). For instance, educated women may know that iron deficiency during pregnancy increases the risk that their unborn child will develop anemia and low birth weight, among other health problems. Also, educated women can better understand nutritional information advertised in publications or broadcast on radio and television (Bishow Parajuli & Ayoya, 2018).

Improved literacy also has an impact on dietary diversity. As mentioned by Makate and Makate (2018a), classroom numeracy and literacy skills help students understand the importance of better nutrition to overall health. Nutrition and diets are taught in Zimbabwe's primary and secondary schools. Food and nutrition were among the many vocational or technical subjects introduced into the secondary school curricula following independence in 1980 (Nherera, 1994; Zimbabwe School Examinations Council, 2012). Thus, reviewing these curricula and implementing school-based health and nutrition initiatives may improve the health and nutrition of the nation. Dietary diversity may also be influenced by household wealth. Several studies have shown that education increases wealth. People with a higher level of education are more likely to generate income and save to build wealth (Emmons & Noeth, 2015).

Our results also indicate that dietary diversity is inversely related to having children under five residing in a household. The number of children in a family may indicate intra-household resource competition, which may reduce dietary diversity (Kaul, 2018), which may reduce dietary diversity. Dietary diversity was also positively associated with urban residence. This finding reinforces that educated people are more likely to relocate to cities for work (Speare &

TABLE 4 Two-stage least squares (2SLS) estimates of the effect of education on dietary diversity outcomes among young women in Zimbabwe

Specifications	Size of bandwidths for above (+) and below (−) the cut point +/−3 years (1)		+/−4 years (2)		+/−5 years (3)		+/−6 years (4)	
Panel I: Vitamin A–rich foods								
Years of completed schooling	0.117**	(0.020)	0.114**	(0.024)	0.102**	(0.010)	0.096**	(0.010)
First stage F-statistic (*p*-value)	84.846	(0.017)	20.079	(0.024)	50.555	(0.030)	62.723	(0.026)
Average (%) outcome variable—Treatment	0.631		0.592		0.614		0.608	
Average (%) outcome variable—Comparison	0.448		0.459		0.476		0.477	
Observations	200		280		367		468	
Panel II: Iron-rich foods								
Years of completed schooling	0.130**	(0.027)	0.096**	(0.029)	0.099**	(0.018)	0.075*	(0.027)
First-stage F-statistic (*p*-value)	84.846	(0.017)	20.079	(0.024)	50.555	(0.030)	62.723	(0.026)
Average (%) outcome variable—Treatment	0.421		0.388		0.421		0.413	
Average (%) outcome variable—Comparison	0.298		0.297		0.329		0.329	
Observations	200		280		367		468	

Notes: Robust standard errors are shown in parentheses and clustered at the region of residence and age in 1980. In all the regressions, we included additional control variables as described in Table 3.
*significant at 5% level, and.
**significant at 1% level.

Harris, 1986). Our empirical analysis demonstrates that education promotes dietary diversity and that educated women are more likely to consume vitamin A- and iron-rich foods and adhere to the minimum recommended diet. The IV estimates we find are significantly larger than the OLS estimates suggest.

5.1 | Larger instrumental variable estimates

The impact of education on dietary diversity is significantly larger in all IV estimates presented in this paper than the OLS estimates suggest. There are several possible explanations for this. If the IV estimate assumes that school reform is the only source of variation in education, the estimated effect of education may be biased and overestimate the actual effect of education reform. The results, however, might also be a result of other countrywide developmental changes or how well the education system has adapted to changes in the curricula that have occurred since independence. First, we draw attention to the possible contribution that political stability and peace after independence may have contributed to improved school enrolment. Second, after independence, adjustments were made to the school curriculum that made it possible to include food and nutrition topics in the primary and secondary school curricula (UZ/MSU, 1990). As a

TABLE 5 Causal mediation estimates exploring the potential mechanisms through which education might influence dietary diversity among young women in Zimbabwe

	Reads newspapers	Final say on household purchases	Final say on large household purchases	Literate	Married
Exposure to the 1980 school reform	0.660 (0.176)***	0.798 (0.189)***	0.796 (0.188)***	0.379 (0.185)*	0.793 (0.186)***
Mediating variable	1.017 (0.186)***	−0.040 (0.205)	−0.023 (0.208)	1.084 (0.153)***	0.011 (0.264)
% of total effect mediated	16.6	−0.60	−0.30	50.7	0.00
Observations	468	468	468	468	468
	Number of prenatal care visits	**Uneducated husband/partner**	**Children under 5 years**	**Low household wealth**	**Urban residence**
Exposure to the 1980 school reform	0.718 (0.184)***	0.766 (0.185)***	0.735 (0.184)***	0.623 (0.177)***	0.509 (0.174)**
Mediating variable	0.102 (0.033)**	−0.304 (0.288)	−0.220 (0.096)*	−0.898 (0.156)***	1.660 (0.211)***
% of total effect mediated	9.30	3.50	7.40	21.98	35.60
Observations	468	468	468	468	468

Notes: The estimates shown here represent the average of 500 draws of a Monte Carlo Simulation. All equations included linear age polynomials (below and above the cut point) and region fixed effects as additional explanatory variables. The column headers represent the variables that are included as mediating variables in each of the models estimated. Estimates are based on the sample using the optimal bandwidth of size (+/−6). The estimates reported here are generated using Equation 6.

*significant at 10% level,

**significant at 5% level, and.

***significant at 1% level.

result, the effect might reflect how well the education system has done over time at teaching nutrition concepts in classrooms (Makate, 2017). Lastly, between 1980 and 2006, several large-scale community-based programs and primary healthcare policies were implemented to address the nation's health and nutrition situation. The Community Food and Nutrition Program (CFNP), a predecessor to the Supplementary Feeding Production Program (SFPP), launched in 1987 and aimed at children living in drought-prone areas with a higher risk of malnutrition, is one example of such initiative (Ismail, Immink, Immink, Mazar, & Nantel, 2003). The CNFP focused heavily on improving food production and access at the local village levels, with the program's success resting on community participation. Given that no data is available to evaluate this program's success, it is not easy to gauge its overall impact on dietary diversity, even though its impact cannot be minimized. Some women in our analysis sample living in areas targeted by the program could have benefited from the program's provisions. However, as the CFNP program was prematurely discontinued due to a lack of fiscal space and dwindling donor support (Tagwireyi, 2006), we have no reason to believe it is driving our results. Moreover, unlike the 1980 school reform, we have no reason to believe that the CNFP would have an age-specific impact that could explain our findings.

5.2 | Limitations of the study

There are a few limitations to this study. First, we acknowledge the possibility of bias due to other developmental initiatives implemented after independence. Second, the cross-sectional nature of the data is a possible shortcoming of the analysis because it precludes us from exploring the dynamic effects of female education on dietary diversity. Third, while our measure of maternal dietary diversity is consistent with standard recommended approaches (FAO, 2016), it is possible that it could still be measured with error given that it is computed using self-reported information regarding the specific food items that the respondent had consumed in the 24 hours leading to the survey. Fourth, the DHS data we use rely on dietary information collected from a sample of women who had at least one surviving child under the age of 3 years and living with them at the time (Zimbabwe Central Statistical Office and Macro International Inc, 2006). Given that female education is known to reduce fertility (Osili & Long, 2008), the composition of our analysis sample may be somewhat different for individuals who were just below the secondary starting age and those who were relatively older or above the secondary school starting age. Regardless of our concerns, this study adds to the growing body of evidence regarding the effect of female education on health and nutrition outcomes in SSA.

6 | CONCLUSION AND POLICY IMPLICATIONS

This paper establishes important empirical evidence that increased female education promotes dietary diversity among young women—an essential correlate of child dietary diversity and nutrition. The results provide insight into possible explanations of the observed education gradient in dietary diversity. We found possible explanations linked to improved access to information, literacy, prenatal care use, asset wealth, location of residence, and intra-household competition for food resources. The findings highlight the importance of increasing schooling opportunities, especially for young girls, to improve population health outcomes. In addition, findings from the causal mediation analysis underscore the need for a multi-sectoral approach to improving dietary quality. Overall, the results demonstrate that expanding access to secondary school opportunities, especially for young girls in developing countries, could be a useful policy strategy to promote healthy eating among women and consequently improve population health outcomes for both women and the next generation.

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DATA AVAILABILITY STATEMENT

The data used in this paper are from the 2005-06 Zimbabwe Demographic and Health Survey (ZDHS) accessed through the Integrate Public Use Microdata Series (IPUMS) DHS platform. This data is available in the public domain but is only accessible following a formal request to MEASURE DHS, see here for more instructions on how to access the data: https://dhsprogram.com/data/Access-Instructions.cfm. All the Stata programs or do-files used to produce the results of this paper are available from the authors upon a reasonable request.

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ENDNOTE

[1] We also estimated our models using higher order polynomials, and the results are somewhat similar and not reported here but are available from the authors upon request.

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SUPPORTING INFORMATION

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