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RESEARCH ARTICLE

## Effects of maternal employment on child malnutrition in Cameroon: an analysis using the recursive bivariate probit model

[Effets de l'emploi maternel sur la malnutrition infantile au Cameroun : une analyse par modèle probit bivarié récursif]

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**Abstract:** This study sought to investigate the effect of maternal employment on child malnutrition via the 2011 and 2018 Cameroonian Demographic and Health Survey (DHS). We employed a recursive bivariate probit model to address potential endogeneity related to the reverse causality of the relationship between maternal employment and child malnutrition and to unobserved heterogeneities. The results revealed that maternal employment significantly increased the probability of stunting (28.4%) and overweight (15.8%) among children, corresponding to the Average Treatment Effect (ATE), that is, the average difference in predicted probabilities across the study population. Heterogeneity analysis revealed that the effects are more prevalent among boys, children aged 1-5 years, those living in rural areas, and mothers with low levels of education. The implications of these results suggest an urgent need for gender-sensitive public policies focused on better support for maternal employment, improved childcare services, and enhanced nutrition education and awareness for mothers.

**Keywords:** maternal employment; child malnutrition; recursive bivariate probit; Cameroon.

**Résumé :** Cette étude visait à examiner l'effet de l'emploi maternel sur la malnutrition des enfants à partir des données des Enquêtes démographiques et de santé (EDS) menées au Cameroun en 2011 et 2018. Nous avons utilisé un modèle probit bivarié récursif afin de tenir compte de l'endogénéité potentielle liée à la causalité inverse entre l'emploi maternel et la malnutrition infantile, ainsi que des hétérogénéités non observées. Les résultats révèlent que l'emploi maternel augmente significativement la probabilité de retard de croissance (28,4 %) et d'insuffisance pondérale (15,8 %) chez les enfants, ce qui correspond à l'effet moyen du traitement (ATE), c'est-à-dire la différence moyenne des probabilités prédites dans l'ensemble de la population étudiée. L'analyse de l'hétérogénéité révèle que ces effets sont plus fréquents chez les garçons, les enfants âgés de 1 à 5 ans, ceux vivant en milieu rural et les mères ayant un faible niveau d'éducation. Les implications de ces résultats suggèrent un besoin urgent de politiques publiques sensibles au genre, axées sur un meilleur soutien à l'emploi maternel, l'amélioration des services de garde d'enfants et le renforcement de l'éducation nutritionnelle et de la sensibilisation des mères.

**Mots clés :** emploi maternel ; malnutrition infantile ; probit bivarié récursif ; Cameroun.

### Introduction

Child malnutrition continues to be a significant public health problem in low- and middle- income countries, despite numerous

worldwide attempts to address it [1]. It has serious and lasting consequences for children's health, physical growth, cognitive ability and socioeconomic trajectory [2-3].

For example, stunting may delay school entry, hinder learning, reduce future economic opportunities and increase the likelihood of living precariously in adulthood. In addition, there are increased risks of morbidity and mortality [4-5], as well as long-term sequelae such as reduced height, impaired metabolism, greater vulnerability to mental disorders and higher blood sugar levels in adulthood [6]. In addition to its individual effects, child malnutrition undermines overall economic development by sustainably affecting population productivity [7].

Global statistics on malnutrition are alarming. More than 144 million children under five years of age suffer from stunting, 47 million from wasting and 20 million from underweight [8]. Sub-Saharan Africa remains the most affected region, with nearly one-third of children under five stunted. Cameroon is not escaping this trend. Indeed, despite a slight improvement between 2011 and 2018, the nutritional status of children remains a major concern. According to the latest Demographic and Health Survey (DHS, 2018), 29% of children under five years of age were stunted, and 11% were underweight, whereas 33% and 6% were underweight, respectively, in 2011. These indicators, which are well below the targets set by the Sustainable Development Goals (SDGs), highlight the importance of appropriate policy measures to improve child nutrition.

The literature highlights several determinants of child malnutrition, including socioeconomic factors such as poverty [9-10], food availability [11], and sanitary conditions [12-13]. Other variables related to household characteristics (size, number of children), children (age, sex, health) or parents (level of education, mother's body mass index) are also associated with children's nutritional outcomes [14-16].

Among these factors, maternal employment has recently attracted considerable attention in the literature, especially in developing countries [1,17-18].

This variable has a potential dual effect on child nutrition. On the one hand, the increase in income generated by mothers' employment can improve access to nutritious food, quality healthcare and healthy housing conditions [19-21]. On the other hand, reduced time spent by employed mothers on childcare and feeding may negatively affect children's nutrition, particularly in the absence of adequate childcare services [22-24]. This dilemma, theorized by Becker [25] as the trade-off between the income effect and the substitution effect, raises an important empirical question: what is the net effect of maternal employment on child nutrition?

Empirical findings on this issue remain highly controversial. Some studies report a negative association between maternal employment and child nutritional status [26-28], whereas others find no significant relationship [16,29], and others suggest a positive effect, particularly when children accompany their mothers to the workplace or benefit from adequate childcare arrangements [30]. These discrepancies can be attributed in part to methodological differences. Indeed, most studies rely on simple correlational analyses without correction for potential biases related to the endogeneity of the maternal employment variable [17]. However, maternal employment may depend on unobserved factors (motivation, health, skills) that also influence child nutritional status or even child health itself, which may lead the mother to enter or leave the labor market.

To address these limitations, more recent studies have adopted causal estimation methods with endogeneity bias correction to better capture the real effect of maternal employment on child nutrition. The results remain mixed: in some contexts, the negative effect persists [1,24], whereas in others, positive effects dominate, particularly for the most vulnerable children or when employment is compatible with childcare [31]. Research thus increasingly converges on the idea that what matters is

not whether mothers work but rather the conditions under which employment takes place, as well as the availability of resources to ensure children's nutritional well-being.

In the case of Cameroon, studies on child malnutrition have focused mainly on demographic and socioeconomic determinants [10,32-35], without explicitly exploring the role of maternal employment. The few studies that address this issue mostly examine reverse causality, namely, the impact of child health on women's participation in the labor market [36-37]. It therefore seems relevant to reverse this perspective and investigate how maternal employment affects child nutrition. However, such an analysis raises significant methodological challenges, notably the need to correct for endogeneity bias.

Against this backdrop, the present study seeks to estimate the causal effect of maternal employment on the nutritional status of children under five years in Cameroon via a bivariate recursive probit model applied to DHS IV and V data. To the best of our knowledge, this specification has not yet been employed in this field in developing countries. It allows for better identification of causal relationships by explicitly addressing the endogeneity of maternal employment.

This study makes several significant contributions. First, it provides robust evidence of the causal effect of maternal employment in a context that has yet to be fully explored. Second, it contributes to the growing body of research on the determinants of child malnutrition, using an appropriate method to estimate direct and indirect effects. Finally, it innovates methodologically using the bivariate recursive probit model while filling an important gap in the Cameroonian literature on the link between maternal employment and child nutrition.

This study also contributes to the health preferences literature by highlighting how maternal time allocation reflects implicit trade-offs between income generation and child health investments. In contexts with limited childcare infrastructure, maternal

employment decisions may reveal constrained preferences, where economic necessity overrides optimal health-related choices for children. Understanding these trade-offs is crucial for designing policies that align labor market participation with child health objectives.

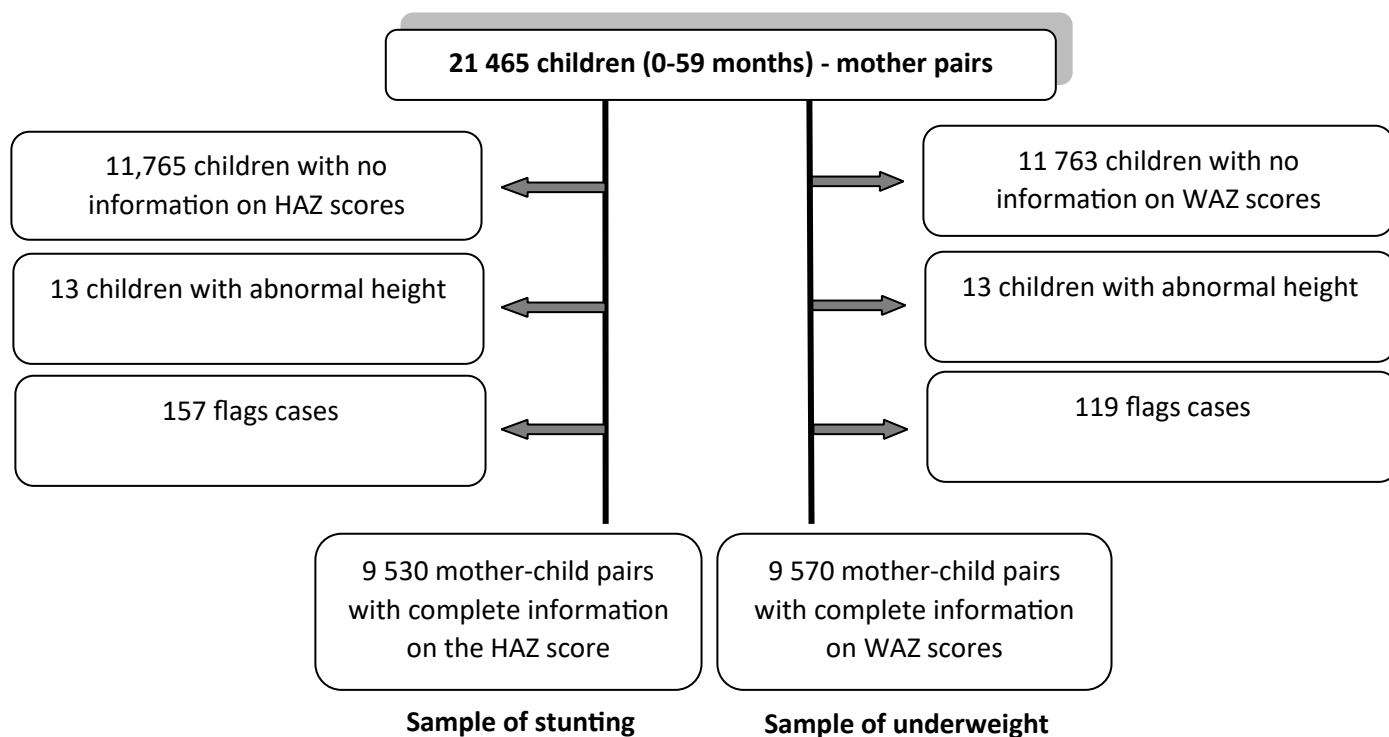
## **Methodology**

### ***Data and sample***

The data used in this article for empirical analysis are drawn from the fourth and fifth rounds of the Cameroon Demographic and Health Surveys (DHS), conducted in 2011 and 2018, respectively, by the National Institute of Statistics (NIS) in close collaboration with the Ministry of Public Health and several development partners.

In total, four types of questionnaires were used to collect data for the DHS: the Household Questionnaire, the Individual Woman Questionnaire, and the Individual Male Questionnaire for DHS IV, to which the Biomarkers Questionnaire for DHS V was added. The Woman Questionnaire specifically generated the child module, which was exploited in this study. In this module, information on the nutritional status of children aged 0-5 years, including the sociodemographic and economic characteristics of the mother, was reported exclusively by women aged 15-49 years who reported having a child within this age range in the household. For the purposes of analysis, this study uses matched records for children under 5 whose mothers were aged 15-49, yielding 11,732 and 9,733 observations for DHS IV and V, respectively. The DHS dataset contains nutritional indicators for children, including height-for-age Z scores (HAZ) and weight-for-age Z scores (WAZ) scores, which are based on World Health Organization (WHO) standards.

Our sample focuses on mothers matched with children under 5 for whom information has been collected. The overall sample for stunting is 9,530, comprising 5,033 in 2011 and 4,497 in 2018; the sample for underweight is 9,570, comprising 5,033 in 2011 and 4,537 in 2018.



**Figure 1.** Flowchart of study sample construction, DHS (2011 and 2018)

### ***Empirical method***

The empirical approach adopted in this study is grounded in the theoretical framework of the domestic production model developed by Becker [25], according to which households arbitrate between working time, income, and time spent on caregiving to maximize their utility. Within this framework, mothers' employment can affect children's nutritional health through two potentially opposing channels: an income effect, which increases household spending on food and health care, and a substitution effect, which reduces the amount of time devoted to childcare.

However, empirically assessing this relationship raises a central issue of endogeneity: the decision to engage in employment may be influenced by unobserved factors (preferences, social or family constraints, initial health status of the child) that simultaneously affect nutritional outcomes. Moreover, the relationship may be bidirectional: poor child health may deter mothers from participating in paid employment. In this context, estimating

causal effects requires a strategy that accounts for both unobserved error correlation and the recursive structure of the system.

To address these methodological challenges, we rely on a recursive bivariate probit model, as proposed by Coban [38], inspired by the implementation of Oyekale [39]. This model simultaneously specifies two structural equations: one explaining the child's nutritional status and the other explaining the mother's employment. Unlike the standard probit model, which assumes the exogeneity of the employment variable and thus ignores selection bias related to unobserved factors, the recursive model explicitly incorporates a correlation structure between the error terms of the two equations. Furthermore, when the standard bivariate probit model considers the two equations symmetrically, the recursive model specifies a causal relationship explicitly linking mothers' employment to children's nutritional health. This structural specification allows rigorous estimation of treatment effects as well as

total marginal effects, distinguishing between the direct and indirect effects of covariates included in both equations.

Furthermore, as Coban [38] noted, while the estimation of structural parameters does not fundamentally differ from the standard bivariate case, the recursive model offers significant analytical advantages: it enables robust postestimation analyses consistent with the causal structure of the system. This specificity allows for better identification of the net effects of mothers' employment on nutrition, correcting for both selection bias and the effects of unobserved omitted variables. The estimates are performed via the `rbiprobit` command in STATA, which is specifically designed for recursive structural models with binary endogenous variables.

The model is formally based on the following two structural equations:

$$\begin{aligned} y_1^* &= \theta y_2 + x'\beta + \varepsilon_1, y_1 = 1[y_1^* > 0] & (1) \\ y_2^* &= z'\alpha + \varepsilon_2, y_2 = 1[y_2^* > 0] & (2) \end{aligned}$$

In equations (1) and (2),  $y_1^*$  represents the expected value of child malnutrition, and  $y_2^*$  is the endogenous treatment variable representing the mother's propensity to obtain employment. These two variables are latent variables that are only observed in the form of their categorical realizations  $y_1$  and  $y_2$ . Furthermore,  $x'$  and  $z'$  are the vectors of the exogenous explanatory variables of the malnutrition equation and the maternal employment equation, respectively. In our study, both equations include the same exogenous explanatory variables, which, according to Han and Lee [40], means that the estimates are, at best, weakly identified.  $\theta$ ,  $\beta$ , and  $\alpha$  are the vectors of the estimated parameters. In addition,  $\varepsilon_1$  and  $\varepsilon_2$  are the error terms of equations (1) and (2), respectively. Notably, the latter are jointly normally distributed according to the following specification:

$$\begin{pmatrix} \varepsilon_1 \\ \varepsilon_2 \end{pmatrix} \sim N \left[ \begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} 1 & \rho \\ \rho & 1 \end{pmatrix} \right]$$

where  $\rho$  is the correlation coefficient between the error terms. The software generates Wald test statistics, which allow us to conclude whether there is a significant

correlation between the error terms. If the null hypothesis of no correlation is not rejected, then the model can be estimated via the conventional probit model, disregarding any considerations of endogeneity.

The recursive structure of the model allows maternal employment to be interpreted as an endogenous treatment variable affecting child nutrition. Within this framework, the estimated parameters can be used to derive treatment effects such as the Average Treatment Effect (ATE), the Average Treatment Effect on the Treated (ATET) and the Average Treatment Effect on Conditional (ATEC). These effects are computed from predicted probabilities generated by the joint distribution of the error terms, thereby incorporating both observed and unobserved heterogeneity. Consequently, the transition from structural equations to treatment effects is ensured through post-estimation procedures that exploit the probabilistic structure of the model.

According to Greene [41], the Average Treatment Effect (ATE) of mothers' employment can be expressed as follows:

$$ATE = \frac{1}{w} \sum_{i=1}^n w_i [\Phi(\theta + x'_i\beta) - x'_i\beta] \quad (3)$$

In equation (3),  $w$  represents the total weight of the sample,  $w_i$  is the sample weight of individual  $i$ , and  $n$  is the total number of observations.

The Average Treatment Effect on the Treated group (ATET) is given by:

$$ATET = \frac{1}{w_2} \sum_{i=1}^{n_2} w_i \left[ \Phi \left( \frac{x'_i\beta + \theta - \rho z'_i\alpha}{\sqrt{1-\rho^2}} \right) - \Phi \left( \frac{x'_i\beta - \rho z'_i\alpha}{\sqrt{1-\rho^2}} \right) \right] \forall y_{2i} = 1 \quad (4)$$

The Average Treatment Effect on Conditional success probabilities (ATEC) is defined by Alrasheed [42] as follows:

$$ATEC = \frac{1}{w_2} \sum_{i=1}^{n_2} w_i \left[ \frac{\Phi_2(x'_i\beta + \theta, \rho z'_i\alpha)}{\Phi(z'_i\alpha)} - \frac{\Phi_2(x'_i\beta - z'_i\alpha, \rho)}{\Phi(-z'_i\alpha)} \right] \quad (5)$$

In equations (4) and (5),  $w_2$  is the total sample weight of treaties, and  $n_2$  is the number of treaty observations.

Moreover, Greene [41] breaks down total marginal effects into their direct and indirect components. The effects of the exogenous variables in equations (1) and (2) can thus be decomposed. Denoting  $x_{1T}$  for equation 1 (treatment equation) and  $x_{1E}$  for equation 2 (endogeneity equation), the decomposition of marginal effects into their direct and indirect components is given by:

- ❖ Case of continuous exogenous variables

$$ME = \frac{\delta Pr}{\delta \begin{pmatrix} x_{1T} \\ x_{1E} \end{pmatrix}} = \underbrace{\frac{\delta Pr}{x_{1T}}}_{\text{effdirect}} + \underbrace{\frac{\delta Pr}{x_{1E}}}_{\text{effindirect}} \quad (6)$$

- ❖ Case of discrete exogenous variables

The decomposition of the discrete variables of both the treatment and the endogeneity equations was performed according to Hasebe [43] and Edwards et al. [44].

$$ME = \frac{\delta Pr}{\delta \begin{pmatrix} x_{1T} \\ x_{1E} \end{pmatrix}} = \underbrace{|Pr|_{x_{1T}=1} - |Pr|_{x_{1T}=0}}_{\text{effdirect}} + \underbrace{|Pr|_{x_{1E}=1} - |Pr|_{x_{1E}=0}}_{\text{effindirect}} \quad (7)$$

To ensure robustness of results, we first estimate the recursive bivariate probit model of the baseline model before proceeding with the sensitivity analysis. An exploratory analysis is performed beforehand and consists of descriptive statistics of the study variables and correlation analysis of key indicators. Analysis was performed via STATA 17 [45].

#### **Presentation and description of variables**

This study includes several groups of variables that should be presented to identify their meaning and the methodology used for their construction. These include the outcome variable and the explanatory variables, including the variable of interest (the endogenous variable in our context) and the control variables.

#### **Outcome variable: Children's nutritional status**

The three standard anthropometric indices

describing children's nutritional status are the height-for-age Z score (HAZ), the weight-for-height Z score (WHZ), and the weight-for-age Z score (WAZ), which are all measured against the World Health Organization (WHO) child growth standard. A child with an  $HAZ < -2$ ,  $WHZ < -2$ , or  $WAZ < -2$  is considered to be stunted, wasted, or underweight, respectively [17]. Stunting reflects the long-term effects of chronic undernutrition in a population and does not vary significantly with recent food intake or illness. Wasting reflects acute or recent nutritional deficiency. Weight-for-age is a composite index of weight-for-height and height-for-age and does not distinguish between acute malnutrition (wasting) and chronic malnutrition (stunting). A child may thus be underweight because of stunting, wasting, or both.

Among these three commonly used indicators, we used stunting (HAZ) and underweight (WAZ) as binary dependent variables. The HAZ score is now widely considered the most relevant indicator of overall nutrition, and the reduction in stunting has become the standard metric for assessing progress toward international nutrition targets [46-47]. HAZ represents a long-term health outcome [48] and is a manifestation of inadequate nutrition, health, and care for the mother before and during pregnancy as well as for the child during the first years of life [6]. Underweight, which measures weight-for-age, is used because of its composite nature and because it incorporates both acute and chronic malnutrition.

#### **Explanatory variable of interest: Maternal employment**

Maternal employment is the key explanatory variable of interest in this study. The DHS-IV (2011) and DHS-V (2018) datasets lack information on mothers' average working hours, which are frequently used in the literature to represent labor supply intensity. Instead, the surveys asked current married women whether they were employed at the time of the interview and/or had been

employed at any time during the 12 months preceding the survey. Based on these questions, we constructed a binary variable indicating maternal participation in the labor market. The variable takes a value of 1 if the mother had worked regularly in the 12 months prior to the survey and 0 otherwise (unemployed) [1,17].

The temporal misalignment between maternal employment, measured retrospectively over 12 months, and cross-sectional nutritional observations introduces indicator-specific interpretive nuances. Height-for-age (HAZ) remains fundamentally consistent with this year-long reference period, as it captures the cumulative effects of chronic deprivation and sustained household resource allocation. Conversely, weight-based metrics like weight-for-age (WAZ) are highly sensitive to acute shocks and short-term environmental shifts, potentially weakening their correlation with annual labor participation. Consequently, while HAZ estimates reflect long-term causal trajectories, effects on acute indicators should be viewed as conservative average associations subject to potential attenuation bias due to measurement timing.

#### **Control variables**

Control variables were selected based on previous research and their relevance to the Cameroonian context. They are organized into three categories: child, parent, and household characteristics.

Child characteristics include sex, which is included to control for possible effects of gender-differentiated cultural or dietary preferences. Age is included as it influences nutritional needs, dietary autonomy, and the frequency of childhood illnesses [17,32].

Parental characteristics include mother's age at childbirth, as it affects biological and social caregiving capacity [1,49-50]. Educational level of both parents is included as a proxy for knowledge about health and nutrition [17,51-52]. Access to information, measured by household television ownership, is also considered. Finally, religion is included because of its potential influence on health, dietary practices and family norms [53-54].

Household characteristics include the wealth index, which is derived from assets and housing conditions and is divided into quintiles ranging from "very poor" (reference category) to "very rich". The number of children in the household affects the allocation of parental time [55]. Access to safe drinking water is measured by the time required to reach an improved source, serving as a proxy of sanitary conditions [56]. The gender of the head of household (binary variable) and the place of residence (urban/rural) are also controlled for, as they influence access to health services and information [17].

Postnatal care utilization and child vaccination status were not included in the baseline specification. Although these variables are important determinants of child health, they may lie on the causal pathway between maternal employment and nutritional outcomes. Including them could therefore lead to over-controlling and biases the estimated effects.

## **Results and discussion**

### ***Descriptive statistics of key study variables***

Table 1 presents descriptive statistics for the main variables. With respect to the outcome variables, the data indicates that 30.04% of the children under five years of age are stunted, whereas 12.95% are underweight, highlighting a significant prevalence of stunting within the sample. Concerning the key explanatory variable of interest, namely, maternal employment, 34.60% are employed, whereas 64.40% are unemployed. This distribution remains almost identical across the subsamples of stunted and underweight children.

Table 2 highlights the association between child, parental and household characteristics and children's nutritional status. The main result reveals that the prevalence of stunting is higher among children of working mothers (31.68%) than those of nonworking mothers (27.04%). We also observe a similar but less pronounced trend for underweight, with rate of 12.26% among children of working mothers versus 11.86% among those of nonworking

mothers. These differences are statistically significant at 1% level, suggesting a positive association between maternal employment and child malnutrition at the bivariate level. However, these results should be

interpreted with caution, as they provide preliminary evidence of an association rather than a causal relationship, which justify the need for a more rigorous econometric analysis.

**Table 1.** Distribution of variables according to study samples, DHS (2011 and 2018)

Variables	Stunting (N = 9,530)		Underweight (N = 9,570)	
	N	Mean/%	N	Mean/%
<b>OUTCOME VARIABLE</b>				
<b>Child stunting</b>				
Yes	2 863	30.04		
No	6 667	69.96		
<b>Child underweight</b>				
Yes			1 160	12.12
No			8 410	87.88
<b>ENDOGENOUS VARIABLE</b>				
<b>Maternal employment status</b>				
Working mother	6 165	64.69	6 189	64.67
Nonworking Mother	3 365	35.31	3 381	35.33
<b>CONTROL VARIABLES</b>				
Child's age	9 530	27.734	9 570	27.708
<b>Child's sex</b>				
Male	4 741	49.75	4 764	49.78
Female	4 789	50.25	4 806	50.22
Number of children under 5	9 530	2.418	9 570	2.419
Access to water (time to get water source)	9 530	19.926	9 570	19.924
<b>Place of residence</b>				
Urban	4 123	43.26	4 132	43.18
Rural	5 407	56.74	5 438	56.82
<b>Wealth quintile</b>				
Poorest	1 898	19.92	1 920	20.06
Poorer	2 203	23.12	2 209	23.08
Middle	2 127	22.32	2 131	22.27
Richer	1 899	19.93	1 903	19.89
Richest	1 403	14.72	1 407	14.70
<b>Sex household head</b>				
Female	1 645	17.26	1 651	17.25
Male	7 885	82.74	7 919	82.75
<b>Household media exposure</b>				
Yes	4 122	43.25	4 132	43.18
No	5 408	56.75	5 438	56.82
<b>Mother's religion</b>				
Christian	7 771	81.54	7 797	81.47
Muslim	1 309	13.74	1 317	13.76
Other	450	4.72	456	4.76
Mother's age (years)	9 530	18.945	9 570	18.945
<b>Mother's education level</b>				
Uneducated	2 166	22.73	2 178	22.76
Primary	3 758	39.43	3 771	39.40
Secondary	3 223	33.82	3 237	33.82
Higher	383	4.02	384	4.01
<b>Father's level of education</b>				
Uneducated	2 926	30.70	2 940	30.72
Primary	2 928	30.72	2 946	30.78
Secondary	3 061	32.12	3 067	32.05
Higher	615	6.45	617	6.45

Note: based on DHS data (2011 and 2018).

**Table 2.** Bivariate relationships between nutrition and individual and household characteristics

Variables	Stunting (N = 9,530)			Underweight (N = 9,570)		
	Yes	Not	P value	Yes	Not	P value
	Average/%	Average/%		Average/%	Average/%	
<b>Mother's employment status</b>						
Working mother	31.68	68.32	0.000	12.26	87.74	0.000
Nonworking mother	27.04	72.96		11.86	88.14	
Child's age	30.761	26.435	0.000	28.421	27.610	0.000
<b>Child's sex</b>						
Male	32.78	67.22	0.000	13.33	86.67	0.000
Female	27.33	72.67		10.92	89.08	
Number of children	2.559	2.357	0.000	2.727	2.376	0.000
Access to water	22.563	18.794	0.000	22.759	19.533	0.000
<b>Place of residence</b>						
Urban	21.42	78.58	0.000	6.75	93.25	0.000
Rural	36.62	63.38		16.20	83.80	
<b>Wealth quintile</b>						
Poorest	43.78	56.22	0.000	24.48	75.52	0.000
Poorer	38.86	61.14		15.98	84.02	
Middle	28.73	71.27		8.96	91.04	
Richer	21.01	78.99		5.31	94.69	
Richest	11.83	88.17		3.20	96.80	
<b>Sex household head</b>						
Female	27.78	72.22	0.028	9.02	90.98	0.000
Male	30.51	69.49		12.77	87.23	
<b>Household media exposure</b>						
Yes	20.01	79.99	0.000	5.69	94.31	0.000
No	37.68	62.32		17.01	82.99	
<b>Mother's religion</b>						
Christian	28.99	71.01	0.000	11.04	88.96	0.000
Muslim	35.60	64.40		18.38	81.62	
Other	32.00	68.00		12.50	87.50	
Mother's age (years)	18.329	19.210	0.000	17.910	19.088	0.000
<b>Mother's education level</b>						
Uneducated	41.74	58.26	0.000	25.80	74.20	0.000
Primary	33.08	66.92		10.98	89.02	
Secondary	21.22	78.78		5.44	94.56	
Higher	8.36	91.64		2.08	97.92	
<b>Father's level of education</b>						
Uneducated	35.20	64.80	0.000	18.71	81.29	0.000
Primary	33.95	66.05		12.83	87.17	
Secondary	25.51	74.49		7.01	92.99	
Higher	9.43	90.57		2.76	97.24	

Note: based on DHS data (2011 and 2018).

### **Results of econometric estimates**

#### **Baseline Model: The recursive bivariate probit model**

Results of the estimates derived from the recursive bivariate probit model are presented in Table 3. This model was used to analyze mothers' employment decisions and children's nutritional status simultaneously, considering the potential endogeneity between these two phenomena. The approach distinguishes between two

anthropometric indicators of malnutrition, stunting and underweight, which are assessed separately in two model specifications. Equations (1) and (3) model the determinants of maternal employment in the stunting and underweight specifications, respectively. Equations (2) and (4) examine the effects of maternal employment, along with other sociodemographic factors, on the two malnutrition indicators.

The joint estimation of the equations via the recursive bivariate probit model yielded statistically robust adjustments. The Wald statistic is significant at the 1% threshold in both models, indicating the overall relevance of the explanatory variables selected. Furthermore, the correlation test between the error terms of the two equations (test of  $\rho = 0$ ) is significant in both cases, at the 1% threshold for the stunting model and 10% for the underweight model. These results confirm the presence of structural interdependence between mothers' employment decisions and child nutrition, thus justifying the use of this recursive approach.

More specifically, the estimated coefficients of  $\rho$  are negative and relatively high: -0.5743 for the stunting model and -0.5624 for the underweight model. This negative correlation between the equation residuals suggests that unobserved factors simultaneously but oppositely influence mothers' employment and their children's nutritional health. In the Cameroonian context, such factors may include maternal ability and efficiency in time allocation, whereby more capable or better-organized mothers are both more likely to engage in income-generating activities and more effective in maintaining adequate child

feeding and care practices. Similarly, unobserved household resilience mechanisms such as access to informal support networks, extended family assistance in childcare, or diversified income sources may facilitate maternal employment while buffering children against nutritional deprivation.

Conversely, adverse household shocks, including income instability, health shocks, or food price volatility, may reduce mothers' capacity to work while simultaneously increasing the risk of child malnutrition. The negative sign of  $\rho$  therefore reflects a selection process in which mothers who are able to combine employment with effective childcare differ systematically from those who are constrained by unobserved vulnerabilities, underscoring the importance of accounting for endogeneity when estimating the causal impact of maternal employment on child nutritional outcomes.

Furthermore, given the potential risk of multicollinearity between maternal and paternal education, due to assortative mating and similar socio-economic backgrounds within households, we assessed their association using Spearman's rank correlation coefficient. The estimated correlation is moderate ( $\rho = 0.437$ ,  $p < 0.001$ ), indicating no strong collinearity.

**Table 3.** Effect of maternal employment on child nutritional status; results of recursive bivariate probit model regressions.

Variables	Recursive bivariate probit model of stunting		Recursive bivariate probit model of underweight	
	Maternal employment (1)	Stunting (2)	Maternal employment (3)	Underweight (4)
<b>Mother's employment status</b>		0.9412*** (0.3587)		0.8838** (0.4012)
<b>Child's age</b>	0.0227*** (0.0031)	0.0462*** (0.0093)	0.0233*** (0.0031)	0.0060 (0.0065)
<b>Child's age squared</b>	-0.0002*** (0.0001)	-0.0007*** (0.0001)	-0.0003*** (0.0001)	-0.0001 (0.0001)
<b>Child's sex (ref. Male)</b>	0.0141 (0.0270)	0.1432*** (0.0321)	0.0110 (0.0270)	0.0997*** (0.0360)
<b>Number of children</b>	-0.0081 (0.0097)	0.0269*** (0.0096)	-0.0084 (0.0097)	0.0415*** (0.0116)
<b>Access to water</b>	0.0007 (0.0005)	0.0002 (0.0005)	0.0008* (0.0005)	-0.0004 (0.0006)
<b>Residence (ref. Rural)</b>	-0.2046***	0.0844*	-0.2008***	0.1298**

	(0.0388)	(0.0486)	(0.0389)	(0.0569)
<b>Mother's religion (ref. Christian)</b>				
<b>Muslim</b>	-0.2509*** (0.0404)	0.1121** (0.0528)	-0.2472*** (0.0403)	0.1158* (0.0601)
<b>Other</b>	0.0896 (0.0657)	-0.0418 (0.0647)	0.0929 (0.0655)	-0.0723 (0.0790)
<b>Wealth quintile (ref. Poorest)</b>				
<b>Poorer</b>	-0.1115** (0.0441)	-0.0304 (0.0463)	-0.1191*** (0.0441)	-0.0917 (0.0568)
<b>Middle</b>	-0.2443*** (0.0527)	-0.1724** (0.0772)	-0.2518*** (0.0528)	-0.2881*** (0.1036)
<b>Richer</b>	-0.4479*** (0.0705)	-0.2604** (0.1283)	-0.4683*** (0.0706)	-0.3917** (0.1644)
<b>Richest</b>	-0.4754*** (0.0800)	-0.4335*** (0.1560)	-0.4899*** (0.0800)	-0.4332** (0.1816)
<b>Sex household head (ref. Male)</b>	-0.0947** (0.0374)	0.0307 (0.0403)	-0.0990*** (0.0374)	0.0912* (0.0493)
<b>Media exposure (ref. No)</b>	0.0234 (0.0440)	-0.0902** (0.0443)	0.0250 (0.0440)	-0.0575 (0.0570)
<b>Mother's age</b>	-0.0022 (0.0252)	-0.0169 (0.0259)	-0.0046 (0.0253)	-0.0491 (0.0320)
<b>Mother's age squared</b>	0.0004 (0.0006)	0.0001 (0.0006)	0.0004 (0.0006)	0.0007 (0.0008)
<b>Mother's education level (ref. Uneducated)</b>				
<b>Primary</b>	0.3669*** (0.0411)	-0.1985*** (0.0591)	0.3695*** (0.0410)	-0.4397*** (0.0513)
<b>Secondary</b>	0.1719*** (0.0472)	-0.2324*** (0.0469)	0.1784*** (0.0472)	-0.4919*** (0.0601)
<b>Higher</b>	0.2293*** (0.0887)	-0.3702*** (0.1071)	0.2335*** (0.0890)	-0.6003*** (0.1572)
<b>Father's education level (ref. Uneducated)</b>				
<b>Primary</b>	0.2524*** (0.0371)	-0.0872* (0.0503)	0.2584*** (0.0370)	-0.1794*** (0.0534)
<b>Secondary</b>	0.1800*** (0.0381)	-0.0987** (0.0443)	0.1875*** (0.0381)	-0.2364*** (0.0492)
<b>Higher</b>	0.1353** (0.0661)	-0.3812*** (0.0818)	0.1413** (0.0662)	-0.2914*** (0.1124)
<b>Constant</b>	0.0127 (0.2658)	-1.1550*** (0.2908)	0.0333 (0.2657)	-0.6162* (0.3695)
<b>Observations</b>	9 530		9 570	
<b>Wald chi2(45)</b>	2465.84		1742.39	
<b>P Value</b>	0.0000		0.0000	
<b>Rho</b>	-0.5743		-0.56240	
<b>Wald test of rho=0</b>	3.6123		2.963	
<b>P value</b>	0.0000		0.0852	

Notes: based on DHS data (2011 and 2018); the table reports the coefficients from the recursive bivariate probit model estimation, with robust standard errors in parentheses. OLS, \*\*\*p < 0.01; \*\*p < 0.05; \*p < 0.1.

Finally, the use of the recursive bivariate probit model, particularly the `rbiprobit` command in Stata, presents a significant methodological advantage. It allows not only the simultaneous estimation of structural equations but also the computation of full marginal effects and average treatment effects (ATE, ATET, ATEC) while accounting for the recursive structure of the model. This level of analytical precision constitutes an important methodological contribution that

is rarely used in the literature on developing countries and, to our knowledge, is still unprecedented in the Cameroonian context. *Determinants of maternal employment* Results in columns (1) and (3) of the recursive bivariate probit model (Table 3) reveal that several variables significantly influence the likelihood of a mother being employed. The age of the child has a positive and significant effect on maternal employment (coefficients: 0.0227\*\*\* and

0.0233\*\*\*), although this effect decreases with the age of the child, as shown by the negative coefficient of the age squared variable (-0.0002\*\*\* and -0.0003\*\*\*). Mothers living in urban areas are less likely to be employed than those in rural settings are (coefficients: -0.2046\*\*\* and -0.2008\*\*\*). Similarly, Muslim mothers are less likely to be employed than Christian mothers are (-0.2509\*\*\* and -0.2472\*\*\*).

Household wealth has a negative effect on maternal employment. The coefficients become increasingly negative as wealth increases, reaching -0.4754\*\*\* for mothers from very wealthy households. Conversely, education has a positive effect because mothers with a primary education (0.3669\*\*\*; 0.3695\*\*\*), secondary (0.1719\*\*\*; 0.1784\*\*\*) or higher (0.2293\*\*\*; 0.2335\*\*\*) education are more likely to be employed than those with no education. Father's education further increases this probability, with significant coefficients for primary (0.2524\*\*\*; 0.2584\*\*\*), secondary (0.1800\*\*\*; 0.1875\*\*\*) and higher (0.1353\*\*; 0.1413\*\*) levels. Finally, when the head of the household is female, the probability of her being employed decreases (-0.0947\*\*; -0.0990\*\*\*), possibly reflecting the dual burden of domestic and economic responsibilities. Overall, these findings highlight the combined influence of demographic, educational, and socioeconomic factors in explaining mothers' employment in Cameroon.

#### *Effect of maternal employment on child nutrition and analysis of treatment effects*

Results of the recursive bivariate probit model (Table 3) indicate that maternal employment has a significant and negative effect on child nutrition. Indeed, the estimated coefficients for the effect of employment on stunting and underweight are 0.9412\*\*\* and 0.8838\*\*, respectively, suggesting that children of employed mothers are more likely to be malnourished than those whose mothers are not employed.

The analysis of the treatment effects confirmed this trend (Table 4). Indeed, the average treatment effect (ATE) shows that children of employed mothers have, on average, a 28.39% higher probability of being stunted and a 15.80% higher probability of being underweight than those of unemployed mothers. Furthermore, the average treatment effect of the treated (ATET) is positive for both malnutrition indicators but statistically significant only for stunting. This result suggests that children of employed mothers are 36 percentage points on average more likely to suffer from stunting than they would have been if their mothers were unemployed. Unlike stunting, which reflects chronic nutritional deprivation, being underweight is a more volatile and composite indicator that captures both acute and chronic conditions. As such, it is more sensitive to short-term shocks which may vary considerably among children of employed mothers, thereby reducing the precision of the estimated effect.

Furthermore, the average treatment effect on conditional probability (ATEC) is not significant for either nutrition indicator, suggesting that maternal employment does not affect malnutrition once a given nutritional state is established. This indicates that its impact operates mainly through baseline risk, while structural constraints, such as limited healthcare access, inadequate childcare, and persistent poverty, continue to drive nutritional outcomes regardless of maternal employment status.

These results should be interpreted in light of the well-known dilemma between the income effect and the substitution effect: while maternal employment can improve a household's standard of living (expected positive effect), it may also reduce the time and quality of care provided to children (negative effect). In the Cameroonian context, the results suggest that the substitution effect dominates. The combination of professional work and

domestic tasks seems to overload mothers, negatively affecting their ability to meet their children's nutritional needs. Furthermore, the use of alternative childcare solutions<sup>1</sup> (nannies, extended family members), which are often informal and

poorly supervised, may not adequately compensate for maternal absence, thereby increasing the risk of malnutrition. These findings are consistent with the work of Rashad and Sharaf [18], Jakaria et al. [17], and Hosen et al. [1].

**Table 4.** Treatment effects of mothers' employment

Treatment effect	Stunting		Underweight	
	Parameter	Standard errors	Parameter	Standard errors
Average treatment effect (ATE)	0.2839***	0.1058	0.1580**	0.0829
Average effect of treatment on people treated (ATET)	0.3600**	0.1417	0.2057	0.1320
Average treatment effect on conditional probability (ATEC)	0.0017	0.0097	-0.0048	.0070

Note: based on DHS data (2011 and 2018)

#### *Direct and indirect marginal effects of exogenous variables*

In addition to the main variable of interest, several determinants related to the characteristics of the child, parents, and household significantly influence malnutrition. Table 5 presents the breakdown of the marginal effects of exogenous variables into their direct and indirect components. Direct effects capture the impact of explanatory variables on child nutrition, holding maternal employment constant, while indirect effects operate through maternal employment. With respect to child-specific factors, the results revealed that younger children are more prone to malnutrition. Children under 37<sup>2</sup> months of age have a 1.47% probability of suffering from stunting, whereas those above this threshold have only a 0.02% probability of not being stunted. For underweight individuals, the threshold is 35 months, with probabilities of 0.28% (below the threshold) and 0.004% (above). Age is therefore a key determinant of child and adolescent malnutrition, confirming the findings of Fambon and Baye [10], Jakaria et al. [17], and Tangwa et al. [57]. Furthermore,

contrary to some previous studies [1,29,32,58], girls are more at risk than boys are, with a 3.58% probability of stunting and a 1.41% probability of being underweight, effects driven mainly by the direct component (3.34% and 1.31%, respectively).

Concerning household and environmental characteristics, the number of children under five constitutes a risk factor. Each additional child increases the probability of stunting by 0.48% and of being underweight by 0.47%. This result is justified by the intuition that the quantity and quality of food available in the household decreases as the number of children increases. This result is consistent with the findings of McKenna et al. [59] and Tchakounté [32]. The effect of place of residence is more nuanced and reveals that living in a rural area indirectly increases the probability of stunting by 3.57% and, overall, by 1.61%, with no significant effect on underweight. Children living in rural areas are thus more prone to chronic malnutrition due to difficulties in accessing basic services (health care, hygiene, and childcare facilities), corroborating the work of Srinivasan et al. [60] and Hosen et al. [1].

<sup>1</sup> In the Cameroonian context, particularly in rural and peri-urban areas, there are no professional or adequately structured childcare facilities available to households.

<sup>2</sup> The threshold age is obtained by differentiating the structural equation of malnutrition with respect to age and setting it equal to zero.

Household wealth significantly influences child nutrition. Indeed, it appears that the wealthier the household is, the lower the likelihood of malnutrition. Specifically, children living in very poor households are 17.97% more likely to be stunted and 9.66% more likely to be underweight than those from very rich households are, which aligns with the findings of Amare et al. [61] and Tchakounté [32]. Although the total effect of the household head's gender is statistically

insignificant, households headed by women indirectly reduce the likelihood of malnutrition (1.65% for stunting and 0.84% for underweight). Conversely, the direct effect of this factor indicates a 1.2% increase in the probability of being underweight. Media exposure, measured by access to television, significantly reduces the probability of stunting by 1.69%, confirming the educational role of the media in child health and nutrition.

**Table 5.** Decomposition of the marginal effects of exogenous variables into their direct and indirect components

Variables	Stunting			Underweight		
	Direct	Indirect	Total	Direct	Indirect	Total
Child's age	0.0108***	0.0039***	0.0147***	0.0008	0.0020***	0.0028***
Child's age squared	-0.0002***	-0.0001***	-0.0002***	-0.00001*	-0.0001***	0.00004***
Child's sex (ref. Male)	0.0334***	0.0025	0.0358***	0.0131***	0.0009	0.0141***
Number of children	0.0063***	-0.0014	0.0048**	0.0055***	-0.0007	0.0047***
Access to water	0.00005	0.0001	0.0002	-0.00005	0.00006	0.00001
Residence (ref. Rural)	0.0197	-0.0357***	-0.0161*	0.0171**	-0.0171***	0.00002
Mother's religion (ref. Christian)						
Muslim	0.0266*	-0.0106***	0.016**	0.0161	-0.0205***	-0.0068
Other	-0.0096	0.0157	0.0052	-0.0090	0.0081	-0.0021
Wealth quintile (ref. Poorest)						
Poorer	-0.0077	-0.0200**	-0.0309***	-0.0144*	-0.0108*	-0.0300***
Middle	-0.0423***	-0.0439***	-0.0923***	-0.0405***	-0.0225**	-0.0690***
Richer	-0.0626***	-0.0800***	-0.1453***	-0.0518***	-0.0400***	-0.0925***
Richest	-0.0992***	-0.0848***	-0.1797***	-0.0559***	-0.0415***	-0.0966***
Sex household head (ref. Male)	0.0071	-0.0165**	-0.0094	0.0120*	-0.0084**	0.0036
Media exposure (ref. No)	-0.0210**	0.0041	-0.0169*	-0.0076	0.0021	-0.0054
Mother's age	-0.0039	-0.0004	-0.0043	-0.0065	-0.0004	-0.0069
Mother's age squared	0.00002	0.0001	0.0001	0.00009	0.00003	0.00012
Mother's education level (ref. Uneducated)						
Primary	-0.0482***	0.0643***	0.0210**	-0.0683***	0.0313***	-0.0301***
Secondary	-0.0561***	0.0300***	-0.0203*	-0.0743***	0.0146**	-0.0500***
Higher	-0.0864***	0.0401**	-0.0415*	-0.0853***	0.0194**	-0.0577***
Father's education level (ref. Uneducated)						
Primary	-0.0208	0.0443***	0.0251***	-0.0253***	0.0220***	-0.0016
Secondary	-0.0235**	0.0315***	0.0097	-0.0323***	0.0158***	-0.0148**
Higher	-0.0842***	0.0237*	-0.0585***	-0.0386***	0.0118*	-0.0245**

Notes: based on DHS data (2011 and 2018); Table contains the decomposed marginal effects, \*\*\*p < 0.01; \*\*p < 0.05; \*p < 0.1.

The analysis of parental variables revealed noteworthy heterogeneity in their influence on child nutritional outcomes. Mothers' religion has a differentiated impact. Children of Muslim mothers are 1.6% more likely to be underweight than are those of Christian mothers. This disparity may largely reflect regional patterns, as predominantly Muslim areas, particularly in northern Cameroon, have long recorded

higher rates of child malnutrition. Parental education significantly affects child nutrition among children under five years of age. Compared with children of uneducated mothers, those whose mothers have achieved higher education levels are 4.15% less likely to experience stunting and 5.77% less likely to be underweight. This finding suggests that maternal education enhances health literacy, nutritional awareness, and

the adoption of appropriate feeding and childcare practices. These results are consistent with those of Jakaria et al. [17] and Tangwa et al. [57].

***Sensitivity analysis: heterogeneity in the effect of maternal employment on child nutrition***

The effect of maternal employment on child nutrition may vary according to certain individual or contextual characteristics. To assess potential heterogeneity, the recursive bivariate probit was re-estimated for several subsamples defined by child and household characteristics, including sex, age, birth order, place of residence, and maternal education. The overall results are reported in the first row of Table 6 for reference. Columns (1) and (3) display the estimated coefficients, whereas columns (2) and (4) indicate the corresponding average treatment effects (ATEs) for stunting and underweight, respectively.

Concerning child characteristics, the results indicate that boys are more adversely affected by maternal employment than girls. The risk of stunting increases on average by 25.7% for girls (coefficient = 0.890\*) and by 31.4% for boys (coefficient = 1.008\*\*). For underweight individuals, only the male category was significantly affected, with an average increase of 26.3% (coefficient=1.260\*\*\*). This pattern may be partly explained by biological factors, as male children tend to exhibit higher metabolic needs and greater susceptibility to infectious diseases during early childhood, making them more sensitive to nutritional stress [19]. Beyond biology, subtle differences in caregiving and feeding practices may also play a role. In some Cameroonian households, girls may receive more consistent feeding or supervision in infancy due to perceptions of greater fragility or compliance, while boys are encouraged to be more physically active at earlier ages, potentially increasing their energy expenditure [32]. These mechanisms can contribute to a higher risk of malnutrition among boys, even in households with similar resource constraints. These findings echo those of

Morrill [4] and Gennetian et al. [19], who reported that boys tend to be more vulnerable to changes in maternal labor conditions, possibly because of biological and behavioral sensitivities during early development.

When disaggregating by age group, the analysis also reveals substantial variation. Among children under 1 year of age, maternal employment appears to have a protective effect against stunting, with a significant negative coefficient of -1.033\* and an ATE of -27.3%. Conversely, this effect becomes detrimental for children aged one year and older, for whom maternal employment significantly increases the risk of stunting by 33.7% (coefficient=1.068\*\*\*). For underweight children, the effect of maternal employment remains insignificant among infants but becomes significant (15.8%, coefficient of 0.881\*\*) for older children. These results suggest that infants often benefit from enhanced caregiving, typically from other family members, when their mothers are employed [4]. In contrast, older children tend to experience the negative consequences of reduced maternal attention due to longer working hours and limited daily availability [17]. Consequently, the net effect of maternal work on child nutrition depends critically on access to reliable childcare options, flexible work arrangements, and the broader social support system.

Birth order also introduces a form of heterogeneity [62]. The results are not statistically significant for firstborn children, but for older children, maternal employment significantly increases the likelihood of being underweight by 17.9% (coefficient of 0.994\*\*\*). This result can be explained by the cumulative burden of childcare and household resource allocation. Indeed, with several dependent children, mothers may be forced to make trade-offs in terms of time and care, often to the detriment of younger or less prioritized children. This finding contrasts with those of Whiteman et al. [63] and Jakaria et al. [17], who observed in their study strong negative effects among firstborn children.

**Table 6.** Heterogeneity of the effect of mothers' employment according to selected characteristics of children, parents and the environment.

Sample	Stunting (1)		N	Underweight (2)		N
	Parameter (1)	ATE (2)		Parameter (3)	ATE (4)	
All children	0.941*** (0.359)	0.284*** (0.105)	9 530	0.884** (0.401)	0.158** (0.083)	9 570
<b>Child's sex</b>						
Female	0.890* (0.490)	0.257* (0.138)	4 789	0.370 (0.630)	0.059 (0.099)	4 806
Male	1.008** (0.473)	0.314** (0.144)	4 731	1.260*** (0.147)	0.263*** (0.044)	4 764
<b>Child's age</b>						
Less than 1 year	-1.033* (0.452)	-0.273* (0.146)	2 166	-0.357 (0.921)	-0.062 (0.172)	2 183
1 year and over	1.068*** (0.283)	0.337*** (0.084)	7 364	0.881** (0.443)	0.158* (0.091)	9 466
<b>Order of birth</b>						
First-born	-0.423 (1.505)	-0.122 (0.438)	2 100	-0.940** (0.455)	-0.171 (0.110)	2 110
2nd born and above	0.750 (0.499)	0.229 (0.145)	7 430	0.994*** (0.344)	0.179*** (0.073)	7 460
<b>Place of residence</b>						
Rural	1.360*** (0.071)	0.422*** (0.020)	5 407	1.089*** (0.303)	0.228*** (0.071)	5 438
Urban	-0.683 (0.985)	-0.194 (0.301)	4 123	-0.090 (0.606)	-0.011 (0.073)	4 132
<b>Mother's education</b>						
Primary or -	0.970** (0.465)	0.316** (0.140)	5 924	1.230*** (0.165)	0.279*** (0.047)	5 949
Secondary or +	0.727 (1.087)	0.184 (0.285)	3 606	0.239 (0.811)	0.023 (0.078)	3 621

Notes: based on DHS data (2011 and 2018); Robust standard errors are recorded in parentheses. All specifications include the set of control variables shown in Table 5 above.  $p < 0.01$ ; \*\* $p < 0.05$ ; \* $p < 0.1$ .

Another source of heterogeneity in the effect of maternal employment on child nutrition is the household's place of residence. Results reveal that the effect of maternal employment is significantly more harmful in rural areas than in urban ones. In rural areas, maternal employment increases the probability of stunting by 42.2% (coefficient of 1.360\*\*\*) and that of underweight by 22.8% (coefficient of 1.089\*\*\*), whereas both effects are not significant in urban areas. The stronger adverse effects observed among rural children reflect structural and occupational characteristics specific to rural settings. Maternal employment in rural Cameroon is largely concentrated in informal and physically demanding agricultural activities, which require long working hours and limit the time available for childcare and appropriate feeding practices. Unlike urban employment, rural work rarely provides stable income flows or access to childcare substitutes, increasing reliance on maternal time. Additionally, rural households face limited access to healthcare facilities and diversified food sources, while being more exposed to seasonal income fluctuations and

climatic shocks. These constraints amplify children's vulnerability to malnutrition and help explain the larger nutritional penalties associated with maternal employment in rural areas. These results are consistent with those of Jakaria et al. [17].

Mothers' educational attainment also plays a decisive role in modulating the effect of maternal employment on child nutrition. Following Jakaria et al. [17], the analysis distinguishes between two educational categories: primary or below and secondary or higher. For mothers with primary education or less, the effect is strongly significant: the probability of stunting increases by 31.6% (coefficient of 0.970\*\*), and the risk of underweight increases by 27.9% (coefficient of 1.230\*\*\*). Conversely, the effects are not significant for mothers with secondary or higher education. This result suggests that less educated mothers, who generally have lower financial resources, limited access to nutritional information, and more physically demanding jobs, are less able to provide adequate nutrition for their children. These findings corroborate those of Hahn and Lee [40] and Jakaria et al. [17].

Finally, the mother's level of education plays a decisive role in shaping the effect of employment on child nutrition. Following Jakaria et al. [17], the analysis distinguishes between two educational categories: primary or below and secondary or higher. For mothers with primary education or less, the effect is strongly significant: the probability of stunting increases by 31.6% (coefficient = 0.970\*\*), and that of underweight mothers increases by 27.9% (coefficient = 1.230\*\*\*). Conversely, the effects are statistically insignificant among mothers with secondary or higher education. This result suggests that less educated mothers, who generally have lower financial resources, limited access to nutritional information, and more physically demanding jobs, are less able to provide adequate nutrition for their children. These findings corroborate those of Hahn and Lee [40] and Jakaria et al. [17].

### **Conclusion**

This study investigated the impact of maternal employment on the nutritional status of children under five years of age in Cameroon, using data from the 2011 and 2018 Demographic and Health Surveys (DHSs) and applying a recursive bivariate probit model. This method accounts for endogeneity bias arising from the complex interplay between maternal employment and child nutrition outcomes. The results consistently indicate that maternal employment increases the likelihood of stunting and underweight among children, suggesting that the substitution effect, linked to reduced time devoted to childcare, dominates the income effect.

The adverse effect of maternal employment on child nutrition observed in this study can be better understood in light of the specific employment conditions faced by women in Cameroon. A substantial proportion of employed mothers are engaged in informal and vulnerable sectors, with 35.34% working in agriculture, 20.55% in petty trade, and 15.62% in manual labor. These activities are typically characterized by long working hours, low and unstable

income, and limited compatibility with adequate childcare [17]. In such contexts, mothers often have no alternative but to take their children with them to the workplace, exposing them to unfavorable environmental conditions and irregular feeding practices. Beyond these structural constraints, the findings also reflect underlying health-related preferences shaped by necessity rather than choice. Faced with limited childcare options and pressing income needs, mothers are compelled to prioritize short-term economic survival over optimal child health investments, revealing a context of constrained health preferences [25]. The absence of accessible and reliable childcare services further limits their ability to align labor participation with recommended childcare practices. These combined factors help explain why the substitution effect dominates, ultimately resulting in a higher risk of malnutrition among children of employed mothers.

Analysis of heterogeneity revealed that this adverse effect is particularly pronounced among boys, children older than one year, those living in rural areas, and those born to mothers with low educational attainment. These findings imply that children's nutritional vulnerability depends not only on whether their mothers are employed but also on the nature of their jobs and the surrounding socioeconomic context. Maternal employment is therefore not inherently detrimental to child nutrition; it becomes so in environments lacking adequate support systems such as childcare services, parental leave, and flexible work arrangements [17].

From a policy standpoint, the findings underscore the need for integrated strategies that reconcile maternal employment with child well-being. This involves promoting decent and stable jobs for women that accommodate parental responsibilities while simultaneously expanding access to quality childcare services, particularly in rural areas. Complementary interventions should focus

on strengthening nutrition education, maternal health programs, and social protection initiatives targeting less-educated mothers to mitigate the time-related constraints that negatively affect child nutrition.

Although this study fills an empirical gap in the Cameroonian context, it also highlights directions for future research. Further studies could explore the quality of maternal employment, including working hours, job stability, and sectoral differences, and examine intrahousehold dynamics in time and resource allocation. Incorporating information on formal and informal childcare arrangements, as well as paternal involvement in caregiving, would provide a deeper understanding of the mechanisms linking maternal work and child nutrition.

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### Competing interest

Authors declared no competing interests

### Ethical standard

No ethical approval or review was required by any ethical review committee.

### Data availability statement

The dataset from the 2018 Cameroon Demographic and Health Survey used for this investigation is accessible online. Data can be accessed by applying through <https://dhsprogram.com/data/Access-Instructions.cfm>.

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