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MATERNAL EMPLOYMENT AND CHILDHOOD MALNUTRITION IN ECUADOR

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JEL Codes: I15, J22, C21

Keywords: Maternal employment, malnutrition outcomes, childhood, Ecuador

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Maternal employment and childhood malnutrition in Ecuador

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Abstract

This paper estimates the causal impact of maternal employment on several childhood malnutrition outcomes in Ecuador, to understand the trade-off between the time mothers devote to work and child-caring activities. We use exogenous regional variation in maternal labour market conditions to account for the potential endogeneity of mothers' employment. Using the Ecuadorian National Health and Nutrition Survey 2018 and the Living Conditions Survey 2014, the instrumental variable estimations indicated that maternal employment increases the probability of having stunted children by between 4.3 and 21 percent, while no significant effect was found on children suffering from wasting, underweight or overweight. We found that children with more educated, richer mothers appeared to be the most negatively affected. The results were robust to several robustness checks.

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1. Introduction

Child malnourishment in its various forms has become one of the most significant health problems in low- and middle-income countries and specifically in Ecuador over the last three decades (INEC, 2019). It is crucial to understand the impact of malnourishment on an individual's lifetime, and on a country's economic prosperity. For instance, childhood stunting leads to infections, developmental deficits, brain damage, adult non-communicable diseases, early mortality, poorer school performance, and lower productivity and income as an adult (Black et al., 2008; Guerrant et al., 2013; Sudfeld et al., 2015). Additionally, children born to stunted adults are more likely to suffer from stunting themselves (Addo et al., 2015).

Although public expenditure on healthcare has increased dramatically in Ecuador in the last decade (The World Bank, 2020a), indicators of child malnutrition have not improved as expected. For instance, the Ecuadorian National Statistical Office reported that 25% of children under five (U5) years old were stunted (children considered too short for their age) in 2012. This figure fell slightly to 23% in 2018. Similarly figures were found for wasting (too thin for height) and underweight (Freire et al., 2014; INEC, 2019), while the percentage of U5-year-old children with overweight increased in the same period. Among Latin American countries, Ecuador has the highest value of stunting after Guatemala (FAO et al., 2018), with rates comparable to sub-Saharan countries (The World Bank, 2020b).

Policy responses to tackle malnutrition in Ecuador have not succeeded in reducing this health issue significantly. Government interventions such as the *Programa Aliméntate Ecuador* (MIES, 2020), which includes provision of micronutrients such as iron, zinc, vitamin A and folic acid, have not been a structural solution to reduce malnutrition among U5-year-old children. Rivera (2019) suggested that potential reasons for this are insufficient institutional coordination (e.g. nutrition and primary health programmes work separately), which limits the efficient allocation of resources and management capacities between central and local governments to tackle childhood malnourishment. In addition, the lack of government policy evaluations has restricted the analysis of the success or failure of these policies, transparency during policy processes and efficient targeting of interventions.

Likewise, women's participation in the labour market has increased remarkably in the last decade (INEC, 2020), which raises the question of whether maternal employment is a detrimental factor in child health (Bravo et al., 2018; Canencia Yanacallo et al., 2017). Maternal employment increases household income, which improves families' wellbeing and therefore children's nutrition and health status. However, mothers who work outside the home have less time available for breastfeeding and childcare activities, which worsens children's nutrition and health outcomes. Therefore, it is crucial to further analyse the trade-off between mothers who generate additional income or provide maternal time, especially in low- and middle-income countries. This analysis can help to understand the net effect of maternal employment on children's nutritional status, which is not obvious a priori, and to design more informed policy interventions. Certainly, adverse socioeconomic conditions may not allow mothers to decide whether to work or take care of their children. They may be forced to work to provide enough income and escape from poverty. For these population groups, it is fundamental to capture the effects of maternal labour supply on child nutrition.

This study aims to estimate the causal effect of maternal labour supply on children's nutritional status in Ecuador (that is, stunting, underweight, wasting and overweight) considering mothers' desire to be involved in the labour force or to remain economically inactive, and accounting for the potential endogeneity of maternal employment – the presence of unobserved characteristics correlated with mothers' decision to work and children's health and simultaneity issues – and the heterogeneous impacts on population subgroups. We estimated this causal effect through an instrumental variable (IV) approach, using as instruments women's employment status and the average number of hours worked at cantonal level. The main IV findings suggest that the probability of having a stunted U5-year-old child is between 4.3 and 21 percent higher for employed mothers than for unemployed or inactive mothers. This estimate is large and significant. No evidence of the impact of maternal employment on wasting, underweight or overweight was reported. Moreover, the effect size of maternal employment was much larger for moderately stunted children than for the severely stunted. Interestingly, we showed that the effect of maternal employment on child malnourishment was greater among mothers with high education and household income. This shows the dominance of the time constraint over the income effect. The IV-2SLS results indicate that, in contrast to the ordinary least squares (OLS) strategy, maternal employment increases a child's risk of experiencing stunting to a larger extent.

This paper contributes to the literature in several ways. Firstly, our findings are one of the first pieces of evidence that quantifies the causal effect of maternal labour supply on childhood malnutrition in the South America region. Secondly, childhood outcomes were constructed by employing more reliable clinically measured anthropometric data that do not suffer from self-reported biases. Thirdly, in contrast to Rashad & Sharaf (2019), this paper used an area-level margin of mothers' work to instrument maternal employment.

The rest of the paper is structured as follows. Section 2 discusses empirical findings from the literature on the effects of the maternal labour supply on children's health status. Section 3 details the empirical specification and methodology. Section 4 outlines the characteristics of the data used. Section 5 presents the empirical results. Section 6 presents some robustness checks, and Section 7 offers some discussion and concluding remarks.

2. Literature review

The empirical literature on how parental supply affects child health status is mostly based on the child health production function (Becker, 1965; Grossman, 1972; Jacobson, 2000). Several studies adopted this theoretical framework and focused on how parental labour supply affects child health indicators. For instance, Morrill (2011) estimated the effects of maternal employment on the health of US school-age children measured as overnight hospitalizations, asthma episodes, injuries and poisonings. Using an instrumental variable (IV) approach, where the instrument for maternal labour supply is the exogenous variation in each child's youngest sibling's eligibility for kindergarten, she found significant large estimates suggesting that maternal employment increases the probability of a child having a negative health episode. Meyer (2016) examined the impact of maternal employment on the risk of childhood overweight in Germany, considering the number of younger siblings in the household as an instrument. She argued that the probability of being overweight increased due to maternal full-time employment. Thus, she attributed unhealthy behaviour in children, in terms of diet and activity, to the reduced amount of maternal time devoted to them. In the same line, Anderson et al. (2003) adopted an IV strategy and found that American mothers who work more hours per week are more likely to have an overweight child. Similarly, Datar et al. (2014) investigated the relationship between mother's working hours, child body mass index (BMI) and obesity in the US using an IV approach. In their study, the instrument was maternal work based on state-level variations in labour market conditions. The results of this research showed a positive association between maternal working hours and child BMI and obesity, particularly in households with higher socioeconomic levels.[1]

In contrast with previous findings, Bishop (2011) used economic conditions in the mother's area as instruments and sibling difference models and found that part-time and full-time maternal work decreased excess bodyweight in youths in Australia. Interestingly, a higher income did not appear to be responsible for this effect. Mocan et al. (2015) studied the impact of maternal earnings, determined by hours of work and wages, on birth weight and gestational age of infants in the US within an IV framework. They concluded that labour earnings had a positive but small effect on birth weight and gestational age of the newborns of low-skilled mothers, while an increase in the earnings of high-skilled mothers did not have any effect on the health of newborns. Finally, Greve (2011) found no effect of maternal working hours on child overweight status in Denmark, which contradicts all the relevant literature from other countries.

However, there is scarce literature on the causal impact of maternal employment on children's nutritional outcomes measured from anthropometric indicators (i.e., stunting, wasting, underweight and overweight). Existing studies have mainly been carried out in developing countries.[2] One of the most similar studies to that conducted herein is Rashad & Sharaf (2019), who investigated the causal impact of maternal employment on malnutrition indicators among U5-year-old children in Egypt, using local employment conditions as the instrument to account for the endogeneity of maternal employment. This current body of evidence points towards a stronger positive causal effect of maternal employment on the probability of being stunted or wasted.

3. Empirical specification and methodology

3.1. Child health production function

The empirical framework of this paper is based on the child health production function, which has its foundations in Becker's (1965) theory on the allocation of time between work and household goods production. Households are consumers and producers. They produce goods by combining inputs and time, in line with microeconomic theory. A more general framework was developed by Grossman (1972) with a model of demand for the commodity "good health". Jacobson (2000) extended Grossman's model by proposing a model where the family is the producer of health. In this sense, each family member will produce their own and other family members' health by investing resources in health until the rate of marginal consumption benefits equals the rate of marginal net effective costs of health capital. From an empirical standpoint, Rosenzweig & Schultz (1983) used exogenous variations in health to conclude that the correlation between health inputs and outcomes of a hybrid health production function may not be considered causal effects. Hence, the effect of endogenous health inputs can be estimated consistently when considering the presence of unobserved heterogeneity. Similarly, the Cebu Study Team (1992) addressed unobserved heterogeneity when it modelled and estimated a child health production function based on health input demands and health outcomes. Thus, longitudinal data methods

[1] Interestingly, Cawley & Liu (2012) showed that maternal employment is associated with less time spent on grocery shopping and on cooking, eating, playing with, supervising and caring for children, particularly among mothers with young children. This sheds light on the causes of childhood obesity when the mother is involved in the labour market.

[2] The broader correlational literature on the association between maternal employment and child malnutrition outcomes in developing countries is generally inconclusive. For instance, Van der Meulen Rodgers (2011) and Rastogi & Dwivedi (2014) found that mothers' labour supply was positively related to child malnourishment in some Asian countries and large cities in India, respectively. In contrast, Tucker and Sanjur (1988) found that maternal employment was not correlated with child nutritional status in Panama. See Glick (2002) for an extensive review.

along with the IV strategy were used to provide consistent estimates and therefore alleviate heterogeneity and endogeneity problems of important explanatory variables.

Several papers used this framework to estimate a child health production function based on an analysis of the relation between parental labour supply and child health (Bishop, 2011; Glick, 2002; Greve, 2011; Ruhm, 2008). In this setting, households invest time and goods (inputs) to produce child health (output) and maximize their utility by allocating parental time and income to produce time-intensive activities, such as devoting quality time to their kids, preparing healthy and nutritious food, and medical care. Since time and income are scarce commodities, two effects emerge when parents, particularly mothers, face the decision to work. If mothers reduce the time invested in taking care of their children, this can be detrimental to the child health capital stock. However, mothers who devote greater time to work can earn income and purchase more and better quality health inputs, which increases child health output. The trade-off between these two effects leads to a theoretically ambiguous net effect (Bishop, 2011).

3.2. Econometric model

We estimated the effect of mothers' labour status on their children's health outcomes according to the following structural equation:

$$ChildH_{ij} = \alpha + \beta MLS_{ij} + \gamma X_{ij} + \varepsilon_{ij} \quad (1)$$

where $ChildH_{ij}$ represents the health outcomes of interest of child i of mother j, i.e., stunting, wasting, underweight and overweight. MLS_{ij} is a vector of maternal labour supply variables, X_{ij} is a vector of social, economic and demographic characteristics of the child, mother and his/her family and ε_{ij} is the error term of the equation. Notice that parameter β in equation (1) is the parameter of interest, which captures the effect of mother's employment on child malnutrition. Estimating this model by OLS or binary response models would yield inconsistent, biased estimates, which only capture a correlation between child health and maternal employment. The presence of omitted variables and simultaneity issues makes MLS a potentially endogenous regressor ($cov(MLS, \varepsilon) \neq 0$). This would produce inconsistent OLS estimates.[3]

To address this problem, we used an instrumental variables (IV) framework on a two-stage least squares (2SLS) model (Wooldridge, 2003). The first stage estimates the endogenous variable MLS on a vector of exogenous instruments Z that do not directly affect children's health and all the remaining exogenous covariates of the structural equation (1).[4]

$$MLS_{ij} = \alpha_1 + \beta_1 Z_{ij} + \gamma_1 X_{ij} + \mu_{ij} \quad (2)$$

The second stage is the OLS regression of equation (1), which includes the prediction ($\widehat{MLS}$) from the first stage estimation and serves as an instrument for the endogenous variable. Thus, the second stage equation is:

$$ChildH_{ij} = \alpha + \beta_{IV} \widehat{MLS}_{ij} + \gamma X_{ij} + \omega_{ij} \quad (3)$$

[3] OLS estimates would appear downward biased if, for instance, employed mothers also had high skills in caring for children, which would act as a protective child health factor. In contrast, upward biased estimates would be observed if working mothers were less inclined to carry out childcare activities.

[4] The validity of instrument Z rests on two assumptions: i) the relevance condition where the correlation between MLS and Z is different from zero [$cov(Z, MLS) \neq 0$]; ii) the instrument and the error term in equation (1) are uncorrelated, $cov(Z, \varepsilon) = 0$. If these assumptions are satisfied, the instrument is considered "exogenous" in equation (1).

In this case, if the aforementioned assumptions about the instrument hold, parameter β_{IV} is consistent and unbiased and becomes the causal effect of maternal labour supply on child malnutrition outcomes. However, if the model in equation (1) is identified exactly, notice that parameter β_{IV} is equivalent to the ratio of the parameter associated with Z of the "reduced form" equation, i.e., the regression of the outcome against the instrument and the X covariates and the parameter β_1 of the first stage ($\beta_{IV} = \beta_{RF}/\beta_1$).

Although the outcome variables analysed in this study are all dichotomous (=1 if the child experienced stunting, wasting, underweight or overweight, and 0 otherwise), regression estimates from linear models would not be inadequate as the sample size employed is moderately large, and the mean of the outcome variables is centred (Cleary & Angel, 1984). In addition, the estimation of causal effects using an IV approach generates similar average effects, whether the model is linear or not.[5]

As a robustness analysis, we ran the seemingly unrelated regression (SUR) bivariate probit model (Greene 2018; Christofides et al., 1997). In this framework, structural and reduced form equations are estimated together, based on a binary choice probit model and allowing the errors of both equations to be correlated (i.e., $E(\varepsilon) = E(\mu) = 0, var(\varepsilon) = var(\mu) = 1, cov(\varepsilon, \mu) = \rho$). Note that if parameter ρ is significantly different from zero, which indicates the existence of unobservable factors related to maternal employment and child malnutrition, then the bivariate probit will provide consistent, unbiased estimates of the effect of maternal labour supply.

3.3. Endogeneity of maternal employment

Maternal labour supply is potentially endogenous when its association with child heath is analysed in equation (1). As mentioned before, this is for two main reasons. The first is associated with the presence of unobserved characteristics correlated with maternal labour supply and child health outcomes. For instance, mothers with low interest/motivation or skills/ability may work less and take less care of their kids than high interest/skilled mothers with more health knowledge (Anderson et al., 2003; Bishop, 2011; Greve, 2011). Similarly, Morrill (2011) argued that unobserved characteristics related to mother's preferences and skills might influence the choice of whether to work or not, which would lead to a selection bias in the mother's sample.

More interestingly, Cavapozzi et al. (2021) argued that unobserved cultural or gender norms may affect female labour market behaviour. For example, a mother's decision to work can be determined by the gender role in her household. Similarly, a woman whose peers present gender-egalitarian behaviour is more likely to work than a woman with more traditional peers. This gender role might also affect children's nutrition status in terms of the time spent by mothers taking care of their children.

Further potential unobserved characteristics correlated with maternal employment and parental inputs could be the decision to continue a pregnancy (wantedness), the taste for risky behaviours or maternal health endowment (Reichman et al., 2009). For instance, pregnant women who take care of themselves promptly because they want to continue with their pregnancy would be more likely to eat more nutritiously, avoid large amounts of stress and potentially harmful substances and engage in appropriate physical exercise. Therefore, the child's nutrition will be better due to pregnancy wantedness. In contrast, the taste for risky prenatal behaviour (i.e., cigarette smoking,

[5] Following Angrist (2001), the binary nature of the key endogenous regressor (maternal employment) would not be of concern as the 2SLS estimates are consistent ,regardless of whether or not the first stage regression is linear.

alcohol consumption and illicit drug use) leads to inadequate parental care and consequently detrimental child health status. Similarly, maternal health endowment prior to and during the pregnancy may affect the decision to work and impact child nutritional status.

The second issue is related to a reverse causality problem that emerges when the mother chooses whether to work or not based on her child's health status, or decide to quit her job to care for a child who suffers detrimental health conditions, inducing a negative correlation between employment and child health (Bishop, 2011). Furthermore, several studies have evidenced how child health status influences parental working decisions (Kuhlthau & Perrin, 2001; Mork et al., 2014; Tambi & Nkwelle, 2013).

3.4. *Instrumental variables*

This paper instrumented maternal labour supply (MLS) using the average working status and the average number of daily hours worked by employed and self-employed women measured at cantonal level. These exogenous instruments were assumed to capture local labour market conditions and the labour demand for women. Higher average employment rates and an intensive margin of maternal employment in a particular area makes it easier for resident mothers to find a job. In contrast, high unemployment or shorter working times in a cluster is a barrier to finding a job. Both instruments are an external factor for mothers and households, and are determined exogenously by local economic conditions (Bishop, 2011; Rashad & Sharaf, 2019).

Previous literature has used such instruments to deal with endogeneity problems in the attempt to find a causal relationship between mothers' labour supply and child health status (Anderson et al., 2003; Bishop, 2011; Datar et al., 2014; Greve, 2011; Rashad & Sharaf, 2019). A total of 206 cantons (ENSANUT 2018) and 213 cantons (ECV 2014) of the country were considered to compute the instruments. Reasonable variation was found among the geographical areas.[6]

Note that among self-employed women, only those working outside the home were included in the definition of maternal employment, to exclude mothers who earn an income and devote time to childcare at the same time. This avoids an in-built association. Moreover, the reference population for calculating the instrument was all working age women (15+), including economically active (employed and unemployed) and inactive women. It was assumed that the chances of finding a job affect not only mothers who are already employed, but also those who are looking for a job and those who have decided to study or stay at home looking for new opportunities to enter into the economically active population.

We considered that both cantonal-level variables were likely to be good instruments once we had controlled for mother's income and education, since there was variation in labour demand conditions and this variation was positively correlated with maternal labour supply in our sample. In addition, to assure the exogeneity of the instruments in the structural equation, other cantonal-level defined covariates were included in the second stage (Equation 1) that should be correlated with child health conditions such as poverty and inequality, as measured by the Gini index.[7] Thus, we addressed potential heterogeneity over cantons that might have confounded the IV estimates. Therefore, the instrument should be correlated with mother's working behaviour, and not directly

[6] Given the large number of cantons considered, disaggregation of the instruments by sector and/or industries was deemed inadequate due to the small number of observations.

[7] Cantonal level covariates included in equation 1 were compared to cantonal level covariates estimated from the Labor Force Survey (*Encuesta Nacional de Empleo, Desempleo y Subempleo*, ENEMDU 2018) to assure the consistency of these covariates. The comparison (Figure A2 in the Appendix) showed that these regional measures are highly similar in both surveys.

affect the probability of a child being malnourished, considering the relevance condition and the exclusion restriction.

4. Data and variables

This study used two sources of Ecuadorian cross-section microdata: the National Health and Nutrition Survey (ENSANUT) 2018 and the Life Conditions Survey (ECV) 2014, which were produced by the National Institute of Statistics and Censuses (INEC). Two datasets were used because the information on the number of hours worked by the interviewed individuals was lacking in the ENSANUT database. Fortunately, both datasets contained the information needed to run the econometric models.[8] They are intended to generate indicators on the main health situation and living standards of the Ecuadorian population (Instituto Nacional de Estadísticas y Censos, 2015, 2019a). The data requirements were extensive and included information on anthropometrics, children and women's health information, parent's socioeconomic variables and other household information.

The sample design implemented in ENSANUT 2018 followed a two-stage probabilistic stratified sample. The first stage stratified the sample through primary sampling units (PSU), while the second stage considered a certain number of dwellings randomly (18 dwellings on average) per PSU (Instituto Nacional de Estadísticas y Censos, 2019b). Thus, the survey investigated a total of 2,591 clusters and 43,311 dwellings. The total sample size was 168,747 individuals (INEC, 2019). Specifically, the ENSANUT 2018 sample used in this study considered 17,587 mothers matched with 20,204 U5-year-old children. Similarly, ECV 2014 was produced under the same sample design as ENSANUT 2018 (Instituto Nacional de Estadísticas y Censos, 2015) and the working sample considered 8,824 mothers matched with 10,837 children. Both surveys include canton identifiers so that regional market conditions can be estimated more precisely.

4.1. Child nutritional outcomes

The key outcomes investigated to characterise child malnutrition were three measures of undernutrition (stunting, wasting and underweight) and overweight status. According to the World Health Organization (2021), stunting or low height-for-age is the result of chronic or recurrent undernutrition and is usually related to deprived socioeconomic status, poor maternal health and nutrition, and inappropriate feeding and care in early life. Stunting appears to hold children back from reaching their physical and cognitive potential. Wasting or low weight-for-height indicates recent and severe weight loss, because of the scarcity of meals and/or the impact of infectious diseases. Children with low weight-for-age are known as underweight. In contrast, overweight status refers to a child who is too heavy for his or her height due to an excessive accumulation of fat, which impairs children's health.

To assess these nutritional outcomes, we used *clinically* measured anthropometric information available in both datasets. Each survey included sex and age in days of children U5 years old and their weight in kilograms (kg) and height/length in centimetres (cm). Using this information and the child growth standards developed by the WHO (de Onis et al., 2006), we estimated the standard deviation (SD) score (Z-score), one of the most common and frequently used indexes in the literature, as the difference between anthropometric value (weight, height/length) for a child and the median value of the reference population (WHO child growth standards) for the same sex and age, divided by the SD of the reference population (O'Donnell et al., 2007; WHO, 1995).

[8] The only exception is potable water control, which was not included in ECV 2014.

Therefore, we used three Z-scores (continuous and normally distributed variables) to define the anthropometric indicators: height-for-age Z-score (HAZ), weight-for-age Z-score (WAZ) and weight-for-height Z-score (WHZ).[9]

Following WHO (2006), a child with a HAZ, WHZ or WAZ lower than -2 SD is considered stunted, wasted or underweight respectively, while an overweight status corresponds to a child with a WAZ greater than +2 SD. Moreover, this study followed WHO recommendations of dropping values that are outside the range of plausible Z-scores (WHO, 2006).[10]

Note that the length, height and weight values of U5 children were clinically measured in survey data following a specific procedure (Instituto Nacional de Estadísticas y Censos, 2018). Each child was measured using a weight scale and stadiometer twice. The medical device was restarted each time.[11] A third measure was taken when the difference between the two first measurements was greater than ±0.5 kg/cm. For the purposes of this study, the final value of length, height and weight was computed as the average of the two first measurements if the differences between them were smaller than ±0.5 kg/cm. In contrast, if the difference between them was larger, the third anthropometric measurement was considered, and the final value would be the average between the two closest measures. The aim of following this strategy was to minimize any possible measurement error due to the specific measurement procedure applied during the interview.

4.2. Data variables

The key independent variable of this analysis was maternal employment (MLS). First, we measured it as a dichotomous variable that equals 1 for wage earning mothers and self-employed mothers working outside the household; and 0 otherwise. Second, we measured it as a continuous variable using the average number of hours worked per day. Note that while the former was calculated using ENSANUT 2018, the latter was computed by means of ECV 2014. As maternal employment is measured contemporaneously, the estimated effects were limited by the influence of current household conditions on health nutritional outcomes.[12]

A large set of controls was used to explain the variability of the nutritional outcomes of children. We included the following characteristics of children: i) gender, ii) age in months and squared age, since the relationship between a child's age and the dependent variable is non-linear, iii) information on early initiation of breastfeeding (equal to 1 if the child was born in the last 12 months and was put to the breast within one hour of birth, due to the first milk produced in the first days, which is called colostrum) as an important source of nutrition and immune protection for the newborn (World Health Organization, 2010). Regarding the mother's characteristics we accounted for: iv) maternal age and age squared, based on the same rationality as child age, v) mother's height and weight to control for potential genetic heritage, vi) maternal educational attainment level, vii) marital status, and viii) cultural origins (ethnicity). The latter is thought to play an important role in childhood nutrition since different ethnic groups in Ecuador may have varying food intake, culture and traditions. Among the family and environmental characteristics of children, the econometric specifications considered: ix) household income in quintiles, a key

[9] $HAZ = \frac{H_i - H_m}{SD_H}$, where H_i represents the height of the child, H_m is the median height of the reference population for the same sex and age and SD_H is the standard deviation of the height in the reference population. All other Z-scores followed the same criterion.
[10] According to WHO (2006), data are excluded if a child's HAZ is below –6 SD or above +6 SD, WAZ is below –6 SD or above +5 SD and WHZ is below –5 SD or above +5 SD.
[11] Children younger than 2 years old were measured in the supine position using a specific stadiometer.
[12] Note that we assumed zero worked hours for the inactive mother.

determinant of the child health production function discussed in section 2, x) paternal employment, xi) number of U5-year-old children in the household, xii) area of residence (urban/rural), xiii) number of women (aged 15 to 65) living in the same household, taken as a substitute for childcare when the mother works outside the home. Finally, we also included xiv) dwelling features such as overcrowding, inappropriate excreta disposal and unsafe drinking water that might be relevant in the nutrition of children, i.e., our preferred specification. Table A1 describes all variables used in this study, while Tables A2 and A3 in the Appendix show the descriptive statistics of these variables.[13]

5. Empirical results

5.1. Descriptive statistics

The data showed that long-term, chronic malnutrition in Ecuador is a public health challenge of the first order compared to short-term malnutrition. For instance, according to ENSANUT 2018, the prevalence of stunting among U5-year-old children living with their mothers is as high as 23.1%, while wasting and underweight have a much lower prevalence of 3.7% and 5.2%. The percentage of overweight children is 14.2%, which makes it an important health issue in the country (see Figure 1).[14] In the South American region, Ecuador has the highest percentage of stunting according to UNICEF. Neighbouring countries have much better nutritional conditions. For instance, in 2015–2016 Colombia reached a 12.7% stunting rate. More recently, in 2019, Peru had a 12.2% stunting prevalence rate (UNICEF et al., 2021).

The socioeconomic characterization of the malnourished children is shown in Table 1. Note that boys have poorer nutritional outcomes than girls, except for overweight, which has the same prevalence rate in both sexes. Further analysis showed that stunting was much greater among children of low educated mothers. Interestingly, the percentage of overweight children was much greater for mothers with a high level of education than those with a low or intermediate level. As expected, the percentage of stunted children in poor households was higher than the percentage in richer households. A significant difference of 16.3 percentage points in the prevalence of stunting was found between the lowest and the highest family income quintile. The phenomenon was the same for underweight children.[15]

High levels of stunting were found mainly in the Amazon region and in the central highlands of Ecuador. Nevertheless, the province that reported the highest percentage of stunting was in the coastal region. The south of the country was more affected by the prevalence of overweight, while the highest prevalence of wasting and underweight was mostly located in one province in the coastal region (Esmeraldas) and in the Amazon (see Figure A1 in the Appendix).

Table 2 shows the structure of the condition of maternal labour activity over relevant socioeconomic background. Note that in Ecuador the economically inactive group of mothers, regardless of place of residence, is larger than the economically active group (employed and unemployed). Interestingly, a considerable number of mothers in the no or primary and secondary education group were inactive, whereas mothers with tertiary education were more likely to be involved in the labour force. For instance, 64.3% of total working age mothers with tertiary education were employed whereas 5.2% were unemployed.

[13] Note that survey weights were utilized for all mean calculations and standard errors were corrected for the clustering effect.
[14] Similar (lower) rates for stunting and underweight (wasting and overweight) were reported by ECV 2014.
[15] Although not shown, this SES gradient of malnutrition was also documented by ECV 2014.

Table 2 reveals that the share of employed (unemployed) mothers rose (or fell) with the level of their family income. This result is unsurprising, since mothers who work earn more money for their households. Finally, single and divorced mothers were more likely to be employed, while married women or women in couples were more likely to be inactive. A highly similar labour market composition was documented in the ECV 2014 dataset.

Overall, the descriptive evidence from both datasets indicated the existence of a social gradient of stunted children concentrated among socioeconomically vulnerable households, particularly those with low education levels and poor mothers. On the contrary, overweight children were more likely to live in households with highly educated mothers. We did not find much variation in wasted children with respect to socioeconomic background. Moreover, most mothers were out of the labour market or inactive, especially those married or in a couple and with low education and medium family income. In contrast, highly educated mothers with the highest family income were more likely to be involved in the labour force as employed mothers.

5.2. OLS estimates

The first step to empirically analyse the relationship between maternal labour supply and nutritional status of children was to run an OLS model estimation (Table 3). The baseline model considered up to four specifications based on a wide set of controls.[16] The estimates reported in Panel A evidence that dummy maternal employment was not significantly related to stunting in children. However, when MLS was proxied by the number of daily work hours (Panel B) the coefficient of maternal employment became, although modestly, positive and significantly related to stunting, i.e., children of working mothers were on average between 0.4 and 0.6 percent more likely to be stunted than children whose mothers did not work (unemployed or economically inactive) for each additional work hour.

Most coefficients of the controls were highly significant with the expected sign (see Table A4). For example, mothers with secondary and tertiary education were less likely to have stunted children than mothers who had no or primary education. Similarly, children of mothers living in richer households were less likely to be stunted than those who lived in poorer households. Interestingly, the higher the number of U5 children residing in the household, the higher the probability of having a stunted child. Moreover, the taller and the heavier the mother (a sign of past health status and prosperity), the lower the probability of having a stunted child. Remarkably, we found that early initiation of breastfeeding decreased the probability of being stunted. Finally, column (4) of Table A4 reveals that mothers in the indigenous ethnic group were more likely to have stunted children than mothers with mestizo cultural origins.[17]

In contrast, the remaining columns of Table 3 evidence that maternal employment was unrelated to underweight, wasting and overweight in children in Ecuador.[18]

As discussed above, it is problematic to estimate the impact of maternal employment on all nutritional outcomes of children based on an OLS framework, since the mother's choice to work could be driven by unobserved characteristics that are also correlated with children's likelihood of being stunted, underweighted, wasted or overweight. In the next section, we test whether the relationship between maternal employment and nutritional outcomes is exogenous.

[16] Standard errors are clustered at the PSU level in all specifications.
[17] The coefficients associated with each control remain highly constant across specifications.
[18] The exception was the significant positive association between dummy maternal employment and overweight when only children's characteristics were controlled for.

5.3. *Testing for maternal labour supply exogeneity*

We used the Durbin-Wu-Hausman (DWH) test to evaluate the consistency of the OLS estimator of maternal employment compared to an alternative, consistently estimated IV model. We found that the DWH test rejected the null hypothesis of exogeneity in both surveys (F-statistic = 36.187, p-value = 0.000 in ENSANUT 2018; F-statistic = 80.879, p-value = 0.000 in ECV 2014). This suggests that maternal employment in the OLS framework for the equation of stunting children is endogenous. Thus, the estimates of maternal labour supply presented in Table 3 regarding the stunting model are inconsistent and biased.

However, the DWH tests did not reject the null hypothesis of exogeneity of maternal employment for the other outcomes.[19] These results indicate that OLS will provide consistent estimates of maternal employment on all nutritional outcomes except for stunting. Notwithstanding, we will address the estimates of the causal effect of maternal labour supply on children's nutrition based on the IV framework.

5.4. *Instrumental variable estimates*

We report in Table A5 the estimates of the first stage regression based on the preferred specification. Using ENSANUT 2018, the findings indicate that the instrument mothers' employment status at cantonal level was strongly correlated with maternal employment. Similarly, as for ECV 2014, our estimates show that cantonal-level average maternal hours worked were significantly related to the endogenous maternal labour supply, although employment status was unrelated. Furthermore, at the bottom of this table the Montiel-Olea & Pflueger (2013) robust effective first-stage F statistic and the corresponding p-value of the test of weak instruments is reported. In both cases, this statistic was well above the critical values and the null hypothesis of a weak instrument was rejected. Therefore, the relevance condition was satisfied.

Tables 4 and A6 show the main findings of the paper. Table 4 displays the two IV-2SLS estimates for the outcome of stunting. Note that each structural form equation includes two cantonal level variables (poverty and inequality) to ensure the exogeneity of the instrument and thus satisfy the validity condition. According to column (1), after controlling for observed characteristics (preferred specification), unobserved characteristics and reverse causality, maternal employment status increases the probability of having a stunted U5-year-old child by roughly 21 percent compared to unemployed and inactive mothers. However when we considered two instruments and ECV 2014, we found that the heteroscedasticity-robust Wooldridge's score test rejected the null hypothesis of overidentifying restrictions at the 5% significance level (Chi2(1) = 9.885, p-value = 0.002). Hence, the estimates presented in column (2) are the just identified model instrumenting the endogenous variable with cantonal-level mothers' average work hours. Interestingly, we found that each additional hour of work per day spent by the working mother raised the probability of a stunted child by 4.3 percent with respect to unemployed and inactive mothers. Table 4 shows that most of the coefficients associated with the controls had the expected sign and were statistically significant.

In contrast, Table A6 shows that the effect of maternal employment on underweight, wasting and overweight children in both surveys was again not significantly different from zero.[20] Overall,

[19] Regarding ENSANUT, the F-statistic for underweight was 0.022 (p-value=0.882), for wasting it was 0.099 (p-value=0.753) and for overweight it was 0.420 (p-value=0.517). For ECV 2014, the F-statistic for underweight was 0.139 (p-value=0.709), for wasting it was 3.050 (p-value=0.081) and for overweight it was 0.012 (p-value=0.913).

[20] The exception was the negative and significant, although very small, impact on wasting.

these findings suggests that maternal employment in Ecuador might have an impact on stunting or chronic long-term malnutrition, but does not affect these other malnourishment outcomes. Hence, the OLS model may underestimate the effect of maternal labour supply on stunted children to a large extent.

5.5 Heterogeneous effects on children stunting

In Table 5 we investigate heterogenous effects and explore whether and to what extent the impact of maternal employment on children's stunting differs by socioeconomic status. With respect to education groups, the IV estimates performed on both surveys confirm a much larger, statistically significant impact on children's stunting among mothers with tertiary education. Similarly, the findings display stronger, statistically significant effects among the group of mothers living in high-income families.

Overall, this evidence would appear to suggest that more educated, richer mothers, hence those with higher skills and ability, tend to incur higher opportunity costs of working outside home. This results in poorer health nutrition outcomes for their children. For these population groups, the time sacrificed in terms of providing better, healthier food and care for children (time constraint) may dominate the extra income ensured by working in better paid jobs (income effect).

6. Robustness checks

Table 6 explores the robustness of the main findings under two different approaches. The first was associated with using a different sample criterion, while the second considered the same sample of the baseline scenario but used a seemingly unrelated regression (SUR) bivariate probit model instead. Note that all these estimates were based on the preferred specification.

We report in columns (1) and (2) of Table 6 the OLS and IV-2SLS estimates, respectively, based on the ENSANUT 2018 survey when the estimation sample was restricted to mothers involved in the labour force only (economically active mothers). This comprised a much-reduced sample size.[21] Therefore, the analysis now focuses on the trade-off between maternal employment and unemployment. Interestingly, the results reported in column (1) are not statistically significant for any malnutrition outcome. Hence, they are similar to the OLS estimates based on the complete sample reported in Table 3. However, when the IV-2SLS model was run we found roughly a similar effect of maternal employment on children stunting. Working mothers appeared to have a 23.8 percent greater probability of having a stunted child than mothers who are unemployed, although this effect was significant at the 10% level.

The second robustness check considered the complete sample of working age mothers or the baseline scenario but ran a binary choice model with endogenous regressors. In particular, a SUR bivariate probit model was fitted. This allowed for the two probit equations with correlated disturbance terms, which were assumed to come from a joint or bivariate normal distribution (Greene, 2018). Column (3) of Table 6 reports the average partial effect of maternal employment on nutritional outcomes. As expected, for stunted children the coefficient was positive and highly significant. This suggests that maternal employment raises the probability of having a stunted child by 18.4 percent, i.e., this coefficient was similar in size to the impact found using the IV-2SLS model. Furthermore, we found that the error terms of both probit equations were correlated ($\rho = -0.3786$; S.E. 0.0584),[22] which reveals that maternal employment is endogenous. There may

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[21] The instrument used in this framework was women employment rate at cantonal level, considering the economically active women as the reference population.

[22] A Wald test of $\rho = 0$ was rejected for the stunting equation (Chi2 = 44.7321, p-value = 0.0000).

be unobservable characteristics of individuals that adversely influence stunted children and employed mothers and the SUR bivariate probit model is recommended to obtain consistent estimates of the structural equation parameters. The estimates for the remaining children's outcomes were neither significant nor consistent.[23]

6.1. Impact of maternal labour supply on severe and moderate stunting

As a further robustness check we present in Table 7 the causal effect of maternal employment on U5 children who suffer from moderate or severe stunting based on ENSANUT 2018.[24] Based on the preferred specification, we show that maternal employment appears to lead to an increase in the probability of having a moderated stunted child by 15.5 percent compared with maternal unemployment and inactivity. To a lower extent, the probability of having a severely stunted child was 5.8 percent higher for working mothers. Further analysis suggests that the probability of being moderately or severely stunted was higher for boys than for girls, as was reported in the baseline model. In contrast, an increase in the number of U5 children living in the household increases the likelihood of being moderately stunted to a larger extent than the probability of being severely stunted. Regarding the socioeconomic characteristics of the mother, children of mothers with no education or primary education have a higher risk of being moderately or severely stunted, with respect to those who live with mothers who have completed tertiary education. In addition, divorced mothers are less likely to have a moderately stunted child than married mothers/mothers in couples. Interestingly, only indigenous mothers are more likely to have a severely stunted child in comparison to the larger ethnic group of mestizos.

7. Discussion and conclusions

This study sought to determine whether maternal employment influences malnutrition in Ecuadorian U5-year-old children. Based on cross-sectional and individual-level data from ENSANUT 2018 and ECV 2014, we report evidence that, after accounting for the endogeneity of maternal employment, working mothers are from 4.3 to 21 percent more likely to have stunted or chronically malnourished children than unemployed and inactive mothers. Interestingly, these impacts are even stronger for children of mothers with high education living in high income families. This result suggests that the maternal time devoted to childcare and feeding activities is far more important for children's health than the additional income that a working mother may obtain (e.g., Datar et al., 2014). These estimated effects of maternal employment on stunted children are in agreement with those found in the literature for low and lower-middle income countries. For instance, Amaha & Woldeamanuel (2021) reported that unemployed mothers were 23 percent (p-value < 0.01) less likely to have a stunted child than employed mothers in Ethiopia. Similarly, Rashad & Sharaf (2019) evidenced that the probability of being stunted is 18.6 percent (p-value < 0.05) higher for children whose mothers are employed in Egypt.

When chronic malnutrition was disaggregated by severity, we found that maternal employment was also an equally remarkable determinant of a child's risk of experiencing both moderate and severe stunting once endogeneity had been accounted for. While a working mother is 15.5 percent more likely to have a moderately stunted child, the likelihood of having a severely stunted child is roughly 6 percent higher. Again, the existence of a socioeconomic gradient relative to child malnourishment is highly prevalent. Mothers in the poorest family income quintile and those with

[23] The null hypothesis of correlation between the error terms for the other outcomes were rejected at conventional levels. This is similar to the evidence shown by the DWH test.

[24] Moderate stunting: height/length for age z-score in the range of -2 SD, -3 SD. Severe stunting: height/length for age z-score in the range of -2 SD, -6 SD. Definitions taken from consensus reached among experts in children health (World Health Organization, 1995).

lower or no education seem to have more chance of raising a child with malnourishment problems. Interestingly, boys have a higher risk of malnourishment than girls. This shows that children's environment influences whether they suffer detrimental effects in their health status due to malnutrition.

However, this paper found no empirical evidence for the impact of maternal employment on children suffering from wasting, underweight or overweight. The effects are negligible in both the OLS and IV-2SLS regression estimations. Our results are robust to an alternative statistical modelling approach and different samples of mothers.

The results of this study highlight the importance of government interventions to tackle the adverse impact of maternal employment on children's nutrition. Effective conditional cash transfers and/or in-kind family policies targeted at poor and low-skilled mothers and intended to reduce the cost of raising children among vulnerable families are supported by our findings. Likewise, the evidence that early initiation of breastfeeding is correlated with the risk of children being stunted should underpin labour policies for mothers after child delivery, such as extending the breastfeeding period or developing laws to encourage breastfeeding in the workplace.

Overall, the findings of this paper reveal the existence of a socioeconomic gradient associated with malnourished children in Ecuador. That is, children whose mothers have lower socioeconomic status are more likely to suffer stunting, as documented in previous literature (Case et al., 2002, 2005; Currie, 2009).

Potential selectivity bias due to the non-observation of U5-year-old children who did not survive until the date of data collection seems to be of no concern in our setting. Nevertheless, the percentage of mothers who reported U5 dead children in the last five years prior to the survey collection was less than the 1% of the sample.

Our results should be interpreted with caution. For instance, the study focuses on how mothers' labour market decisions impact children's nutrition in the short run. Any long-term effects are disregarded. In fact, maternal employment could lead to a positive net effect on children's health stock in the long run. A deeper understanding of this effect is needed. Notably, the estimates documented here are based on a cross-section analysis. Consequently, some time-invariant unobserved characteristics may not be accounted for, despite the fact that an IV strategy is used.

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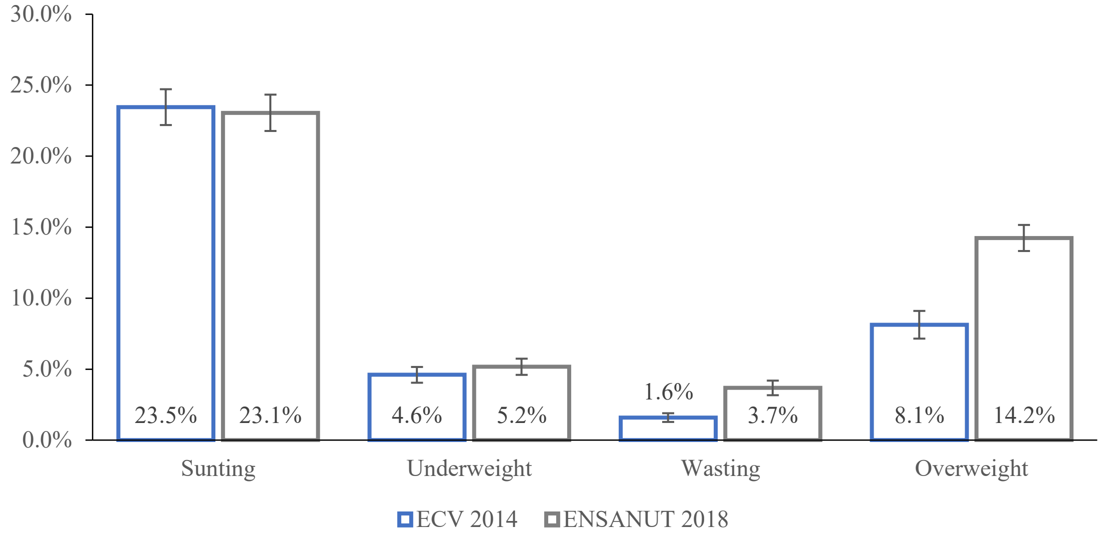
Figures

Figure 1. Nutritional and health outcomes for children U5 years old



Notes: Figures presented here are estimated considering the sample of U5-year-old children living with their mothers and survey weights.

Tables

Table 1. Child malnutrition by socioeconomic characteristics. ENSANUT 2018

Variable	Stunting	Underweight	Wasting	Overweight
Child's sex				
Male	24.8%	6.0%	4.2%	14.2%
Female	21.2%	4.3%	3.1%	14.3%
Mothers' education				
No education/primary	29.0%	6.4%	3.1%	13.3%
Secondary	22.7%	5.3%	4.1%	13.5%
Tertiary	17.1%	3.6%	3.1%	17.1%
Family income (in quintiles)				
Lowest	32.3%	7.1%	3.6%	15.2%
Second	23.9%	5.7%	3.7%	12.7%
Middle	22.6%	5.7%	3.5%	13.9%
Fourth	20.2%	5.1%	4.9%	13.6%
Highest	16.1%	2.2%	2.6%	15.7%

Source: ENSANUT 2018

Table 2. Maternal activity condition by socioeconomic characteristics. ENSANUT 2018

Variable	Employed	Unemployed	Inactive
Region of residence			
Urban	41.5%	3.4%	55.1%
Rural	44.6%	5.3%	50.2%
Mothers' education			
No education/primary	38.6%	2.8%	58.6%
Secondary	37.5%	5.3%	57.2%
Tertiary	64.3%	5.2%	30.5%
Household income quintile			
Lowest	34.3%	7.5%	58.3%
Second	26.9%	5.6%	67.5%
Middle	30.8%	4.6%	64.6%
Fourth	50.5%	3.8%	45.7%
Highest	75.3%	2.0%	22.7%
Mothers' marital status			
Married or union	40.8%	3.2%	55.9%
Divorced	56.6%	8.7%	34.7%
Single	52.2%	12.0%	35.7%

Source: ENSANUT 2018

Table 3. OLS estimation: Impact of mother's employment on malnutrition among U5-year-old children

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
	Stunting				Underweight				Wasting				Overweight			
	Model 1	Model 2	Model 3	Model 4	Model 1	Model 2	Model 3	Model 4	Model 1	Model 2	Model 3	Model 4	Model 1	Model 2	Model 3	Model 4
PANEL A																
ENSANUT 2018																
Maternal employment	-0.001	0.013*	0.008	0.001	-0.006*	-0.003	-0.004	-0.004	0.000	0.001	-0.001	-0.000	0.010**	0.009	0.009	0.006
	(0.007)	(0.007)	(0.007)	(0.007)	(0.003)	(0.004)	(0.004)	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.005)	(0.006)	(0.006)	(0.006)
Observations	18,902	18,369	18,060	18,060	18,947	18,410	18,099	18,099	18,268	17,747	17,450	17,450	18,891	18,355	18,046	18,046
R-squared adjusted	0.0102	0.0496	0.0579	0.0605	0.00163	0.00876	0.0108	0.0109	0.00748	0.00708	0.00739	0.00741	0.0103	0.0134	0.0143	0.0150
PANEL B																
ECV 2014																
Maternal working hours	0.005***	0.005***	0.006***	0.004***	-0.001*	-0.001*	-0.001	-0.001	0.000	0.000	0.000	0.000	-0.000	-0.000	-0.000	-0.000
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
Observations	10,557	10,441	10,428	10,428	10,595	10,479	10,466	10,466	10,459	10,343	10,330	10,330	10,577	10,461	10,448	10,448
R-squared adjusted	0.0187	0.116	0.131	0.136	0.00194	0.0182	0.0253	0.0260	0.0102	0.0112	0.0116	0.0115	0.00548	0.0159	0.0176	0.0176
Child char.	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mother char.	No	Yes	Yes	Yes	No	Yes	Yes	Yes	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Househ. cha.	No	No	Yes	Yes	No	No	Yes	Yes	No	No	Yes	Yes	No	No	Yes	Yes
Ethnic char.	No	No	No	Yes	No	No	No	Yes	No	No	No	Yes	No	No	No	Yes

Notes: Standard errors in parentheses clustered by primary sampling units. *** p<0.01, ** p<0.05, * p<0.1

Table 4. IV – 2SLS: Impact of mother's employment on children stunting

Variables	(1) ENSANUT 2018	(2) ECV 2014
Maternal employment status	0.209***	-
	(0.039)	
Maternal working hours	-	0.043***
		(0.006)
Child's sex (male)	0.039***	0.035***
	(0.007)	(0.009)
Child's age in months	0.001	0.009***
	(0.001)	(0.001)
Child's age in months square	-0.000***	-0.000***
	(0.000)	(0.000)
Early breastfeeding	-0.069***	-0.082***
	(0.014)	(0.018)
Mother's age	-0.013***	-0.015***
	(0.004)	(0.005)
Mother's age square	0.000**	0.000**
	(0.000)	(0.000)
Mother's height	-0.008***	-0.014***
	(0.001)	(0.001)
Mother's weight	-0.002***	-0.002***
	(0.000)	(0.000)
Education (no or primary education)		
Secondary education	-0.042***	-0.035***
	(0.009)	(0.012)
Tertiary education	-0.104***	-0.063***
	(0.014)	(0.016)
Marital status (married or in couple)		
Divorced	-0.044***	-0.047
	(0.017)	(0.030)
Single	-0.005	-0.053*
	(0.017)	(0.030)
Paternal employment	0.016	0.022
	(0.013)	(0.026)
Children in the household	0.039***	0.037***
	(0.008)	(0.009)
Women in the household	0.001	-0.010
	(0.005)	(0.007)
Area of residence (urban)	-0.014*	-0.000
	(0.009)	(0.013)
Overcrowding	0.021**	0.019
	(0.009)	(0.012)
Inappropriate excreta disposal	0.027***	-0.021
	(0.009)	(0.014)
Safe drinking water	0.005	
	(0.007)	

Family income quintile (poorest)		
Second	0.018	-0.031**
	(0.013)	(0.015)
Middle	-0.006	-0.076***
	(0.013)	(0.017)
Fourth	0.003	-0.135***
	(0.013)	(0.021)
Highest	-0.008	-0.188***
	(0.012)	(0.026)
Cultural origins (mestizo)		
Indigenous	0.018	0.037*
	(0.014)	(0.019)
Black	-0.023	0.005
	(0.015)	(0.020)
Montubio	-0.003	0.001
	(0.017)	(0.023)
White	-0.017	0.039
	(0.030)	(0.034)
Cluster poverty	0.074**	-0.027
	(0.031)	(0.034)
Cluster Gini index	0.011	-0.000
	(0.047)	(0.099)
Constant	1.633***	2.605***
	(0.116)	(0.162)
Observations	17,193	10,428

Notes: Standard errors in parentheses clustered by primary sampling units; *** p<0.01, ** p<0.05, * p<0.1. Safe drinking water is not available in ECV 2014.

Table 5. IV – 2SLS: Impact of mother's employment on children stunting. Heterogenous effects by education and family income

Heterogeneity	(1) Maternal labour supply	(2) Maternal working hours
	ENSANUT 2018	ECV 2014
Education of the mother		
No or primary	0.182***	0.031***
	(0.052)	(0.007)
Secondary and tertiary	0.221***	0.056***
	(0.053)	(0.010)
Family income quintile		
Poorest and second	0.189***	0.033***
	(0.053)	(0.009)
Middle	0.116	0.046***
	(0.075)	(0.010)
Fourth and highest	0.310***	0.053***
	(0.075)	(0.010)

Notes: Regressions run based on the preferred specification including cantonal level poverty and Gini indexes. Standard errors in parentheses clustered by primary sampling units; *** p<0.01, ** p<0.05, * p<0.1.

Table 6. Impact of mother's employment on malnutrition outcomes. ENSANUT 2018

Malnutrition outcomes	(1)	(2)	(3)
	Labour force sample		Working age sample
	OLS	IV-2SLS	Bivariate Probit
Stunting	0.017	0.238*	0.184***
	(0.016)	(0.131)	(0.001)
Underweight	0.002	-0.066	-0.011
	(0.008)	(0.077)	(0.000)
Wasting	0.003	-0.028	-0.009
	(0.007)	(0.052)	(0.000)
Overweight	0.009	0.038	0.035
	(0.014)	(0.094)	(0.001)

Notes: Regressions run based on the preferred specification including cantonal level poverty and Gini indexes. Standard errors have been clustered by primary sampling units. The coefficient obtained from the SUR bivariate probit model has been computed as the average partial effect. Standard errors for the average partial effects are computed using bootstrapping method with 100 repetitions. *** p<0.01, ** p<0.05, * p<0.1.

Table 7. IV – 2SLS: Impact of mother's employment on moderate and severe children stunting. ENSANUT 2018

Variables	(1) Moderate	(2) Severe
Maternal employment	0.155***	0.058**
	(0.033)	(0.027)
Observations	17,194	17,194

Notes: Regressions run based on the preferred specification including cantonal level poverty and Gini indexes. Standard errors in parentheses clustered by primary sampling units; *** p<0.01, ** p<0.05, * p<0.1.

Appendix

Table A1. Variable description

Variable	Description
Stunting	Equals 1 if HAZ is lower than -2 SD, 0 otherwise.
Wasting	Equals 1 if WHZ is lower than -2 SD, 0 otherwise.
Underweight	Equals 1 if WAZ is lower than -2 SD, 0 otherwise.
Overweight	Equals 1 if WAZ is higher than +2 SD, 0 otherwise.
Maternal employment status	Equals 1 for wage earners mothers and self-employed mothers working outside the household; and 0 otherwise.
Maternal hours worked	Average number of hours worked per day by the employed and self-employed mother
Child's sex	Equals 1 if child is male, 0 otherwise.
Child's age in months	Age in months since birth.
Child's age in months square	Square of child's age in months.
Early breastfeeding	Equals 1 if the child had an early initiation of breastfeeding, 0 otherwise.
Mother's age	Age in years since birth.
Mother's age square	Square of mother's age in years.
Mother's height	Height of the mother in centimetres.
Mother's weight	Weight of the mother in kilograms.
Education of the mother	Dummies for educational attainment levels: Primary, secondary, and tertiary.
Marital status of the mother	Dummies for marital status of the mother: Married or in couple, divorced and single.
Paternal employment	Equals 1 for working fathers; and 0 otherwise.
Children in the household	Number of all under-five-year-old children living in the household.
Women in the household	Number of all women between 15 to 65 years old living in the same household of the children.
Area of residence	Equals 1 if area of residence is urban, 0 otherwise.
Overcrowding	Equals 1 if the household is overcrowded, 0 otherwise.
inappropriate excreta disposal	Equals 1 if the household does not have appropriate excreta disposal, 0 otherwise.
Safe drinking water	Equals 1 if the household has safe drinking water, 0 otherwise.
Family income quintile	Per capita household income in quintiles without the labour mother's income.
Cultural origins	Dummies for the ethnicity of the child: Mestizo, indigenous, black, montubio and white.
Cantonal-level poverty	Poverty headcount rate measured at cantonal level.
Cantonal-level Gini index	Gini index measured at cantonal level.

Table A2. Descriptive statistics, means. ENSANUT 2018

Variable	Obs.	Weight	Mean	Std. Error	Min	Max
Child characteristics						
Stunting	18,921	1243697	0.231	0.006	0.000	1.000
Underweight	18,966	1244480	0.052	0.003	0.000	1.000
Wasting	18,286	1203377	0.037	0.003	0.000	1.000
Overweight	18,910	1240977	0.142	0.005	0.000	1.000
Sex (male)	20,204	1311403	0.514	0.006	0.000	1.000
Age in months	20,204	1311403	29.406	0.207	0.000	59.000
Early initiation breastfeeding	20,204	1311403	0.138	0.004	0.000	1.000
Mother's characteristics						
Age	20,204	1311403	28.565	0.091	11.000	71.000
Height	19,117	1240977	154.147	0.108	110.500	199.000

Weight	19,117	1240977	64.318	13.605	29.500	160.000
Household characteristics		Safe				
Paternal employment	18,921	1243697	0.747	0.007	0.000	1.000
Children in the household	18,339	1204357	1.267	0.008	0.000	4.000
Women in the household	18,921	1243697	1.452	0.010	0.000	8.000
Area of residence (urban)	18,921	1243697	0.683	0.007	0.000	1.000
Overcrowding	18,921	1243697	0.254	0.007	0.000	1.000
inappropriate excreta disposal	18,921	1243697	0.236	0.009	0.000	1.000
Safe drinking water	18,828	1239204	0.342	0.008	0.000	1.000

Table A3. Descriptive statistics, proportions. ENSANUT 2018

Variable	Obs.	Weight	Mean	Std. Error
Activity condition				
Employed	20,204	1311403	0.435	0.007
Unemployed	20,204	1311403	0.047	0.003
Inactive	20,204	1311403	0.517	0.007
Level of education				
Non-education/primary	20,204	1311403	0.244	0.006
Secondary	20,204	1311403	0.540	0.007
Tertiary	20,204	1311403	0.216	0.006
Marital status				
Married or union	20,201	1311350	0.798	0.006
Divorced	20,201	1311350	0.102	0.005
Single	20,201	1311350	0.100	0.004
Familiar income quintile				
Lowest	20,199	1310726	0.201	0.006
Second	20,199	1310726	0.204	0.006
Middle	20,199	1310726	0.196	0.005
Fourth	20,199	1310726	0.200	0.005
Highest	20,199	1310726	0.200	0.006
Cultural origins				
Indigenous	20,204	1311403	0.090	0.004
Black	20,204	1311403	0.042	0.003
Montubio	20,204	1311403	0.051	0.003
Mestizo	20,204	1311403	0.803	0.006
White	20,204	1311403	0.014	0.002

Figure A1. Malnutritional outcomes over province, ENSANUT 2018

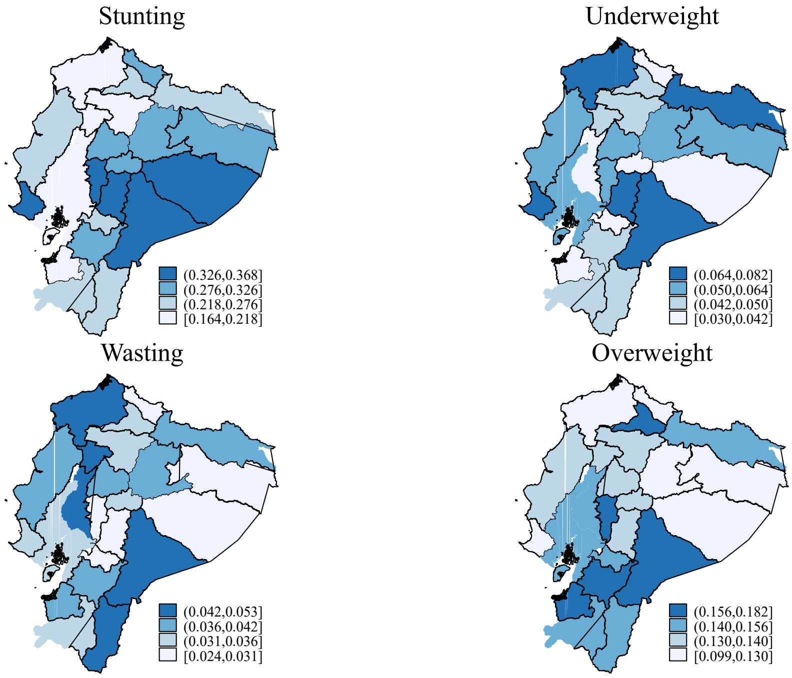
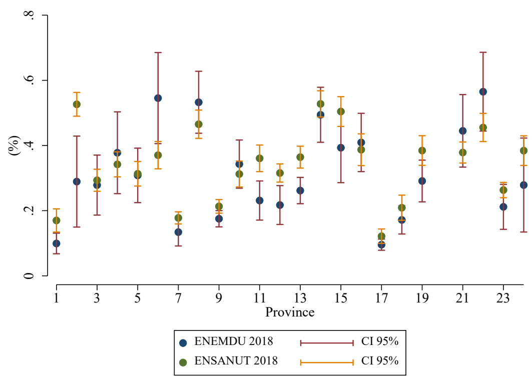
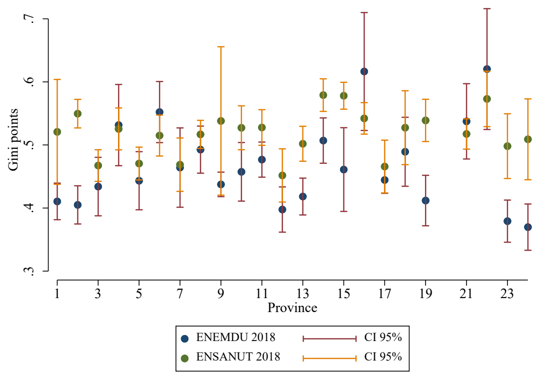


Figure A2. Provincial-level covariates comparison ENSANUT 2018 vs ENEMDU 2018

Panel A: Employment rate



Panel B: Gini coefficient index



Panel C: Poverty rate

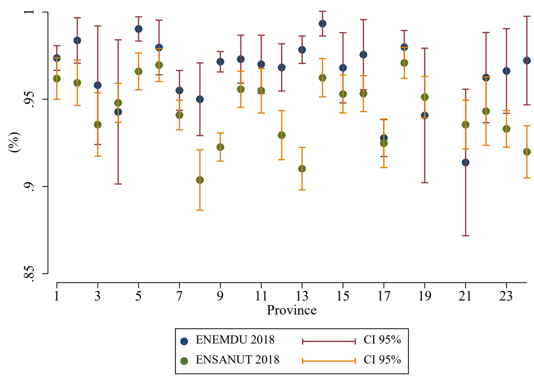


Table A4. OLS estimation: the impact of mother's employment on U5-year-old child nutritional status

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
	Stunting				Underweight				Wasting				Overweight			
	Model 1	Model 2	Model 3	Model 4	Model 1	Model 2	Model 3	Model 4	Model 1	Model 2	Model 3	Model 4	Model 1	Model 2	Model 3	Model 4
Maternal employment	-0.001	0.013*	0.008	0.001	-0.006*	-0.003	-0.004	-0.004	0.000	0.001	-0.001	-0.000	0.010**	0.009	0.009	0.006
	(0.007)	(0.007)	(0.007)	(0.007)	(0.003)	(0.004)	(0.004)	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.005)	(0.006)	(0.006)	(0.006)
Child's sex (male)	0.042***	0.043***	0.042***	0.042***	0.013***	0.013***	0.012***	0.012***	0.007**	0.006**	0.006**	0.006**	0.017***	0.018***	0.018***	0.017***
	(0.006)	(0.006)	(0.006)	(0.006)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.005)	(0.005)	(0.005)	(0.005)
Child's age in months	0.002*	0.002*	0.002**	0.002**	-0.001	-0.001	-0.001	-0.001	-0.003***	-0.002***	-0.002***	-0.002***	-0.006***	-0.006***	-0.006***	-0.007***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Child's age in months square	-0.000***	-0.000***	-0.000***	-0.000***	0.000	0.000	0.000	0.000	0.000***	0.000***	0.000***	0.000***	0.000***	0.000***	0.000***	0.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Early breastfeeding	-0.053***	-0.062***	-0.066***	-0.069***	0.006	0.006	0.005	0.005	0.002	0.006	0.006	0.006	0.000	-0.001	0.001	0.000
	(0.014)	(0.014)	(0.014)	(0.014)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.012)	(0.012)	(0.012)	(0.012)
Mother's age		-0.002	-0.003	-0.003		-0.001	-0.001	-0.001		-0.001	-0.001	-0.001		-0.003	-0.001	-0.000
		(0.003)	(0.004)	(0.004)		(0.002)	(0.002)	(0.002)		(0.001)	(0.002)	(0.002)		(0.003)	(0.003)	(0.003)
Mother's age square		0.000	0.000	0.000		0.000	0.000	0.000		0.000	0.000	0.000		0.000	0.000	0.000
		(0.000)	(0.000)	(0.000)		(0.000)	(0.000)	(0.000)		(0.000)	(0.000)	(0.000)		(0.000)	(0.000)	(0.000)
Mother's height		-0.008***	-0.008***	-0.007***		-0.001***	-0.001***	-0.001***		0.001**	0.001**	0.001**		0.001**	0.001**	0.001**
		(0.001)	(0.001)	(0.001)		(0.000)	(0.000)	(0.000)		(0.000)	(0.000)	(0.000)		(0.000)	(0.000)	(0.000)
Mother's weight		-0.002***	-0.002***	-0.002***		-0.001***	-0.001***	-0.001***		-0.000	-0.000	-0.000		0.001***	0.001***	0.001***
		(0.000)	(0.000)	(0.000)		(0.000)	(0.000)	(0.000)		(0.000)	(0.000)	(0.000)		(0.000)	(0.000)	(0.000)
Education (non or primary education)																
Secondary education		-0.061***	-0.037***	-0.036***		-0.012**	-0.007	-0.007		0.001	0.002	0.002		0.001	0.003	0.003
		(0.009)	(0.009)	(0.009)		(0.004)	(0.005)	(0.005)		(0.003)	(0.004)	(0.004)		(0.007)	(0.007)	(0.007)
Secondary education		-0.103***	-0.062***	-0.059***		-0.024***	-0.015***	-0.014***		-0.001	0.003	0.003		0.019**	0.016*	0.017*
		(0.010)	(0.011)	(0.011)		(0.005)	(0.006)	(0.005)		(0.004)	(0.005)	(0.005)		(0.008)	(0.009)	(0.009)
Marital status (married or in union)																
Divorced		-0.027**	-0.018	-0.014		0.006	0.006	0.006		-0.006	0.001	0.001		-0.009	0.006	0.007
		(0.011)	(0.015)	(0.015)		(0.006)	(0.009)	(0.009)		(0.004)	(0.007)	(0.007)		(0.009)	(0.012)	(0.012)
Single		0.007	0.020	0.021		0.011*	0.013	0.012		-0.001	0.006	0.006		0.014	0.024*	0.025**
		(0.011)	(0.015)	(0.015)		(0.006)	(0.008)	(0.008)		(0.005)	(0.007)	(0.007)		(0.009)	(0.013)	(0.013)
Paternal employment			0.013	0.010			0.001	0.001			0.011*	0.011**			0.016	0.015
			(0.012)	(0.012)			(0.007)	(0.007)			(0.006)	(0.006)			(0.010)	(0.010)
Children in the household			0.036***	0.032***			0.005	0.006			-0.006**	-0.005*			0.005	0.003
			(0.007)	(0.007)			(0.004)	(0.004)			(0.003)	(0.003)			(0.006)	(0.006)
Women in the household			-0.001	-0.002			-0.001	-0.001			0.001	0.001			0.006	0.006
			(0.005)	(0.005)			(0.002)	(0.002)			(0.002)	(0.002)			(0.004)	(0.004)

Area of residence (urban)			-0.032***	-0.022***			0.000	-0.001			-0.002	-0.004			-0.010	-0.006
			(0.008)	(0.008)			(0.004)	(0.004)			(0.003)	(0.003)			(0.007)	(0.007)
Overcrowding			0.017**	0.017**			0.004	0.004			-0.001	-0.001			-0.008	-0.008
			(0.008)	(0.008)			(0.004)	(0.004)			(0.004)	(0.004)			(0.007)	(0.007)
Inappropriate excreta disposal			0.031***	0.027***			0.015***	0.015***			0.002	0.003			-0.006	-0.007
			(0.008)	(0.008)			(0.004)	(0.004)			(0.003)	(0.003)			(0.007)	(0.007)
Safe drinking water			0.001	0.002			-0.001	-0.001			-0.000	-0.000			-0.008	-0.007
			(0.007)	(0.007)			(0.003)	(0.004)			(0.003)	(0.003)			(0.006)	(0.006)
Family income quintile (poorest)																
Second			-0.026**	-0.017*			-0.012**	-0.012**			-0.002	-0.002			0.016*	0.019**
			(0.010)	(0.010)			(0.006)	(0.006)			(0.004)	(0.005)			(0.008)	(0.008)
Middle			-0.053***	-0.045***			-0.006	-0.006			-0.003	-0.003			0.013	0.016*
			(0.010)	(0.011)			(0.006)	(0.006)			(0.005)	(0.005)			(0.009)	(0.009)
Fourth			-0.034***	-0.027**			-0.009	-0.009			-0.010**	-0.010**			-0.007	-0.004
			(0.011)	(0.011)			(0.006)	(0.006)			(0.005)	(0.005)			(0.009)	(0.009)
Highest			-0.036***	-0.029***			-0.014**	-0.014**			-0.006	-0.007			0.015	0.018*
			(0.011)	(0.011)			(0.006)	(0.006)			(0.005)	(0.005)			(0.010)	(0.010)
Cultural origins (mestizo)																
Indigenous				0.064***				-0.001				-0.005				0.026***
				(0.012)				(0.006)				(0.004)				(0.010)
Black				-0.030**				0.020**				0.010				-0.033**
				(0.015)				(0.010)				(0.008)				(0.013)
Montubio				-0.033**				-0.000				-0.006				-0.008
				(0.016)				(0.008)				(0.007)				(0.013)
White				-0.026				0.012				0.004				0.014
				(0.028)				(0.015)				(0.014)				(0.028)
Constant	0.253***	1.757***	1.645***	1.561***	0.057***	0.316***	0.298***	0.309***	0.075***	0.006	0.002	0.012	0.229***	0.025	-0.018	-0.060
	(0.015)	(0.100)	(0.106)	(0.106)	(0.008)	(0.048)	(0.054)	(0.055)	(0.008)	(0.040)	(0.043)	(0.043)	(0.013)	(0.083)	(0.087)	(0.088)
Observations	18,902	18,369	18,060	18,060	18,947	18,410	18,099	18,099	18,268	17,747	17,450	17,450	18,891	18,355	18,046	18,046
R-squared adjusted	0.0102	0.0496	0.0579	0.0605	0.00163	0.00876	0.0108	0.0109	0.00748	0.00708	0.00739	0.00741	0.0103	0.0134	0.0143	0.0150

Notes: Standard errors in parentheses clustered by primary sampling units, *** p<0.01, ** p<0.05, * p<0.1

Table A5. First stage of the IV-2SLS: baseline specification

Variables	Employment status (ENSANUT 2018)	Daily Hours Worked ECV (2014)
Cantonal-level average employment status	0.918***	0.332
	(0.039)	(0.607)
Cantonal-level average daily working hours		0.703***
		(0.079)
Observations	17,193	10,428
Adj. R-squared	0.1960	0.2329
Partial R-squared	0.0420	0.0555
MOP Effective F statistic	514.305	177.513

Notes: Standard errors in parentheses clustered by primary sampling units. These estimations include all the exogenous covariates including both cantonal-level poverty and Gini indexes. Safe drinking water is not available in ECV 2014. MOP stands for the Montiel-Olea & Pflueger (2013) robust first stage statistic. *** p<0.01, ** p<0.05, * p<0.1

Table A6. IV – 2SLS: Impact of mother's employment on underweight, wasting and overweight status

	(1)	(2)	(3)	(4)	(5)	(6)
Variables	ENSANUT 2018			ECV 2014		
	Underweight	Wasting	Overweight	Underweight	Wasting	Overweight
Maternal employment status	-0.001	-0.005	0.024			
	(0.020)	(0.015)	(0.031)			
Maternal working hours				-0.004	-0.003**	-0.001
				(0.003)	(0.002)	(0.003)
Constant	0.302***	0.024	-0.020	0.470***	0.034	-0.001
	(0.057)	(0.045)	(0.091)	(0.084)	(0.045)	(0.083)
Observations	17,230	16,616	17,180	10,466	10,330	10,448

Notes: Standard errors in parentheses clustered by primary sampling units. These estimations include all the exogenous covariates including both cantonal-level poverty and Gini indexes. Safe drinking water is not available in ECV 2014. *** p<0.01, ** p<0.05, * p<0.1