Africa Development, Volume XLIX, No. 1, 2024, pp. 125-142

(ISSN: 0850 3907) https://doi.org/10.57054/ad.v49i1.5618

Does Having more Children Reduce Women's Labour Market Participation? Evidence from Kenya

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Abstract

Using the Kenya Demographic Health Survey 2014, this article estimates the motherhood penalty in Kenya's labour market. To control for endogeneity, the study used mixed-sibling sex preference to explain the exogenous variation in fertility. The results from the auxiliary regression show the existence of 'mixed-sibling' sex preference in Kenya. The probability of having an additional child increases by 24.77 per cent for women whose first two children are the same sex. The instrumental variable model shows that the exogenous variation in fertility afforded by mixed-sibling sex preference significantly reduces the probability of women's labour supply for decent work by 4.29 per cent. The effect of fertility is heterogeneous across age groups. The article finds that labour supply for decent work reduces by 6.08 per cent and 8.29 per cent for women in the age groups 15–24 and 24–34, respectively. Policy incentives such as providing access to affordable childcare services are critical in reducing the motherhood penalty.

Keywords: endogeneity; fertility; Kenya.; labour supply; women

Résumé

À l'aide de l'Enquête démographique et de santé du Kenya de 2014, cet article évalue le désavantage lié à la maternité sur le marché du travail du Kenya. Aux fins de contrôle de l'endogénéité, l'étude a utilisé la préférence pour une mixité frères-sœurs pour expliquer la variation exogène de la fécondité. Les résultats de la régression auxiliaire montrent l'existence d'une préférence pour une mixité « frères et sœurs » au Kenya. La probabilité d'avoir un enfant supplémentaire augmente de 24,77 pour cent pour les femmes dont les deux premiers enfants sont du même sexe. Le modèle à variables instrumentales

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montre que la variation exogène de la fécondité induite par la préférence pour une mixité frères-sœurs réduit considérablement la probabilité pour les femmes d'accéder à un travail décent (4,29 pour cent). L'effet de la fécondité est hétérogène selon les groupes d'âge. Le document révèle que l'offre de main-d'œuvre pour un travail décent baisse respectivement de 6,08 pour cent et de 8,29 pour cent pour les femmes des tranches d'âge de 15 à 24 ans et de 24 à 34 ans. Les incitations politiques telles que l'accès à des services abordables de garde d'enfants sont essentielles dans la réduction des désavantages liées à la maternité.

Mots-clés : endogénéité ; fertilité ; Kenya ; offre de main-d'œuvre ; femmes

Introduction

Recently, much progress has been made in recognising the rights of women. However, inequalities are still prevalent particularly in the labour market. The International Labour Organization (ILO) statistics for 2019 show that the labour market global gender gap is 27 per cent (ILO 2019) and the wage gender gap is 20 per cent (ILO 2018). Studies that have investigated the labour-market gender differentials have presented contrasting results. Some argue that labour-market gender differentials can be explained by human-capital differentials, in that women who are well-educated have equal chances of entering the labour market (Siphambe and Thokweng-Bakwena 2001; Nordman and Wolff 2009; Totouom, Mboutchouang and Kaffo Fotio 2018). On the other hand, some studies show that despite human capital improving labour-market outcomes, it does not eliminate gender differentials (Nicita and Razzaz 2003; Nordman, Robilliard and Roubaud 2011; Kuépié 2016; Mulwa and Gichana 2020). Other studies claim that labour-market gender differentials emanate from the fact that women disproportionately bear a great burden in household production associated with childbearing and care responsibilities (Waldfogel 1998; Angrist and Evans 1998; Agüero and Marks 2008).

In the context of sub-Saharan Africa, where the total fertility rate is approximately five children per woman (Westoff, Bietsch and Koffman 2013), fertility can generate penalties, especially in the labour market. From a theoretical perspective, Becker (1991) argued that the procreation role of women can affect women's labour productivity and supply. Gronau (1977) argued that labour supply is a choice variable that involves choosing between leisure, unpaid care, and market work. The presence of children, the price of market substitutes for childcare, and potential wage, among other factors, determine the trade-off for time allocation between the three uses of time. Thus, an increase in the number of children implies increased time for childcare and reduced time for market work and leisure. Alternatively, if the price of substitute for childcare is affordable, then the decision for market work increases and that of household production decreases.

Empirically, studies that have analysed fertility and labour supply show that fertility significantly reduces women's labour-market participation (Angrist and Evans 1998; Jacobsen, Pearce and Rosenbloom 1999; Chun and Oh 2002; Cruces and Galiani 2007; He and Zhu 2016; Kuépié 2016; Goldin 2014; Adda, Dustmann and Stevens 2017). However, the effect of fertility goes beyond labour-market participation. Women with children earn less, work for less (paid) hours and work in less prestigious jobs (Browning 1992; Waldfogel 1998). Similarly, studies that have estimated the effect of childbearing on men find that men's labour-market participation increases with the number of children (Angrist and Evans 1998; Kuépié 2016). Therefore, households respond to an increased number of children by either reducing the mother's labour supply or by purchasing childcare services in the market.

A key policy question is, to what extent do more children generate penalties for women's labour supply? Even though the question is relevant within the policy space, its answer is not straightforward because fertility is plausibly endogenous in the labour supply equation. Children are a choice variable that might be affected by a woman's labour-market participation status (Agüero and Marks 2008). Reverse causality is plausible, where women's participation in the labour market affects the fertility decision. Similarly, there are likely to be omitted variables that are unobservable, such as career ambitions, that affect both fertility and labour-supply decisions (Rosenzweig and Wolpin 1980; Agüero and Marks 2008). For instance, women in high-earning careers might decide to have fewer or no children due to unobserved factors brought about by the nature of their jobs. Failure to account for this endogeneity might result in biased estimates.

Most studies use the instrumental variable (IV) technique to deal with the potential endogeneity bias. However, the challenge arises in finding a valid IV that is strongly correlated with fertility and is plausibly exogenous in the labour supply equation. Some studies have used twins as IV for fertility (Rosenzweig and Wolpin 1980; Jacobsen *et al.* 1999; He and Zhu 2016), whereas others construct an IV from the sex composition of children (Angrist and Evans 1998; Cruces and Galiani 2007). Other exogenous variations of family size are son preference (Chun and Oh 2002), abortion legislation (Bloom *et al.* 2009) and infertility shocks (Agüero and Marks 2008).

Rosenzweig and Wolpin (1980) argue that the exogenous variation of fertility is affected by twins because if families have an unexpected child, this makes the control and treatment groups randomly assigned. However, the drawback in this identification is the scarcity of twins at first birth (Nakamura and Nakamura 1992) and that twins might not be entirely random, with multiple births being common among women undergoing fertility treatment (He and Zhu 2016). For sex composition of children, Angrist and Evans (1998) argue that parents prefer a mixed-sex composition as opposed to children of the same sex. Parents whose first children are of the same sex have a higher probability of conceiving an additional child, and given that the sex of a child is randomly assigned, it becomes a good IV for fertility for those with at least two children. For the son preference IV, Chun and Oh (2002) provide evidence in the Korean context that originates from differentials in the labour-market performance between women and men. Families who have a daughter at first birth are likely to have another child until they conceive a son, and given that the sex of a child is randomly assigned, this becomes a good IV for fertility. For abortion legislation, Bloom *et al.* (2009) argues that even though the legislation to do with abortion reflects broader societal trends that might correlate with women's labour supply, the timing of such legislation is random and exogenous. Agüero and Marks (2008) propose that alternative exogenous variation in fertility is generated by infertility shocks. Infertility affects childbearing and, besides increasing with age, is almost randomly assigned. Regardless of the approach used to determine the exogenous variation of family size, virtually all studies show that women's labour-market participation decreases with the number of children.

Background Information on Kenya

In Kenya, there are gender inequalities in access to opportunities especially in the labour market. The gender gap stands at 5 per cent, with women's labour force participation rate at 72.1 per cent compared to 77.3 per cent for men (Baumann, 2021) and more women working part-time than men, accounting for 36.2 per cent compared to 21.1 per cent of men (KNBS 2018). Further, women are underrepresented in wage employment. In 2021 they made up only 39.3 per cent compared to men (60.7 per cent), which is a slight improvement from 2019, when women accounted for 35.5 per cent of the wage employment (KNBS 2022). In addition, women earn less than men: the wage gap is estimated at 68 per cent (WEF 2019). Women find it hard to secure their first employment and oftentimes are overrepresented in the agricultural and informal sectors (Mulwa and Gichana 2020). Women account for 68 per cent of those in vulnerable employment compared to 39 per cent of men (KNBS 2020b). In addition, women total 70 per cent of the workers in the agricultural sector, which is categorised by low earnings and productivity (KNBS 2019). For those women employed in the agricultural sector, 43 per cent are not paid wages and only 9 per cent receive in-kind payments (KNBS 2014). The low representation of women in the labour market could be due to the traditional role assigned to women. In 2019, women spent an average of 11.1 hours on unpaid work compared to 2.9 hours by men (KNBS 2019). The implication of this is that women have less command of economic resources. The female Gross National Income (GNI) is estimated at KES 3,666 compared to KES 4,829 for men (Baumann 2021).

Women's participation in the labour market in Kenya could potentially be linked to fertility rates. In Kenya, total fertility rates have decreased sharply from 7.2 births per woman in 1989 to 3.9 in 2014 (DHS 2014). With decreased fertility, women's labour force participation has increased by 5 per cent from 1989 to 2015 (KNBS 2018). With the fertility rate at 3.9 births per woman, this implies that a woman is expected to have at least four children in her lifetime (KNBS 2014). According to the Demographic Health Survey (DHS) report (2014), the highest fertility rate is among women between the ages of 25 and 34 (331 per 1000) and 15 to 24 (302 per 1000). At the same time, women's labour-force participation is lowest among the age group 15 to 19, at 18.8 per cent, and 20 to 24, at 53.1 per cent. This compares with men's labour force participation at 34.7 per cent for the age group 15–19 and 75.8 per cent for the age group 20–24, and over 70 per cent for the age group 25–49 in 2014 (KNBS 2014). On average, by her late twenties a woman in Kenya has at least two children and in her thirties she has more than four children. The high fertility rate of women in these age brackets could be as a result of high marriage rates. Marriage, which is linked to childbearing, starts early in Kenya. DHS (2014) shows that 29 per cent and 48 per cent of women were married by age 18 and 20, respectively. On average, women who marry at an early age are more likely to have their first child at a young age and have more children in their lifetime. According to DHS (2014), men marry relatively late compared with women – on average, the mean age of marriage for men is 25.3 years. Childbearing linked to marriage imposes an additional burden on women due to increased time spent on unpaid care, which might cause a delay in labour-market entry, withdrawal from the labour market or suboptimal labour participation.

Despite the decline in fertility coinciding with women's increased labour-force participation, there is a scarcity of household studies that estimate the relationship between fertility and women's labour-market participation in Kenya. The existing studies on labour-market outcomes in Kenya (Kabubo-Mariara 2003; Wambugu 2011; Mulwa and Gichana

2020) have not specifically estimated the childbearing burden within the Kenyan labour market. Kabubo-Mariara (2003) and Mulwa and Gichana (2020) present evidence of a gender penalty in the Kenyan labour market, with women highly discriminated against. Therefore, the purpose of this article is to estimate the causal effect of fertility on women's labour supply and offer policy insights that could help reduce the motherhood penalty.

The novelty of the study is that it investigates whether the gender penalty witnessed within the Kenyan labour market could be explained by the fertility burden imposed on women. Second, instead of focusing on fertility and overall labour market participation, the article pays particular attention to the effect of fertility on decent jobs. Third, unlike previous studies that present evidence of the effect of the fertility burden on labour supply from developed countries, this article provides evidence from a developing country. Finally, the article contributes to the causal discussion of childbearing and women's labour supply (Angrist and Evans 1998; Chun and Oh 2002; Cruces and Galiani 2007; Xiaobo and Rong 2015). Cruces and Galiani (2007) have argued that the 'mixed-sibling sex preference' IV can be generalised in developing countries context, quantitatively and qualitatively. This article explores to what extent the identification strategy proposed by Angrist and Evans (1998) can be applied in the developing country context. The identification of the instrument is drawn from parents' preference for the mixed-sex composition of their children rather than a same-sex composition.

The rest of this article is arranged in the following order. The next sections give the data and summary statistics, explain the estimation strategy used and discuss the results before inferring conclusions and policy implications.

Data and Summary of Statistics

The article used cross-sectional national representative data obtained from the DHS, which was conducted in 2014. The dataset contains information for 31,079 women. The data captures information about fertility, sociodemographic factors and labour-market participation for women aged 15 to 49. Fertility is measured by the total number of children who had not reached 18 years at the time of the survey. Although the main concern of the DHS is not labour-market conditions, this article makes use of information on groups of occupation. The survey groups occupations in seven categories: professional/technical/managerial; clerical; services; skilled workers; self-employed in agriculture; household and domestic workers; and low-skilled workers. The article uses the first four groups to plausibly infer 'decent jobs in the wage sector', which is an indicator of integration in the labour market. For individuals who are self-employed in agriculture it is difficult to distinguish the factors of a decent job; therefore, they are not included in the decent job category. For this reason, it is plausible that the indicator of decent job used in this article underestimates the population who are in the decent job category. It is important to note that this misspecification might lead to an attenuation bias in the estimation of the fertility burden in the labour market. Therefore, the results need to be interpreted as the lower bound of the actual estimate.

The article restricts itself to married women, given that they have the highest fertility rate compared to those who are unmarried (DHS 2014). Due to this modification on the initial dataset, 15,126 observations were used to estimate the effect of fertility on women's labour supply. The summarised statistics from the sample dataset are presented in Table 1. On average, 8.9 per cent of married women are in decent jobs. The average number of children is 3.2, with the average age of the mother at first birth being 19. The average years of schooling are 9.1. On average, 84 per cent of the women are Christian. Given that fertility is endogenous in the labour-force equation, the same sex variable is used as the IV. On average, 41 per cent of women have children of the same sex, 21 per cent of women have the first two children as boys and 19 per cent of women have the first two children as girls.

Table 1: Summarised statistics

Variable	**Measurement**	**Mean**	**Standard Deviation**
Labour supply	1 if working in high-end job, 0 otherwise	0.0892	0.2851
Children	The number of children	3.2970	1.7473
Age at first birth	Years	19.5902	3.5750
Age	Age in years	31.3353	7.6818
Education	Years of schooling	9.1030	4.8275
Husband's occupation	Husband's occupation in the labour market	0.4647	0.4988
Christian	1 if Christian, 0 otherwise	0.8462	0.3607
IV			
Same sex	1 if the first two children are of the same sex, otherwise 0	0.4148	0.4927
Two_boys	1 if the first two children are boys, 0 otherwise	0.2179	0.4128
Two_girls	1 if the first two children are girls, 0 otherwise	0.1969	0.3977
Observations		15,126	

Empirical Strategy

Empirical model

The study uses the two-stage least square model (2SLS) to control for endogeneity of fertility in the labour-market participation decision. In the first stage, fertility is regressed against the IV and the control variables to predict the number of children.

The fertility decision equation is given as:

$$X_i = \partial V_i + \gamma S_i + \mu_i \quad (1)$$

where V_i is a vector of sociodemographic factors; S_i is an indicator of the sex composition, which equals 1 if the first two children are the same sex, otherwise zero.

In the second stage, the labour-market participation equation is expressed as follows:

$$Y_i = aV_i + \beta Xi + \varepsilon_i \quad (2)$$

where *Yi* is a binary variable for women's labour-market participation, which equals 1 if working (decent job), otherwise zero; *Xi* represents the number of predicted children obtained in the first stage analysis. Given equation 1, that is, the fertility equation has an additional exogenous variable S_i, the system of equations is exactly identified.

Strength of the instrument

The auxiliary (first stage) regression shows that the *same-sex* variable is a relevant and strong instrument for fertility (see Table 2). The F statistic for same sex is 155.69 and is statistically significant at 1 per cent and the F statistic for two boys and two girls is 78.09 and is statistically significant at 1 per cent.

The results show that the probability of having an additional child increases by 24.77 per cent for women whose first set of children are of the same sex, by 25.87 per cent for those who have two girls and by 23.78 per cent for those who have two boys, all at a 1 per cent significance level. The 2.1 percentage point difference might be due to a cultural preference for boys but there is no substantial evidence of son preference. Cruces and Galiani (2007) in Argentina and Mexico, and Angrist and Evans (1998) in United States, document a similar cultural preference for boys. Generally, the results indicate a mixed-sex composition preference in Kenya with a slight bias for boys. The mixed-sex preference exists in the United States (Angrist and Evans 1998) and in Argentina and Mexico (Cruces and Galiani 2007).

Table 2: Auxiliary regression

Independent Variables	**Children**	
Age (years)	0.4752*** (0.0101)	0.4751*** (0.0101)
Age at first birth	-0.1853*** (0.0030)	-0.1853*** (0.0030)
Age*age	-0.0049*** (0.0002)	-0.0049*** (0.0002)
Education	-0.0811*** (0.0023)	-0.0811*** (0.0023)
Husband's occupation	0.0114 (0.0191)	0.0114 (0.0192)
Christian	-0.1667*** (0.0301)	-0.1668*** (0.0301)
Same sex	0.2477*** (0.0199)	
Two boys		0.2378*** (0.0243)
Two girls		0.2587*** (0.0253)
Constant	-2.0433*** (0.1607)	-2.0423*** (0.1607)
Note: Parentheses indicate standard errors; ***, ** and * indicate significance at 1%, 5% and 10% respectively		

Exclusion restriction

After establishing 'mixed-sex' preference, the next step is to determine the validity of the *same sex* variable in instrumenting for the number of children within the Kenyan context. This identification strategy is drawn from parents' preference for a mixed-sex composition of children rather than a same-sex composition. Arguably, the sex composition is randomly assigned and does not affect women's labour supply directly, which provides a plausible instrument to explain for exogenous variation in fertility. Although the exclusion restriction is not directly testable, its plausible exogeneity must be discussed (Conley, Hansen and Rossi 2012). In this article's context, the possible treats to exogeneity of the *same sex* instrument are discussed.

First, it could be argued that within the developing country context, and particularly in Kenya, families would have a greater preference for sons, which could potentially alter children's sex composition, through selective abortion (or stopping rules), hence threatening the validity of the IV. However, the data shows the contrary, that the sex ratio of infants (of boys to girls aged 0–1) is almost identical to the biological ratio, which is 1.02 and 1.01, respectively (DHS 2014), and therefore rules out the possibility of son preference. Moreover, the infant sex ratio shows no evidence of a son preference effect on the mortality rates for girls. Actually, the DHS (2014) shows that male children have a higher mortality rate than female children (44 deaths vs. 37 deaths per 1000).

In addition, son preference might emanate from expected future contribution, where sons are expected to support their families; when daughters are married, they have to take care of their own families and their husband (Basu and Gupta, 2001). If this form of son preference existed then it would imply that families invested more in sons in terms of human-capital accumulation as it is expected that they would earn more in the future and therefore contribute more to the parents' wellbeing. However, the data indicates that this seems not to be the case. For instance, the primary school enrolment for boys and girls is virtually the same, at 50.69 per cent and 49.30 per cent, respectively (KNBS 2022). Secondary school enrolment is slightly higher for girls than boys (50.12 per cent and 49.88 per cent, respectively). Enrolment rates in tertiary education for girls almost match those of boys. This evidence rules out the concern that due to son preference parents would invest in the human capital of sons more than daughters.

Another concern about the validity of the *same-sex* variable relates to the 'hand-me-down' effect identified by Rosenzweig and Wolpin (2000). In their study, they show that households with *same-sex* siblings in India have a lower per-child expenditure than those with mixed-sex siblings. They argue that this is because of the 'hand-me-down' savings on clothing and footwear items that are likely to be witnessed in households with *same-sex* siblings. They argue that given the significant share of these items to the total household expenditure, the sex composition might alter labour supply through other channels other than through fertility. Therefore, if sex composition directly affects labour supply, then the validity of the *same-sex* variable is threatened.

Although data on per-child expenditure in Kenya is not available, the data shows that it is implausible that sex composition has a significant effect on total expenditure. In Rosenzweig and Wolpin's (2000) study they found that the share of expenditure on clothing and footwear for children under 18 years was 11 per cent. According to the Kenya Statistical Abstract, the share of clothing and footwear is 3 per cent for all household members (KNBS 2020a). This share is significantly lower because it represents the clothing and footwear share of expenditure for all members of the household, not only for children under 18 years. Further, Rosenzweig and Wolpin (2000) estimate the 'hand-me-down' savings to be 1.3 per cent of total income. Assuming that these savings existed in the Kenyan context, their share would be too small to account for any significant effect on the mother's labour-supply decision.

Thus, the evidence suggests that son preference in Kenya is almost non-existent and the evidence of a bias towards boys presented in the first-stage regression is mainly due to cultural preference and not strong evidence of discrimination against girls. Sex composition is unlikely to affect the consumption pattern especially for items such as clothing and footwear. Therefore, the *same-sex* variable is plausibly exogenous and is a valid instrument for fertility in the labour-supply equation.

Results

Main results

Table 3 presents the main results of the study while controlling for demographic factors such as age, age at first birth, years of schooling, husband's occupation and religion. In the first column, we present the OLS estimates; the second column presents the IV model with the *same-sex* variable; the third column presents the results when *two boys* and *two girls* variables are used as instruments for the number of children.

The OLS model estimates show that for every additional child a woman's labour market participation decreases by 0.87 per cent at a 1 per cent significance level. The OLS estimates are not reliable given the endogeneity bias that arises when estimating fertility and labour supply (Chun and Oh 2002). The IV model is estimated to correct for the potential endogeneity.

After correcting for endogeneity using the *same-sex* variable as the IV, the results show that the OLS underestimates the childbearing effect on women's labour-market participation. The difference is due to the exogenous variation in family size explained by the sex composition. The results show that women's labour supply (that is, in decent jobs) decreases by 4.29 per cent for every additional child at a 1 per cent significance level. The result is consistent with the results from developed countries (Angrist and Evans 1998; Jacobsen *et al.* 1999; Goldin 2014; Adda *et al.* 2017) and developing countries (Chun and Oh 2002; Cruces and Galiani 2007; Agüero and Marks 2008; He and Zhu 2016; Kuépié 2016) that show a similar negative effect of childbearing on a mother's labour supply.

Table 3: Main results

Independent Variables	**Ols (I)**	**IV Model (ii)**	**(ii)**
Age (years)	0.0107**** (0.0024)	0.0274*** (0.0182)	0.0271*** (0.0092)
Age*age	-0.0001**** (0.0000)	-0.0003*** (0.0001)	-0.0003*** (0.0001)
Age at first birth	0.0011 (0.0007)	-0.0053 (0.0035)	-0.0052 (0.0035)
Education	0.0110*** (0.0005)	0.0081*** (0.0015)	0.0082*** (0.0016)
Husband's occupation	0.1854*** (0.0043)	0.1858*** (0.0043)	0.1858*** (0.0043)
Christian	-0.0298*** (0.0068)	-0.0355*** (0.0075)	-0.0354*** (0.0075)
Number of children	-0.0087*** (0.0018)		
Same-sex		-0.0429*** (0.0182)	
Two boys and two girls			-0.0423
Observations F statistic	15,126 373.79***	15,126 2440.87***	15,126 2542.79
Overidentification test Sargan-test			X^2 =0.3516 (p=0.5532)
Basmann test			X^2 =0.3514 (p=0.5533)
Note: Parentheses indicate the standard errors; ***, ** and * are significant at 1%, 5% and 10% respectively			

The third column decomposes the *same sex* IV into the specific sex of the children. The results show that each additional child decreases women's labour supply (in decent jobs) by 4.23 per cent. Angrist and Evans (1998) argue that bias from any secular effects of a child's gender on labour supply would be different for the two IVs, whereas the effect of children on the labour supply seems independent of whether the *same-sex* variable equals that of *two boys* or *two girls.* Given that the endogenous variable is one with two IVs, an overidentification test must be performed. The Sargan-Basmann test rejects the null hypothesis at a 10 per cent significance level that the *two boys* and *two girls* IVs are overidentified in the labour-supply equation, and therefore they are valid instruments.

To ensure the robustness of the results, the article accounts for the heterogeneity of age groups in the childbearing effect on women's labour supply. The article considers the following three age groups: 15–24; 25–35; and 35–49.

Table 4: Robustness analysis: checking for heterogeneity across age groups

Independent Variables	**Sub-Sample Age Group 15–24 (i)**	**(ii)**	**Sub-Sample Age Group 25–34 (iii)**	**(iv)**	**Sub-Sample Age Group 35–49 (v)**	**(vi)**
Age	0.0134 (0.0421)	0.0134 (0.0421)	0.1026*** (0.0337)	0.1004*** (0.0335)	0.0116 (0.0260)	0.0116 (0.0259)
Age*age	0.0002 (0.0010)	0.0002 (0.0010)	-0.0015*** (0.0005)	-0.0014*** (0.0005)	-0.0001 (0.0003)	-0.0001 (0.0003)
Age at first birth	-0.0183*** (0.0061)	-0.0183*** (0.0061)	-0.0152* (0.0081)	-0.0143* (0.0080)	0.0024 (0.0055)	0.0024 (0.0054)
Education	0.0057*** (0.0013)	0.0057*** (0.0013)	0.0054** (0.0028)	0.0057** (0.0028)	0.0139*** (0.0042)	0.0139*** (0.0042)
Husband's occupation	0.1283*** (0.0085)	0.1282*** (0.0085)	0.2040*** (0.0067)	0.2039*** (0.0067)	0.1937*** (0.0076)	0.1937*** (0.0076)
Christian	-0.0109 (0.0125)	-0.0108 (0.0125)	-0.0378*** (0.0130)	-0.0370*** (0.0129)	-0.0626*** (0.0141)	-0.0626*** (0.0140)
Number of children						
Same sex	-0.0608*** (0.0227)		-0.0829*** (0.0349)		-0.00006 (0.0038)	
Two boys and two girls		-0.0607*** (0.0226)		-0.0790** (0.0345)		-0.00007 (0.0385)

Note: Parentheses indicate the standard errors;
***, ** and * are significant at 1%, 5% and 10% respectively

The results from Table 4 show that after accounting for heterogeneity across age groups, the effect of childbearing is higher for women in the age group 25–34 compared with women in the age group 15–24 and those in age group 35–49. When considering the same sex instrument, for women aged 15 to 24 the results show that every additional child decreases women's labour supply by 6.08 per cent at a 1 per cent significance level. For women aged 25–34, an additional child decreases labour supply by 8.29 per cent at a 1 per cent significance level. When the same sex instrument is decomposed into two boys and two girls, the effect of childbearing is still higher for women aged 25–34. For women aged 15–24, for every additional child, labour supply decreases by 6.07 per cent at a 1 per cent significance level, whereas for women aged 25–34, for every additional child, labour supply decreases by 7.9 per cent at a 5 per cent significance level. The results show that the effect of childbearing for women in the age group 35–49 is non-existent and statistically insignificant. This result is in line with the fact that childbearing is a positive function of age (see Table 2) and the labour-supply effect increases with age.

Discussions and Policy Implications

Discussions

The article analysed the causal effect of fertility on women's labour supply in Kenya. With fertility being endogenous in the labour-supply equation, the article corrects for this endogeneity using IV. The article uses 'mixed-sex' preference to draw an exogenous variation in fertility. The identification of the instrument is drawn from parents' preference for mixed-sex composition rather than *same-sex* composition. Arguably, the sex composition of the first two children is randomly assigned, which provides a plausible instrument to explain variation in family size and labour-supply decision.

The results from auxiliary regression show the existence of 'mixed-sex' preference for families in Kenya, with women more likely to have additional children if the first two children are of the same sex. The IV results show that women's labour-market participation for decent jobs decreases by 4.29 per cent for every additional child at a 1 per cent significance level. For women in the age bracket 15–24 years, labour-market participation decreases by 6.08 per cent, and for women aged 25 to 34 years it decreases by 8.29 per cent. One possible explanation for this result is the high fertility rate – the DHS (2014) shows that the highest fertility is among women between the ages of 25 and 34 (331 per 1000) and 15 to 24 (302 per 1000). On average, by her late twenties a woman in

Kenya has had at least two children and, in her thirties, she has had more than four. Another possible reason could be that women in this age group (15–24) are mostly in school, whereas those aged 25 to 34 are mostly starting their career and motherhood imposes significant costs on their career trajectory. For women in the age group 35–49, the results show that they do not significantly face the motherhood penalty in the labour market. It is plausible that the effect on women in the age group 35–49 is non-existent given that most women give birth at very early ages and the effect of childbearing disappears as they get older. A further possible explanation for this result is that women between the age of 35 and 49 years face extra costs in raising children, such as education costs, which might result in an increased labour-market participation. Rosenzweig and Wolpin (1980) found a similar result when using twins to instrument for exogenous variation of family size – that women within the ages of 35 and 44 years do not experience a motherhood penalty. The results show that the duration of the effect of childbearing is incurred during the ages of 15 to 24 and 24 to 34, then disappears in ages 35 to 49 years. This implies that the motherhood penalty is not necessarily permanent but rather occurs for specific age groups and disappears once women reach the age of 35 to 49 years.

In general, these results imply that an increased number of children reduces women's time for paid work and increases the time they allocate for childcare. This could be taken to mean that much of the gender differentials witnessed in Kenya (Kabubo-Mariara 2003; Mulwa and Gichana 2020) could be explained by fertility. The huge penalty witnessed could be because market work is substitutable by time for household production (Willis 1987). An additional child implies an increased marginal value of time in the household as a result of childcare. Within the Kenyan context, women are primary caregivers and are responsible for spending a large portion of their time on household production and unpaid care. Therefore, a mother's time for childcare increases with the number of children, causing a withdrawal from the labour market.

With an underdeveloped market for childcare, the high cost of formal childcare services in Kenya (Lokshin, Glinskaya and Garcia 2004) and the lack of gender-transformative policies that facilitate women's paid work and motherhood, women who have children are likely to drop out of the labour market to take care of their children. These institutional constraints signal to women that it might be difficult to combine motherhood and market work, and thereby lead to a delay in their entry into the labour market, their withdrawal from the labour market or suboptimal labour market participation. The availability of a well-developed market for childcare services plays a critical role in labour supply, and if the market price for these services is lower than the wage rate for women, then mothers would choose to participate in rather than delay or withdraw from the labour market. Finally, although fertility imposes a penalty on mothers, the reduction in labour-market participation might produce some positive externalities. As mothers devote more time to childcare services than market activities, this might make some children better off (Blau and Grossberg 1992).

Policy Implications

This article shows that childbearing reduces women's labour market participation especially in decent jobs, which could have implications on household income, gender inequalities in the labour market and reduced investment in human capital for children in those households, leading to poverty traps and increased economic inequalities. From a policy point of view, policy incentives that encourage women's labour-market participation are critical to help reduce the motherhood penalty. First, in the Kenyan context, it would be important for government to consider formalising childcare services and providing fiscal incentives to make childcare services more accessible and affordable. This would enable women to create time for market work, especially decent work, and reduce the time allocated for childcare, which in turn could help to reduce gender inequalities in the labour market and social exclusion more broadly. Also, beyond increasing the likelihood of women returning to work after maternity leave, it would act as a signal for future workers – particularly those considering having children. Second, working collaboratively with the private sector, government should consider introducing a policy directive that would ensure that all workers not only have the right to request flexible working arrangements but also that there is an enabling environment for flexible working to be adopted to allow mothers to care for their children without hindering their career progression. Such policies could play a significant role in reducing women's labour-supply withdrawal and delay in entry into the labour market.

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