**DOCUMENTOS DE TRABAJO**
ISSN 2174-4912

**SON PREFERENCE, GENDER DISCRIMINATION AND MISSING GIRLS IN RURAL SPAIN, 1750-1950**

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DT-AEHE Nº 2007
www.aehe.net



July 2020

DOCUMENTOS DE TRABAJO
ISSN 2174-4912

**PREFERENCIA POR EL VARÓN, DISCRIMINACIÓN DE GÉNERO Y NIÑAS DESAPARECIDAS EN LA ESPAÑA RURAL, 1750-1950**

**Francisco J. Marco-Gracia** [‡] **y Francisco J. Beltrán Tapia**[§]

DT- 2007, July 2020
JEL: I14, I15, J13, J16, N33

## RESUMEN

Utilizando microdatos longitudinales pertenecientes a una zona rural española entre 1750 y 1950 (casi 35,000 trayectorias vitales), este artículo muestra que conductas discriminatorias influyeron en las tasas de mortalidad infantil por sexo. Aunque es muy probable que las familias discriminaran a las niñas durante el primer año de vida, el exceso de mortalidad femenina es especialmente visible en el grupo de edad comprendido entre 1-5 años. En este sentido, mientras la lactancia parece haber mitigado temporalmente los efectos de la discriminación de género, las tasas de mortalidad por sexo se comportaron de forma marcadamente distinta después del destete. Estas familias por tanto priorizaron a los niños durante la infancia en términos de alimentos o cuidados con el objetivo de incrementar sus posibilidades de supervivencia.

**Palabras clave:** Mortalidad infantil, Discriminación de género, Exceso de mortalidad femenina, Salud.

## ABSTRACT

Relying on longitudinal micro data from a Spanish rural region between 1750 and 1950 (almost 35,000 life courses), this article evidences that discriminatory practices affected sex-specific mortality during infancy and childhood. Although it is likely that families also discriminated girls during the first year of life, the female excess mortality was especially visible in the 1-5 age-group. In this regard, while breastfeeding seemed to have temporarily mitigated the effects of gender discrimination, sex-specific mortality rates behaved markedly different once children were weaned. Parents therefore prioritised boys during infancy and childhood in the allocation of food and/or care in order to enhance their survival chances.

**Keywords:** Infant and child mortality, Gender discrimination, Female excess mortality, Health.

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# SON PREFERENCE, GENDER DISCRIMINATION AND MISSING GIRLS IN RURAL SPAIN, 1750-1950[1]

## 1.- Introduction

Son preference is a common feature of traditional societies where girls are considered of lesser value than boys (Williamson 1976; Sen 1990; Das Gupta et al. 2003). Less equal gender roles tend to arise from particular beliefs and values that expect women to be in charge of domestic tasks and therefore discourage their participation in the labour market (Boserup 1989; Alesina et al. 2013; Giuliano 2015, 2018). Property and inheritance rules usually favoured males who would then take over the family farm, provide parents with old-age security and ensure the continuity of the family name (Knodel and De Vos 1980). Moreover, although daughters could help with younger siblings and take care of their parents in old age (Sandström and Vikström 2015, 58; Lynch 2011, 258-260), strict dowry systems could make them a burden for the family resources (Bhalotra et al. 2018).

Gender discrimination against girls, resulting in female excess mortality, have long been practiced in societies characterised by strong patriarchal traditions favouring males (Sen 1990; Das Gupta 2003; Bhaskar and Gupta 2007; Drixler 2012; Gupta 2014). Although European women historically enjoyed a considerable better status than their counterparts in other regions, they were nonetheless discriminated in many dimensions, especially in Southern and Eastern Europe (Szoltysek et al. 2017; Kok 2017; Carmichael and Rijpma 2017; Dilli et al. 2019). There is, however, little evidence that gender discriminatory practices increased female mortality in infancy and childhood in historical Europe. According to Lynch (2011), household formation patterns, as well as cultural and religious values, limited excess female mortality and missing girls in historical Europe (also in Derosas and Tsuya 2010).

Several studies have though challenged this view and suggested that gender discrimination was more important than previously assumed. On the one hand, son preference influenced the propensity to have additional children in Sweden, Germany, Italy and Spain (Kolk 2011; Manfredini et al. 2016, Sandström and Vikström 2015, Reher and Sandström 2015, Marco-Gracia, 2020)[2]. In this regard, families with only female offspring were more likely to continue childbearing. In the Italian case, this was

[1] This research received financial support from the Research Council of Norway (Project 301527) and the Spanish Ministry of Science and Education (Projects HAR2015-64076-P ECO2015-65582).

[2] See also Bohnert et al. (2012) for the US.

more visible among those households that depended on sons to take over the family farm. On the one hand, there is scattered evidence suggesting that families resorted to female infanticide as a means of controlling the size and sex composition of their offspring (Bechtold 2001; Hynes 2011; Hanlon 2016; Beltrán Tapia and Raftakis 2019; Beltrán Tapia and Marco-Gracia 2020)[3]. Lastly, it appears that gender-discriminatory practices were affecting girls' net nutritional status and unduly increasing female mortality rates during infancy and childhood via an unequal allocation of food, care and/or workload (Tabutin 1978; Johansson 1984; Pinnelli and Mancini 1997; Baten and Murray 2000; McNay et al. 2005; Horrell and Oxley 2016; Beltrán Tapia and Gallego-Martínez 2017; 2020).

Unveiling patterns of gender discrimination in infancy and childhood is, however, especially challenging. For biological reasons, males are more vulnerable, and their mortality rates are naturally higher, especially during the first year of life. This frailty was especially visible in the high-mortality environments that characterised pre-industrial Europe due to poor living conditions, lack of hygiene and the absence of public health systems (Beltrán Tapia 2019). High-quality data is therefore needed to distinguish between mortality risks resulting from intra-household discrimination and those arising from other factors (Kok 2017; 44).

Relying on longitudinal micro data from a rural region in North-eastern Spain between 1750 and 1950 (almost 35,000 individuals), this article evidences that discriminatory practices affected sex-specific mortality rates during infancy and childhood. In this regard, although it is likely that gender discrimination also increased female mortality during the first year of life, the gender mortality gap especially widened during the 1-5 age-group. While breastfeeding seemed to have temporarily mitigated the effects of gender discrimination (probably because it is a non-competitive resource), sex-specific mortality rates behaved significantly different once the children were weaned. The female penalty is even more visible in children born at high parities, when additional children further constrained the limited household resources. Moreover, having no male siblings is related to lower male mortality rates in early childhood. It appears that parents prioritised boys in the allocation of food and/or care in order to enhance their survival chances and secure at least one male heir. Girls' mortality rates remained indeed significantly higher than those of boys until age 5, a feature that clashes with the female biological advantage.

This article supports previous studies that challenged the idea that there were no missing girls in historical Europe (Beltrán Tapia and Gallego-Martínez 2017; Beltrán Tapia 2019). In this regard, it shows that the high sex ratios in infancy and childhood found in previous studies are not driven by problems with the quality of the registers, but by female excess mortality. In addition, the micro-data used here sheds light on how these societies ended up showing unbalanced number of boys and girls by illustrating

[3] Interestingly, Hynes (2011) and Hanlon (2016) consider that families could also target boys depending the circumstances.

not only when these discriminatory practices were taking place, but also the mechanisms at play. Discriminatory practices during childhood seem to have been part of a generalised cultural system that privileged boys in terms of access to food and/or care that unduly increased female mortality rates. These practices were shared both by farmers and landless labourers and proved to be quite persistent. Although definitely more important during the 19$^{th}$ century, discriminatory patterns affecting sex-specific mortality rates were still visible during the first decades of the 20$^{th}$ century, thus evidencing that son preference continued to be a strong cultural norm within these societies.

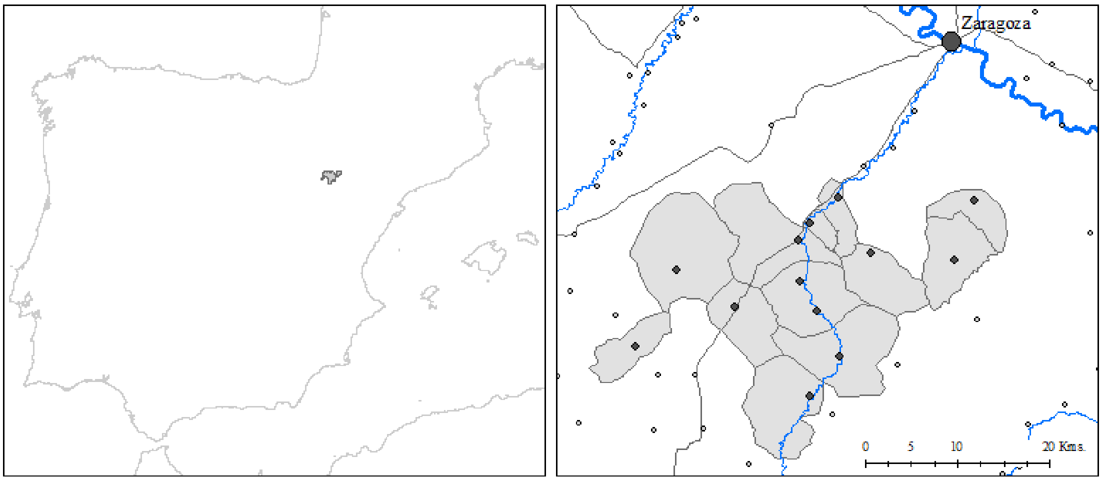
Our findings are especially relevant because they take place in an area where nuclear households prevailed and inheritances were equally distributed among all children, regardless of their sex. Likewise, women maintained full control of the resources they brought to the marriage and were entitled to freely dispose of their patrimony through wills at their death (Jarque Martínez and Salas Ausens 2007, 126-127). Moreover, although it is true that women did not enjoy the same status in the labour market than men (lower salaries, less workdays, etc.), female waged labour was widespread and their contribution was crucial to sustain the household economy (Borderías et al. 2018; Germán Zubero 2009, Lana Berasain 2007). These features are not particularly related to patriarchal societies (Szoltysek et al. 2017), so these results open up the possibility of finding similar or even more extreme manifestations of son preference in other European regions. In this regard, female excess mortality during childhood was also present in 19$^{th}$-century samples from Italy, Sweden and Belgium (Alter et al. 2004, 334-337; Oris et al. 2004, 366).

Lastly, the existence of female excess mortality derived from discriminatory practices in infancy and childhood also puts into question the implications derived from historical life tables (Coale and Demeny 1968). The observed gender gap in mortality rates cannot be longer considered solely the result of the interaction between biological dimensions and different disease environments. The way societies treated their children affected their chances of survival and, if gender discrimination was in place, life tables are likely over-estimating female mortality rates and/or underestimating male ones. Routinely relying on life tables at face value perpetuates the notion that some gender gaps in mortality are the result of "natural" factors when they are actually mediated by differential treatment. This is especially problematic in areas exhibiting relatively high child sex ratios, so this investigation stresses the need to consider behavioural factors when analysing sex-specific mortality patterns in these contexts. Moreover, as Harris and Ross (1987, 155) argue, failing to consider these practices distorts our understanding of the demographic transition. In this regard, part of the decline in infant and child mortality, especially of females, from the late-19$^{th}$ century onwards resulted from gradual changes on how parents treated their sons and daughters as the trade-off between child-rearing costs and benefits evolved.

## 2.- Data and historical background

This study focuses on a small rural area in Aragon, in North-Eastern Spain, that is located around 19-40 kilometres away from Zaragoza, the regional capital (see Map 1). This area, a combination of plains and foothills near the Huerva river, comprises 13 small municipalities[4]. Their total population was approximately 5,525 inhabitants in 1750, 8,315 in 1857 and 9,556 in 1950. The statistical analysis relies on the complete Church registers of these villages, whose records provide high-quality information on all births, marriages and deaths that occurred from 1575 onwards (although the starting date varies by location; more details about the 'Alfamén and Middle Huerva Database' in Marco-Gracia 2017; 2019; 2020). In total, the whole dataset contains information on 63,175 individuals (name, sex, place and date of birth, parents' names, date of death, etc.) born between 1750 and 1950, thus allowing reconstituting their complete life histories (and their families). This longitudinal dataset has also been complemented with information on occupation and literacy contained in population lists (1747-1830), population censuses (1857, 1860) and electoral rolls (1890-1955)[5].

**Map 1. Area of study: Middle Huerva (Aragón, Spain)**




Note: Dark dots refer to the localities studied here (except Zaragoza, the provincial capital) and the corresponding shaded areas to their municipal boundaries. Apart from rivers (in blue) and main roads (grey), the map also depicts neighbouring villages (white dots).

Given that we rely on local records, we do not have all the information on those individuals who migrated in/out our area of study. Children may have moved out with their families and die somewhere else, but we cannot observe their age of death. Similarly, an individual may have moved into our area of study, so we can have information on his/her death but not on his/her birth. The analysis of infant and child

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[4] The studied localities are: Alfamén, Aylés, Botorrita, Cosuenda, Jaulín, Longares, Mezalocha, Mozota, Muel, Torrecilla de Valmadrid, Tosos, Valmadrid, and Villanueva de Huerva.

[5] Due to the lack of information, data on occupations and literacy is missing for around one-quarter of the individuals. This scarcity is more important in the earlier periods and very low in the final years analysed here.

mortality should therefore be restricted to those individuals from who we have complete information. In this regard, we consider all observations that we know both their birth and age of death. In addition, we also consider those who got married and therefore were alive during their infancy and childhood. Given that migration was predominantly male, we are excluding a larger proportion of males. This is crucial because, although we don't know their age of death, it is likely that these individuals did not die before age 10. Excluding them from the sample thus artificially increases male mortality rates because they are no longer in the denominator and makes finding female excess mortality even more challenging.

Moreover, twins not only suffered extremely high mortality rates, but also generated an unexpected shock to the household resources that can distort our analysis. We are interested in unveiling what happens within regular families, so we have excluded from the analysis those children born in families who raised twins. In addition, we should bear in mind however that registration quality greatly improved from 1750 onwards. Infant and child mortality rates before that date are too low, so under-registration of deaths is likely to be an issue. We will therefore restrict our analysis to the period 1750-1950. The restricted sample contains 33,924 individuals. Table A1 in the appendix shows that these restrictions hardly affect the socio-economic composition of our sample. As mentioned before, however, the proportion of males is now lower, thus increasing male mortality rates and thus biasing our research strategy against the likelihood of finding gender discrimination[6].

The area of study, 13 villages covering around 500 kms$^2$, hosted a population who mostly lived in nuclear households and was essentially devoted to agriculture (mostly wheat and some wine) and sheep grazing. Our records show that around 85 per cent of the male working population was engaged in the agricultural sector between 1800 and 1950. Average marital fertility of complete families (both spouses reaching 49 years old) was relatively stable around 6-7 children up to 1900 and declined rapidly thereafter following the demographic transition. Infant and child mortality rates were very high though and only around half of the children survived to their fifth birthday.

Mortality rates began declining consistently in the last third of the 19th century due to increasing living standards. The decline firstly benefited children in their late childhood and spread later to younger cohorts. Infants were the last one to join this trend and their survival chances only significantly increased from 1900 onwards when hygienic conditions and their mothers' health improved, an improvement that was especially visible during early childhood. Anthropometric evidence also indicates that standards of living were extremely low: the average male height was around 160 centimetres in mid-19th-century, well below their European counterparts or their fellow Spanish in other regions (Martínez-Carrión et al. 2016; Hatton and Bray 2020). In an

[6] Analysing a less restrictive sample including those individuals whose age of death is unknown makes the female death penalty even more visible. Results available upon request.

area where most of the population enjoyed living standards close to subsistence levels, choices mattered and discriminatory practices could have had lethal consequences.

Although the quantitative information provided by the sources is exceptional, qualitative evidence on how parents treated their children is scarce or non-existent. Although the inheritance system prescribed that all property was divided equally among all children, son preference seems to have formed part of the shared norms in this area during the period of study. In this regard, Spanish women did not enjoy the same status as men: legally subordinated to their fathers and husbands, women were expected to remain within the domestic realm and those who did work in paid jobs received significantly lower wages (Camps 1998; Sarasúa 2002; Borderías et al. 2010; Borderías and Muñoz 2018). Fertility decisions seemed to have indeed been related to the sex composition of the surviving children (Reher and Sanz-Gimeno 2007, Reher and Sandström 2015, Marco-Gracia 2020). Those families who did not have previous sons not only were more likely to continue bearing more children but also had shorter birth intervals, so fertility control strategies where used to control fertility and ensure at least a male heir. Reher and González-Quiñones (2003) also found that the probability of dying of daughters increased more than that of boys after the death of the mother, thus suggesting that son preference was stronger for fathers. It also appears that women were discriminated through an unequal allocation of resources within the household, both in terms of nutrition and educational investments (Sarasúa 2002; Borderías et al. 2014). As well as to the relative backwardness of this area, literacy rates also testify to how unequally parents treated boys and girls: while around 40 per cent of men were literate in 1860, less than 5 per cent of women were able to read and write. Crucially, Beltrán Tapia and Marco-Gracia (2020) found that families in this region were neglecting a significant fraction of their female babies during the 19th century and the first decades of the 20th century. Next sections address whether these underlying attitudes towards boys and girls also translated into their respective survival chances during infancy and childhood.

## 3.- Mortality rates in infancy and childhood

In contexts characterised for low standards of living and high mortality rates, son preference may generate an unequal allocation of resources and thus have deleterious effects on girls' health. However, it is crucial to stress that finding patterns of gender discrimination against girls is specially challenging due to the female biological advantage. Males are more vulnerable, and their mortality rates are thus higher both at birth and during the first year of life. This frailty was especially visible in the high-mortality environments that characterised pre-industrial Europe due to poor living conditions, lack of hygiene and the absence of public health systems[7]. This is illustrated

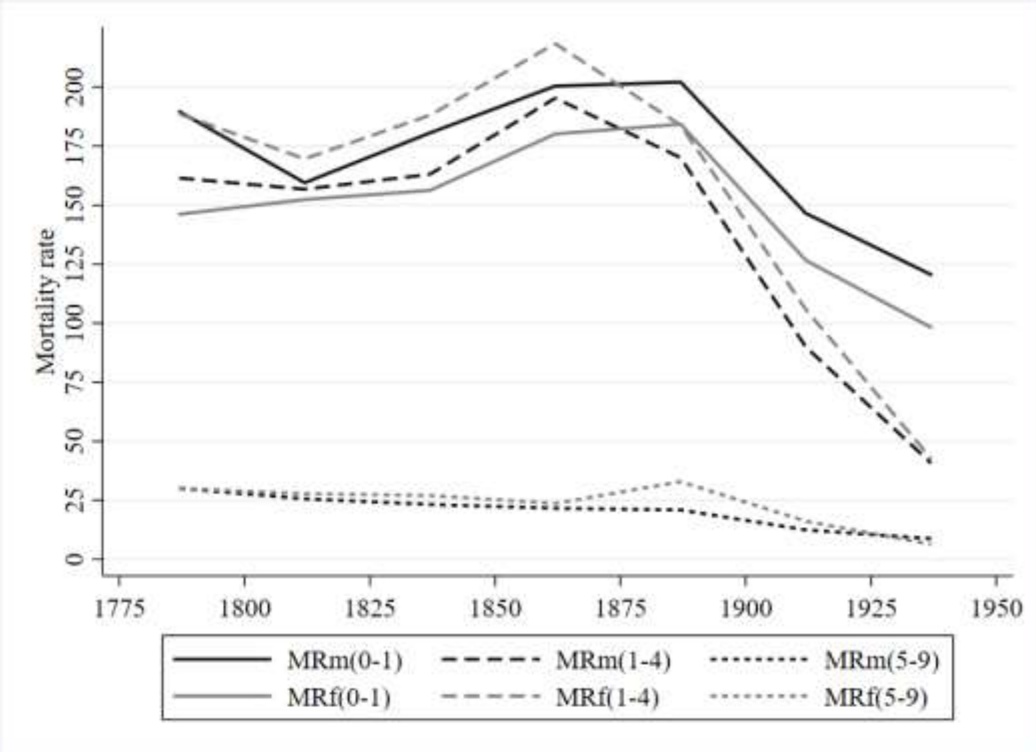
[7] On infant and child mortality in Spain, see Gómez Redondo (1992), Dopico and Reher (1998), Ramiro-Fariñas and Sanz-Gimeno (2000), Cussó and Nicolau (2000), Reher and Sanz-Gimeno (2004) and Llopis Agelán et al. (2015).

in figure A1 in the Appendix using information from as many European countries and periods as possible (taken from the Human Mortality Database). These graphs show that higher average mortality rates are associated with a wider gap between male and female mortality rates, especially during the first year of life.

On the other hand, registering deaths required paying the funeral fee, so under-reporting of deaths might be an issue. However, if anything, this would affect girls more than boys, so female mortality rates would then actually be even higher. Therefore, gender gaps reported here would only provide a lower threshold of the impact of gender discrimination. This is more likely to have affected infants, so it is plausible that our results underestimate the importance of female deaths during the first days or weeks.

Figure 1 plots the evolution of sex-specific mortality rates for different age-groups between 1775 and 1950. As expected, due to the female biological advantage, more boys than girls were dying during the first year of life. In contrast, female mortality rates were higher at older ages, and especially between their first and the fifth birthday. This is striking because male frailty continues throughout childhood, especially during the second year of life (Waldron 1998; WHO 2019). Therefore, either more girls or less boys were dying than it should be expected, thus suggesting than discriminatory practices were unduly widening the gender gap in mortality during the 1-4 age-group. Infant and especially child mortality rates dropped significantly since the late 19$^{th}$ century but the gender gap favouring boys was still visible in the first decades of the 20$^{th}$ century. Mortality rates at late childhood (aged 5-9) were much lower and kept slowly declining throughout our period of study. It is nonetheless remarkable that girls at this age-group suffered a mortality spike in the second half of the 19$^{th}$ century.

**Fig. 1. Mortality rates in infancy and childhood (by sex), 1775-1925**

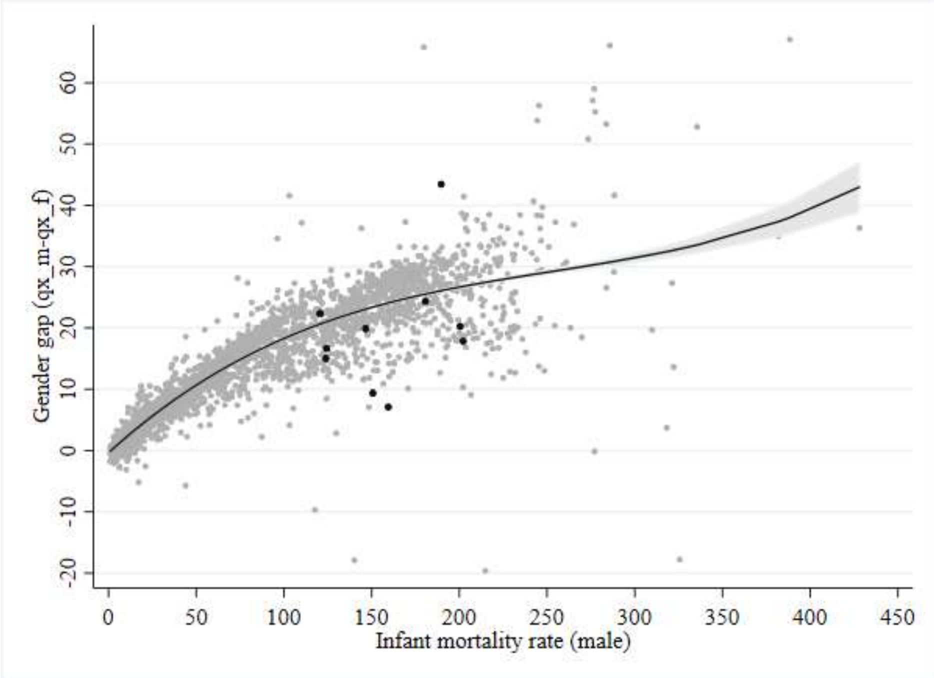



Note: MRm and MRf refer to male and female mortality rates, respectively. Age-groups between brackets. Source: AMHDB

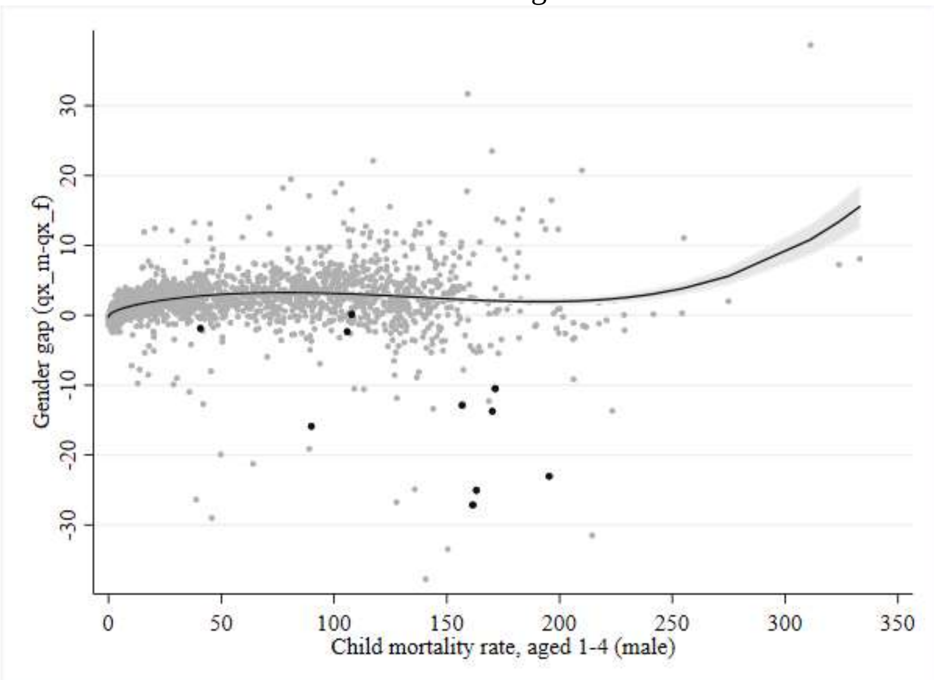
As explained above, the sex-specific mortality rates during early childhood (aged 1-4) stand in marked contrast to what it should be expected considering that female are less vulnerable during this age. The gender mortality gap during the first year of life favoured girls but this was expected for the same reasons. It is though still possible that gender discrimination may have also played a role during the first year of life despite not being able to fully compensate girls' biological advantage. Figure 2 puts our case study in international perspective using information from the Human Mortality Database. These figures, that plot historical mortality rates by age-group (0-1, 1-4 and 5-10) and the corresponding gender gap for as many European countries and periods as possible, further confirm that the gender gaps that we observe in our area of study do not conform to what it should be expected. The black dots depicting the corresponding figures from our data are always well below the international trend: the gender gap is subsequently smaller and therefore more girls (or less boys) are dying than expected according to those mortality rates. The disparity is widest for the 1-4 age-group, but it is also significant during the first year of life. This means that, even though we observe more boys than girls dying at that age, the gap is smaller than what it should be, thus suggesting that some sort of discrimination was affecting sex-specific mortality rates.

**Fig 2. Mortality rates and the gender gap in infancy and childhood, 1750-2016**

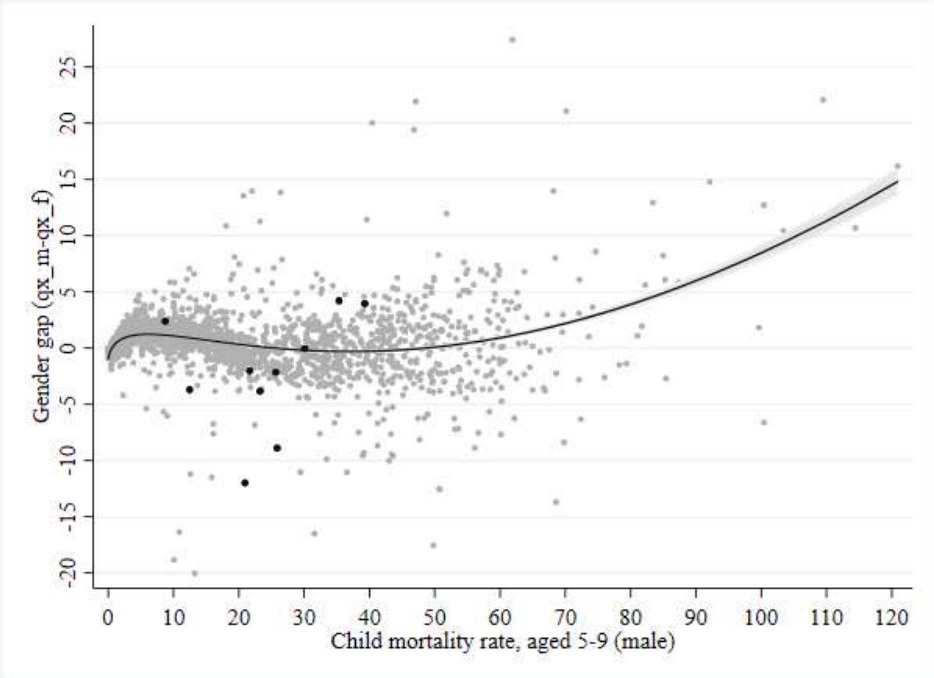
Aged 0-1




Aged 1-4




Aged 5-9




Source: Human Mortality Database and AMHDB. Coverage varies by country. The black dots refer to the observations in our Spanish sample.

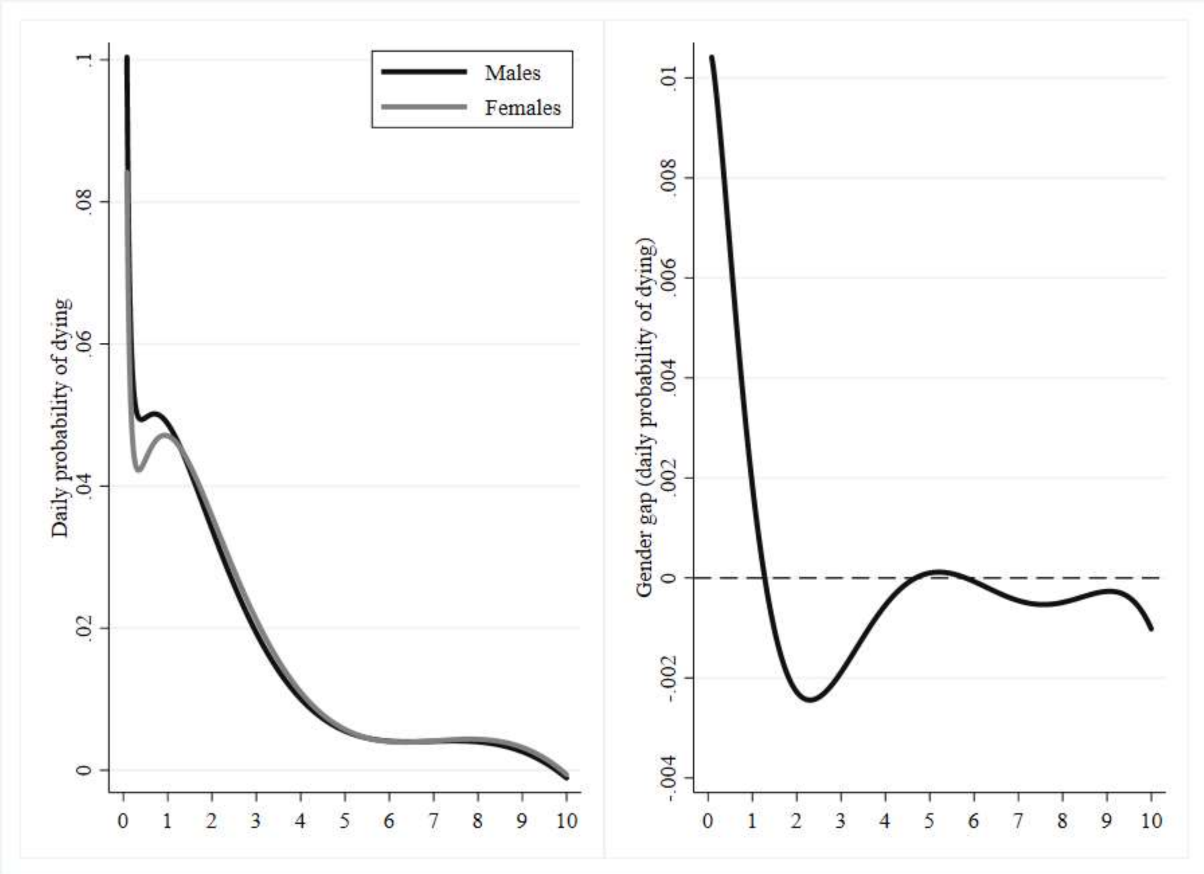
In order to shed more light on the nature of these patterns, figure 3 displays both daily mortality rates during the first 10 years of life and the subsequent gender mortality gap between 1750 and 1900, when gender discrimination was potentially more important. Given that the mortality is highest during the first days/weeks, including them in the visualization prevents noticing any pattern because it completely distorts the scale of the y-axis. We have therefore excluded the first month in order to be able to detect what it is going on[8]. Regardless of what may have been happening during the first four weeks of life, this plot suggests that weaning had especially deleterious effects on girls' health. While male and female mortality rates fell dramatically following a very similar trend during the first months of life, something changed abruptly around the 6-7th month for both sexes alike. Breastfeeding seems to have protected boys and girls equally but, as soon as they were weaned, girls began to die at a higher rate[9]. In this regard, while breastfeeding does not imply competition for resources because there is only one infant that can benefit from it, the introduction of solid food seems to have unleashed discriminatory practices in the quantity or quality of the food that was given to boys and girls. Alternatively, once children were weaned and deployed from the protective effect of breastfeeding, infants and toddlers were more likely to contract gastro-intestinal diseases due to the prevailing hygienic conditions (Guinnane and Ogilvie 2014; Pérez Moreda et al. 2015). It is therefore also plausible that, when ill, parents devoted more attention and care to their sons than their daughters. In high-

[8] Figure A2 in the Appendix provides the whole picture.

[9] Given that the abrupt change in mortality happens simultaneously for both boys and girls suggest that there were no differences in the duration of breastfeeding.

mortality contexts as the one present in our area of study, minor differences in how these children were treated were likely to have had lethal consequences. Girls continued indeed dying in higher numbers as children grew older, especially during the 2-5 age-group but also at later ages.

**Fig. 3. Sex-specific daily probability of dying and gender mortality gap (age 0-10), 1750-1900**




Note: In order to help detecting these patterns, the first month of life has been excluded from the visualization.

Source: AMHDB.

The patterns observed before are statistically significant, even after controlling for individual characteristics. Table 1 reports the results of estimating a logit model assessing whether being female affected the probability of dying at different stages of infancy and childhood between 1750 and 1900. This exercise controls for the number of surviving siblings at the beginning of those stages (at birth, 1, 2 or 5 years old), as well as village and period fixed effects (odd columns). An additional set of potential confounding factors, mother's age, father's occupation and literacy are also considered (even columns). Summary statistics for the variables employed in the statistical analysis are reported in Table A2 in the Appendix. This analysis basically confirms the shape of the graphs above: the survival female advantage dramatically decreased during the first year of life and completely disappeared between age 1-2. Girls in the 2-5 age-group actually suffered higher mortality levels than boys.

**Table 1. Probability of dying (by age), 1750-1900**

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | At birth | | 1st week | | 2-4 weeks | | 1-6 months | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Female | -0.347*** | -0.418 | -0.221** | -0.137 | -0.196** | -0.077 | -0.174*** | -0.058 |
| | (0.134) | (0.282) | (0.098) | (0.231) | (0.084) | (0.179) | (0.059) | (0.101) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 26,768 | 6,083 | 26,306 | 5,849 | 25,700 | 5,912 | 24,803 | 5,742 |
| Pseudo R2 | 0.0237 | 0.0436 | 0.0130 | 0.0165 | 0.0136 | 0.0260 | 0.0104 | 0.0175 |
| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
| | 6-12 months | | 1-2 years | | 2-5 years | | 5-10 years | |
| | (9) | (10) | (11) | (12) | (13) | (14) | (15) | (16) |
| Female | -0.125*** | -0.159* | 0.036 | 0.049 | 0.065 | 0.140* | 0.087 | 0.079 |
| | (0.040) | (0.095) | (0.039) | (0.097) | (0.053) | (0.074) | (0.076) | (0.121) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 22,908 | 5,309 | 21,076 | 4,852 | 17,796 | 4,070 | 14,896 | 3,463 |
| Pseudo R2 | 0.0080 | 0.0187 | 0.0092 | 0.0155 | 0.0125 | 0.0180 | 0.0166 | 0.0278 |

Coefficients estimated using a logit regression model. Robust standard errors in parentheses (clustered at the village level); *** p<0.01, ** p<0.05, * p<0.1; For simplicity, the intercept is not reported. The first set of controls include child order, number of children alive at the beginning of the period (at birth, 0, 1, 2 or 5 years), village and period fixed-effects. The second set of controls include mother's age, father's occupation and father's literacy.

The fact that the gender gap in mortality is hardly visible even during the first year of life when the additional set of controls are included (even columns) further suggests that discriminatory practices were either increasing female mortality rates or decreasing males ones, a pattern that was also evident when comparing the gender gap existing in our area of study with the international figures (figure 2 above). The female penalty is even clearer when analysing those children born at high parities, when additional children further constrained the limited household resources. Given the female biological advantage, adverse circumstances should inflict a deadlier tall on boys. The empirical analysis however evidences that the opposite was true, especially after weaning (see Table A3 in the Appendix that restricts the analysis to those children born at parity 4 or higher). Male excess mortality is not visible now when children are aged 6-12 months and the female penalty increases during the 2-5 age-group.

Rather than a conscious family decision that affected a small number of families, gender discriminatory practices that affected sex-specific mortality rates during infancy and childhood seem to have been part of a generalised cultural system that privileged boys in terms of access to food and/or care. On the one hand, there are very little differences between socio-economic groups: table A4 in the appendix shows that the gender mortality gap hardly changes between children born in families of landowners and landless and semi-landless. On the other hand, this widely shared attitudes persisted over time: table A5 in the appendix evidences that the patterns found here did not change during the first half of the 20th century. If anything, the female penalty became even more pronounced during the second year of life. We should bear in mind though that, by slicing the sample into smaller units, these analyses are noisier and the statistical results less precise.

## 4.- Son preference and sex-specific mortality rates during infancy and childhood

In order to shed more light on the underlying mechanism behind these patterns, we now investigate whether individual characteristics, such as the number of siblings and the sex composition of those children, could accentuate discriminatory practices that translated into differential mortality rates for boys and girls. Tables 2A and 2B report the results of estimating the effect of the number of surviving children, as well as whether there were no males or females among them, on the sex-specific probability of dying at different stages during infancy and childhood (while panel A assesses the effect on male mortality, panel B does the same for females). This model controls for parity, village and period fixed effects (odd columns) and then adds and an additional set of other potential confounding factors: mother's age, father's occupation and literacy (even columns). This analysis focuses on the period 1750-1900 when gender discrimination was potentially more important. Given that discriminatory practices are probably more visible in large families due to the pressure that additional children exerted over the limited household resources, tables A6 (a) and (b) in the appendix replicate the analysis focusing only on children born at high parity (4 or above). We should bear in mind though that the female biological advantage implies that boys should suffer more under adverse circumstances, so not observing mortality differences between boys and girls would be an indirect evidence of gender discriminatory practices.

**Table 2A. Probability of dying (by age), 1750-1900**

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| PANEL A: BOYS | At birth | | 1st week | | 2-4 weeks | | 1-6 months | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Children alive | 0.054 | 0.158 | 0.032 | -0.400* | -0.071 | -0.339** | -0.050 | -0.173* |
| | (0.085) | (0.192) | (0.058) | (0.205) | (0.056) | (0.132) | (0.035) | (0.100) |
| No males | 0.082 | 0.235 | 0.276** | -0.578 | -0.049 | -0.110 | -0.123 | -0.232 |
| | (0.144) | (0.465) | (0.124) | (0.423) | (0.125) | (0.313) | (0.079) | (0.276) |
| No females | 0.197 | 0.314 | 0.414*** | -0.152 | 0.128 | -0.158 | -0.074 | -0.488*** |
| | (0.177) | (0.331) | (0.139) | (0.245) | (0.087) | (0.136) | (0.108) | (0.100) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 13,659 | 3,008 | 13,332 | 2,961 | 13,030 | 2,957 | 12,535 | 2,966 |
| Pseudo R2 | 0.0263 | 0.0619 | 0.0182 | 0.0330 | 0.0167 | 0.0489 | 0.0086 | 0.0174 |

| PANEL A: GIRLS | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | At birth | | 1st week | | 2-4 weeks | | 1-6 months | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Children alive | 0.053 | 0.283* | -0.021 | -0.062 | 0.008 | -0.006 | 0.009 | -0.092 |
| | (0.114) | (0.171) | (0.082) | (0.091) | (0.087) | (0.146) | (0.062) | (0.117) |
| No males | 0.477*** | 1.053** | -0.028 | -0.045 | 0.064 | -0.244 | 0.046 | -0.248 |
| | (0.167) | (0.498) | (0.167) | (0.283) | (0.108) | (0.282) | (0.109) | (0.329) |
| No females | 0.325 | -0.302 | -0.084 | -0.070 | 0.085 | -0.413 | 0.061 | 0.049 |
| | (0.223) | (0.383) | (0.190) | (0.355) | (0.154) | (0.308) | (0.089) | (0.103) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 13,109 | 2,791 | 12,869 | 2,801 | 12,670 | 2,847 | 12,268 | 2,761 |
| Pseudo R2 | 0.0272 | 0.0638 | 0.0181 | 0.0378 | 0.0134 | 0.0402 | 0.0160 | 0.0409 |

Coefficients estimated using a logit regression model. Robust standard errors in parentheses (clustered at the village level); *** p<0.01, ** p<0.05, * p<0.1; For simplicity, the intercept is not reported. The number of children alive, as well as the fact that none of them are male or female, refer to the beginning of the period analysed (at birth, 0, 1, 2 or 5 years). The first set of controls include parity, village and period fixed-effects. The second set of controls include mother's age, father's occupation and father's literacy.

In contrast to what happened around birth when having no surviving brothers is associated with the neglect of female infants (as reported in the previous section), the sex composition of the siblings did not affect sex-specific mortality rates during the remaining first year of life. The number of siblings, however, seems to have reduced male mortality during the first six months. The mechanism behind this association is unclear and it is difficult to separate whether it is produced by biological or behavioural considerations. Note however that this effect is visible even controlling for parity and mother’s age, thus suggesting that nature is not the only driving force here. The analysis of those children born at high parities confirms that boys were somewhat prioritised before turning 6 months: not only was the effect of number of surviving children even stronger, but having no surviving siblings clearly benefited their chances of survival, an effect that is not visible for girls. As mentioned above, the fact that boys are biologically more vulnerable makes detecting discriminatory practices more difficult. Moreover, it is very likely that the non-competing nature of breastfeeding protected both boys and girls alike during the first months. Although girls were clearly neglected right after birth, it seems that, once the child is accepted, the mother started breastfeeding, thus making discrimination less visible until weaning starts.

**Table 2B. Probability of dying (by age), 1750-1900**

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| PANEL A: BOYS | 6-12 months | | 1-2 years | | 2-5 years | | 5-10 years | |
| | (9) | (10) | (11) | (12) | (13) | (14) | (15) | (16) |
| Children alive | -0.050 | -0.089 | -0.008 | -0.066 | -0.015 | 0.053 | -0.006 | 0.049 |
| | (0.040) | (0.065) | (0.028) | (0.067) | (0.038) | (0.068) | (0.037) | (0.075) |
| No males | -0.056 | -0.193 | -0.314*** | -0.664*** | -0.359*** | -0.320* | -0.510** | -0.737 |
| | (0.098) | (0.154) | (0.091) | (0.257) | (0.130) | (0.173) | (0.199) | (0.558) |
| No females | -0.262*** | -0.270 | -0.131*** | -0.256** | -0.396*** | -0.344** | -0.479*** | -0.118 |
| | (0.097) | (0.226) | (0.049) | (0.122) | (0.088) | (0.169) | (0.149) | (0.276) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 11,505 | 2,746 | 10,534 | 2,497 | 8,920 | 2,104 | 7,505 | 1,805 |
| Pseudo R2 | 0.0106 | 0.0402 | 0.0116 | 0.0300 | 0.0195 | 0.0399 | 0.0313 | 0.0405 |

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| PANEL A: GIRLS | 6-12 months | | 1-2 years | | 2-5 years | | 5-10 years | |
| | (9) | (10) | (11) | (12) | (13) | (14) | (15) | (16) |
| Children alive | 0.001 | 0.055 | -0.059** | -0.079 | -0.027 | -0.096** | -0.028 | 0.093 |
| | (0.036) | (0.079) | (0.029) | (0.062) | (0.031) | (0.044) | (0.029) | (0.092) |
| No males | -0.172 | -0.108 | -0.108 | -0.212 | -0.453*** | -0.644*** | -0.387*** | -0.259 |
| | (0.173) | (0.281) | (0.093) | (0.275) | (0.081) | (0.220) | (0.099) | (0.427) |
| No females | 0.076 | 0.239 | -0.430*** | -0.267 | -0.399*** | -0.386* | -0.826*** | -0.088 |
| | (0.128) | (0.214) | (0.071) | (0.176) | (0.099) | (0.198) | (0.128) | (0.247) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 11,403 | 2,544 | 10,542 | 2,348 | 8,876 | 1,960 | 7,391 | 1,651 |
| Pseudo R2 | 0.0084 | 0.0184 | 0.0145 | 0.0228 | 0.0188 | 0.0253 | 0.0277 | 0.0468 |

Coefficients estimated using a logit regression model. Robust standard errors in parentheses (clustered at the village level); *** $p<0.01$, ** $p<0.05$, * $p<0.1$; For simplicity, the intercept is not reported. The number of children alive, as well as the fact that none of them are male or female, refer to the beginning of the period analysed (at birth, 0, 1, 2 or 5 years) The first set of controls include parity, village and period fixed-effects. The second set of controls include mother's age, father's occupation and father's literacy.

Discriminatory practices clearly resurfaced during the second year of life. While having no male or female siblings alive (regardless of their sex) reduced male mortality for the 1-2 age-group, this feature did not influence female mortality during that age. Notice also that the coefficients on having no male siblings are much higher than those on having no female siblings. It appears that families were clearly favouring boys when no other siblings were alive (especially males) in order to secure a male heir. The effect of discriminatory practices favouring boys is less visible from age 2 onwards, probably because children at those ages are more robust anyway, thus difficulted that this behaviour translated into higher mortality rates. Although analysing those children born at parity 4 or higher reduces sample size and thus reduces the accuracy of the estimates, it shows that the gender mortality gap created by the absence of male siblings during childhood is even larger at higher parities.

Taken together, this section evidences that boys were clearly prioritised, both during infancy and even more clearly as soon as children were weaned. In contrast to what happened at birth and during the first day of life when female neglect was linked to individual characteristics that accentuated that kind of behaviour, discriminatory practices during infancy and childhood seems to be rooted within the cultural norms that shaped how parents treated their children. Discriminatory behaviour generally translated into differential mortality rates by sex, an outcome that is more visible in larger families due to limited resources.

## 5.- Conclusion

This study documents that gender discrimination increased female mortality (or reduced male mortality) during infancy and childhood in pre-industrial Spain. On the one hand, although the fact that boys are biologically more vulnerable during the first year of life makes detecting discriminatory practices more difficult, our evidence suggests that boys were somewhat prioritised during this early stage. On the other hand, breastfeeding appears to have protected boys and girls alike, thus mitigating the effect of discriminatory behaviours. Sex-differences in mortality rates, however, became clearly visible as soon as children were weaned.

In this regard, stopping breastfeeding usually translates into higher mortality rates due to the increased incidence of gastro-intestinal diseases, the main cause of children's deaths, associated with bad hygiene practices. Moreover, access to solid foods also means beginning competing for scarce resources. This turning point indeed dramatically altered the sex-specific mortality patterns in our sample. Firstly, while female mortality rates here behaved as expected and increased around the $6^{th}$-$8^{th}$ month, male mortality rates did not suffer such penalty. Parents seem to have favoured boys in terms of food and/or care and clearly mitigated the negative effects of weaning, to the point that female mortality rates overcame those of males. Secondly, although overall mortality rates decreased as children grew older, girls continued dying in higher numbers during childhood due to an unequal allocation of food, care and/or workload withing the household[10]. The female penalty was even clearer in those children born in already large

[10] Child abuse was also pervasive and it sometimes resulted in children dying (Tausiet 2001), so this could also have differentially affected boys and girls.

families, when additional children further constrained the limited household resources. Therefore, the allocation of resources within families clearly harmed girls as soon as infants and toddlers began competing for resources. In societies close to subsistence levels where mortality is very high, minor differences in how these children were treated had indeed lethal consequences.

The mortal effects of discrimination are visible regardless whether families had access to land or not, thus evidencing that these practices formed part of a widely shared norm of behaviour. Moreover, individual-level information shows that parents specially prioritised boys when there were not other male siblings alive in order to enhance his survival chances and secure a male heir. The results reported here actually underestimate the deadly effect of discriminatory practices. In this regard, our research strategy has excluded those individuals whose age of death was unknown, mostly due to migration. Given that these migrants were predominantly male, our conservative approach over-estimates male mortality rates and therefore biases our results against the possibility of finding female excess mortality. Moreover, in case under-registration of deaths were infra-estimating mortality rates, this issue would be more important for girls, especially very early in life.

Son preference in pre-industrial Spain therefore generated discriminatory practices that affected sex-specific mortality rates both around birth and during childhood. While the neglect of girls around birth was possibly a conscious family decision that affected a small number of families under certain circumstances (Beltrán Tapia and Marco-Gracia 2020), discriminatory practices during childhood seem to have been part of a generalised cultural system that privileged boys in terms of access to food and/or care and altered sex-specific mortality patterns. Although the effect of these practices was more visible during the 19th century, it persisted into the first decades of the 20th century, thus evidencing that son preference continued to be a strong cultural norm within these societies. These findings not only illustrate the mechanisms behind excess female mortality, but also confirm that the high child sex ratios found in Spain and other regions in Southern Europe are not an artefact of the quality of the registers (Beltrán Tapia and Gallego-Martínez 2017; Beltrán Tapia 2019).

Although this article strongly challenges the notion that there were no missing girls in historical Europe, the analysis focuses on a small region in North-eastern Spain, so more research is needed to assess whether this behaviour was also shared in other European regions. The evidence provided in other studies nonetheless suggest that these patterns were indeed more pervasive than it had been traditionally acknowledged (Tabutin 1978; Johansson 1984; Pinnelli and Mancini 1997; Alter et al. 2004; Oris et al. 2004; Beltrán Tapia and Raftakis 2019). Accounting for the effect of potential discriminatory practices on mortality rates thus becomes of paramount importance for understanding the demographic transition. Not only female mortality rates in some regions seem to have been "unnaturally" higher, but their decline from the later 19th-century onwards could be partly explained by gradual changes on how parents treated their sons and daughters.

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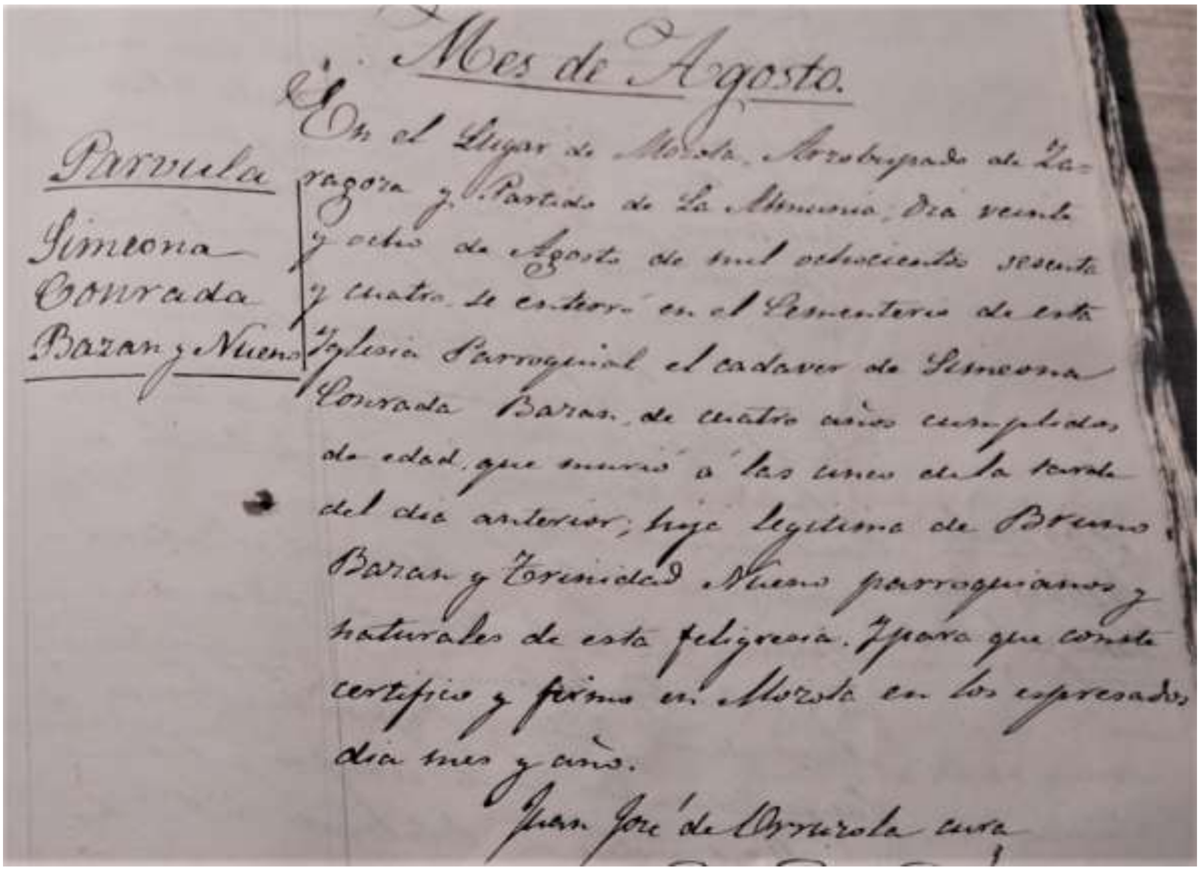
# APPENDIX

**Document 1. Death register**



Mes de Agosto.

Parvula
Simeona
Conrada
Bazan y Nueno

En el Lugar de Mozota, Arzobispado de Zaragoza y Partido de La Almunia, dia veinte y ocho de Agosto de mil ochocientos sesenta y cuatro, se enterró en el Cementerio de esta Iglesia Parroquial el cadaver de Simeona Conrada Bazan, de cuatro años cumplidos de edad, que murió á las cinco de la tarde del dia anterior, hija legitima de Bruno Bazan y Trinidad Nueno parroquianos y naturales de esta feligresía. Y para que conste certifico y firmo en Mozota en los expresados dia mes y año.

Juan José de Ormazola cura

Translation: In the village of Mozota, archbishopric of Zaragoza and judicial area of La Almunia, in August 28 (1864), the dead body of Simeona Conrada Bazan Nueno was buried; she was four years old and died at 5 pm of the previous day. She was the daughter of Bruno Bazan and Trinidad Nueno, parishioners and born in this village. And for the record, I certify and sign it in Mozota on the above day, month and year.

Source: AMHDB.

**Table A1. Frequency statistics, 1750-1950**

| | Whole sample | | Restricted sample | |
|---|---|---|---|---|
| **Sex** | Obs. | % | Obs. | % |
| Female | 30,478 | 48,24 | 16,577 | 48,87 |
| Male | 32,697 | 51,76 | 17,347 | 51,13 |
| Total | 63,175 | 100,00 | 33,924 | 100,00 |
| **Period** | Obs. | % | Obs. | % |
| 1750-1799 | 14,971 | 23,70 | 7,727 | 22,78 |
| 1800-1849 | 15,581 | 24,66 | 8,049 | 23,73 |
| 1850-1899 | 17,840 | 28,24 | 10,843 | 31,96 |
| 1900-1949 | 14,783 | 23,40 | 7,305 | 21,53 |
| Total | 63,175 | 100,00 | 33,924 | 100,00 |
| **Father's occupation** | Obs. | % | Obs. | % |
| Farmers | 8,737 | 36,38 | 5,277 | 35,23 |
| Labourers | 10,721 | 44,64 | 6,857 | 45,77 |
| Shepherds | 1,448 | 2,29 | 897 | 5,99 |
| Artisans | 1,911 | 7,96 | 1,268 | 8,46 |
| Elites | 294 | 1,22 | 186 | 1,24 |
| Other | 908 | 3,78 | 495 | 3,30 |
| Total | 24,019 | 100,00 | 14,980 | 100,00 |
| **Father's literacy** | Obs. | % | Obs. | % |
| Illiterate | 11,012 | 50,15 | 6,957 | 52,68 |
| Literate | 10,948 | 49,85 | 6,250 | 47,32 |
| Total | 21,960 | 100,00 | 13,207 | 100,00 |

Note: The restricted sample excludes those individuals whose age of death is not observed due to migration. Given that we are interested in infant and child mortality, we have kept those individuals who got married and therefore did not die while growing up. Children born in families who raised twins are also excluded. Source: AMHDB.

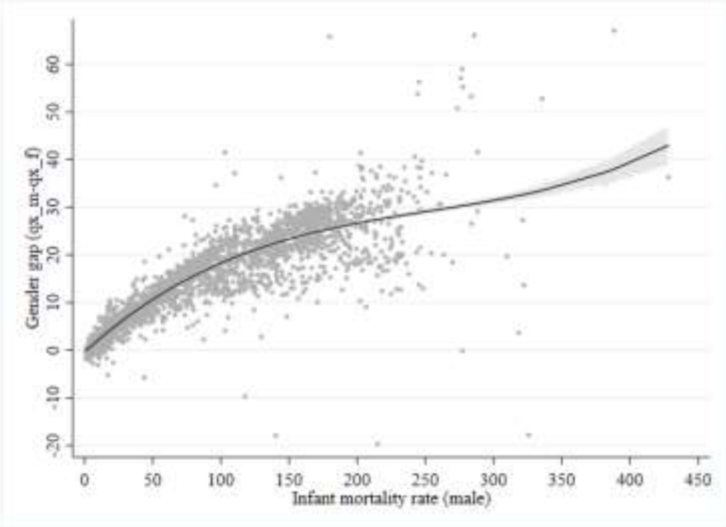
**Table A2. Summary statistics, 1750-1950**

| | Obs. | Mean | St. Dev. | Min | Max |
|---|---|---|---|---|---|
| Sex (Male=0) | 33,924 | 0.511 | 0.500 | 0 | 1 |
| Age of death (years) | 25,773 | 23.16 | 30.42 | 0 | 105.0 |
| Dead during the first day of life | 33,924 | 0.020 | 0.141 | 0 | 1 |
| Dead between 1-7 days | 33,237 | 0.022 | 0.146 | 0 | 1 |
| Dead between 8-30 days | 32,514 | 0.034 | 0.180 | 0 | 1 |
| Dead between 1-6 months | 31,423 | 0.077 | 0.267 | 0 | 1 |
| Dead between 6-12 months | 29,003 | 0.078 | 0.269 | 0 | 1 |
| Dead between 1-2 years | 26,734 | 0.142 | 0.349 | 0 | 1 |
| Dead between 2-5 years | 22,934 | 0.134 | 0.341 | 0 | 1 |
| Dead between 5-10 years | 19,854 | 0.056 | 0.231 | 0 | 1 |
| Children alive at birth | 33,924 | 1.65 | 1.715 | 0 | 14 |
| Children alive at age 1 | 33,924 | 1.68 | 1.755 | 0 | 13 |
| Children alive at age 2 | 33,924 | 1.77 | 1.792 | 0 | 14 |
| Children alive at age 5 | 33,924 | 2.10 | 1.922 | 0 | 13 |
| No males at birth | 33,924 | 0.498 | 0.500 | 0 | 1 |
| No males at age 1 | 33,924 | 0.499 | 0.500 | 0 | 1 |
| No males at age 2 | 33,924 | 0.479 | 0.500 | 0 | 1 |
| No males at age 5 | 33,924 | 0.417 | 0.493 | 0 | 1 |
| No females at birth | 33,924 | 0.513 | 0.500 | 0 | 1 |
| No females at age 1 | 33,924 | 0.516 | 0.500 | 0 | 1 |
| No females at age 2 | 33,924 | 0.495 | 0.500 | 0 | 1 |
| No females at age 5 | 33,924 | 0.438 | 0.496 | 0 | 1 |

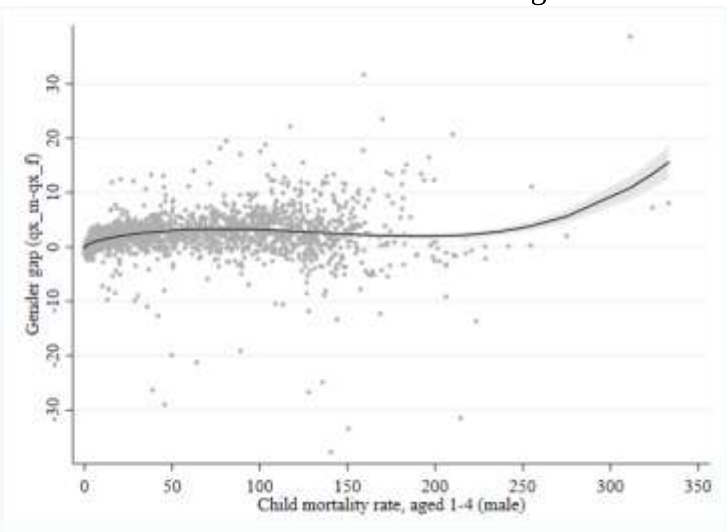
Source: AMHDB.

**Fig. A1 Mortality rates and the gender gap in infancy and childhood, 1750-2016**

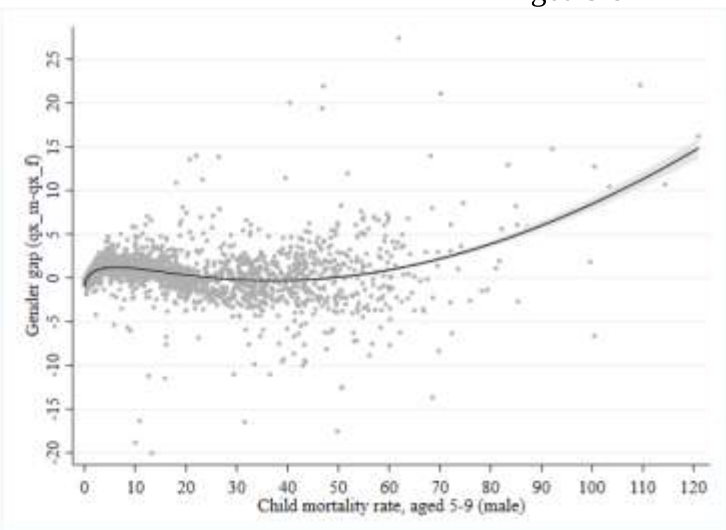
Aged 0-1




Aged 1-4

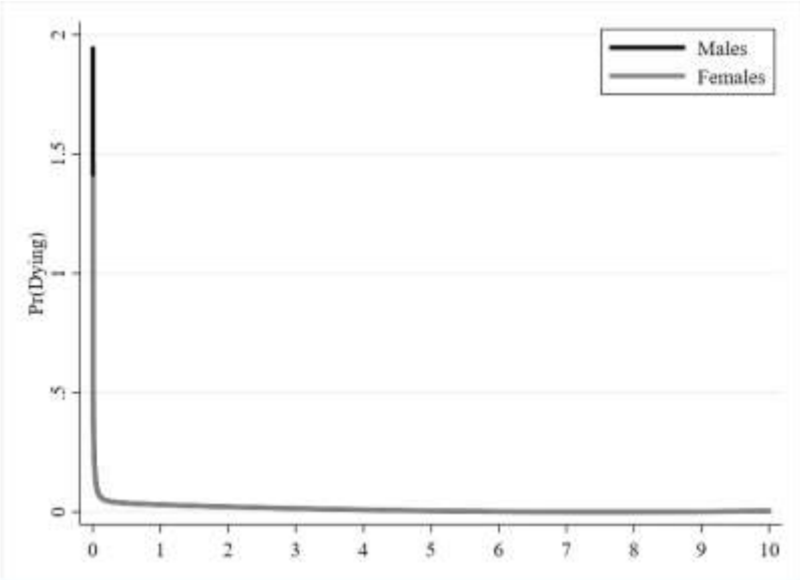



Aged 5-9




Source: Human Mortality Database and AMHDB. Coverage varies by country.

**Fig. A2. Daily probability of dying (by sex and age), 1750-1900**




Source: AMHDB

**Table A3. Probability of dying at high parities (by age), 1750-1900**

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | At birth | | 1st week | | 2-4 weeks | | 1-6 months | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Female | -0.407*** | -0.473 | -0.223 | -0.095 | -0.248*** | -0.101 | -0.178*** | -0.180 |
| | (0.132) | (0.436) | (0.136) | (0.349) | (0.064) | (0.292) | (0.067) | (0.140) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 12,656 | 2,928 | 12,465 | 2,883 | 12,206 | 2,822 | 11,868 | 2,909 |
| Pseudo R2 | 0.0179 | 0.0293 | 0.0141 | 0.0172 | 0.0091 | 0.0174 | 0.0086 | 0.0207 |

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | 6-12 months | | 1-2 years | | 2-5 years | | 5-10 years | |
| | (9) | (10) | (11) | (12) | (13) | (14) | (15) | (16) |
| Female | -0.087 | 0.051 | 0.013 | 0.007 | 0.112*** | 0.195* | 0.046 | 0.021 |
| | (0.079) | (0.186) | (0.046) | (0.141) | (0.041) | (0.116) | (0.091) | (0.208) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 11,027 | 2,725 | 10,148 | 2,485 | 8,574 | 2,069 | 7,201 | 1,748 |
| Pseudo R2 | 0.0088 | 0.0262 | 0.0119 | 0.0206 | 0.0199 | 0.0268 | 0.0238 | 0.0528 |

Coefficients estimated using a logit regression model. Robust standard errors in parentheses (clustered at the village level); *** p<0.01, ** p<0.05, * p<0.1; For simplicity, the intercept is not reported. The first set of controls include child order, number of children alive at the beginning of the period (at birth, 0, 1, 2 or 5 years), village and period fixed-effects. The second set of controls include mother's age, father's occupation and father's literacy.

**Table A4. Probability of dying (by age), farmers vs labourers 1750-1900**

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | At birth | | 1st week | | 2-4 weeks | | 1-6 months | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Female | -0.328** | -0.526** | -0.144 | 0.103 | 0.261* | -0.013 | -0.245 | -0.080 |
| | (0.161) | (0.216) | (0.279) | (0.285) | (0.147) | (0.237) | (0.221) | (0.231) |
| Fem*Labourers | -0.241 | 0.038 | 0.017 | -0.280 | -0.411 | -0.253 | 0.355 | 0.135 |
| | (0.209) | (0.345) | (0.253) | (0.371) | (0.262) | (0.250) | (0.251) | (0.251) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 7,810 | 4,801 | 7,666 | 4,730 | 7,495 | 4,613 | 7,274 | 4,469 |
| Pseudo R2 | 0.0297 | 0.0472 | 0.0106 | 0.0157 | 0.0224 | 0.0300 | 0.0136 | 0.0153 |

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | 6-12 months | | 1-2 years | | 2-5 years | | 5-10 years | |
| | (9) | (10) | (11) | (12) | (13) | (14) | (15) | (16) |
| Female | -0.115 | -0.155 | 0.086 | 0.122 | 0.019 | 0.177 | -0.111 | -0.083 |
| | (0.149) | (0.182) | (0.113) | (0.146) | (0.086) | (0.164) | (0.236) | (0.317) |
| Fem*Labourers | -0.081 | 0.004 | -0.111 | -0.097 | 0.165 | 0.128 | 0.173 | 0.045 |
| | (0.213) | (0.290) | (0.149) | (0.186) | (0.101) | (0.198) | (0.338) | (0.453) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 6,703 | 4,142 | 6,160 | 3,789 | 5,149 | 3,179 | 4,360 | 2,713 |
| Pseudo R2 | 0.0128 | 0.0171 | 0.0086 | 0.0168 | 0.0136 | 0.0147 | 0.0225 | 0.0299 |

Coefficients estimated using a logit regression model. Robust standard errors in parentheses (clustered at the village level); *** p<0.01, ** p<0.05, * p<0.1; For simplicity, the intercept is not reported. The first set of controls include father's occupation, child order, number of children alive at the beginning of the period (at birth, 0, 1, 2 or 5 years), village and period fixed-effects. The second set of controls include mother's age and father's literacy.

**Table A5. Probability of dying (by age), 1800-1950**

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | At birth | | 1st week | | 2-4 weeks | | 1-6 months | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Female | -0.310 | -0.238 | -0.280** | -0.165 | -0.346** | -0.450 | -0.090 | -0.090 |
| | (0.200) | (0.459) | (0.131) | (0.215) | (0.164) | (0.303) | (0.095) | (0.183) |
| *1850-1900 | -0.031 | -0.175 | 0.132 | 0.003 | 0.202 | 0.474 | -0.022 | 0.061 |
| | (0.261) | (0.432) | (0.170) | (0.199) | (0.240) | (0.316) | (0.091) | (0.170) |
| *1900-1950 | -0.128 | -0.522 | 0.254 | 0.031 | 0.376** | 0.509* | -0.074 | -0.043 |
| | (0.288) | (0.508) | (0.298) | (0.451) | (0.178) | (0.260) | (0.108) | (0.283) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 26,053 | 9,946 | 25,649 | 9,525 | 25,109 | 9,650 | 24,276 | 9,368 |
| Pseudo R2 | 0.0235 | 0.0430 | 0.0168 | 0.0197 | 0.0162 | 0.0170 | 0.0088 | 0.0137 |

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | 6-12 months | | 1-2 years | | 2-5 years | | 5-10 years | |
| | (9) | (10) | (11) | (12) | (13) | (14) | (15) | (16) |
| Female | -0.037 | -0.116 | -0.025 | -0.206 | 0.009 | 0.229 | 0.075 | 0.032 |
| | (0.098) | (0.172) | (0.084) | (0.154) | (0.084) | (0.244) | (0.113) | (0.347) |
| *1850-1900 | -0.081 | -0.065 | 0.064 | 0.281** | 0.082 | -0.103 | 0.199 | 0.067 |
| | (0.101) | (0.159) | (0.114) | (0.140) | (0.121) | (0.248) | (0.162) | (0.388) |
| *1900-1950 | -0.147 | 0.016 | 0.261* | 0.523*** | 0.031 | -0.061 | 0.062 | -0.096 |
| | (0.136) | (0.199) | (0.140) | (0.184) | (0.143) | (0.207) | (0.216) | (0.344) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 22,363 | 8,665 | 20,546 | 7,985 | 17,515 | 6,932 | 15,391 | 6,209 |
| Pseudo R2 | 0.0075 | 0.0137 | 0.0186 | 0.0260 | 0.0335 | 0.0367 | 0.0299 | 0.0280 |

Coefficients estimated using a logit regression model. Robust standard errors in parentheses (clustered at the village level); *** p<0.01, ** p<0.05, * p<0.1; For simplicity, the intercept is not reported. The first set of controls include child order, number of children alive at the beginning of the period (at birth, 0, 1, 2 or 5 years), village and period fixed-effects. The second set of controls include mother's age, father's occupation and father's literacy.

**Table A6 (a). Probability of dying at high parities (by age), 1750-1900**

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| PANEL A: BOYS | At birth | | 1st week | | 2-4 weeks | | 1-6 months | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Children alive | 0.054 | 0.123 | 0.005 | -0.473*** | -0.082 | -0.342** | -0.053 | -0.260*** |
| | (0.097) | (0.215) | (0.058) | (0.164) | (0.067) | (0.160) | (0.038) | (0.077) |
| No males | 0.036 | 0.039 | 0.078 | -0.996* | -0.116 | -0.212 | -0.029 | -0.366* |
| | (0.186) | (0.601) | (0.174) | (0.585) | (0.209) | (0.357) | (0.072) | (0.209) |
| No females | 0.150 | 0.120 | 0.387* | -0.317 | 0.205 | -0.022 | -0.199 | -0.728*** |
| | (0.274) | (0.417) | (0.204) | (0.349) | (0.207) | (0.499) | (0.148) | (0.153) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 6,211 | 1,434 | 6,331 | 1,485 | 6,197 | 1,417 | 6,003 | 1,491 |
| Pseudo R2 | 0.0208 | 0.0680 | 0.0177 | 0.0669 | 0.0154 | 0.0541 | 0.0104 | 0.0315 |

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| PANEL A: GIRLS | At birth | | 1st week | | 2-4 weeks | | 1-6 months | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Children alive | 0.049 | 0.270* | -0.014 | 0.010 | 0.010 | -0.108 | 0.021 | -0.092 |
| | (0.135) | (0.138) | (0.085) | (0.083) | (0.082) | (0.142) | (0.066) | (0.132) |
| No males | 0.329 | 0.997* | -0.185 | 0.144 | -0.036 | -0.783* | 0.061 | -0.324 |
| | (0.300) | (0.605) | (0.241) | (0.638) | (0.172) | (0.423) | (0.189) | (0.327) |
| No females | 0.291 | -0.630 | 0.134 | 0.117 | 0.147 | -0.157 | 0.103 | 0.061 |
| | (0.288) | (0.680) | (0.221) | (0.448) | (0.173) | (0.568) | (0.084) | (0.156) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 6,130 | 1,168 | 6,021 | 1,359 | 6,009 | 1,299 | 5,853 | 1,408 |
| Pseudo R2 | 0.0205 | 0.0811 | 0.0261 | 0.0392 | 0.0058 | 0.0635 | 0.0125 | 0.0632 |

Coefficients estimated using a logit regression model. Robust standard errors in parentheses (clustered at the village level); *** p<0.01, ** p<0.05, * p<0.1; For simplicity, the intercept is not reported. The number of children alive, as well as the fact that none of them are male or female, refer to the beginning of the period analysed (at birth, 0, 1, 2 or 5 years). The first set of controls include parity, village and period fixed-effects. The second set of controls include mother's age, father's occupation and father's literacy.

**Table A6 (b). Probability of dying at high parities (by age), 1750-1900**

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| PANEL A: BOYS | 6-12 months | | 1-2 years | | 2-5 years | | 5-10 years | |
| | (9) | (10) | (11) | (12) | (13) | (14) | (15) | (16) |
| Children alive | -0.034 | -0.052 | -0.014 | -0.030 | 0.015 | 0.152** | 0.003 | 0.030 |
| | (0.035) | (0.065) | (0.031) | (0.079) | (0.034) | (0.077) | (0.045) | (0.084) |
| No males | 0.048 | -0.094 | -0.627*** | -1.215*** | -0.470*** | -0.134 | -0.777** | -1.553** |
| | (0.085) | (0.260) | (0.110) | (0.278) | (0.162) | (0.341) | (0.344) | (0.650) |
| No females | -0.220 | -0.248 | -0.158* | -0.130 | -0.437*** | -0.223 | -0.724* | -0.052 |
| | (0.180) | (0.279) | (0.094) | (0.211) | (0.146) | (0.330) | (0.410) | (0.426) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 5,544 | 1,393 | 5,086 | 1,275 | 4,302 | 1,063 | 3,644 | 910 |
| Pseudo R2 | 0.0153 | 0.0611 | 0.0219 | 0.0524 | 0.0290 | 0.0493 | 0.0469 | 0.0900 |

| | Dep. Variable: Probability of dying, by age-group | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| PANEL A: GIRLS | 6-12 months | | 1-2 years | | 2-5 years | | 5-10 years | |
| | (9) | (10) | (11) | (12) | (13) | (14) | (15) | (16) |
| Children alive | -0.008 | 0.023 | -0.068** | -0.069 | -0.042 | -0.063 | -0.042* | 0.080 |
| | (0.041) | (0.082) | (0.028) | (0.053) | (0.041) | (0.047) | (0.024) | (0.101) |
| No males | -0.050 | 0.090 | -0.333** | -0.685 | -0.578*** | -0.728** | -0.491*** | -0.672 |
| | (0.188) | (0.264) | (0.167) | (0.429) | (0.103) | (0.346) | (0.184) | (0.498) |
| No females | -0.097 | -0.277 | -0.586*** | -0.244* | -0.687*** | -0.615* | -1.086*** | -0.175 |
| | (0.174) | (0.291) | (0.115) | (0.147) | (0.143) | (0.364) | (0.218) | (0.353) |
| Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Additional controls | NO | YES | NO | YES | NO | YES | NO | YES |
| Observations | 5,483 | 1,323 | 5,062 | 1,201 | 4,272 | 1,004 | 3,549 | 775 |
| Pseudo R2 | 0.0086 | 0.0268 | 0.0216 | 0.0420 | 0.0352 | 0.0308 | 0.0427 | 0.0886 |

Coefficients estimated using a logit regression model. Robust standard errors in parentheses (clustered at the village level); *** p<0.01, ** p<0.05, * p<0.1; For simplicity, the intercept is not reported. The number of children alive, as well as the fact that none of them are male or female, refer to the beginning of the period analysed (at birth, 0, 1, 2 or 5 years). The first set of controls include parity, village and period fixed-effects. The second set of controls include mother's age, father's occupation and father's literacy.