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Missing women in the United Kingdom

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Abstract

This paper investigates the gender-selection decisions of immigrants in the United Kingdom, using data from the 1971–2006 General Household Survey. We examine sex-selection in the UK among immigrant families and the gender composition of previous births, conditional on socio-economic characteristics. Our key result is that better-educated immigrants balance their family after the birth of two sons, by having a daughter thereafter. Our study also is the first to estimate the number of missing women among Asian immigrants in a European country, contributing to research on the US and Canada that missing women are also a phenomenon of the developed world.

JEL codes: J13, J15, O52, Z13

Keywords: Sex-selection, Gender bias, UK

1. Introduction

Emerging economies depict striking demographic patterns: during the past four decades, the sex ratio at birth (SRB), measured as the annual number of male births per 1000 female births, has followed an increasing trend, predominantly in China and India. 'Missing women' were investigated first by Visaria (1967) and Sen (1992), to designate the conspicuous gender imbalances in the sex ratio at birth in India, attributed mainly to the long-standing cultural value and economic necessity of sons. Currently, estimates show that as many as 89 million women are demographically missing in various Asia-Pacific countries, with China and India jointly accounting for 84 million (UNDP 2010). Moreover, in the developed world which is *absolutely* more prosperous than the developing world, it is interesting to examine if missing women also exist here (Dubuc and Coleman 2007; Abrevaya 2009; Almond et al. 2009).

A concern is that parents preferring sons to daughters may resort to prenatal sex-selection techniques to establish the birth of a son (Hu and Schlosser 2012). One important study in the field shows that sex-selection incentives become more pertinent as a household approaches its own predetermined family size limit, as the opportunity cost of having a child of the less preferred gender increases (Abrevaya 2009). Therefore the predictions are that son-biased gender selection tendencies most likely manifest themselves through abnormally high male-birth likelihoods at later birth parities, when family size limitations start to bite, as well as at births following daughters.

This paper extends the scarce literature on missing women in the developed world by investigating whether these two irregularities in male births are present among Indian and Pakistani immigrants in the UK. We also present the determinants of a household's sex-selection choice. To the best of our knowledge, there has been one important study which examines missing women in the UK (Dubuc and Coleman 2007). However, by employing national birth data, this study does not analyze the underlying causes of sex-selection at the household level by observing the gender composition of previous births.

Our study addresses this gap by examining the implied incidence of sex-selection amongst Indian and Pakistani immigrants in the UK, based on data from the General Household Survey 1971–2006. We find higher male-birth likelihoods for Indian families, relative to English families, both at later birth parities and when following the birth of daughters. Moreover, the results suggest that higher levels of education increase the propensity to sex-select. We also provide evidence that a larger family size among Pakistani families potentially mitigates their tendency to sex-select as early as the third parity. One of our main findings for better educated Indian families both in the UK and India is a desire to balance the family after the birth of two sons by having a daughter thereafter. Additionally, this study disentangles the 'explained' and 'unexplained' components that constitute the average differences in male-birth likelihoods between Indian immigrants and English natives, using the Oaxaca-Blinder decomposition method, and finds that 87% of that difference can be attributed to the effect of ethnicity. Lastly, our study quantifies the number of missing women among Indian immigrants in 1993–2006 and calculates this value as 914.

2. Sex ratios at birth and missing women

The existing literature on missing women is diverse and is mainly focused on developing countries (Anderson and Ray 2010). In this section, we outline why the issue of missing women has captured the attention of economists, and in particular why the issue may be of relevance in developed countries as well, even though the scale on which we observe this in the developed world is still small, relative to countries such as India and China.

Demographically, as many as 200 million women and girls around the world are missing according to the United Nations. Put simply, 'missing women' are those whom *ceteris paribus* should be alive, but are not. Boys are naturally more likely to die in infancy than girls and more boys are born than girls. In all societies that record births, 103–105 boys are normally born for every 100 girls (see Austin and Edwards 1981; Johannson and Nygren 1991; Gini 1955; James 2000; Pollard 1969). The biological SRB had long been stable at roughly 1.05 in all societies that record births (Austin and Edwards 1981). This has changed in the past 25 years when ultrasound technology (and to a lesser extent amniocentesis technology) became available and affordable to women. The use of ultrasound techniques for sex determination and sex-selective abortion has been reported since the early 1980s in South Korea and China (Park and Cho 1995; Zeng et al. 1993) and since the late 1980s in India (Bhat and Zavier 2003; Das Gupta and Bhat 1997), especially at higher birth orders. According to UN estimates, the sex ratio at birth has increased globally from a stable 1.05 in the early 1970s to a recent peak of 1 (Ganatra 2008 and Sen 2003).

In China, despite the country's official ban on prenatal sex determinations since 1989 and on sex-selective abortion since 1994, the sex ratio for the generation born between 2000 and 2004 was more than 1.20 (Shuzhuo 2007). According to the Chinese

Academy of Social Sciences, the ratio today has reached 123 boys per 100 girls. In India, according to the 2001 Census, the ratio of boys to girls in the 0–6 age group was 1.08, up from 105.8 in 1991. In 1998 the sex ratio in Pakistan reached 108.5 in the country as a whole and 112.2 in urban areas, reflecting a 2.7% increase from 1981.

There is a small body of literature on the existence of 'missing women' in developed countries. The nationally observed SRB in modern economies does not suggest sex-selection. There appears to be a general consensus that, on average, parents in developed countries do not exhibit a preference for children of particular sex by this measure. For example, the evidence offered in Angrist and Evans (1998) for the United States and McDougall et al. (1999) for Canada reveals that parents in these counties have a preference for having a child of each sex. That said, son preference has been found in some immigrant and ethnic populations in these countries. Examining the sex ratio of births at different parities among mothers who have yet to give birth to a son, Abrevaya (2009) and Almond et al. (2009) report evidence of son preference within the East and South Asian communities of the US and Canada respectively. The likelihood of a male birth has remained at just above 51% in the US (Abrevaya 2009). However, when disaggregating 1971–2004 California birth data by race, Abrevaya (2009) finds that for Indian families, a third-born child is approximately 7% more likely to be male than the first-born child, conditional on all previous births being female. Similarly, for Canada, Almond et al. (2009) uncover missing women among first generation South East Asian-born immigrants, by analyzing 2001 and 2006 Canadian Census data. One very recent study of immigrant groups in Canada shows that multiparous women born in India were significantly more likely to have a male infant than their counterparts born in Canada (Ray et al. 2012). Related work has also documented in Canada and the US, gender differences in the care of young children and early child inputs that are important for child development (Baker and Milligan 2011).

To the best of our knowledge, only one study, that by Dubuc and Coleman (2007), focuses on the existence of missing women amongst immigrant populations in the UK. This study documents the overall upward trend in the SRB to India-born mothers living in the UK since the 1980s. Using time-series analysis, they find a sharp rise in the SRB, from 1.04 during 1969–1989 to 1.14 for third and later births after 1990. These figures closely mimic the experience of India. According to India's 2006 National Family Health Survey-1, the average SRB obtained for the overlapping period of 1978–1992 stands at 1.06, while in 2000, at 1.12. This escalating trend in the SRB coincides with the public availability and affordability of contemporary prenatal gender-selection technologies (Scholly et al. 1980; Dubuc and Coleman 2007).

2.1. Economic and cultural motives underlying male-biased sex-selection

Institutional and socio-economic factors may prompt parents to select sons. Current research emphasizes the failure of capital markets: the inability to save, to take on insurance and the lack of a coherent national pension system in most Asian countries which implies that the poor rely heavily on children for support in old age, especially sons (Chung and Das Gupta 2007). Patrilineality may be another factor: core productive assets, such as land, are traditionally passed on through the male line, while women may be given movable goods in the form of a dowry, placing families with daughters at an economic disadvantage. This makes women dependent on their sons as a form of informal familial insurance (Das Gupta et al. 2003). Yet, what is curious both in India and in China is that it is the affluent rather than poorest states that may discriminate most against girls, casting doubt on whether male-biased sex-selection is a consequence of economic necessity alone (Almond et al. 2009).

There is some research which examines the cultural value accorded to sons in some Asian societies. For example, Mencius, a Chinese philosopher wrote that *"A woman is to be subordinate to her father in youth, her husband in maturity and her son in old age"*. In Hinduism only a son can light a parent's funeral pyre (see Iyer 2002, pp. 40–42 for a more full discussion). Qian (2008) reveals the apparent inability of women in China to influence the household's sex-selection decision. Moreover, the persistence of fertility-related cultural norms within particular communities might cause both high fertility and gender selection. If the economic motive is the principal factor explaining gender-selection in the home country, then this practice should be significantly limited in the UK. If, however, cultural norms do indeed persist, it is not clear whether economic development would suppress this tendency.

2.2 Family size and gender composition of previous births

Abrevaya (2009) emphasizes that the incentives for gender selection depend not only on gender preferences but also on preferences for family size. Gender-selection incentives become stronger as a family approaches its own size limit. For instance, if a family has strong son-preference and wants only two children, then the data should show a higher percentage of boys among second births as families use stopping rules and possibly gender selection. More generally, the incentives for gender selection increase as the desired family size is reached (Retherford and Roy 2003) and the opportunity cost of having a child of the less-preferred gender increases. This suggests that son-biased gender selection will most likely manifest itself in unusually high male birth rates at later births and unusually high male-birth rates following daughters.

The sex-composition of previous children and the family's predetermined optimal size jointly determine its subsequent fertility behavior. This is known in the literature as the family's 'fertility stopping rule' (Ebenstein 2007; Larsen et al. 1998). Jha et al. (2006) reveal that in India, unnaturally high SRB occur for children born at higher birth parities, solely when the gender of all previous births is female. Further, recent studies show fewer missing women among Muslim households than among Hindu ones (Borooah et al. 2009; Almond et al. 2009).

We anticipate, therefore, that families with strict stopping rules - son preference compounded by a small desired family size - may resort to sex-selection earlier in the birth order, conditional on previous births being female (Retherford and Roy 2003). In this study we investigate whether the irregularities of high male birth rates at higher birth parities, and high male birth rates following daughters are present among different groups in the United Kingdom as well as whether skewed sex ratios present in a particular ethnic group are associated with their desire for smaller families. The use of higher parity and male birth percentages conditional upon previous gender has been considered in several previous studies of Asian countries (see, for example, Das Gupta and Bhat 1997, Gu and Roy 1995, Park and Cho 1995, Retherford and Roy 2003, Jha et al. 2006, and Ebenstein 2007) but has not been used before for the United Kingdom.

3. Data

We employ data from the General Household Survey (GHS), a nationally representative survey of individuals living in private households in the United Kingdom. It includes family and individual-level information such as wealth, economic status, educational attainment, country of birth, area of residence and family structure. We construct an independent cross-sectional dataset, for the years 1971–2006, [1] as the number of observations for immigrant families in each year is too small to yield precise estimates.

In the data, the year and month of birth is reported for each family member. Using this information, we code variables for the sex and birth parity of all children born. These are then linked to their parents, for whom we know the country of birth. These data provide us with both dependent and independent variables for the analysis. The full dataset consists of 1,091,629 observations at the individual level but we convert these to the household level, using household identifiers unique to each individual. When we include only English, Indian and Pakistani households, this comprises a total of 251,667 households in the dataset. Among them, 102,430 households have at least one child.

We define family size as completed fertility. We use this to proxy desired family size which is asked in demographic surveys to ascertain the number of children that people want rather than the number they actually have. Our survey includes questions on cohabitation prior to marriage, previous marriages and all live births, and was addressed to all women aged 16 to 49, except non-married women aged 16 and 17. Our key variable of interest is the sex at first birth, second birth, third birth and so forth and refers to all live births. If there are step-children living in the household, then this is not reflected in the birth variable. Adopted and stepchildren are included in the same family unit as their adoptive/stepparents. Foster-children, however, are not part of their foster-parents' family (since they are not related to their foster-parents) and are counted as separate non-family units. If the child is not living in the household currently or is deceased, it would still be included in the 'all live birth data' variable.

Table 1 presents descriptive statistics on the data. The first three columns present information on the gender of the first three children of the household, by parents' ethnicity. For example, among the 99,691 first children whose parents are both English, we observe 51,586 males and 48,105 females. The last two columns present the gender of the third child given the sex-composition of the previous two children. Although the parents' ethnicity and the gender of previous children capture some heterogeneity among households, they do not allow for further heterogeneity among the subgroups.

A first look at the dataset shows that parents' education and the family's predetermined optimal size seem to be determinants for the gender of the child: for example, 48.8% of Indian households with two daughters gave birth to their first son. However, this percentage increases to 60.9% for the best educated Indian households. Determinants such as education are included in the estimation procedure which we present in the next section. For that section, the ethnicity categories are reduced to three due to the low number of mixed-household observations. In Table 2, we present

Table 1 Descriptive statistics of children's gender by ethnicity

Ethnicity	First child: Male dummy	Second child: Male dummy	Third child: Male dummy	Third child: Male/First two children female	Third child: Male/One child male and one female
English	0.517	0.510	0.511	0.504	0.516
	99,691	76,351	30,660	7549	14,258
Indian	0.489	0.475	0.530	0.488	0.559
	1400	1131	598	205	272
Pakistani	0.541	0.475	0.507	0.504	0.508
	614	549	432	115	199
English-Indian	0.517	0.524	0.525	0.692	0.522
	503	382	158	39	67
English-Pakistani	0.533	0.460	0.482	0.571	0.455
	184	139	83	21	44
Indian-Pakistani	0.474	0.594	0.478	0.667	0.333
	38	32	23	3	15
All	0.517	0.510	0.512	0.505	0.516
	102,430	78,584	31,957	7932	14,855

Source: The means are calculated from data in the UK general household survey.

the distribution of children by ethnicity. There are 89.5% of English households which have three children or less, whereas these percentages for Indian and Pakistani households are 78% and 50.3% respectively.

In this paper we compare our findings to those from a different dataset taken from the National Family Health Survey of India (NFHS), a nationally representative survey of people living in private households in India from 2005–2006. We excluded twins, triplets and quadruplets from the dataset. The distributions of children of Indian residents in India and Indian residents in the UK do have important differences; for example Indian residents in India typically have larger families. If this is the case, it is possible that Indian immigrants in the UK may have a stronger preference for gender selection compared to Indian residents at a given parity. This can happen if and only if their preferences do not change due to relocation. However, it is possible that their preferences may change due to cultural contagion experienced in the host country.

In the analysis of our data from both the UK and Indian datasets, we use the following variables, whose definitions are listed in Table 3.

Table 2 Distribution of children by ethnicity

Number of children	English	Indian	Pakistani	English-Indian	English-Pakistani	Indian-Pakistani	Total
1	23,340	269	65	121	45	6	23,846
2	45,688	533	117	224	56	9	46,627
3	20,187	291	127	101	39	7	20,752
4	6914	165	111	34	27	8	7259
5	2220	86	73	10	9	6	2404
6+	1342	56	121	13	8	2	1542
Total	99,691	1400	614	503	184	38	102,430

Source: The numbers are calculated from data in the UK general household survey.

Table 3 Variable definitions

Variable	Definition
	Dependent variables
Boy	Indicator variable: Child born at a given parity is male
Family size	Proxy for desired family size: Number of children born per household
	Previous children and ethnicity variables
All girls	Indicator variable: All previous children are girls
Gender mix	Indicator variable: Previous children are a mix of boys and girls
Indian	Ethnicity dummy: Household is Indian (0,1): At least one parent born in India
Pakistani	Ethnicity dummy: Household is Pakistani (0,1): At least one parent born in Pakistan
English	Ethnicity dummy: Household is English (0,1): At least one parent born in the UK
	Individual and household variables
University	Dummy variable: A household is educated at the university level (0,1): At least one parent has first degree, higher degree or higher education qualifications
Secondary	Dummy variable: A household is educated at the secondary education level (0,1): At least one parent has achieved at least one of GCSE, O-level or A-Level qualifications
Never school	Dummy variable: A household never went to school (0,1): At least one parent never attended primary school
Employed	Dummy variable: A household is employed (0,1): At least one parent is employed
Wealthy	Dummy variable: A household is wealthy (0,1): Household owns a house
London	Region dummy that household resides in London (0,1)
West Midlands	Region dummy that household resides in the West Midlands (0,1)
East Midlands	Region dummy that household resides in the East Midlands (0,1)
North	Region dummy that household resides in the North (0,1)

Note: Most of the immigrants in our dataset live in the four regions listed above. All the other regions are included in the omitted variable category.

4. Methodology

4.1 The Determinants of male-biased sex selection

The focus of our study is on the determinants of male-biased sex selection. Since sex selective abortion is a latent variable, similar to other studies in this field, we proxy this with the likelihood of male births (Jha et al. 2006). The dependent variable is an indicator variable capturing the likelihood of a male birth; hence a Logit or Probit model of discrete choice is appropriate. Under the logistic distribution, the general form of the function linking child gender with a set of the determinants of sex selection is:

$$\Pr(Y = 1|\mathrm{x}) = \frac{e^{x'\zeta}}{1 + e^{x'\zeta}} = \Lambda(x'\zeta) \tag{1}$$

Where Y is a child gender dummy variable and x is a vector of the determinants of male-biased sex selection. ζ is the set of parameters that need to be estimated and $\Lambda(.)$ indicates the logistic cumulative distribution function. Such non-linear binary choice models are estimated based on the method of Maximum Likelihood and have several advantages over linear probability models (LPM), such as overcoming the 0–1 interval restriction imposed on fitted probabilities by the LPM, and constant marginal effects. It is acknowledged that the LPM can be an alternative model since the fitted probabilities are very close to 50 percent. However, it yields nearly identical results, so we choose to use the logit model. Abrevaya (2009) found nearly identical results by using LPM and probit estimation. For this comparison, he assumed that the gender of a child conditional on the previous sex composition is estimated using a probit model. Similarly, we assume that the gender of a child conditional on the previous sex composition is modelled estimated using a logit model. Finally, our prediction procedure gives almost identical results with logit and the LPM[2].

The descriptive statistics shown in Section 3 provide evidence that the interactions of socio-economic variables with gender composition of previous births should also be included, thus the term x'ζ in equation (1) can be written as follows:

$$x'\zeta = \sum\nolimits_{j=1}^{J} \beta_j X_{ij} + \sum\nolimits_{k=1}^{K} \sum\nolimits_{j=1}^{J} \gamma_{jk} \mathrm{P}reviousBirths_{ik} X_{ij} + \sum\nolimits_{t=1}^{T} \tau_t D_t \tag{2}$$

X_{ij} is a set of J-1 socio-economic covariates plus a constant vector which takes the value one for all observations which determine the likelihood of a male birth at a given parity, for household *i*. The J-1 variables capture the heterogeneity among individuals and households and their definition is given in Tables 4, 5 and 6. β_j is a set of J coefficients which need to be estimated. $PreviousBirths_{tk}$ is a set of K gender composition of previous births dummy variables. The omitted case is all previous children are male. In the case of the second child, this is a single dummy variable that takes the value 1 if the first child is female. In the case of the third child, this term is a set of two dummy variables: all previous children are female; and previous children are one male and one female. In this case 2J γ coefficients need to be estimated. D_t is a set of T time dummy variables included to adjust for the fact that the population may have different distributions in different time periods and τ_t is a set of T coefficients. Equation 1 (and 2) should be estimated for each ethnic group. Thus we reduce the ethnic groups to three and we present 3KJ+T coefficients.

In keeping with our discussion above, we estimate a logit model based on equation 2 for each ethnic group, but with the cost of a lack of a formal test of whether the coefficients differ among the groups. A frequent strategy in overcoming this issue is to use the full sample and interact all the equation variables with the ethnic dummy variables. Thus we can test for the difference between two coefficients across ethnic groups. Applying this strategy, equation 2 becomes:

$$x'\zeta = \sum\nolimits_{m=1}^{3} \sum\nolimits_{j=1}^{J} \beta_{jm} X_{ij} Eth_{im} + \sum\nolimits_{m=1}^{3} \sum\nolimits_{k=1}^{K} \sum\nolimits_{j=1}^{J} \gamma_{jkm} \mathrm{P}reviousBirths_{ik} X_{ij} Eth_{im} + \sum\nolimits_{t=1}^{T} \tau_t D_t \tag{3}$$

Where Eth_{im} is a set of three variables: a constant vector which takes the value one for all observations and the Pakistani and Indian dummy variables as defined in Table 2. The omitted variable is the English dummy variable and we can test if the differences in coefficients are statistically significant or not.

Given our previous discussion of the link between religion and optimal family size, we argue that sex selection may be greater among Indian rather than among Pakistani immigrants, as the proportions of Hindus relative to Muslims are 45% compared to 13% for Indian immigrants, while they are 1% of Hindus compared to 92% of Muslims for Pakistani immigrants (UK Census 2001). We expect this because some kinds of abortion are religiously prohibited in Islam, so Muslims might not use this as much (see Iyer 2002 for a detailed discussion). Second, Muslim families may be more tolerant of having daughters and larger families than their Hindu counterparts due to their lower practice of giving dowries at the time of marriage (Borooah et al. 2009). As most of the Pakistanis are Muslims (92%) and most of the Indians are Hindu (45%), we

Table 4 Estimation results from logit and poisson regression models

Dataset used:	*UK general household survey*			*India NFHS*
Dependent variable:	*Third child: Boy (1)*	*Second child: Boy (2)*	*Family size (3)*	*Third child: Boy (4)*
University	−0.021**	−0.012***	−0.134***	
	(0.0087)	(0.0036)	(0.0014)	
Secondary	0.017***	−0.0054	−0.244***	
	(0.0049)	(0.0034)	(0.010)	
Never-schooled	−0.059*	0.024	−0.0048	
	(0.035)	(0.016)	(0.018)	
Employed	0.0040	−0.0031	−0.168***	
	(0.013)	(0.0042)	(0.013)	
Wealthy	−0.015	−0.0026	−0.172***	
	(0.027)	(0.0025)	(0.013)	
Constant	−0.051	−0.185**		
	(0.091)	(0.081)		
Mixed gender interactions				
University	0.031			
	(0.020)			
Secondary	−0.016***			
	(0.0059)			
Never-schooled	0.072**			
	(0.036)			
Employed	0.0058			
	(0.019)			
Wealthy	0.015			
	(0.027)			
Constant	0.0028			
	(0.0060)			
All girls interactions				
University	0.042***	−0.0018		
	(0.015)	(0.0038)		
Secondary	−0.012*	0.010**		
	(0.0069)	(0.0048)		
Never-schooled	0.083**	−0.033		
	(0.036)	(0.035)		
Employed	0.0084	−0.0013		
	(0.013)	(0.0061)		
Wealthy	0.032	0.0017		
	(0.028)	(0.0084)		
Constant	−0.032***	−0.010		
	(0.0095)	(0.011)		
Number of observations	31,950	78,555	102,430	40,316
Likelihood ratio test: null hypothesis: No overdispersion			0.4	

Notes: Tables 4, 5 and 6 are one table but we separate the results into three tables. For columns one and two, we estimate the Logit model using the general household survey dataset. In column three, we estimate the Poisson Regression Model using the same dataset. In column four, we examine the preferences of Indian households that live in India using the Logit model and information obtained from the NFHS dataset. Regional and time dummy variables are included but not reported for brevity. Marginal effects are reported for all models (and standard errors associated with the marginal effects). Ethnicity-Region-clustered (for columns 1–3), heteroskedastic -robust standard errors are given in parentheses. *, ** and *** denote significance at 10%, 5% and 1% level respectively.

Table 5 Estimation results: coefficients of Indian interactions with the independent variables

Dataset Used:		***UK general household survey***		***India NFHS***
Dependent variable:	*Third child: Boy (1)*	*Second child: Boy (2)*	*Family size (3)*	*Third child: Boy (4)*
University	−0.234**	0.019	−0.486***	−0.091*
	(0.109)	(0.046)	(0.113)	(0.044)
Secondary	−0.095	0.062**	0.011	−0.0076
	(0.094)	(0.027)	(0.035)	(0.017)
Never-schooled	0.288**	0.028	0.349***	−0.0042
	(0.137)	(0.046)	(0.121)	(0.014)
Employed	−0.065	0.019	−0.070	
	(0.116)	(0.037)	(0.090)	
Wealthy	0.122	0.065*	0.500***	0.014
	(0.098)	(0.038)	(0.078)	(0.011)
Constant	−0.071	−0.076**	−0.040	0.012
	(0.068)	(0.034)	(0.128)	(0.014)
Mixed gender interactions				
University	0.214			0.119**
	(0.148)			(0.052)
Secondary	0.062			0.021
	(0.136)			(0.020)
Never-schooled	−0.327**			0.013
	(0.160)			(0.018)
Employed	0.046			
	(0.155)			
Wealthy	−0.027			−0.0086
	(0.142)			(0.014)
Constant	0.013			0.0019
	(0.098)			(0.017)
All girls interactions				
University	0.440***	0.032		0.223***
	(0.112)	(0.078)		(0.054)
Secondary	0.178	−0.086		0.037*
	(0.178)	(0.062)		(0.022)
Never-schooled	−0.124	−0.0091		0.011
	(0.213)	(0.066)		(0.020)
Employed	0.052	−0.031		
	(0.091)	(0.051)		
Wealthy	−0.261*	−0.035		0.023
	(0.151)	(0.071)		(0.016)
Constant	`0.203**	0.031		0.0059
	(0.103)	(0.060)		(0.018)

Notes: The coefficients for columns 1–3 should be interpreted as the difference in the probability of a male birth at a given parity between an Indian and an English household, conditional on all previous births. It is important to note that for column 4 the coefficients represent directly the preferences and not the difference from other ethnic groups. Regional and time dummy variables are included but not reported for brevity. Marginal effects are reported for all models (and standard errors associated with the marginal effects). Ethnicity-Region-clustered (for columns 1–3), heteroskedastic-robust standard errors are given in parentheses. *, ** and *** denote significance at 10%, 5% and 1% level respectively.

Table 6 Estimation results: coefficients of Pakistani interactions with independent variables

Dataset Used:	***UK general household survey***			***India NFHS***
Dependent variable:	*Third child: Boy (1)*	*Second child: Boy(2)*	*Family size (3)*	*Third child: Boy (4)*
University	0.310**	−0.034	−0.041	
	(0.091)	(0.044)	(0.106)	
Secondary	0.210***	0.032	−0.190***	
	(0.069)	(0.091)	(0.067)	
Never-schooled	0.199**	−0.039	0.474***	
	(0.088)	(0.071)	(0.082)	
Employed	−0.071	−0.123***	−0.039	
	(0.089)	(0.043)	(0.078)	
Wealthy	0.046	−0.047	0.420***	
	(0.060)	(0.060)	(0.056)	
Constant	−0.141	0.113**	0.476***	
	(0.092)	(0.052)	(0.135)	
Mixed gender interactions				
University	−0.435***			
	(0.119)			
Secondary	−0.156			
	(0.129)			
Never-schooled	−0.300***			
	(0.102)			
Employed	0.064			
	(0.060)			
Wealthy	0.032			
	(0.093)			
Constant	0.046			
	(0.073)			
All girls interactions				
University	−0.480***	0.088		
	(0.158)	(0.101)		
Secondary	−0.143	0.035		
	(0.107)	(0.110)		
Never-schooled	−0.219**	0.084		
	(0.108)	(0.081)		
Employed	−0.028	0.133		
	(0.096)	(0.097)		
Wealthy	0.129	0.030		
	(0.083)	(0.103)		
Constant	−0.047	−0.179		
	(0.162)	(0.113)		

Notes: The coefficients for columns 1–3 should be interpreted as the difference in the probability of a male birth at a given parity between a Pakistani and an English household, conditional on all previous births. Regional and time dummy variables are included but not reported for brevity. Marginal effects are reported for all models (and standard errors associated with the marginal effects). Ethnicity-Region-clustered (for columns 1–3), heteroskedastic-robust standard errors are given in parentheses. *, ** and *** denote significance at 10%, 5% and 1% level respectively.

expect a positive coefficient for the *IndianXAllgirls* variable. We thus expect a positive and significant relationship between *IndianXAllgirls* and $Pr(Boy_t{=}1)$ at the third or even second parity, depending on the strictness of their stopping rules: an Indian household, endowed with son preference and having only female children, may possibly ensure that the birth of their next child is male through sex-selective abortion. Due to their presumably larger desired family size, the coefficient on *PakistaniXAllgirls* should not be suggestive of sex-selective abortion at such parities. Moreover, we expect that neither *IndianXGendermix* nor *PakistaniXGendermix* will be positively related to the probability of a male birth, as both ethnicities are less likely to resort to sex-selective abortion at any parity, if they already have at least one child of the preferred gender.

We control for the average household education level and employment status, as highly educated and employed families tend to have fewer children. Further, highly educated households are typically better able to access and implement these new technologies (Chung and Das Gupta 2007).We thus anticipate high sex ratios for such son-preferring families. We further control for income effects such as wealth, measured by whether the household owns a house. Following Clark (2003), we further propose that home ownership for an immigrant household perhaps represents a step towards cultural assimilation within the British community, and which might possibly also lead to the erosion of certain practices from the home country including son preference.

As discussed above, a larger family size is likely to provide the desired number of sons, alleviating the need to resort to sex-selection as early as the third parity. Sex-selection among Pakistani immigrant families is therefore more likely to show up at higher than third order births. Hence we expect a statistically insignificant coefficient on *PakistaniXAllgirls.* If this is indeed the case, there is a way to test if Pakistani families have a larger desired family size than Indian families, when both are compared to the English base group.

Following the methodology of Borooah and Iyer (2004), we can examine the relationship between family size and ethnicity, using a Negative Binominal Regression Model and a Poisson Regression model. We can test if Pakistani families have a larger desired family size than Indian families, when both are compared to the English base group. Allowing for different coefficients for various ethnic groups, we estimate the following simple Poisson Regression Model:

$$P(Y|x) = \frac{e^{-\lambda i}\lambda i^{y}}{Y!}\,;\, \lambda_i = \exp\left(\sum_{m=1}^{3}\sum_{j=1}^{J}\beta_{jm}X_{ij}Eth_{im} + \sum_{t=1}^{T}\tau_t D_t\right) \tag{4}$$

4.2 Oaxaca-blinder decomposition: the decomposition of inter-ethnic differences in the likelihood of a male birth

The original Oaxaca-Blinder method of linearly decomposing inter-group differences in means into an ‘explained’ and ‘residual’ component has been extended to decomposing inter-group differences in probabilities, derived from models of discrete choice (Nielsen 1998). We use this method to disentangle the ‘explained’ and ‘unexplained’ components accounting for the difference in the average probability of a male birth at a given parity between ethnic groups.

Defining the likelihood at a given parity of a male birth under a Logit model as in equations 1 and 2, the average probability of a male birth for ethnicity r is:

$$\overline{Male^r} = \bar{P}\left(Z_i^r, \theta^r\right) = N_r^{-1}\sum\nolimits_{i=1}^{Nr}\Lambda\left(Z_i^r, \theta^r\right)$$

where Z_t^r is a vector of all variables in equation 2 and θ^r its associated vector of coefficients. N_r is the number of households per ethnic group r.

Taking the difference in the average probabilities of a male birth between group 1 and group 2, we define:

$$\overline{Male^1} - \overline{Male^2} = \left[\bar{P}\left(Z_i^1, \theta^1\right) - \bar{P}\left(Z_i^2, \theta^1\right)\right] + \left[\bar{P}\left(Z_i^2, \theta^1\right) - \bar{P}\left(Z_i^2, \theta^2\right)\right] \tag{5}$$

The first term in square brackets, in equation 5, represents the 'Ethnic' effect: it is the difference in the average probabilities of a male birth between the two groups resulting from inter-ethnic differences in unobserved endowments (as exemplified by differences in the coefficient vectors) conditional on all observable covariates. The second term in square brackets represents the 'Attributes' effect: it is the difference in average outcomes between two groups resulting from inter-group differences in attributes, when these attributes are evaluated using a common coefficient vector. The Oaxaca-Blinder decomposition results allow us to distinguish the weight of the 'Ethnic' and 'Attributes' effects in explaining outcome differences between the two groups.

4.3 Estimating the number of missing women in the United Kingdom

One contribution of our study is to estimate the number of missing women among a sample of immigrants in the UK, employing a methodology not previously used in the literature, which follows nicely from the Oaxaca-Blinder decomposition. First, we estimate equation 2 for English families, using a logit model with robust standard errors and generate the probability of a male birth at a given parity for each English household. Summing these probabilities over all English households gives us the actual number of male births at a given parity for English natives. We perform the same exercise for the sample of ethnic immigrants. However, if we estimate equation 2 for an ethnic group of immigrants using a logit model and we apply these estimates in the same equation but use the English covariates instead, this will generate the probability of a male birth at a given parity, if an English household behaved as if it were an immigrant household:

$$\Pr^{EI}\left(Male_i^{EI} = 1\right) = \frac{\exp\left(Z_i^{ENG}, \theta^{IMM}\right)}{1 + \exp\left(Z_i^{ENG}, \theta^{IMM}\right)} \tag{6}$$

where EI corresponds to English covariates and Immigrant coefficients. By summing the probabilities over all English households, we can find the predicted number of male births for a given parity among English families if they behaved as immigrants. The predicted percentage change in the number of male births for English families is:

$$\frac{\sum_{i=1}^{N_{ENG}}\left\{\Pr^{EI}\left(Male_i^{EI} = 1\right)\right\} - \sum_{i=1}^{N_{ENG}}\left\{\Pr^{ENG}\left(Male_i^{ENG} = 1\right)\right\}}{\sum_{i=1}^{N_{ENG}}\left\{\Pr^{ENG}\left(Male_i^{ENG} = 1\right)\right\}} x\,100$$

By a similar methodology, one can estimate the relevant prediction for immigrant families, had they behaved as the native English. It is worth noting that the main difference between this approach and the Oaxaca decomposition is that we assume that only the preferences (captured by the coefficients) will change for one of the two groups suggesting that the household attributes will remain the same as before. Hence we are careful to use the phrase "behaved as immigrants" rather than "had been immigrants" to describe their behavior.

5. Results

In Tables 4, 5 and 6 we present our econometric results based on equations 1 and 3 (or 1 and 2) for the Logit model and equation 4 for the Poisson Regression Model. For all the models, we report marginal effects and standard errors associated with the marginal effects. We control for regional and time effects but for brevity do not report them.

The standard errors are clustered at the level of ethnicity and geographic area. Clustering at the level of ethnicity allows for the possibility that the behavior of households originating from the same home country may be correlated due to ethnic-specific norms (Clark 2003). However, as cluster robust standard errors require the number of groups to go to infinity in order to be consistent, we choose to cluster on the interactions of the ethnicity groups reported in Table 1 and geographic area. This strategy allows for the English (or any other ethnic group) that live in London to behave differently compared to the English (or any other ethnic group) that live in other regions. Additionally, this increases our group numbers to 31 rather than 3.

For columns one and two, we estimate the Logit model based on equation 1 and 3 using the General Household Survey dataset. In column one, the model is estimated with the dependent variable being the third-child-male indicator variable for families with at least three children, whereas, in column two we use the male indicator variable at the second parity for families with at least two children. In column three, we estimate the Poisson Regression Model using the same dataset.

In column four, we estimate equations 1 and 2 using the information obtained from the NFHS dataset. Hence we examine the preferences of Indian households that live in India. The dependent variable for this case is the third-child-male indicator for families with at least three children. It is important to note that the coefficients represent directly the preferences and not the differences from other ethnic groups. So the first three columns in each table are the results for the UK; the last column is for India.

As English is the omitted category, the coefficients in Table 4 should be interpreted as: The probability of a male birth at a given parity for an English household, conditional on all previous births being male, mixed-gender or female. The coefficients in Tables 5 and 6 should be interpreted as: The difference in the probability of a male birth at a given parity between an Indian (or Pakistani) and an English household, conditional on all previous births. Similarly, the omitted variable in all three tables is all previous births being male. Consequently, the all girls (or mixed gender) interactions should be interpreted as: The difference in the probability of a male-birth at a given parity between households with all previous children being female (or a mixed set of males and females) and household with all previous children being male, depending on ethnicity.

Looking at the results of column 3, employed, wealthy and more educated English households prefer smaller families. Pakistani families have on average more children than English families. Uneducated and wealthier Pakistani households tend to have even more children. These results and the descriptive statistics we present in Table 2 provide evidence that the practice of sex-selective abortion among Pakistani immigrants may be prevalent among higher order births but unlikely to show up at second and third births as their 'male-preferring stopping rule' implies relatively low restrictions on family size. On the other hand, Indian households deviate from English households as follows: Uneducated households prefer larger families, whereas the best educated households prefer smaller families. Owning a house boosts the preferences for larger families. These results provide evidence that the practice of sex-selection among uneducated and wealthy Indian immigrants may be prevalent at the fourth parity. However, as we want to examine the differences in coefficients between ethnic groups, we give emphasis to the second and third parity since 89.5% of English household have three or less children. Regarding the choice of the PRM, the likelihood ratio test fails to reject the null hypothesis of no overdispersion, suggesting that the PRM is an improvement to the NBRM.

Comparing column 1 and column 2, we observe higher coefficients in absolute values for the first column which supports the hypothesis that households are engaging in sex-selection when they have their last child. Thus, it is reasonable to focus on the column 1 results for all English households and non-uneducated Indian households. The results in column 1 possibly provide evidence that Indian families are engaging more in sex-selection compared to English families. On average, uneducated English families seem to engage in sex-selection of male births in case they have two male children. Again, only uneducated English households engage in sex-selection of female births when they have two girls or a mixed group of boys and girls. However, these coefficients are relatively small.

Since we reduce our ethnic groups, it is possible that the English sex-selection arises from the fact that we define “English” as a family where at least one parent was born in the UK. To test that, we apply the model presented in column 1 on the “pure” English families only, and we find that for the uneducated English households the coefficient of the interaction of the university dummy with the dummy of households with two female children is 0.088 (near 0.083) even with the reduced sample. The standard error is 0.039 and it is again statistically significant at the 5% level. In contrast, the best-educated Indian households with two boys seem to undertake sex-selection compared to English households possibly in order to give birth to a daughter. Both uneducated English and best-educated Indian households seem to prefer a balanced household. This is consistent with the results in column 4 for Indians in the home country. For Indian families with two girls, all prefer a third male child. The best educated households have much higher incentives to indulge in sex-selection. These results are in line with the results in column 4 for the best educated households, but this is not evident for the remainder of households. It is possible that living in a developed country with publicly-provided health-care provides all residents with enough information about sex-determination techniques so that even the less educated families have this information.

Finally, although the best educated Indian families with a mixed-gender set of previous children in the home country use the techniques we describe above to have a second boy, this does not occur in the host country possibly because of varied cultural effects. Overall, the effect of education on the incentives for sex-selection among son-preferring families is clearly positive, endorsing our initial conjecture and the findings by Jha et al. (2006) that educated households are generally more aware of the availability of gender determination techniques and so more likely to use them [3]. Lastly, for the Pakistani coefficients reported in Table 6, we need to mention that they might not be in line with the results in Table 1, but this occurs because we do not present the regional dummy coefficients; for example, the interaction of *IndianXtwogirlXLondon* is 0.37, statistically significant, and over 60% of Indian households with the first two being female children live in London.

Regarding the Oaxaca-Blinder decomposition using equation 5, the difference in average male birth probabilities between the English (group 1) and the Indians (group 2) in the UK is −0.0166. The ethnic effect is found to be 86.95%. Intuitively, this represents the proportion of the average difference in male birth likelihoods which are explained by giving Indians an English ethnicity, conditional on both groups having the same attributes. The balance percentage is the attribute effect which represents the proportion of the relevant difference that is explained by giving Indians English attributes, conditional on both groups having the same ethnicity.

By using equation 6, we estimate the number of ‘missing women’ in the UK among Indian immigrants. The estimation is carried out for two time periods: 1970–2006 and 1993–2006. For our main results provided in Tables 4, 5 and 6 column 1, we assume that the coefficients are the same for all time periods. However, this is not true as we know that sex-determination was not possible before the 25th week of pregnancy until the late 1980s (Dubuc and Coleman 2007). Also in the period 1980–1992 consumers were uncertain about the quality of the “new product” and the prices of sex-determination procedures were significantly higher. In the last period consumers overcame their hesitation regarding the safety of the procedure and prices fell significantly. Consequently, the latter period 1993–2006 captures the peak of the use of the technology for sex-selective abortion.

We construct Table 7 where Panel (A) contrasts the numbers of boys and girls that Indian families actually gave birth to (Actual), with the corresponding number in the case where they are endowed with English characteristics (Predicted), noting how many Indian households would have essentially changed their fertility decision from boys to girls, holding the number of children constant. Equivalently, Panel (B) depicts the relevant figures for English families.

Panel (A) suggests that, for the whole period, there would have been 3.00% fewer boys among Indian families, had they been given English preferences or equivalently 3.37% more girls. In the subset, these percentages are slightly higher, consonant with enhanced sex-selective abortion availability after 1990.

From Panel (B), we can infer that, had English families had the Indian son-preferring characteristics, they would have had 5.39% less boys in the whole sample but a striking 11.42% additional boys during the final period (or equivalently 751 missing women). The results for the full sample come from the fact that the Indian coefficients were estimated for earlier periods, and because 9170 English households have two boys. This is about 15.6% more households compared to English households with two girls, as shown in Table 1.

Table 7 Estimating the number of Indian 'missing women'

Panels:	**Indian families: Panel (A)**				**English families: Panel (B)**			
Periods:	1970-2006		1993-2006		1970-2006		1993-2006	
Gender:	Boys	Girls	Boys	Girls	Boys	Girls	Boys	Girls
Actual:	400	356	192	176	15,758	15,060	6579	6319
Predicted:	388	368	186	182	14,908	15,910	7330	5568
% Changed:	−3.00	3.37	−3.13	3.41	−5.39	5.64	11.42	−11.88

Notes: The prediction percentages for the subsample (1993–2006) and their associated standard errors for Panel A and Panel B are: 50.60638 (0.57334) and 56.82986 (2.22453) respectively. The associated 95% confidence intervals for the boys of the subsample for Panel A and Panel B are: [182, 190] and [6768, 7892] respectively.

By accepting the evidence of Dubuc and Coleman (2007) regarding the *national* number of Indian third-parity male births (about 8000 for our critical period), we conclude that there are 914 missing women for that period [4]. Abrevaya (2009) finds "a conservative estimate" of 1,300 missing women among Indian families in the USA during a similar period 1992–2004. Our results seem consistent with the Abrevaya study if we take into account the difference between immigrant population sizes: there are approximately 1.5 times as many Indian Americans as there are British Indians (US Census and UK Census). Abrevaya's corresponding number for our case is 867, which belongs to the closed interval of [751,914].

Two important caveats that affect the validity of our findings need to be acknowledged. Firstly, despite the efforts made to expand our sample size by pooling over all survey years, any causal and quantitative claims that concern the determinants of sex-selection decisions deduced from the marginal effects of three-variable interaction terms, must be advanced very cautiously, as the sample size becomes rather small. Indeed, the best we can do is to provide qualitative conclusions about these effects. Secondly, even though we include all controls available in the dataset and account for time, region and ethnic fixed effects, it is still not possible to control for unobserved heterogeneity across households, individuals and time.

6. Conclusion

In this paper, data from the General Household Survey 1971–2006 are used to explore a gap in the literature: the relationship between sex-selection in the UK among immigrant families and the gender composition of previous births, conditional on various socio-economic characteristics.

This study briefly outlines the literature on missing women in the context of fully effective gender-determination technologies and son-preferring cultural attitudes. We contribute to this literature with new evidence of elevated male-birth likelihoods both at later births and conditional on female-dominated previous births, amongst Indian immigrant families. We also provide evidence to suggest that the larger desired family size among Pakistani families possibly mitigates their tendency to sex-select at relatively early parities. We also show that higher education levels enhance the incentives for sex-selection for immigrants while lower education levels enhance the incentives for sex-selection for the English. Furthermore, by using two independent datasets, we provide evidence that better educated Indian households whose first two children are male do have strong preferences for a balanced family. As expected, Indian households whose first two children are female do have stronger preferences for a son.

We show the role of cultural norms in fuelling the sex-selection decision among Indian immigrants, illustrating this with the Oaxaca-Blinder decomposition. We quantify the cultural differences between English and Indian families and show that they are instrumental in explaining the divergence in their average probabilities of male births. We attempt to quantify missing women in the UK among Indian immigrants during the period 1993–2006, using a methodology hitherto not employed in the literature and find that approximately 914 women might be missing. Since we take into account the balanced family preferences of better educated Indian households whose first two children are male, we avoid overestimating the predicted number of missing women.

In terms of the policy implications of our findings, a developed country like the UK can invest either through the formal education system or in traditional media channels in order to minimize the number of missing women. However, this is costly, especially during a global economic crisis. Our results suggest that the government can possibly focus on the subgroup of Indian university-educated households with two daughters. This would lower the cost to government and lower the number of missing women. The government may also consider imposing taxes on sex-selection techniques. Perhaps future research should focus on investigating this question when data on prices of sex-selection techniques become available to researchers.

In the meantime, our study of the UK contributes to existing research on other developed countries such as the US and Canada that missing women are not merely a phenomenon of the developing world.

Endnotes

[1]The data excludes 1997–1998 and 1999–2000, as we do not have enough information about immigrants for these years.

[2]For example, the number 387 in Table 7 becomes 388 when using the LPM results.

[3]Recall that we ignore the results of never-schooled households since column 4 provides evidence that fully uneducated Indian households have significantly larger families.

[4]We note that in our UK dataset, the national number of Indian third-party make births is 6579, and so we estimate that the number of missing women in our specific dataset is 751. However, we report above our missing women estimates relative to the Dubuc and Coleman number for ease of comparison between the two studies.

Competing interests
The IZA Journal of Migration is committed to the IZA Guiding Principles of Research Integrity. The authors declare that they have observed these principles.

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References
Abrevaya J (2009) Are there Missing Girls in the United States? Evidence from Birth Data. American Economic Journal: Applied Economics 1(2):1–34
Almond D, Edlund L, Milligan K (2009) Oh Sister, Where Art Thou? Evidence from Immigrants to Canada. National Bureau of Economic Research Working Paper No, The Role of Son Preference and Sex Choice, p 15391
Anderson S, Ray D (2010) Missing Women: Age and Disease. Review of Economic Studies 77:1262–1300
Angrist JD, Evans WN (1998) Children and their Parent's Labor Supply: Evidence from Exogenous Variation in Family Size. American Economic Review 88:450–77
Austin CR, Edwards RG (1981) Mechanism of Sex Differentiation in Animals and Man. Academic Press, New York
Baker M, Milligan KS (2011) Maternity Leave and Children's Cognitive and Behavioral Development. National Bureau of Economic Research Working Paper No. 17105. NBER, Cambridge
Bhat PNM, Zavier AJF (2003) Fertility Decline and Gender Bias in Northern India. Demography 40(4):637–657
Borooah V, Iyer S (2004) Religion and Fertility in India: The Role of Son Preference and Daughter Aversion, Cambridge Working Papers in Economics, 0436. Faculty of Economics, University of Cambridge
Borooah V, Do Q, Iyer S, Joshi S (2009) Missing Women and India's Religious Demography. World Bank Policy Research Working Papers, No. 5096, November
Chung W, Das Gupta M (2007) The Decline of Son Preference in South Korea: The Roles of Development and Public Policy. Population and Development Review 33(4):757–783

Clark W (2003) Immigrants and the American Dream: Remaking the Middle Class. Guilford Press, New York
Das Gupta M, Bhat PNM (1997) Fertility Decline and Increased Manifestation of Sex Bias in India. Population Studies 51 (3):307–31
Das Gupta M, Zhenghua J, Bohua L, Zhenming X, Chung W, Hwa-Ok B (2003) Why is Son Preference so Persistent in East and South Asia? Journal of Development Studies 40(2):153–187
Dubuc S, Coleman D (2007) An Increase in the Sex Ratio of Births to India-born Mothers in England and Wales: Evidence for Sex-Selective Abortion. Population and Development Review 33(2):383–400
Ebenstein AY (2007) Fertility Choices and Sex Selection in Asia: Analysis and Policy. Working Paper, Department of Economics. University of California, Berkeley, Available at: http://papers.ssrn.com/sol3/papers.cfm?abstract_id=965551
Ganatra B (2008) Maintaining Access to Safe Abortion and Reducing Sex Ratio Imbalances in Asia. Reproductive Health Matters 16(31 Suppl):90–98
Gini C (1955) Sulla Probabilita Chextermini di Una Serie Erratica Sieno Tutti Crescenti (o non decrescenti) Ovvero Tutti Decrescenti (o non crescenti) con Applicazioni ai Rapporti dei Sessi Nascite Umane in Intervalli Successivi e alle Disposizioni dei Sessi Nelle Fratellanze Umane. Metron 17(3–4):1–41
Gu B, Roy K (1995) Sex Ratio at Birth in China, With Reference to Other Areas in East Asia: What We Know. Asia-Pacific Population Journal 10(3):17–42
Hu L, Schlosser A (2012) 'Trends in Prenatal Sex Selection and Girls' Nutritional Status in India' CESifo Economic Studies. Univ. Press, Oxford, ISSN 1610-241X, ZDB-ID 20987614. - Vol. 58.2012, 2, p. 348–372
Iyer S (2002) Demography and Religion in India. Oxford University Press, Delhi
James WH (2000) Analysing Data on the Sex Ratio of Human Births by Cycle Day of Conception. Human Reproduction 15:1206–1207
Jha P, Kumar R, Vasa P, Dhingra N, Thiruchelvan D, Moineddin R (2006) Low Female[corrected]-to-Male[corrected] Sex Ratio of Children Born in India: National Survey of 1.1 million households. Lancet 367(9506):211–218
Johannson S, Nygren O (1991) The Missing Girls of China: A New Demographic Account. Population and Development Review 17:35–51
Larsen U, Chung W, Das Gupta M (1998) Fertility and Son Preference in Korea. Population Studies 52:317–325
McDougall J, DeWit DJ, Ebanks GE (1999) Parental Preferences for Sex of Children in Canada. Sex Roles 41:615–626
Nielsen HS (1998) Discrimination and Detailed Decomposition in a Logit Model. Economics Letters 61:115–120
Park CB, Cho N-H (1995) Consequences of Son Preference in a Low-Fertility Society: Imbalance of the Sex Ratio at Birth in Korea. Population and Development Review 21(1):59–84
Pollard GN (1969) Factors Influencing the Sex Ratio at Birth in Australia, 1902–65. Journal of Biosocial Science 1:125–144
Qian N (2008) Missing Women and the Price of Tea in China: The Effect of Sex-Specific Earnings on Sex Imbalance. Quarterly Journal of Economics 123(3):1251–1285
Ray JG, Henry DA, Urquia ML (2012) Sex Ratios among Canadian Liveborn Infants of Mothers from Different Countries. Canadian Medical Association Journal. June 12; 184(9):E492–E496.
Retherford RD, Roy TK (2003) Factors Affecting Sex-selective Abortion in India and 17 Major States, *National Family Health Survey Subject Reports,* No. 21. IIPS, Mumbai
Scholly TA, Sutphen JH, Hitchcock DA, Mackey SC, Langstaff LM (1980) Sonographic Determination of Foetal Gender. American Journal of Roentgenology 135:1161
Sen A (1992) Missing Women. British Medical Journal 304:586–587
Sen A (2003) Missing women revisited. British Medical Journal 327:1297–8
Shuzhuo L (2007) Imbalanced Sex Ratio at Birth and Comprehensive Intervention in China. Prepared for 4th Asia Pacific Conference on Reproductive and Sexual Health and Rights, Hyderabad, India, 29-31 October
UK Census (2001) Ethnic Groups by Religion 2001., Available at: www.statistics.gov.uk/statbase/expodata/spreadsheets/d6891.xls
Visaria PM (1967) The Sex Ratio of the Population of India and Pakistan and Regional Variations during 1901–61. In: Bose A (ed) Patterns of Population Change in India 1951–61. Allied Publishers, Bombay, pp 334–371
Zeng Y, Ping T, Baochang G, Yi X, Bohua L, Yongping L (1993) Causes and Implications of the Recent Increase in the Reported Sex Ratio at Birth in China. Population and Development Review 19(2):283–302

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