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Working Paper

Public Goods Access and Juvenile Sex Ratios in Rural India: Evidence from the 1991 and 2001 Village Census Data

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Contents

Abstract		v
I.	Introduction and Objective	1
II.	Trends	2
III.	Interstate Differences	8
IV.	Analytical Framework	13
V.	Data	14
VI.	Availability of Health and Water Infrastructure	15
VII.	Empirical Model	17
VIII.	Empirical Results	20
IX.	Empirical Results: Quantile Regression Estimates	24
X.	Conclusion	29
References		31

Abstract

We use village level data from the 1991 and 2001 Indian Censuses to examine how the availability of health facilities and safe drinking water at the village level affect juvenile sex ratios. In addition to controlling for village fixed effects in our estimating equation of the juvenile sex ratio, we also allow villages to be heterogeneous in terms of how their juvenile sex ratios respond to the availability of health facilities and safe drinking water. A key result we obtain is that although the presence of public health facilities does not exert a positive, significant effect on juvenile sex ratios on average, they do so in villages where the problem of discrimination against girls is most acute, i.e., in villages at the 0.10 and 0.25 quantiles of the conditional juvenile sex ratio distribution. Thus public policy can be an effective tool in improving gender balance in cases where it is most needed.

I. Introduction and Objective

A key and unusual demographic feature of India is an imbalance in the sex ratio. The ratio of females to males has been steadily declining for much of the last century. The juvenile sex ratio—the ratio of females to males aged 0–6 years—has been declining even more sharply. While most countries around the world have a small imbalance in their juvenile sex ratios for biological reasons (i.e., there is a biological tendency for more male than female babies to be born to compensate for the slightly higher risk of mortality among newborn boys), the imbalance in India is acute, and is indicative of prenatal selection and excess female infant and child mortality. Both in turn reflect a strong cultural preference for sons over daughters.[1] Some estimates put the number of "missing females" (i.e., unborn girls) in India as high as 37 million (Sen 2003). The low and declining juvenile sex ratio in the country is a matter of grave policy concern, not only because it violates the human rights of unborn and infant girls but also because it deprives the country of the potential economic and social contribution of these "missing women."

To combat this scourge, policy makers need to have a better understanding of the socioeconomic determinants of juvenile sex ratios. Our paper contributes to the existing literature on gender imbalance in several ways. Most existing studies at the national level have used data from the 1981 or the 1991 Census, aggregated to either the state or the district level. Since juvenile sex ratios have changed dramatically in the last two decades, it is important for policy purposes to use the most current data at the most disaggregated level. In this paper, we first use village-level data from the 1991 and 2001 Censuses to examine the variation in juvenile sex ratios across more than 500,000 villages in the country over two census years.[2, 3] Second, we focus explicitly on interventions that are

[1] Son preference can also lead to differential contraceptive use depending on the sex composition of existing children. While this will affect the relationship between family size and sex composition of children (for example, of the families with only one child, a larger proportion can be expected to have a male child as opposed to a female child), it will not affect the sex ratios in the population at large.

[2] We limit ourselves to village-level analysis primarily because of data constraints. While data on juvenile sex ratios are available for both villages and towns from the Indian Census, detailed information on health facilities are only available for all the villages in the country. Moreover, the presence of a heath facility provides a measure of access to health care services that is much more meaningful in the case of villages as opposed to towns.

[3] However, the discussion on juvenile sex ratio patterns and trends in India and in other countries in Sections I–III, which is obtained from secondary published data, relates to both rural and urban areas.

amenable to policy manipulation, in particular, the availability of health facilities and health infrastructure at the village level. Finally, we allow for considerable heterogeneity across villages in our estimates of the impact of policy variables on juvenile sex ratios. We find that the largest effects of new health facilities are in the lower quantiles of the initial sex-ratio distribution. The implication is that public policy can be an effective tool in improving gender balance in cases where it is most needed.

II. Trends

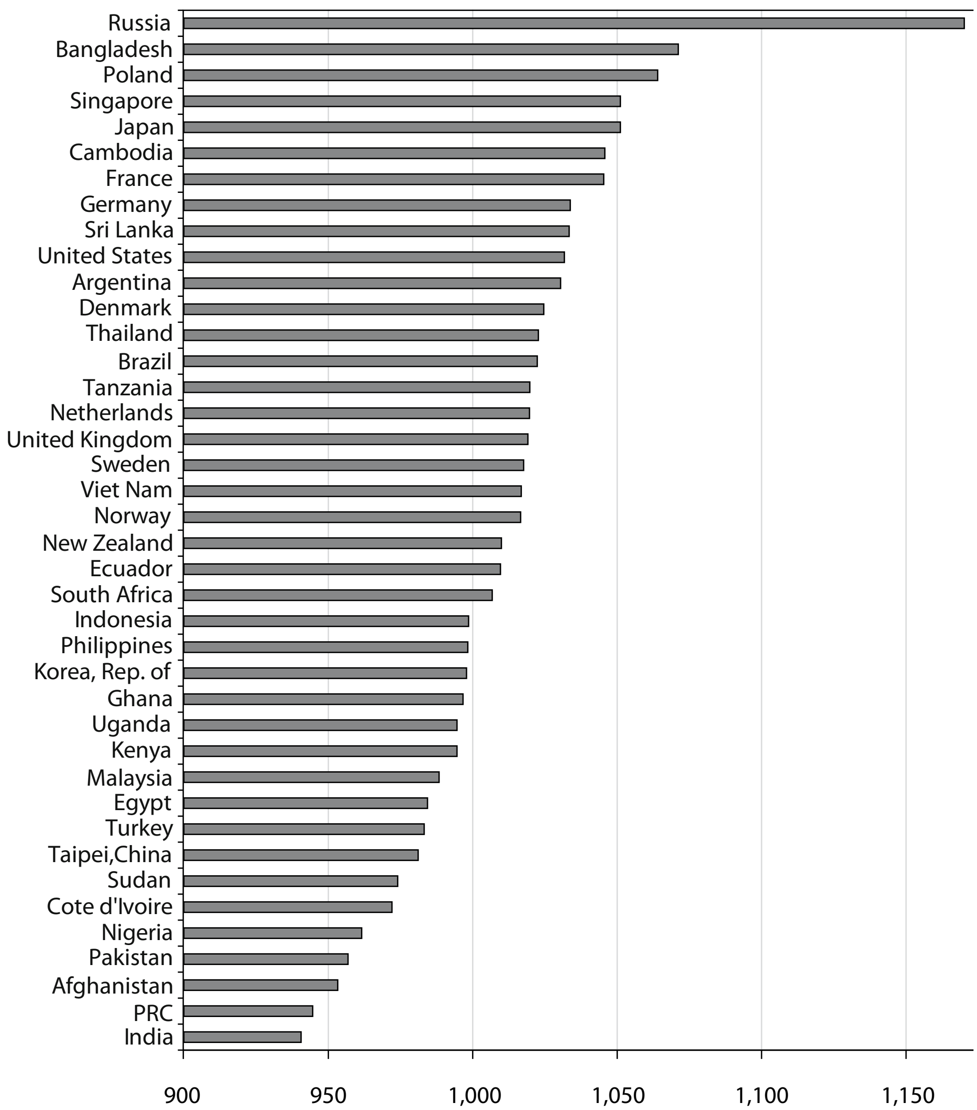
A low and declining sex ratio is an endemic feature of several countries in East and South Asia. The problem is most pronounced in India, which has the lowest sex ratio of any country in the world, lower than that in the People's Republic of China (PRC) and Pakistan (Figure 1). Further, India has seen its aggregate sex ratio decline through much of the 20th century—from 963 females per 1,000 males in 1901 to 933 females per 1,000 males in 2001 (Figure 2).

The overall sex ratio is determined by several factors such as age-specific differences in mortality and migration rates across males and females. For this reason, sex ratios for infants and children are better measures of the differential treatment of males and females in a society. No country in the world has a sex ratio at birth that is 1,000 or more, since there is a biological tendency for more boys to be born than girls. It is thought that this is nature's way of compensating for the slightly higher risk of death among newborn boys relative to newborn girls. As with the aggregate sex ratio, India has one of the lowest sex ratios at birth and one of the lowest juvenile sex ratios in the world (Figures 3 and 4).

The *Census of India* (Registrar General of India, various years) reports the juvenile sex ratio as the number of girls per 1,000 boys in the 0–6 age group. Census figures indicate that this number has declined even more sharply than the overall sex ratio—from 964 in 1971 to 927 in 2001 (Figure 2). Even in decades when the overall sex ratio increased slightly (1971–1981 and 1991–2001), the juvenile sex ratio continued to fall. These trends probably result from three factors: sex-selective abortions (feticide) based on prenatal ultrasounds, excess female (relative to male) infant and child mortality, and differential contraceptive use depending upon the sex composition of existing children. [4] All three in turn reflect a strong cultural preference for sons over daughters.

[4] In a widely cited paper, Oster (2005) argued that biology—not parental preference for boys—was largely responsible for low juvenile sex ratios, at least in the PRC. According to her, the high ratio of male to female births in that country were the result of a high population incidence of the hepatitis B virus, which leads to a higher likelihood of a male birth. However, based on further research, Oster et al. (2008) subsequently retracted her initial findings and concluded that hepatitis B could not explain the skewed sex ratio in the PRC.

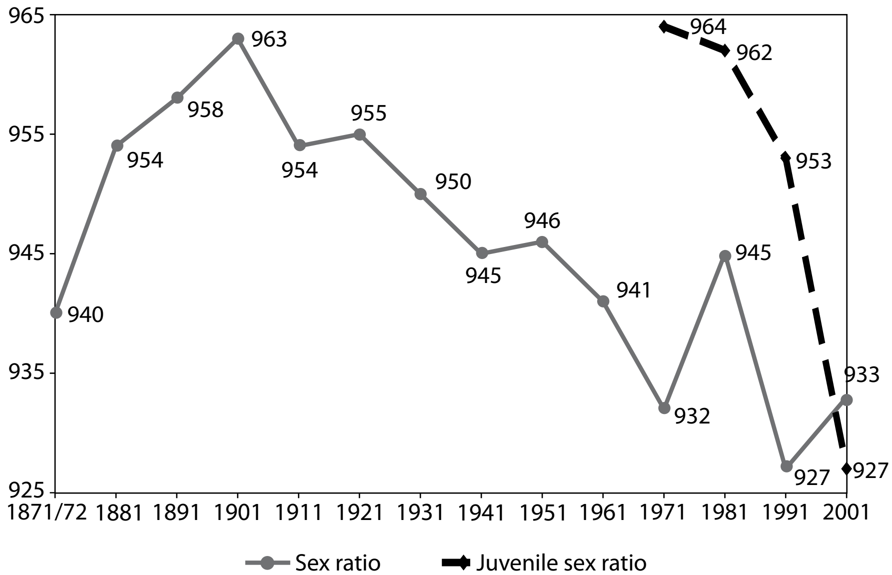
Figure 1: Number of Females per 1,000 Males of all Ages in Selected Economies, 2008 Estimates



Source: International Data Base (online) (US Census Bureau 2008), available: www.census.gov/ipc/www/idb/faq.html, updated December 2008.

Figure 2 is based on national averages and masks the enormous variation across different population subgroups. The 1991 and 2001 Censuses publish village-level juvenile sex ratios. Kernel density plots using village-level data for these years shows that the entire distribution of sex ratios shifted to the left during the decade of the 1990s (Figure 5). While 60% of the villages in the country had a juvenile sex ratio of less than 1,000 in 1991, as many as 65% of the villages were below this threshold in 2001. The median value of the village-level juvenile sex ratio declined from 947 to 933 over the same period.

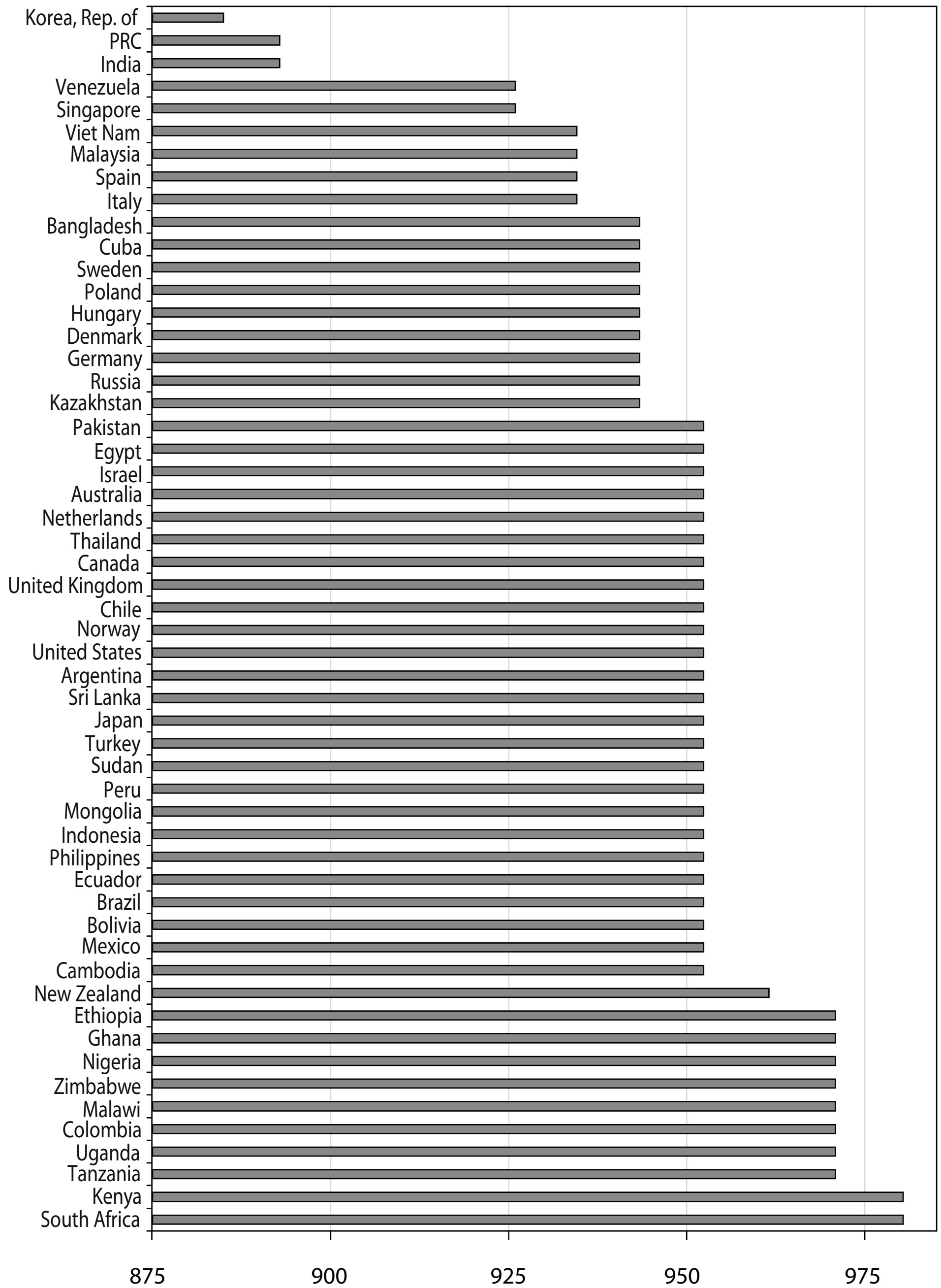
Figure 2: Aggregate and Juvenile Sex Ratio in India, 1871–2001



Note: The Census of 1871/72 was confined to the Old British Provinces and the former Princely State of Mysore. The figures for 1881–2001 pertain to the territory as existing at the 2001 Census.

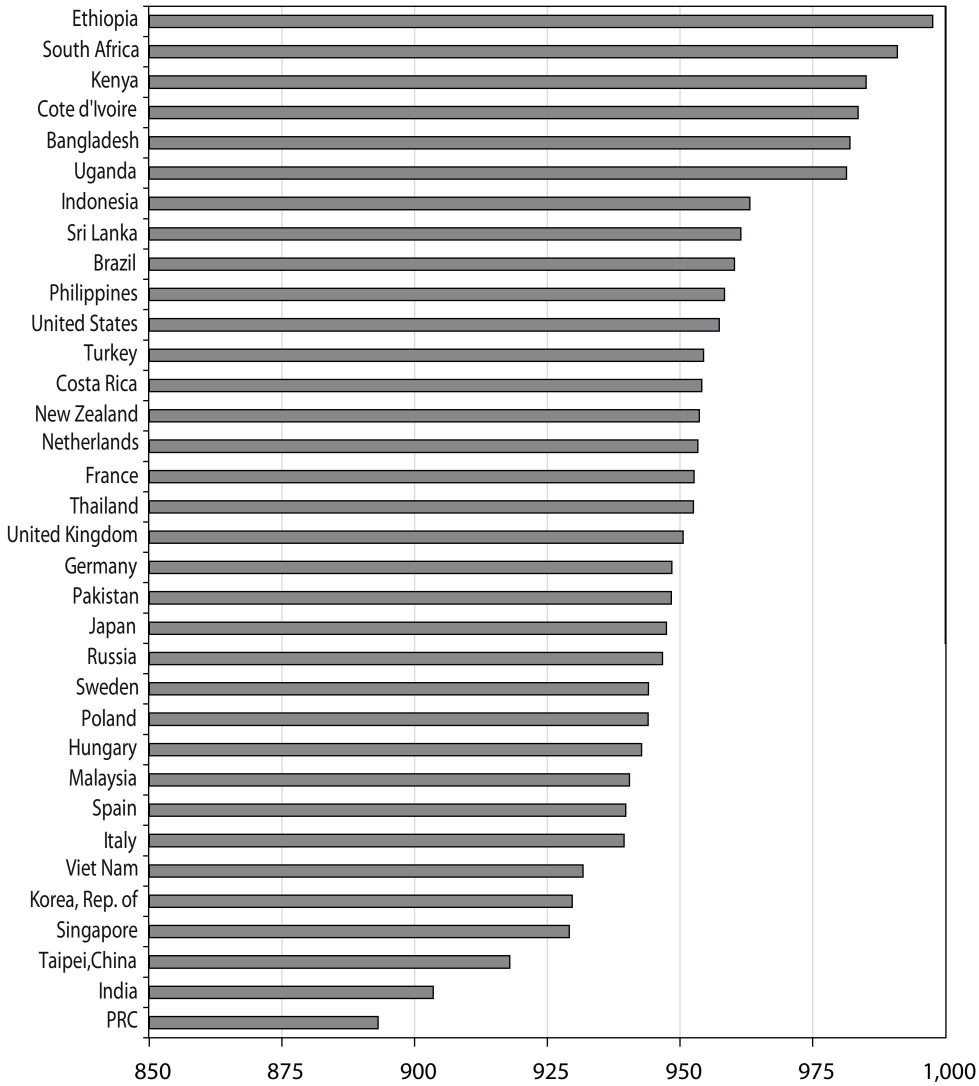
Sources: Hatti and Shekhar (2004), *Census of India* (Registrar General of India 1991 and 2001).

Figure 3: Number of Females Born for Every 1,000 Males Born, by Country, 2001



Source: *The World Factbook 2006* (CIA 2006).

Figure 4: Number of Females Aged 0–4 per 1,000 Males Aged 0–4 Years in Selected Economies, 2008 Estimates



Source: International Data Base (online) (US Census Bureau 2008), available: www.census.gov/ipc/www/idb/faq.html, updated December 2008.

Figure 6, which shows the kernel density plot for *changes* in the village-level juvenile sex ratio between 1991 and 2001, confirms the overall decline in juvenile sex ratios during the decade of the 1990s. The distribution has mean and median values of about –13, with 52.4% of villages in the country having experienced a decline between 1991 and 2001.

Figure 5: Juvenile Sex Ratios, Indian Villages, 1991 and 2001 (females per 1,000 males aged 0–6 years)

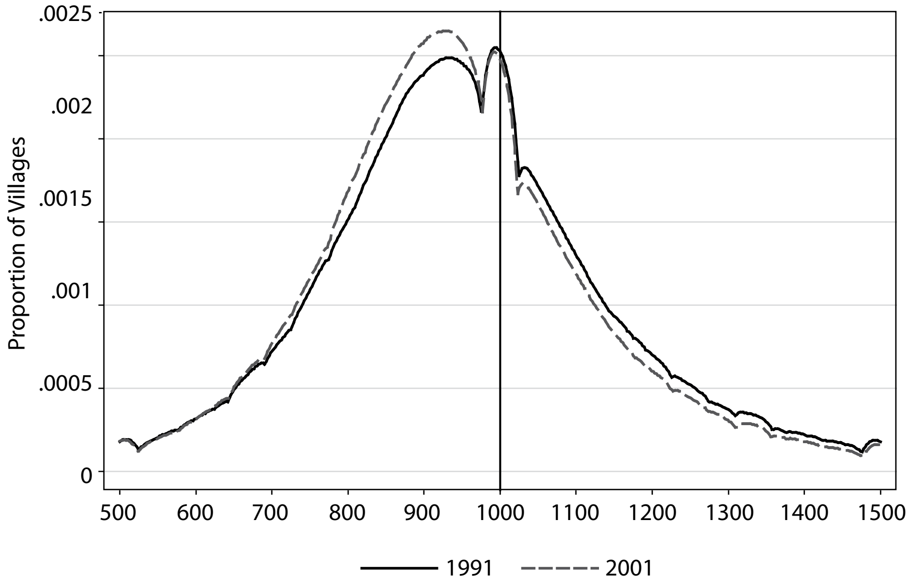


Figure 6: Distribution of Village-level Changes in Juvenile Sex Ratio, 1991–2001

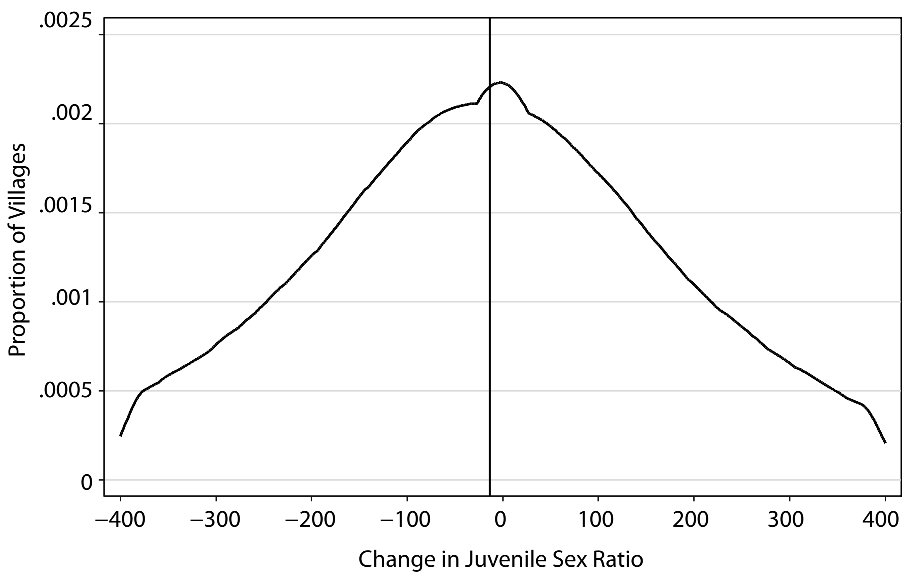
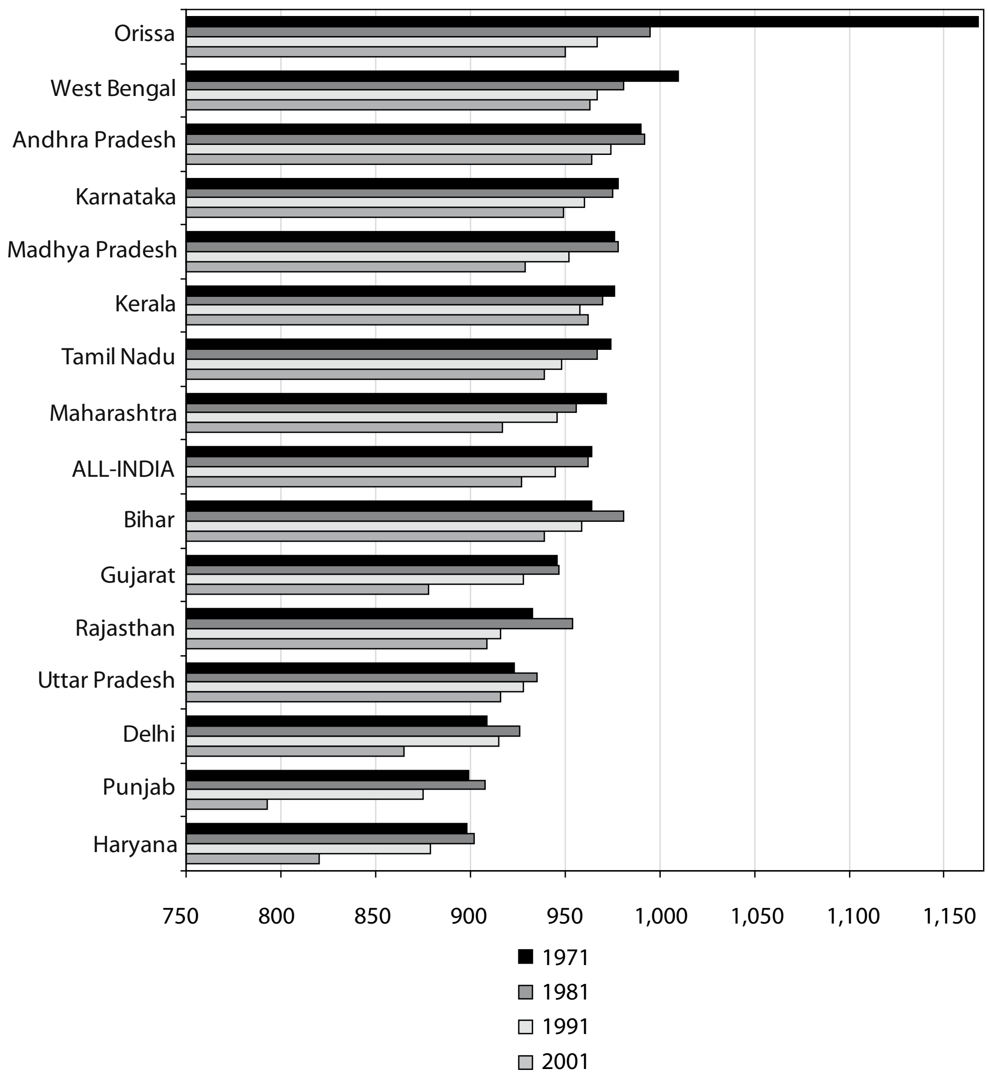


Figure 7: Juvenile (0–6 years) Sex Ratio, by State, 1971–2001



Source: Datanet India Pvt. Ltd. (2009), available: www.indiastat.com/demographics/7/sexratio/251/stats.aspx.

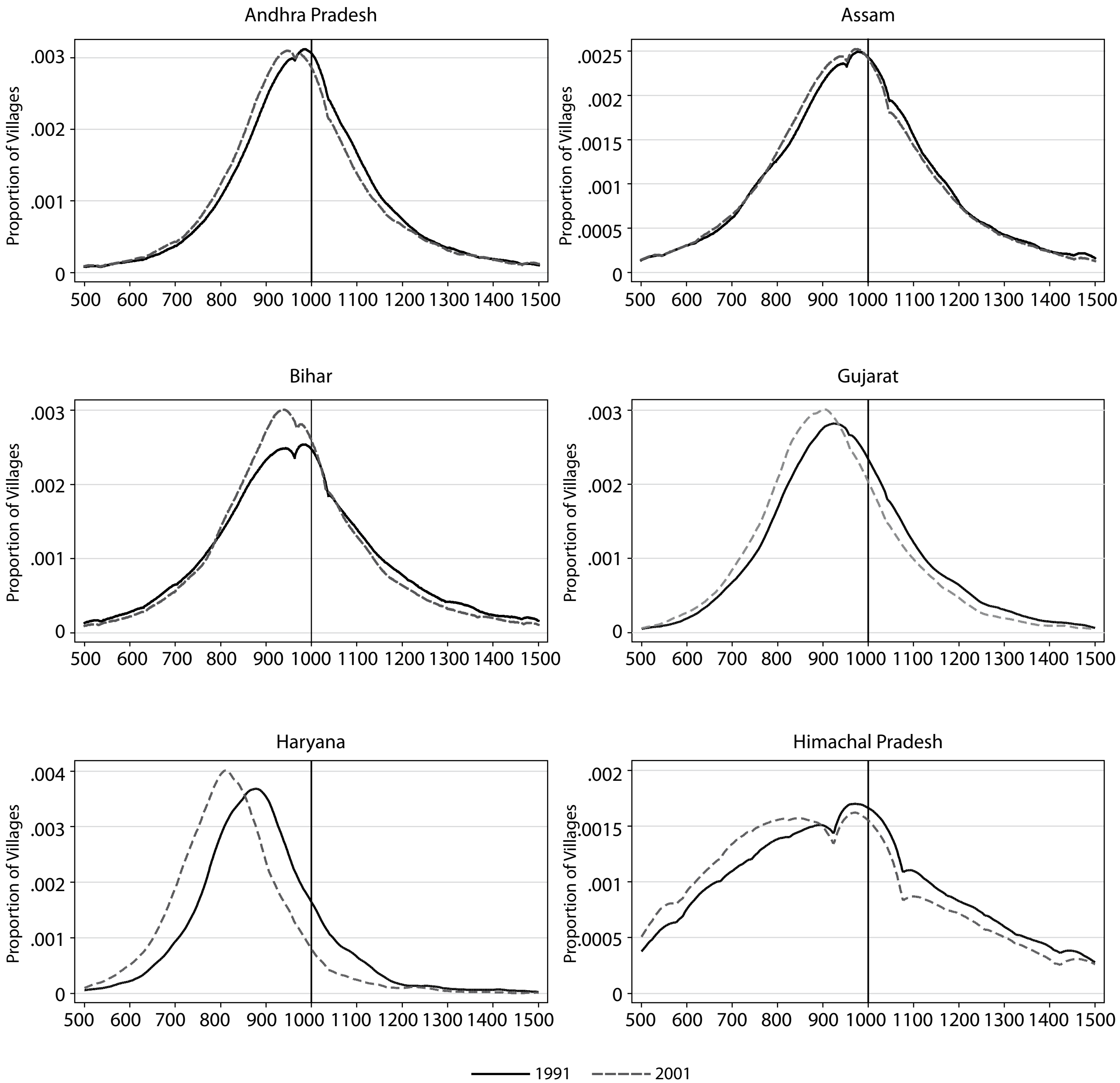
III. Interstate Differences

Much of the literature on Indian sex ratios has focused on a North-South dichotomy in sex ratios (Sopher 1980, Dyson and Moore 1983, Agarwal 1986, Sen 2003). This literature has argued that the northern and western regions of the country not only have more rigid norms of female seclusion, particularly related to participation in economic activities outside the home, but also have marriage customs such as large dowry payments that make girls an economic burden on the family. In contrast, women in the South and the East enjoy greater autonomy and higher social status within the family. These cultural differences manifest themselves in the form of lower sex ratios at birth in the northern and western states than in the southern and eastern states. The data are broadly consistent with these predicted patterns, but the picture is more nuanced, as we discuss below.

Figure 7 shows the juvenile sex ratio for 15 major states in the country at four points in time: 1971, 1981, 1991 and 2001. We find that Punjab and Haryana consistently had the lowest juvenile sex ratio of any state in the country over this entire period despite experiencing some of the largest relative declines in the ratio (of 9–12%) between 1971 and 2001. Andhra Pradesh and West Bengal had the highest juvenile sex ratios in the country, with little change over the period 1971–2001. We also see that the phenomenon of low juvenile sex ratios has become more pervasive over time. States such as Orissa, Gujarat, Maharashtra, and even Tamil Nadu, which had more balanced sex ratios than the northern states, have seen large declines between 1971 and 2001. Orissa saw the sharpest decline in the ratio (about 20%) of any state in the country. Indeed, every major state in the country other than Kerala shows a declining juvenile sex ratio between 1991 and 2001.

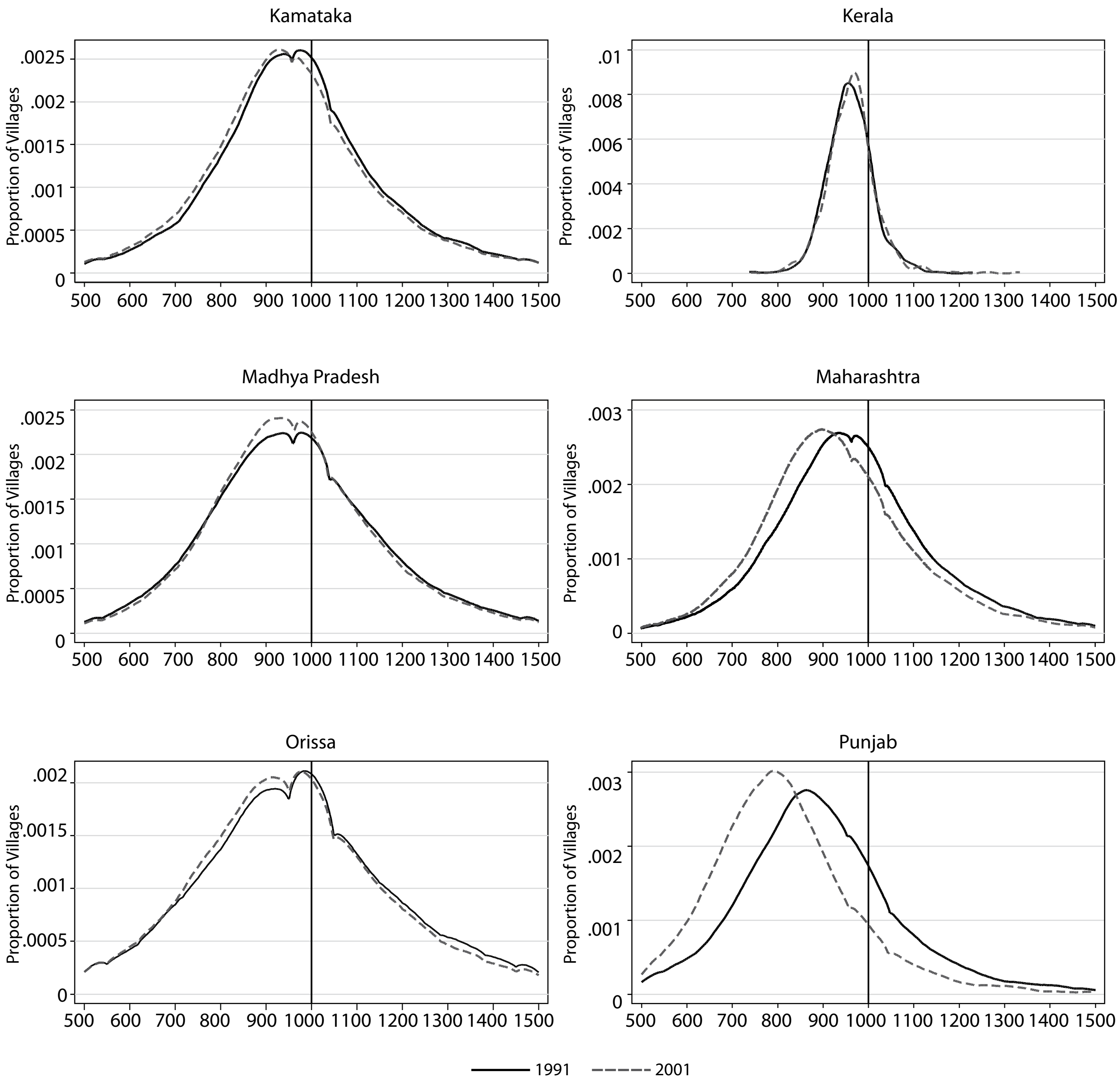
Figure 8 displays state-wise kernel density plots of village-level juvenile sex ratios in 1991 and 2001. These demonstrate the very large variations within each state. Notice, for instance, that while the median juvenile sex ratio in Punjab was less than that in Andhra Pradesh by 157 females per 1,000 males (Figure 9), the distribution in Punjab has a long right tail, which means that there were more than 1,500 villages in Punjab (about 14% of all villages in the state) where the juvenile sex ratio in 2001 was greater than 973 (the median value of the juvenile sex ratio in Andhra Pradesh). Likewise, there were 3,266 villages in Andhra Pradesh (also 14% of all villages in the state) that had a juvenile sex ratio less than 816 (the median value of the juvenile sex ratio in Punjab). Thus, juvenile sex ratios are marked by large variations both within and across states. In the following sections we examine the determinants of juvenile sex ratios using village-level data for all the Indian states.

Figure 8: Kernel Density Plots of Juvenile Sex Ratios, by State, 2001 (females per 1,000 males aged 0-6 years)



continued.

Figure 8b: *continued.*



continued.

Figure 8c: *continued.*

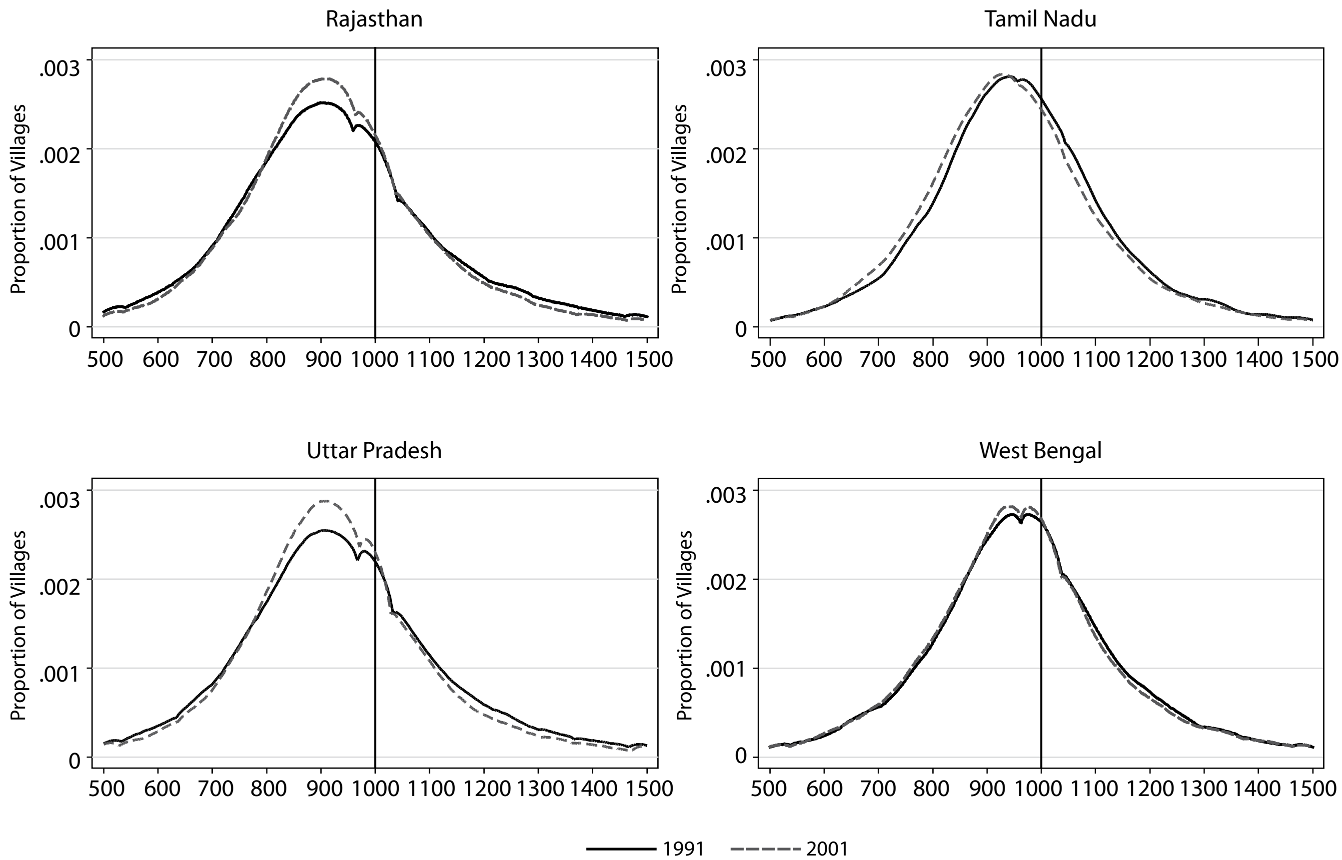
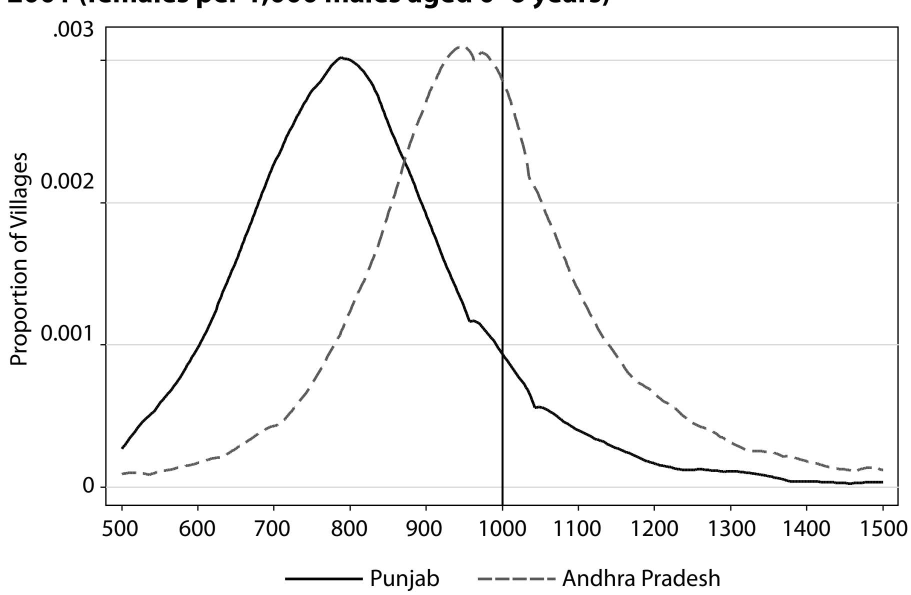


Figure 9: Juvenile Sex Ratios, Punjab, and Andhra Pradesh Villages, 2001 (females per 1,000 males aged 0–6 years)



IV. Analytical Framework

We consider a unitary model of the household in the sense that all decision makers share the same preferences. Utility depends on the number of healthy girls and boys and consumption of other goods:

$$U(c, b, g) = u(b, g) + v(c) \tag{1}$$

The consumption good, *c*, is used as a numeraire, and *b* and *g* refer to the numbers of healthy boys and girls.

Household utility is increasing, and marginal utility decreasing in all three arguments. We assume households preferences are biased toward male children.[5] In particular, we assume that, if a household has *n* children, then for any integers n_1, n_2 and *n*, with $n_1 + n_2 = n$ and $n_1 > n_2$, $u(n_1, n_2) > u(n_2, n_1)$.

Household income and other parent characteristics (such as education) are exogenously given. We denote these characteristics by the vector **x** and the costs of raising children are given by $f(\mathbf{x})(b+g)$, where *f* is some increasing function. This captures the idea that higher income households or those with more education may invest more in child quality (Becker and Tomes 1976). If this quality increases rapidly with income, we could observe lower fertility rates for these households.

For simplicity, we assume two periods. In the first period, couples determine family size (this may take into account average rates of morbidity and mortality). In the second period, there is some probability π that each child will fall sick. A sick child can be treated at some cost *t*. Given treatment costs, families take treatment decisions in the second period to determine the composition of children in their family. The gender composition of children in the family in the two periods is given by (b_1, g_1) and (b_2, g_2).

Family access to public health programs and health infrastructure can lower both the probability of falling sick and the treatment costs. Clean water lowers morbidity and principally influences π, while health centers and village health workers (in addition to providing health information, birth control and immunization services) lower *t*. Roads, bus services, and telephones can also lower health care costs through improved communication and better access to health facilities.

Morbidity rates and treatment costs are likely to interact in complex ways. If treatment costs are prohibitive, high levels of π will tend to increase fertility, and gender bias is likely to further increase family size. The gender composition of children across the

5 The cultural preference for sons might be driven by the fact that sons, not daughters, are largely responsible for the care of elderly parents in India. Or it could be shaped by traditional property inheritance laws that favor sons over daughters.

two periods in this case is likely to be similar. As treatment costs fall, gender bias may cause parents to treat male children more frequently than female children, and the sex composition of children across the two time periods may differ significantly, with lower sex ratios in period 2. At zero treatment costs, it becomes more likely that both types of children are treated and the juvenile sex ratio improves.

It should be noted that, by focusing only on differential treatment (and the resulting mortality differential) between male and female children, the analytical model sketched above abstracts from other factors that also contribute to lower juvenile sex ratios, such as the parental decisions to selectively terminate pregnancies that might have resulted in the birth of a girl. It is certainly possible that the availability of health facilities in a village reduces the parental cost of sex-selective abortions and thus causes the juvenile sex ratio to worsen. An analytical model that captures these different parental decisions is likely to be quite complex and is beyond the scope of this paper. Moreover, empirically disentangling the effects of health facilities on juvenile sex ratios through the different parental decisions would require very fine data that are not readily available.[6] However, it is unclear that our inability to do so is much of a disadvantage. Of the health facilities that we consider in this paper, only the community health centers are likely to possess ultrasound (sex determination) equipment, and fewer than 2% of the villages in our sample have a community health center.

The following section estimates empirically the relationship between public good access and juvenile sex ratios. We use village-level data for infrastructural facilities that are located within a village. These are proxies for the actual cost of treatment and we examine whether the presence of these goods within a village improves the ratio of girls to boys.

V. Data

Although the discussion so far has pertained to the juvenile sex ratio in both rural and urban India, the analysis undertaken in this paper is based on villages owing to data availability, and because a key relationship we want to examine—that between access to health facilities and juvenile sex ratios—makes sense only in the context of villages where the presence of health facilities represent a meaningful measure of access to health care services. The paper combines two types of village-level census data: the Village-Level Amenities Data (VLAD) and the Primary Census Abstract (PCA). As its name indicates, the VLAD include information on the availability and distance to over 100 different amenities in a village, including several types of health, education and physical

[6] One admittedly crude way to do so would be to work with juvenile sex ratios for different ages. For example, the juvenile sex ratio at age 0 could be used to focus on the effects of health facilities on prenatal sex selection. However, this would still confound the effects of female mortality of girls less than a year old.

infrastructural facilities. The PCA reports data on population by age, caste, and working status. Both of these data sets were merged, village by village, for every state in the country.

In the next stage, villages were matched across the 1991 and 2001 Censuses to create a panel data set at the village level. Although there are over 600,000 villages in India, several of these are not inhabited, and complete VLAD and PCA data are available only for about 500,000 villages. A little fewer than 400,000 of these villages could be matched accurately across 1991 and 2001. Our panel consists of these matched villages.

VI. Availability of Health and Water Infrastructure

Since this paper is primarily concerned with measuring the effect of social infrastructure on village-level juvenile sex ratios, we begin with some background discussion of health infrastructure and safe water availability in Indian villages.

Primary health care is delivered to the rural Indian population via a network of government-owned and operated health centers. There are three tiers of centers: community health centers (CHCs), primary health centers (PHCs), and primary health sub-centers (PHSCs). The PHSCs form the backbone of the rural health system; they are often the first point of contact for ill individuals and are intended to serve as a link between the rural population and the national health-care system within the referral system. They also serve an important function in providing preventive health services, such as immunizations, pregnancy care, and communicable disease control; as well as health education, promotion of nutrition, and basic sanitation. The typical staffing for a PHSC is one male and one female health worker, with the latter being an auxiliary nurse midwife.

The PHCs constitute the next tier in the primary health network. Each PHC serves as a hub or referral unit to 5–6 subcenters. A PHC has 4–6 beds and offers curative, preventive, and family welfare services.

The third and top tier in the rural health system is represented by the CHC, which serves as a referral unit to 4–5 PHCs. Each CHC has four specialists—a physician, surgeon, gynecologist, and pediatrician—supported by 21 paramedical and other staff members. It has (or is meant to have) 30 indoor beds, one operation theater, X-ray and labor rooms, and laboratory facilities. It provides emergency obstetrics care and specialist consultation.

Current norms require one PHSC to be present for every 5,000 persons; one PHC for 30,000 persons; and one CHC for 120,000 people in the plains. These norms are lower for populations in tribal, hilly, and difficult to reach areas.

The village-level Census data provide the most comprehensive look at the *actual* availability of these three tiers of health facilities across Indian villages. These data demonstrate the near-complete absence of health facilities in most villages. For instance, the 2001 Census indicates that a mere 1.5% of villages in the country had a functioning CHC within the village, up from 1% in 1991. A slightly higher, but still very small, proportion of villages had PHCs (3.9%). Village PHSCs were somewhat less scarce though still far below government published figures. There are large variations in the local availability of these across states. In 2001, as many as 45% of villages in Kerala, but only 3.6% of villages in Uttar Pradesh, had a PHC.

In 1977, the Government of India launched a new form of health service to the rural population. These were community health workers, who were often local volunteers provided with short-term training and a small financial incentive. Community health workers, also referred to as village health guides, are meant to provide preventive care and outreach services to rural households. They are also expected to provide appropriate referral linkages to government and private providers, and to beneficiaries outside the traditional coverage area. While the expectation was that the program would expand rapidly so that almost every village would have its own community health worker, the program stalled in the 1990s, and gains in coverage have been limited in most states since then. At the all-India level, village coverage of community health workers slipped from 16.6% in 1991 to 14.5% in 2001. However, in a few states, such as Tamil Nadu, Maharashtra, and Madhya Pradesh, community health worker coverage increased over the same period. For instance, while 29.8% of villages in Tamil Nadu had a community health worker in 1991, as many as 59.4% did in 2001.

The health infrastructure described above refers to public health facilities. There is also a large private sector in the country, with an estimated 1.4 million registered private medical practitioners. In addition to the availability of public health facilities, the Census also publishes information on the presence and number of registered private medical practitioners in each village. While medical practitioners can be registered with the government only if they hold a medical degree, the degree can be in medical systems other than allopathy, e.g., *ayurveda*, homeopathy, *unani*, *siddha*, and naturopathy medicine. The Medical Council of India estimates that more than one half of all registered private medical practitioners in the country practice nonallopathic forms of medicine. Government regulation of registered private medical practitioners and of the entire private health sector is very weak in India, resulting in significant unevenness in the quality of registered private medical practitioners. Many rural medical practitioners often have no formal qualifications, not even a high school diploma.

The Census data indicate that 9.4% of villages in the country had a registered private medical practitioner in 2001, up from 6.5% in 1991. Thus, registered private medical practitioners appear to be less common in Indian villages than PHSCs, although they are more ubiquitous than PHCs or CHCs. As with almost all of the health infrastructure variables, there are wide variations across states. In Kerala, as many as 62.2% of villages had registered private medical practitioners in 2001, but in Uttar Pradesh the comparable figure was only 6.8%.

Combining all the various modes of health service delivery, the Census data indicate that only 27.1% of villages in the country had any form of health infrastructure (i.e., PHSC, PHC, CHC, community health worker, or registered private medical practitioner) in 2001. This number varied from 10% to 11% in Assam and Bihar, to 75.6% in Tamil Nadu, and 83.2% in Kerala.

Contaminated water poses one of the most significant risks to child health in India. It is estimated that nearly one fifth of all communicable diseases in India are water-related. Of these diseases, diarrhea alone is estimated to take 700,000 lives each year. The highest mortality from diarrhea is in children under the age of five, highlighting an urgent need for focused interventions to prevent diarrheal disease in this age group.

The village-level Census data obtain and report information on the availability of a drinking water facility in the village. The facilities that are identified include taps, wells, tanks, tube-wells, hand-pumps, rivers, canals, springs and lakes. As tap water is usually considered the safest form of drinking water, we include the availability of tap water in a village in our analysis.[7]

Data from the 1991 and 2001 Village Level Census show a large expansion in the availability of tap water. While 17% of villages in the country had access to tap water in 1991, this nearly doubled to 33% by 2001. As would be expected, there are large interstate variations in the availability of tap drinking water. In 2001, a mere 3.2% of villages in Bihar had access to tap drinking water, while the ratio was 86% in Kerala.

VII. Empirical Model

In this paper, we focus on the reduced-form demand for female versus male children (i.e., the juvenile sex ratio). As a first-order approximation to this demand equation, one can write a linear form:

$$JSR_{ijt} = \alpha_{ij} + \beta_{jt} + cW_{ijt} + \mathbf{d}\,\mathbf{H}_{ijt} + \mathbf{e}\,\mathbf{Z}_{ijt} + \mu_{ijt}, \tag{2}$$

[7] While water from hand-pumps can potentially be included as safe drinking water, the ground water in many rural areas of the country is seriously contaminated because of agricultural irrigation runoffs (and arsenic in the Eastern states of India). In addition, there is the problem of data. While a third of the villages in the country have tap drinking water, more than three quarters of the villages obtain their drinking water from hand-pumps. Thus, the variable formed by combining drinking water available from taps and hand-pumps would have no discriminating power since all the villages in the sample would have access to this form of drinking water.

where the subscripts *i* , *j,* and *t* refer to villages, states, and time, respectively*; JSR* is the juvenile sex ratio (i.e., number of females under 6 years of age to 1,000 males under 6); α is an unobserved village-specific fixed effect; β is an unobserved state-specific effect (that varies over time); *W* is the availability of tap water; ***H*** is a vector of health facilities or services available in the village (e.g., any public health facility such as CHC, PHC, or PHSC;[8] community health worker; and a registered private medical practitioner); ***Z*** is a vector of control variables (e.g., adult male and female literacy rates, village population, access to electricity and to paved road, proportion of scheduled caste/tribe population, etc.); and μ is an iid disturbance term.

The term α_i is an unobserved (time-invariant) village-specific endowment that could reflect cultural differences in parental preferences for boys versus girls across villages or geographical differences in physical endowments (e.g., soil quality, weather, etc.) that might influence the economic returns to male and female labor. Failure to control for these village effects can lead to biased estimates of equation (2), if public health (or other, more general) infrastructure is systematically targeted to villages where cultural preferences for boys versus girls happen to be stronger.[9] In addition to unobserved village heterogeneity, equation (2) also allows for state effects that vary over time to account for policy changes that may have occurred differentially in states over time. For instance, during the 1990s, a few states, such as Andhra Pradesh, Karnataka, Maharashtra, and Tamil Nadu changed their laws to provide women the right to inherit ancestral property. It is widely believed that some of the worst manifestations of gender discrimination in India, such as female feticide and dowry, can be traced to biased inheritance laws favoring sons.

Taking first differences, equation (2) can be rewritten as:

$$\Delta JSR_{ij} = (\beta_{jt} - \beta_{j,t-1}) + c\,\Delta W_{ij} + \mathbf{d}\,\boldsymbol{\Delta H}_{ij} + \mathbf{e}\,\boldsymbol{\Delta Z}_{ij} + \Delta\mu_{ij}, \quad (3)$$

where Δ is the first-difference operator (e.g., $\Delta JSR_{ij} = JSR_{ij,t} - JSR_{ij,t-1}$). Note that only the *difference* in state effects between *t* and *t-1* – viz., $(\beta_{jt} - \beta_{j,t-1})$ and not the absolute value of the state effect in either year can be identified in equation (3).

It is important to address the issue of measurement error in the juvenile sex ratio. Since the number of children in each family is the result of a binomial process, there will be some randomness in the variation in juvenile sex ratios across villages. The smaller the village, the more acute the measurement error in the juvenile sex ratio will be. Further, first-differencing will tend to exacerbate this measurement error. Two points are worth

8 Although it might have been intuitively appealing to include each of the public health facility types separately in the regression (to allow for the possibility that, say, community health centers have different effects on the juvenile sex ratio than health subcenters), it was difficult in practice to implement this, as there were very few villages in the sample with a CHC (merely 1.5%) or a PHC (3.9%). When aggregated, 16.5% of villages in the sample had one of the three types of public health facilities.

9 This is the case of endogenous program placement, first discussed by Rosenzweig and Wolpin (1986).

noting, however. First, random measurement in a dependent variable leads to imprecisely estimated (not biased) estimates. Second, it may be important to control for village size in the regression, as the measurement error in the juvenile sex ratio is likely to be systematically greater in smaller villages. Controlling for size is also important because it can affect the type of facility available (for example, CHCs are usually located in very large villages while PHSCs are located in small ones).

In equations (2) and (3), W is represented by the availability of tap water in a village in our empirical estimation. The variables included in H are the presence of any public health facility (i.e., CHC, PHC, or PHSC); a community health worker; and a registered private medical practitioner in a village. The choice of control variables included in Z is constrained by the type of data that are available from the village Census files. We use male and female literacy rates in a village as measures of adult education, which is an important determinant of parental preferences for sons versus daughters.[10] We also use control for the proportion of scheduled castes and scheduled tribes in a village to allow for the possibility that these groups may differ from the mainstream population in terms of the relative value they place on sons and daughters.[11]

In addition, we use measures of general infrastructure—village-level availability of electricity, telephone connection, and a paved approach road—as these are likely to influence the actual or opportunity cost of obtaining health services. Measures of household income are not available at the village level; instead, we proxy for income by the amount of cultivable land per cultivator and the percentage of cultivable land that is irrigated in the village. Finally, we include village size (population) in the regression to control for the possibility that measurement error in juvenile sex ratios is systematically larger in villages with relatively few children.

While the estimates in equation (3) control for unobserved heterogeneity across villages, they do so linearly. They implicitly assume that the marginal effects of health infrastructure—indeed, of all the independent variables—on the juvenile sex ratio are constant. What are then estimated are the mean effects on the juvenile sex ratio of infrastructure and other explanatory variables. Such estimates miss a point that is crucial for policy makers, namely, that policy interventions may affect the juvenile sex ratio differently at different points of the conditional juvenile sex ratio distribution. For example, while some types of health infrastructure may not matter for the juvenile sex ratio in the "average" village, they may matter a great deal in villages at the bottom of the conditional juvenile sex ratio distribution (i.e., in villages where girls are at the greatest risk of feticide or childhood neglect relative to boys).

As is well known, the quantile regression technique is a means of allowing estimated marginal effects to differ at different points of the conditional distribution of the dependent

10 The PCA only reports data on adult literacy, not on measures of schooling attainment among adults.

11 While it would have been interesting to also examine the effects of religion, information on religion is not available from the Census at the village level.

variable. Quantile regressions were initially developed as a robust regression technique that would allow for estimation where the typical assumption of normality of the error term might not be strictly satisfied (Koenker and Bassett 1978). However, they are now used extensively to analyze the relationship between the dependent and independent variables over the entire distribution of the dependent variable—not just at the conditional mean (Buchinsky 1994, Buchinsky 1995, Eide and Showalter 1997).

In this paper, we attempt to address both types of heterogeneity. We use the fixed-effects quantile regression technique to analyze the socioeconomic and policy determinants of juvenile sex ratios at different points of the conditional distribution of the juvenile sex ratio. This allows us to address not only the question, Can policy influence juvenile sex ratios? More importantly, we address the question, For which types of communities do policy interventions matter the most?

VIII. Empirical Results

Village fixed-effects regressions of the juvenile sex ratio are reported in Table 1. As a comparison, we also report pooled (over 1991 and 2001) ordinary least squares (OLS) estimates as well as OLS estimates for each of the two Census years. The OLS equations include individual state dummies, but obviously do not control for village effects. Since the OLS estimates are likely to be biased and are not our preferred estimates, there is little point in discussing them extensively. The most noteworthy finding in the OLS estimates is that the presence of any health facility or practitioner in a village serves to decrease the juvenile sex ratio, with the negative "effect" of a registered private medical practitioner being the largest in magnitude.[12] In contrast, the availability of tap water has a positive "effect" on juvenile sex ratios in 1991 but a negative effect in 2001.

Turning to the fixed-effects results, the estimated effects of a public health facility and of a community health worker are insignificant (reported in the last column of Table 1). *Prima facie* this suggests that, within states, public health facilities tend to be located in villages where (unobserved) son preference is strong. Since son preference is associated with lower juvenile sex ratios, the observed inverse relationship between public health facilities and juvenile sex ratios disappears after controlling for unobserved village heterogeneity (in son preference). The same is true of private health workers, since the village fixed-effects estimates of the effect of a registered private medical practitioner on the juvenile sex ratio are considerably less negative than the OLS estimates.

[12] Because the OLS estimates do not control for unobserved heterogeneity, they cannot be interpreted as causal effects.

Table 1: OLS and Fixed-Effects Regressions of the Juvenile Sex Ratio, Indian Villages, 1991 and 2001

	OLS – Pooled Data		OLS – 1991		OLS – 2001		Village Fixed-Effects	
Independent Variable	**Coef.**	**T-ratio**	**Coef.**	**T-ratio**	**Coef.**	**T-ratio**	**Coef.**	**T-ratio**
Male literacy rate	**2.34**	95.88	**1.50**	46.92	**3.66**	93.89	**5.67**	105.16
Female literacy rate	**−2.79**	−109.13	**−2.02**	−54.15	**−3.91**	−106.65	**−5.62**	−99.47
Scheduled caste (percent of population)	**0.08**	6.98	**0.06**	3.49	**0.10**	6.35	**0.29**	4.21
Scheduled tribe (percent of population)	**0.43**	50.87	**0.43**	34.83	**0.40**	34.42	**0.32**	3.48
Presence in village of:*								
any public health facility**	**−1.75**	−3.67	**−2.38**	−3.23	**−1.77**	−2.76	0.20	0.23
community health worker	**−3.30**	−6.62	**−3.13**	−3.74	**−1.60**	−2.35	1.36	1.61
registered private medical practitioner	**−7.81**	−14.19	**−5.17**	−6.04	**−6.89**	−9.34	**−0.66**	−2.40
Availability in village of:*								
tap water	**−1.33**	−2.77	**1.31**	1.69	**−3.00**	−4.80	1.08	1.25
power supply	−0.58	−1.05	**−3.03**	−4.00	**2.15**	2.67	**−5.33**	−5.21
at least one phone connection	**−5.22**	−10.44	**−4.08**	−4.21	**−3.00**	−5.08	**4.15**	5.01
paved approach road	**−3.17**	−7.50	**−3.16**	−5.23	**−3.40**	−5.70	0.81	1.03
Log village population	**2.32**	8.94	**2.88**	7.69	**1.60**	4.39	3.34	1.56
Acres of cultivable land per cultivator	**0.07**	2.22	**−0.09**	−1.84	**0.17**	4.52	−0.04	−0.68
Percent of area irrigated	**−0.02**	−3.89	0.01	0.85	**−0.10**	−12.82	−0.02	−1.24
Dummy variable for 2001	0.10	0.22						
Intercept	**892.15**	457.87	**911.06**	334.19	**862.00**	296.17	**−20.92**	−11.11
Number of observations	965,194		479,206		485,988		370,989	
F-ratio	1,225		326		1,027		517	
R-squared	0.04		0.02		0.06		0.05	

* Dichotomous variable.

** Includes primary health sub-center, primary health center, or community health center.

OLS = ordinary least squares, coef = coefficient.

Notes: Estimation employs village-level data from the 1991 and 2001 Censuses. Every equation includes dummy variables for each state, the estimated coefficients for which are not shown due to space considerations. Asymptotic t-ratios are corrected for heteroscedasticity using the Huber-White method. Figures in bold indicate statistical significance of the estimated coefficient at the 10% or lower level.

The estimated effect of tap drinking water on juvenile sex ratios is also insignificant in the fixed-effects regression. However, interestingly, two types of general infrastructure have highly significant effects; the availability of a phone connection serves to increase, while the availability of electricity serves to reduce, juvenile sex ratios.

Perhaps the most unusual finding is the strong negative effect of female literacy rates on juvenile sex ratios, with a one percentage point increase in female literacy reducing the juvenile sex ratio in a village by as much as 6 points. This finding is robust in that it is observed across all estimates (single-year OLS, pooled OLS, and fixed-effects). While the finding may appear to be counterintuitive, it has been observed in other studies for India. For instance, using a large household dataset from a special fertility and mortality survey conducted in India in 1998, Jha et al. (2006) found that the sex ratio among children was considerably more skewed (toward males) among educated mothers than among illiterate mothers. Basu (1999), too, has noted that the core level of son preference in India may actually be strengthened by “modernization”’ (e.g., education) of women.

The perverse association between female literacy and juvenile sex ratios may reflect two facts: first, that female literacy is strongly associated with lower fertility in India, as in other developing countries, and, second, that fertility decline is typically associated with lower juvenile sex ratios. There is a large literature that shows female education and literacy to be one of the strongest determinants of fertility decline in India (Rosenzweig and Evenson 1977, Dreze and Murthi 2001, International Institute for Population Sciences 2007). But a number of studies have also documented that as fertility has declined in India, a culture with a strong son preference, there has been a steep rise in the sex ratio at birth and in the juvenile sex ratio (Dasgupta and Bhat 1997, Basu 1999).

Dasgupta and Bhat (1997) identify two opposing effects of fertility decline on sex ratios: a positive "parity" effect whereby fertility decline results in fewer births at higher parities where discrimination against girls is strongest, and a negative "intensification" effect, whereby parity-specific discrimination against girls becomes more pronounced at lower levels of fertility. The intensification effect reflects the fact that the total number of children desired by couples typically falls faster than the total number of desired sons. In India, at least during the recent past, the intensification effect has dominated the parity effect, with the sex ratio among children worsening as fertility has declined.

The strong relationship between fertility decline and a lower juvenile sex ratio is observed in other countries with a strong son preference, such as the PRC and Republic of Korea (Park 1983, Park and Cho 1995, Croll 2002). The imbalance in the juvenile sex ratio is realized through several different parental actions, such as prenatal sex determination using ultrasound technology and sex-selective abortions (female feticide), female infanticide, parental neglect of female children (leading to their lower rate of survival relative to male children), and differential contraceptive use depending upon the sex composition of existing children (with women who bear daughters early in their reproductive years continuing child-bearing while those who bear sons early halting their child-bearing).

In the fixed-effects regression, the estimated effect of male literacy is exactly the opposite to that of female literacy. A one percentage point increase in the male literacy rate in a village is observed to increase the village's juvenile sex ratio by about 6 points. Borooah and Iyer (2005) also find a similar positive effect of male literacy on the sex ratio, using household data from a national survey of over 33,000 rural households in India. Why would male literacy have a beneficial effect on juvenile sex ratios when female literacy is observed to have a deleterious effect? One possibility is that male literacy, unlike female literacy, is not associated with reduced fertility (Dreze and Murthi 2001).[13] As such, an

[13] Indeed, it is possible that male literacy might even be associated positively with fertility. This might occur if the effect of female literacy (as a proxy for female wage rates and the opportunity cost of time for women) was capturing the "price effect" on fertility demand, while the male literacy effect was representing the "income effect." Thus, as female literacy and female wage rates increase, the shadow price of children rises (as children are very intensive in mother's time) and the parental demand for children falls. On the other hand, an increase in male literacy and male wages largely results in increased household income, and leads parents to desire more children.

increase in male literacy does not result in the same pressure to have fewer daughters and more sons as does an increase in female literacy.

Another interesting, although not unexpected, finding is that, *ceteris paribus*, villages with larger populations of scheduled (backward) castes and tribes have higher juvenile sex ratios. This may reflect the higher social status enjoyed by tribal women relative to women in mainstream Hindu society (Basu 1993). Many tribal societies are matrilineal and women in such social structures enjoy inheritance rights and privileges, autonomy, and economic independence that few women in patriarchal societies enjoy. In addition, the fact that many scheduled tribes have managed to keep their social and physical distance from the mainstream Hindu society may have helped them maintain their culture and social norm of not discriminating against girls. However, this behavior is not characteristic of the scheduled castes, which have typically tended to emulate the son-preference norms of the dominant castes, which is why the finding that scheduled castes have higher juvenile sex ratios is surprising.

The estimated coefficients on the state dummy variables are also of interest (although they are not reported in Table 1 due to space considerations). These coefficients are estimates of $(\beta_{jt} - \beta_{j,t-1})$, namely, the state-specific declines in the juvenile sex ratio that would have occurred between 1991 and 2001 *in the absence of any changes in the right-side variables* (i.e., changes in health and general infrastructure, male and female literacy, and population composition) and *after controlling for village effects*. In Table 2, we rank the different states by the magnitude of the estimated state dummy coefficient. The results indicate that the states of Punjab, Gujarat, Assam, and Haryana experienced the largest *ceteris paribus* declines in the juvenile sex ratio, while Madhya Pradesh, Kerala, West Bengal, and Uttar Pradesh experienced the smallest *ceteris paribus* declines. This ranking of state differs from a ranking based on the simple (unconditional) mean of state-level changes in the juvenile sex ratio between 1991 and 2001. In particular, Uttar Pradesh comes out looking much better and Gujarat much worse in terms of its performance on increasing the juvenile sex ratio between 1991 and 2001 once there is control for changes in the infrastructure and other variables.

Table 2: Ranking of States by Actual Change in the Juvenile Sex Ratio and by the Size of the Estimated State Effect in the Fixed-Effects Regression

	Actual Change in the Juvenile Sex Ratio from 1991 to 2001	Estimated State Effect in the Fixed-Effects Regression (Table 2)
	Punjab	Punjab
	Haryana	Gujarat
	Himachal Pradesh	Assam
	Gujarat	Haryana
	Maharashtra	Andhra Pradesh
	Andhra Pradesh	Maharashtra
Worst	Tamil Nadu	Himachal Pradesh
offenders	Karnataka	Rajasthan
	Orissa	Bihar
	Bihar	Orissa
	Assam	Karnataka
	Uttar Pradesh	Tamil Nadu
	Rajasthan	Madhya Pradesh
	West Bengal	Kerala
	Madhya Pradesh	West Bengal
	Kerala	Uttar Pradesh

IX. Empirical Results: Quantile Regression Estimates

As noted earlier, the fixed-effects results discussed above control for unobserved heterogeneity across villages, but only of a linear sort. In other words, we allow each village to have its own (time-invariant) intercept. But villages could be heterogeneous in terms how their juvenile sex ratios respond to the availability of health services and safe drinking water. It is likely that the estimated effects of public health facilities or female literacy on the conditional mean of juvenile sex ratios are not necessarily indicative of the size and nature of these effects on the lower tail of the sex ratio distribution (where the problem of discrimination against girls is most acute). Below, we go beyond estimating simple "mean" effects, and allow for heterogeneity in the estimated marginal effects by estimating fixed-effects quantile regressions.

The quantile regression results (Table 3) are revealing. They indicate that the presence of any public health facility in a village has a significant and large positive effect on the juvenile sex ratio at the 0.10 and 0.25 quantiles.[14] At the 0.5 quantile, the estimated effect is insignificant, and turns increasingly negative at successively higher quantiles. Exactly the same results hold true of the presence of a community health worker, the presence of a registered private medical practitioner, and the availability of tap water and a paved approach road in a village. The case of electricity is similar across the different quantiles, but with the signs flipped. Figure 10 shows how monotonically the estimated

[14] The 0.25 quantile includes villages with a juvenile sex ratio of less than 840 females per 1,000 males and villages that experienced a decline in the juvenile sex ratio of 160 females per 1,000 males between 1991 and 2001.

marginal effects decline at the upper tails of the juvenile sex ratio distribution. The figure drives home the point that median (or least-squares) regressions that produce a small estimated effect of health services or safe water on juvenile sex ratios at the median or mean of the dependent variable mask very large positive effects at the lower tail of the distribution and very large negative effects at the upper tail.

Table 3: Quantile Fixed-Effects Regressions of the Juvenile Sex Ratio, Indian Villages, 1991 and 2001

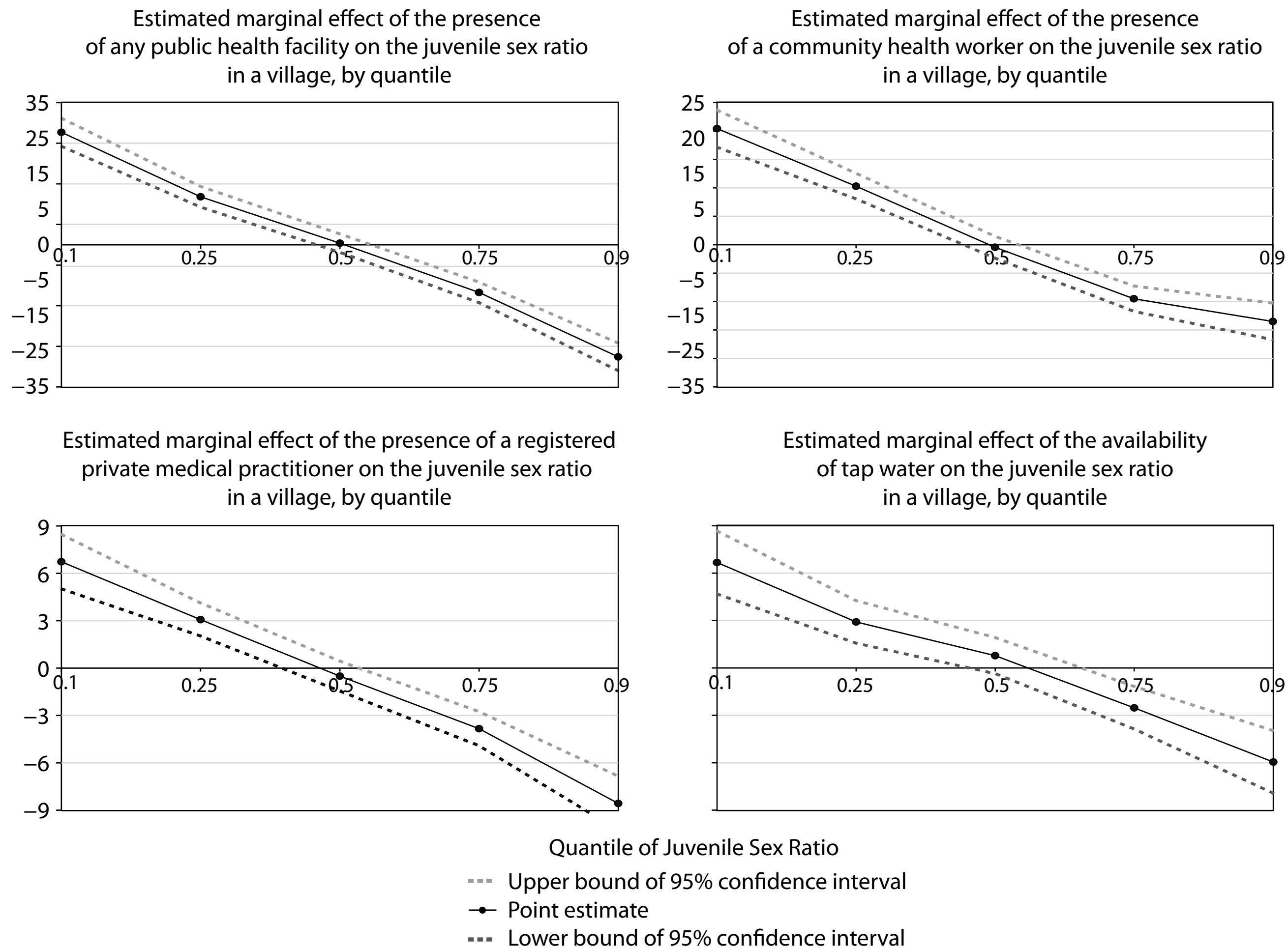
Independent Variables	0.10 Quantile		0.25 Quantile		0.50 Quantile		0.75 Quantile		0.90 Quantile	
	Coef.	T-ratio	Coef.	T-ratio	Coef.	T-ratio	Coef.	T-ratio	Coef.	T-ratio
Male literacy rate	**5.21**	50.83	**5.30**	86.90	**5.61**	118.64	**6.21**	108.49	**6.60**	71.45
Female literacy rate	**−5.59**	−54.31	**−5.50**	−87.29	**−5.52**	−109.29	**−5.83**	−92.81	**−6.21**	−60.56
Scheduled caste (percent of population)	**0.41**	3.19	**0.36**	4.75	**0.35**	5.73	**0.39**	5.14	**0.22**	1.73
Scheduled tribe (percent of population)	**0.64**	3.66	**0.46**	4.59	**0.28**	3.65	**0.18**	1.97	0.01	0.05
Presence in village of:*										
any public health facility**	**27.70**	15.67	**11.82**	9.10	0.39	0.34	**−11.73**	−9.09	**−27.61**	−15.95
community health worker	**20.41**	12.22	**10.29**	8.96	−0.45	−0.46	**−9.50**	−8.33	**−13.49**	−8.23
registered private medical practitioner	**6.73**	7.65	**3.07**	5.75	−0.51	−1.05	**−3.84**	−6.96	**−8.57**	−9.77
Availability in village of:*										
tap water	**11.12**	6.55	**4.85**	4.23	1.29	1.33	**−4.21**	−3.67	**−9.93**	−5.90
power supply	**−20.49**	−10.78	**−13.68**	−11.00	**−4.99**	−4.84	**4.00**	3.26	**9.73**	5.28
at least one phone connection	**38.86**	23.93	**19.35**	17.62	**3.85**	4.12	**−9.51**	−8.69	**−32.10**	−19.97
paved approach road	4.64	3.11	**2.17**	2.19	**1.75**	2.11	−0.82	−0.84	**−3.91**	−2.64
Log village population	0.88	0.22	**5.20**	2.21	**8.87**	4.79	2.84	1.23	0.46	0.12
Acres of cultivable land per cultivator	**−0.25**	−2.40	**−0.22**	−3.28	−0.07	−1.27	0.06	0.90	**0.20**	1.77
Percent of area irrigated	**0.18**	7.23	**0.07**	4.39	−0.02	−1.62	**−0.10**	−6.02	−0.16	−6.63
Intercept	**−296.91**	−76.57	**−150.93**	−59.80	**−23.00**	−10.97	**107.48**	42.92	**256.05**	67.23
Number of observations	370,989		370,989		370,989		370,989		370,989	
Pseudo R-squared	0.04		0.03		0.03		0.03		0.04	

* Dichotomous variable.

** Includes primary health sub-center, primary health center, or community health center.

Note: Estimation employs village-level data from the 1991 and 2001 Censuses. Every equation includes dummy variables for each state, the estimated coefficients for which are not shown due to space considerations. Asymptotic t-ratios are corrected for heteroscedasticity using the Huber-White method. Figures in bold indicate statistical significance of the estimated coefficient at the 10% or lower level.

Figure 10: Estimated Marginal Effects of Health-Related Facilities and Infrastructure



What could account for these results? As noted earlier in Section IV, the lower tail of the juvenile sex ratio distribution includes villages and communities where there is a strong cultural preference among parents for sons. In these communities, parents probably view health care for boys—but not for girls—as a "necessity", which in turn means that the parental demand for treating boys at health facilities is likely to be more price-inelastic than their demand for treating girls. Thus, boys would be better protected than girls from increases in health care prices. However, this inelasticity would also mean that girls' treatment would respond more strongly than boys' treatment to greater local availability of health facilities; in other words, a decline in the price of health care. Thus, the net result of improved health services in such a setting would be a higher survival rate for girls relative to boys and a higher juvenile sex ratio.

The one odd empirical result is the finding that the local availability of electricity reduces juvenile sex ratios in the lower quantiles but increases them in the higher quantiles of the sex ratio distribution. Perhaps the availability of electricity facilitates prenatal sex determination, thereby reducing the juvenile sex ratio in villages with strong son preference. However, this is only speculation at best.

The quantile regression results are consistent with the least-squares results for male and female literacy as well as for affiliation to scheduled castes and tribes. Indeed, the estimated effects of male and female literacy on juvenile sex ratios are surprisingly similar across the entire distribution of the juvenile sex ratio. The consistently negative effect of female literacy on juvenile sex ratios is disturbing, but, as noted earlier, probably reflects the fact that female literacy is associated with lower levels of fertility, which in turns increases the pressure on couples to have boys in a strong son preference culture such as India's.

While the pernicious effects of public health facilities and other infrastructure on the juvenile sex ratio at the upper tail of the juvenile sex ratio distribution are of concern, it is reassuring from a public policy viewpoint to find that in villages where the juvenile sex ratio is especially imbalanced, most public interventions (such as increased availability of public health facilities, community health workers, tap drinking water, agricultural irrigation, telephone connections, and paved approach roads) serve to increase the juvenile sex ratio. The only interventions that reduce the juvenile sex ratio in the lower tail of the sex ratio distribution are the supply of electricity and the availability of cultivable land per cultivator.

The coefficients on the state dummy variables are not reported in Table 3 due to space considerations. However, in Table 4, we rank the different states by the magnitude of the estimated coefficient on the state dummy variables. It is observed that the ranking of states varies significantly across the quantiles. For instance, at the lower tail of the juvenile sex ratio distribution (among villages that experienced a large decline in the juvenile sex ratio), Himachal Pradesh performed significantly worse than Kerala after controlling for changes in all the right-side variables. However, at the upper tail of the distribution (among villages that experienced a large increase in the juvenile sex ratio), Kerala was the worse performer relative to Himachal Pradesh. Orissa and Madhya Pradesh, which do not appear to be problem states based on the results of the simple (i.e., nonquantile) fixed-effects regression, appear to have a major problem with declining juvenile sex ratios in the lower quantile regression results. Since policy makers ought to be primarily concerned about villages at the lower tail of the juvenile sex ratio distribution, our empirical results suggest a stronger need to target policies combating imbalanced sex ratios to states such as Himachal Pradesh, Assam, Punjab, Orissa and Madhya Pradesh.

Table 4: Ranking of States by Size of Estimated State Effect in the Quantile Regressions (reported in Table 3)

Worst offenders

Ranking of States by Magnitude of Estimated State Coefficient in Quantile				
0.1	**0.25**	**0.5**	**0.75**	**0.9**
Himachal Pradesh	Himachal Pradesh	Punjab	Gujarat	Kerala
Assam	Punjab	Gujarat	Haryana	Gujarat
Punjab	Assam	Assam	Kerala	Haryana
Orissa	Orissa	Haryana	Punjab	Maharashtra
Madhya Pradesh	Madhya Pradesh	Maharashtra	Maharashtra	Punjab
Karnataka	Karnataka	Andhra Pradesh	Assam	Assam
Rajasthan	Rajasthan	Himachal Pradesh	Andhra Pradesh	Andhra Pradesh
Tamil Nadu	Gujarat	Orissa	Bihar	Bihar
Bihar	Andhra Pradesh	Rajasthan	Rajasthan	West Bengal
Andhra Pradesh	Bihar	Bihar	West Bengal	Rajasthan
Uttar Pradesh	Tamil Nadu	Karnataka	Tamil Nadu	Uttar Pradesh
West Bengal	Haryana	Madhya Pradesh	Uttar Pradesh	Tamil Nadu
Haryana	Uttar Pradesh	Tamil Nadu	Karnataka	Karnataka
Maharashtra	West Bengal	West Bengal	Madhya Pradesh	Madhya Pradesh
Gujarat	Maharashtra	Uttar Pradesh	Orissa	Orissa
Kerala	Kerala	Kerala	Himachal Pradesh	Himachal Pradesh

It is important to be cognizant of the limitations of the analysis presented in this paper. The key explanatory variables we have used on the availability of health facilities at the village level do not accurately reflect true population access to quality health services. The quality of health services available at public health facilities is uneven across the country, with poorer regions having high rates of absenteeism (of health workers from health facilities) than better-off regions. While the fixed-effects estimates control for this type of unobserved heterogeneity in service quality, they only control for variations in service quality that are time-invariant (between 1991 and 2001). Another limitation of the analysis—driven largely by the availability of data—is that it excludes the urban areas of the country, where the problem of juvenile sex ratios is even more pronounced than in the rural areas. A third limitation is the omission of variables that measure the differential economic contribution of male versus female children across villages—a factor that could drive the excess demand for sons relative to daughters among rural Indian parents. Of course, the fixed-effects estimates control for any time-invariant differences across villages in the "economic value" of male versus female children. But they are unable to account for time-varying changes in the value of sons versus daughters.

X. Conclusion

In this paper, we have used village-level data from the 1991 and 2001 Censuses to analyze the policy determinants of juvenile sex ratios in India. In particular, we have focused on the availability of health facilities and safe drinking water at the village level —interventions that are amenable to policy manipulation—and estimated their impacts on juvenile sex ratios. To do so, we have employed an empirical strategy that entails estimating a reduced-form demand for surviving female versus male children (in other words, the juvenile sex ratio).

We control for village-level heterogeneity in unobserved son preference in two ways. First, we include village fixed effects in our estimating equation of the juvenile sex ratio. Second, we also allow villages to be heterogeneous in terms how their juvenile sex ratios respond to the availability of health facilities and safe drinking water. We do this by estimating fixed-effects quantile regressions that allow the impact of the different health interventions to vary at different points along the conditional distribution of juvenile sex ratios.

Our results suggest that both types of controls for village-level heterogeneity matter. For example, our OLS estimates of the juvenile sex ratio indicate that the presence of a health facility or practitioner in a village serves to *decrease* the juvenile sex ratio. This result, however, is driven by the fact that public health facilities tend to be located in villages where (unobserved) son preference is strong. Once unobserved village heterogeneity is controlled for using village fixed effects, the estimated effects of a public health facility and of a community health worker are not significant.

However, this does not mean that establishing public health facilities in villages would have little impact on the juvenile sex ratio. Results from our fixed-effects quantile regressions indicate that the presence of any public health facility (or tap water) in a village has a significant and large positive effect on the juvenile sex ratio in villages where the problem of discrimination against girls is most acute, i.e., in villages at the 0.10 and 0.25 quantiles of the conditional juvenile sex ratio distribution. At the median of the conditional distribution of sex ratios, however, the estimated effect is not significantly different from zero, and becomes negative at successively higher quantiles.

What drives these results? In all likelihood, villages included in the lower tail of the juvenile sex ratio distribution are those with a strong cultural preference for sons. Since parents in these villages probably view health care for boys as a “necessity” relative to girls, boys may be better protected than girls from increases in the price of health-related services. The flip side of this inelastic (elastic) demand for treatment of boys (girls) is that greater local availability of health facilities (i.e., effectively a decline in the price of health care) increases the likelihood of medical treatment and survival for girls relative to boys.

At upper tails of the juvenile sex ratio distribution (i.e., in villages where there are more girls relative to boys) exactly the opposite effect is obtained.

From a public policy perspective, it is reassuring to find that in villages where the juvenile sex ratio is especially imbalanced, most public health interventions, such as increased availability of public health facilities, community health workers, and tap drinking water increase the juvenile sex ratio. Moreover, since policy makers ought to be primarily concerned about villages at the lower tail of the juvenile sex ratio distribution, our results suggest a stronger need to target policies combating imbalanced sex ratios to states such as Himachal Pradesh, Assam, Punjab, Orissa and Madhya Pradesh where many such villages are to be found.

Other than the public health interventions we have focused on here, what could such imbalanced sex ratio-combating policies be? A typical response may be to increase female literacy. But an unusual, yet robust, finding of our empirical work is that higher female literacy rates in a village exert a strong negative effect on juvenile sex ratios. A possible explanation has to do with the different effects of female versus male literacy on fertility decisions. In particular, female literacy is widely believed to drive fertility declines in India. In the context of a strong son-preference culture, it should not be surprising if the lower levels of fertility lead to greater pressure on couples to have boys.

While outlawing prenatal sex-determination tests and sex-selective abortions could be considered an obvious policy response, this has already been accomplished in India. Sex selection tests have been illegal in India under the 1994 Pre-conception and Pre-natal Diagnostic Techniques (Prohibition of Sex Selection) Act. Unfortunately, this has done little to stop sex-selection abortions from taking place. Enforcement of the law is weak, and there are loopholes in the law that allow clinics and doctors to continue to perform sex-determination tests. As noted throughout this paper, the root cause of imbalanced sex ratios is the strong preference for sons among Indian parents. While there are many reasons for this, one is that, historically, inheritance laws in the country, especially among Hindus, have favored sons over daughters. Hindu inheritance customs were codified into law in a bill enacted in 1956 that provided the right of inheriting ancestral property only to males. It is widely believed that some of the worst manifestations of gender discrimination in India, such as female feticide and dowry, can be traced to biased inheritance laws favoring sons.

During the last two decades, a few stalwart states, such as Andhra Pradesh, Karnataka, Maharashtra, and Tamil Nadu, have changed their laws to provide women the right to inherit ancestral property. (Kerala was the first state to change its inheritance law in 1975.) In 2004, the Indian Parliament introduced and passed the Hindu Succession (Amendment) Bill, which removed the discriminatory provisions of the 1956 Act and allowed parents to bequeath their property to their daughters. In future work, we plan to examine whether these changes have had an effect on juvenile sex ratios.

There may also be an economic reason for the strong preference for sons among Indian parents. Son preference may be driven by the higher perceived value (by parents) of male relative to female children. Given marriage systems and cultural traditions, sons and not daughters are responsible for the care and upkeep of their parents in old age. One way in which policy can respond is by providing old-age benefits to couples that have only female children. Some governments (e.g., the PRC and a few Indian states) have instituted such schemes, albeit on a very small scale. Evaluation of such programs presents a good area for future research.

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About the Paper

Anil B. Deolalikar, Rana Hasan, and Rohini Somanathan use village-level data from the 1991 and 2001 Indian Censuses to examine the variation in juvenile sex ratios across more than 500,000 villages; and determine whether this variation is related to the availability of health facilities and health infrastructure at the village level.

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