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### **Health and the Political Agency of Women** Sonia Bhalotra and Irma Clots-Figueras

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## Health and the Political Agency of Women

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### Abstract

*We investigate whether women's political representation in state legislatures improves public provision of antenatal and childhood health services in the districts from which they are elected, arguing that the costs of poor services in this domain fall disproportionately upon women. Using large representative data samples from India and accounting for potential endogeneity of politician gender and the sample composition of births, we find that a 10 percentage point increase in women's representation results in a 2.1 percentage point reduction in neonatal mortality, and we elucidate mechanisms. Women's*

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*political representation may be an under-utilised tool for addressing health in developing countries.*

Keywords: political identity, global health, gender, mortality, health, social preferences, India.

JEL codes: H41, I18, O15

## **Health and the Political Agency of Women**

### **I. Motivation and Summary**

This paper investigates the political economy of public health provision, focusing upon the relevance of political identity and, in particular, the gender of politicians. The widespread failure of public health provision in developing countries is commanding increasing attention in the global arena (Black et al. 2003) and in India (Chaudhury et al. 2006, DasGupta 2007, Muralidharan et al. 2011). Poverty is not a compelling explanation for the scale of (maternal and) child death given that most deaths can be averted with low cost interventions (Cutler et al. 2006, Deaton 2006). The evidence points to political failure (Kiefer and Khemani 2003, Besley and Kudamatsu 2006) but the idea that public health services may suffer from the under-representation of women in government has not received dedicated consideration. In our sample, in India, 10% of children die before the age of one and less than 4% of state legislators are women. We investigate whether these phenomena are causally related

Preference differences between the sexes identified in numerous settings suggest that women are more likely than men to (i) invest in children and (ii) favour redistribution.<sup>2</sup> This suggests that women may be more likely to

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<sup>2</sup> On (i), see Thomas 1990, Thomas and Welch 1991, Hoddinott 1995, Lundberg, Pollak and Wales 1997, Case 1998, Svaleryd 2002, Besley and Case 2003). One possible explanation of this is paternity uncertainty (Alexander 1974). On (ii), see Besley and Case 2000, Andreoni and Vesterlund 2001, Edlund and Pande 2002, Edlund, Haider and

deliver public health improvements in developing countries because there are sharp positive age and income gradients in health and the essential policy challenge lies in addressing the early life health and survival of *children* in *poor* families.<sup>3</sup> Moreover, the benefits of interventions in this area flow differentially to women because they disproportionately bear the costs of weak reproductive and child health services. Women carry the burden of high fertility and child mortality: they care for the sick, witness their children die, have further births, and risk dying in child birth.<sup>4</sup> In line with our hypothesis, the enfranchisement of women in early 20<sup>th</sup> century America led to a sharp rise in state health expenditure and a steep drop in infant mortality (Miller 2008). And results of the British Election Survey of 2001 show that the single most important concern for men is low taxes while, for women, it is the quality of the National Health Service (Campbell 2004).<sup>5</sup>

We investigate whether women's political representation in state legislatures in India influences policy outcomes for individuals in the districts from which they are elected. Data on health outcomes, health-related behaviours and village-level access to health facilities for a large representative sample of individuals are matched to the gender composition of state legislators around the time of their birth and in the district of their birth.

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Pande 2005, Warner and Steel 1999, Washington 2008, Oswald and Powdthavee 2010. This has been linked to women being, on average, less likely to pay taxes and more likely to receive benefits (e.g. Alessina and La Ferrara 2005).

<sup>3</sup> Early childhood death accounts for 30% of all deaths in poor countries compared with 1% in richer countries (Cutler et al. 2007). There are vast inequalities in health within poor countries and most childhood deaths occur in poor households.

<sup>4</sup> In the sample period, almost one in ten children died in India before the age of one. Structural estimates on data from the Indian state of Uttar Pradesh indicate that for every neonatal death (which we analyse here), women have an additional 0.37 births (Bhalotra and van Soest 2008). Maternal mortality as a fraction of births is estimated at 0.5% (UNICEF).

<sup>5</sup> At the turn of the century, fertility and infant mortality rates in America were similar to those in our estimation sample for India. In Britain in 2001, both were dramatically smaller. Even so, most women expect to engage with public health services at least once through pregnancy and this makes them heavier users, other conditions equal. The greater concern of women for public health in contemporary Britain may also reflect their representation of children and of the poor.

The main identification challenge is that the gender of a competitively elected leader is likely to be correlated with voter preferences and hence endogenous. A second potential source of bias is that the composition of births is endogenous to the political regime. We address these problems using a fixed effects instrumental variables estimator that exploits constituency-level information on close elections between men and women on the one hand and information on siblings in a large sample survey on the other. We test our identifying assumptions.

Consider the size of the estimated impacts on neonatal and infant mortality, which are markers of the state of public health in developing countries. We find that a one standard deviation increase in women's political representation in the birth year (and similarly in the two years preceding birth) results in a 1.5 percentage point (24%) reduction in neonatal mortality, which is a 0.06 standard deviation change. An alternative expression of effect sizes is that a 10 percentage point increase in the share of women elected from the index district to state assembly seats results in a 2.1 percentage point (33.5%) reduction in neonatal mortality. A 10 percentage point increase in women's representation is close to a shift from zero to one woman in the median district in our sample, which has nine seats. The estimated changes in neonatal mortality drive similarly large impacts on infant mortality. To put the effect size in perspective, note that the causal impact of democratization on infant mortality in sub-Saharan Africa in the post-Cold War period is estimated to be a 1.2 percentage point drop in infant mortality, which is 12% of the sample mean (Kudamatsu 2012).

We find limited evidence of "elitism" amongst women elected to state legislatures. For instance, the point estimates suggest larger impacts for low caste children and amongst children of families in the middle as compared with the top of the wealth distribution. There are no significant differences in the impact of women politicians who belong to lower *vs* higher castes. These results are topical given that a bill proposing quotas for women in state

assemblies is currently pending in the Indian Parliament, with one of the axis of debate being that women elected to state assemblies are unlikely to represent the interests of the poor (Rai 2002; section 2 below).

We show that a number of constituency and candidate characteristics are uncorrelated with the gender of the winner in the close election sample and the estimates stand up to a battery of checks including variation of the vote margin that identifies close elections and controls for party and a polynomial in the vote margin. They are robust to a rich set of controls for within-mother and district-level unobservable trends, and to three falsifying tests. We report several extensions, looking at the impacts of decentralisation, political competition, turnout, media activity and party alignment. We find no evidence that devolutionary reforms in 1993, by sharing power with district and village councils, weakened the influence of women in state assemblies. The relative influence of women in state legislatures is greater when the chief minister is a woman. However, belonging to the same party as the chief minister does not enhance the influence of women legislators, which supports our identification of gender as distinct from party effects.<sup>6</sup>

We are able to identify mechanisms, and this underlines the plausibility of our findings. We find that women politicians are more likely to invest in the village level public health infrastructure while men appear more likely to invest in the financial (and telecommunications) infrastructure. Raising female political representation results in substantially improved probabilities of attending antenatal care, taking iron supplements during pregnancy, giving birth in a government facility as opposed to at home (note: we establish no change in the probability of giving birth in a *private* facility), and early initiation of breastfeeding. By analysing outcomes such as breastfeeding that are not dependent upon infrastructure and by comparing first with further births within mother, we argue that the evidence is also consistent with

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<sup>6</sup> The point estimates indicate that turnout and newspaper circulation in the regional language enhance the policy influence of women while party competition tends to lower it but these estimates are imprecise.

politician-led information campaigns. Our finding that the estimated impacts are larger for first-births also suggests a role for information as information obtained for a first birth tends to persist within mother.

Our results resonate with a literature which suggests that women exhibit a stronger propensity to invest in children and have more progressive social preferences than men. Possibly our main contribution is to the literature describing persistence of the global disease burden and, in this context, the failure of public health delivery. We propose reasons that women will tend to have stronger preferences over health provision than men and produce robust estimates of causal effects using several large representative samples of data from India. Our finding that incremental changes in women's representation in state assemblies have sizeable impacts on infrastructure, behaviour and outcomes is of potential policy significance not only in relation to motivating quotas but also insofar as it extends our understanding of how improvements in public health may be achieved.

## **II. Women in Politics: Theoretical Background and Evidence**

In the benchmark model of democracy, characterised by complete policy commitment, policy choices reflect voter preferences and are independent of legislator identity (Downs 1957). However, more recent citizen candidate models admit a role for legislator preferences (Besley and Coate 1997, Levitt 1996, Osborne and Slivinski 1996). This introduces the possibility that women politicians make different policy choices than men but this rests upon women having different preferences from men. In this way, this paper binds two streams of the literature, one concerned with political identity and the other with differences in preferences between the sexes. We have identified public health as a domain over which we expect women have stronger preferences. Identifying the empirical relevance of preference differences and domains in which they are most active is key to informing debates over quotas for women in government, debates which refer not only to representation but also to principles of fairness and diversity (Squires 1996),

but which suffer from a scarcity of causal evidence of the impact of representation on policy outcomes.

Much of the recent evidence in economics on the influence of women in government has emerged from India following a constitutional amendment which came into force in 1993 following which the federal government devolved power to district and village level governments and introduced reservations for women at these levels. In particular, it was mandated that a third of the seats in every village council and, by random assignment, a third of the posts of head of village council be reserved for women. A recent crop of studies analyses policy and process in village-level governance after 1993, comparing villages in which council headship was reserved for women with villages in which it was not (Duflo and Chattopadhyay 2004, Beaman et al. 2007, 2009, Deininger et al. 2011, Iyer et al. forthcoming). Since every village had a third of its seats reserved for women after the amendment, these post-1993 cross-village comparisons identify the impact of council headship rather than women's representation.<sup>7</sup>

Our analysis of women's representation in state legislative assemblies is moot given that a Women's Reservation Bill proposing that a third of seats in state assemblies and parliament be reserved for women is currently under debate. The Bill was introduced in September 1996 and passed by the *Rajya Sabha* (Upper House) in March 2010 but is currently pending in the *Lok Sabha* (Lower House). The introduction of reservations for women in village councils in 1993 is regarded as a victory for India's women's movement as it allowed the grassroots involvement of village women in village-level development programmes. The more limited support for the move to introduce reservations in state and national parliaments reflects, in part, a concern that women who

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<sup>7</sup> An exception is Iyer et al. (2012) who analyse the combined influence of reservation of headship and of a third of the seats of the village council for women. Rather than a cross-village comparison of local outcomes they study cross-state variation in the reporting of crimes against women as a function of variation in the year in which different states implemented the 1993 village-level reform.



claim those seats will tend to be elite and therefore unlikely to represent the interests of the majority of women citizens (Mishra 2000, Rai 2002). As indicated in section 1, we effectively test the hypothesis that the policy choices of women in state assemblies are detached from the interests of the common woman.

We differ from most previous work in studying women politicians in state legislatures rather than village councils, acting as members [of the state legislature] rather than as heads [of the village council] and elected against men rather than appointed to reserved seats.<sup>8</sup> Women in India occupy a small share of state assembly seats and it is conceivable that incremental changes in their share will have limited policy influence even if their headship at the local level counts. In contrast to previous studies, we investigate not only the composition of village level public goods but also the effectiveness of state interventions. We analyse a longer period of India's democratic history that includes pre-1993 data. This allows us to test whether the devolutionary reforms in 1993, by sharing power with district and village councils, weakened the influence of women in state assemblies. This is the first paper to investigate whether women's say varies with indicators of political competition, and with the gender and party-similarity of the state chief minister. Alongside a range of indicators of public health, we investigate fertility and the sex of births as outcomes.

Two previous studies investigate the role of women in state legislatures. One concludes that raising the share of female politicians results in higher education but only in urban areas (Clots-Figueras 2012). The other finds improvements in a wider set of policy variables but shows that these improvements flow essentially from women who hold seats reserved for the

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<sup>8</sup> While randomisation of the headship of the village council in India created quasi experimental data, a limitation of generalizing results from analysis of reservations is that they can have direct effects on policy outcomes by changing the nature of political competition (Franceschet *et al.* 2009). We discuss identification of the influence of competitively elected leaders below.

lower castes, that is, from “non-elite” women politicians (Clots-Figueras 2011). However it looks at law, expenditure and infrastructure aggregated to the state-level and it may be that the influence of higher caste women is apparent at a more disaggregate level, in the allocation of resources towards their constituencies. Alternatively, their influence may lie not in expenditure and infrastructure but in the effectiveness of provision arising, for example, from the conduct of information campaigns. We therefore look at effectiveness of politicians at the local level and motivate a focus on public health provision as a domain in which gendered preferences may be exhibited and also a domain in which the beneficiaries are relatively poor.

### **III. Data and Descriptive Statistics**

**Health Outcomes and Facilities:** We avoid the use of state-level health expenditure given that its effectiveness in developing countries at large has been debated (Filmer and Pritchett 1999, Filmer, Hammer and Pritchett 2000). Raising the share of women in US state legislatures has been shown to raise state health spending (Rehavi 2007) but, in India, a disproportionate share of the health budget is dedicated to (curative) hospital services which are less pro-poor than (preventive) public health expenditures (Peters et al., 2002). While Indian data suggest some impact of health expenditure on infant mortality, this impact is weakest in the most vulnerable groups (Bhalotra 2007). We select neonatal and infant mortality as indicators of public health provision that measure the effectiveness of state intervention, and at a local level, namely, the district from which the politician is elected. To investigate what it is that women politicians actually do, we investigate the influence of politician gender on the village level public health infrastructure. However public health facilities in India are often poorly staffed (Das and Hammer 2007, Banerjee, Deaton and Duflo 2003) and uptake of publicly provided services such as immunization low (Banerjee et al. 2011). So we also investigate a range of indicators of the utilization of publicly provided antenatal and early postnatal care to indicate whether politician gender

influences health-seeking behaviour. There is a sharp wealth gradient in all of the outcomes.

**Health Data:** Individual data on child survival are obtained from retrospective fertility histories reported in the second round of the Indian National Family Health Survey (NFHS) by a representative sample of ever-married women aged 15-49 in 1998-99.<sup>9</sup> This makes it straightforward to identify siblings. There is an average of 3.8 births per mother, conditional on at least 2. The histories provide information on the year of birth and death for births that occur across three decades, so these data can be matched to the time variation in the political data. The data indicate the district of current residence and the years that the mother has lived in her current location. We drop mothers who had births in a location other than the current residence, but we investigate this restriction below. To allow for full exposure to neonatal [infant] mortality risk we exclude children born less than a month [a year] before the date of the survey. We also draw from the NFHS information on health-seeking behaviours in the four years preceding the survey date and on village facilities at the time of the survey. These are defined in the data appendix and discussed further in the results section. The state of public health is very poor. For instance, in our sample period, 10% of children die before the age of one, almost 70% of births occur at home and only 14% of villages have a primary health centre. All descriptive statistics are in Tables 1A, 1B, and 1C.

[Tables 1A, 1B and 1C here]

**Political Structure:** India is the largest and oldest democracy in the developing world. It has a federal structure in which the constitution devolves significant control over their own government to the 28 states and 7 union territories, and population health falls under the purview of State Legislative Assemblies. States and union territories are divided into single-member constituencies in which candidates are elected in first-past-the-post elections.

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<sup>9</sup> For further information, including sampling design, see IIPS and ORC Macro (2000). For a discussion of the potential issues arising with the use of retrospective data, see the Data Appendix.

The boundaries of assembly constituencies are drawn to make sure that there are, as near as practicable, the same number of inhabitants in each constituency. Thus, state assemblies vary in size with the state population. Districts are an important level of local government between state and village governments. Most decisions pertaining to public health provision were made at the state level until 1993 (Rao and Singh 1998). Even after 1993, when powers over the implementation of developmental projects were devolved to the district and village level, state legislators continued to have substantial say and to play a potentially important role in negotiation with and coordination across layers of local government.

**Electoral Data:** We have constituency-level data on the gender and votes won by the winner and runner-up for every election to State Legislatures in India during 1967-1998. Our sample contains information on 76732 constituency elections in the 16 larger states. In order to match these data with health data, they are aggregated to the district level (see the data appendix). Districts have between 1 and 37 electoral constituencies and the median district has 9. Elections are routinely every five years but states often hold mid-term elections at shorter intervals. The mean proportion of seats in a district won by women is 3.6% (s.d. 7.4%). The median is zero and the 75<sup>th</sup> and 90<sup>th</sup> percentiles are, respectively, 5.5 and 14.3%. So the distribution is highly skewed (Figure A1 in the Appendix). At least one woman was elected in 26.4% of district-electoral years and in this sub-sample, the share of seats held by women is 13.8%. There is an upward trend in the share of seats held by women (Figure 1) and substantial differences in both the level and trend by state (Figure 2).

[Figures 1 and 2 here]

We define close elections as elections in which the winner and the runner-up are of opposite gender and the vote margin between them is less than 3.5%. Across the sample period, only 7.45% of district-years had at least one close election, but they have become more common over time. Fewer than

1 in 400 seats had close elections between women and men in the 1960s and 1970s, and that figure jumped to more than 1 in 100 for the 1980s and 1990s.

**Merged Data:** The health data are merged with the political data by district and (lagged) birth year to produce a micro panel of births within mother nested within a district-year panel. The estimation sample contains more than 70000 individual births to some 18000 mothers that occur across 246 districts (in 16 states) over the 32 year period, 1967-1998.<sup>10</sup> A non-parametric (lowess) plot of the relationship between neonatal mortality risk and the share of women leaders at the district level shows that the unconditional relationship is negative (Figure A2 in the Appendix). We now investigate whether a negative relationship persists after allowing for endogeneity and heterogeneity.

#### IV. Empirical Strategy

We are interested in the relationship between an individual's access to public health services and the share of women elected to state assemblies from the individual's district of birth. We focus upon antenatal and early postnatal services targeted at women and newborns and hence at the gender mix of politicians in the year before and the year of birth. In this section we explain how we address the three potential sources of bias that arise. First, women's representation tends to be correlated with voter preferences. Second, the composition of births may vary endogenously. Third, there may be district-year varying changes in other determinants of health that coincide with the election of women.

To address the problem that common unobservables may drive both electoral preferences for women and health-related behaviours we instrument the overall share of women politicians with the share of women who win in close elections against men. Close elections are defined as those in which the winner and the runner up are of opposite gender and the margin of victory is

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<sup>10</sup> The number of election years is 26 and, since different states had elections in different years, the estimation sample contains 1826 district\* years.

small. The premise is that the identity (gender) of the winner of a close election is quasi-random. To see this, note that in a first-past-the-post electoral system, the probability of being elected is a function of the vote difference between the winner and the runner-up and this function has a discontinuity at zero. As the vote difference approaches the discontinuity, constituencies in which a woman wins by a small vote margin are increasingly similar to constituencies in which a man wins by a small margin (Lee 2008, Pettersson-Lidbom 2001). Since we aggregate to the district level over discontinuities at the constituency level, our specification is similar to a fuzzy regression discontinuity design (e.g. Angrist and Pischke 2008: chapter 6). We investigate robustness of the estimates to controls for varying polynomials in the constituency-specific margins of victory or defeat of each man-woman close election in the district.

The identification driving our approach is nicely illustrated in Figure 3, which plots the fraction of seats (constituencies) in the district won by female politicians against the margin of victory between the female and the male candidate (a positive margin means that the woman won). Figure 3A uses all the elections between a male and a female politician, while Figure 3B restricts the sample to those districts that had a single election between a male and a female politician. The data are aggregated into one percentage point bins, as suggested by Imbens and Lemieux (2008), and we plot a lowess smoothing line on each side of the discontinuity. The size of the jump in Figure 3A indicates that a woman winning a close election in a district increases the fraction of constituencies in the district won by female politicians by roughly 10 percentage points. This is what we would expect given that the average district has 9-10 constituencies.

[Figure 3 here]

Although the gender of the winner in a close election may be considered random, the existence of close elections between women and men may not, for example, it may depend on the number of female candidates in the district and

this, in turn, may be correlated with voter preferences. To allow for this, we control for the fraction of seats in the district that had close elections between female and male candidates in the first and second stage of the instrumental variables procedure. Close elections between women and men have been previously used for identification by Rehavi (2007) and Clots-Figueras (2011, 2012). We investigate the assumptions underlying the use of close elections in section 5.2 below, where we also assess the external validity of our estimates.

Since we focus upon newborns, we need to address a further potential source of bias arising from endogenous heterogeneity in the timing of fertility: through selection into either conception or foetal survival, the composition of live births in the sample may vary systematically with the gender composition of politicians.<sup>11</sup> For instance, some women may be more likely than others to consciously advance or defer fertility in response to the policy [political] regime. Alternatively, foetal death (whether through miscarriage or (female) foeticide), may vary systematically as a function of reproductive health facilities and campaigns discouraging foeticide and these facilities and campaigns may in turn be influenced by the share of women politicians. If, for instance, the incidence of female foeticide declines under women politicians then, given that foeticide is more common amongst educated mothers (Bhalotra and Cochrane 2010), the composition of (live) births that we observe under women politicians will tend to be lower risk than the sample observed under male politicians. If we do not control for this, we may attribute what is essentially a selection effect to differences in the behavior of female *vs* male politicians. We limit this problem by using mother fixed effects. These will purge all fixed mother-level unobservables including the time-invariant component of their political preferences and any tendency they may have to compensate or reinforce state investments in children (crowding-out or

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<sup>11</sup> See Dehejia and Lleras-Muney (2004) and Bhalotra (2010) who argue that selection may arise from the conscious timing of fertility and the selective survival of the foetus respectively.

crowding-in). This is a considerable improvement relative to using district fixed effects.

The third potential bias is addressed by controlling for unobservable trends. In the specification illustrated below we control for state-specific trends but in extensions of this we demonstrate robustness to district-specific trends and district by mother-cohort fixed effects. The first set of estimated models is

$$\begin{aligned}
(1) \text{ mortality}_{imdt} &= \alpha_m + \theta_t + \beta \text{ women}_{dt-1} + \sum_{j=1}^N \partial_{1j} I_{jdt} * F(m_{jdt}) + \sum_{j=1}^N \partial_{2j} I_{jdt} + \\
&\quad \lambda \text{ totalclose}_{dt-1} + \mathbf{X}_{imdt} \boldsymbol{\zeta} + \mathbf{Z}_{dt} \boldsymbol{\mu} + \boldsymbol{\varphi}_s t + \varepsilon_{imdt} \\
(2) \text{ women}_{dt} &= \chi_m + \omega_t + \tau \text{ womenclose}_{dt} + \sum_{j=1}^N \partial_{3j} I_{jdt} * F(m_{jdt}) + \sum_{j=1}^N \partial_{4j} I_{jdt} \\
&\quad + \psi \text{ totalclose}_{dt} + \\
&\quad + \mathbf{X}_{imdt} \boldsymbol{\rho} + \mathbf{Z}_{dt} \boldsymbol{\pi} + \mathbf{v}_s t + e_{imdt}
\end{aligned}$$

The subscripts denote the individual child  $i$  of mother  $m$  born in district  $d$  (in state  $s$ ) in year  $t$ . By construction, mothers do not migrate across districts between births (we investigate this restriction in section 5.2). Hence mother fixed effects nest (or imply) district fixed effects. *Mortality* is an indicator for the index child dying by the age of one month (neonatal), *women* is the district-level share of state assembly constituencies to which a woman was elected in the year before the child was born (we investigate timing below) and  $\beta$  is the parameter of interest. The other variables in equation (1), in order, are mother and (child) cohort fixed effects, a third-order polynomial ( $F$ ) in the victory margin ( $m$ ) between the winner and the runner up for every (close or non-close) man-woman election  $j$ , interacted with  $I_{jdt}$ , an indicator for whether such an election exists, *totalclose* is the fraction of constituencies in the district that experienced a close election,  $\mathbf{X}$  is a vector of child or mother-specific controls,  $\mathbf{Z}$  is a vector of district-year varying controls and  $\boldsymbol{\varphi}$  are state-specific



trends. We examine sensitivity to the definition of close elections, inclusion of victory margins, district trends, flexible trends in mother's age and cohort, and individual controls. Equation (2) represents the first stage of the IV procedure. It shows the potentially endogenous regressor, *women*, regressed upon the instrument, *womenclose*, and the same set of controls. The instrument is the fraction of constituencies in the district in which a woman won against a man in a close election.

The cohort dummies control for aggregate time-variation associated with, for example, secular improvements in health technology, episodic shocks like famines, floods and epidemics and any aggregate economic or political regime changes. The state (and, later, district) specific trends allow for regional omitted trends, for example, income or trends in the status of women. The district level controls *Z* include the shares of the population that are female, urban and low caste and male and female literacy rates. Controls *X* include dummies for the gender, birth order, rural/urban location, religion, and caste of the child and an indicator for whether the child is one of a multiple birth.

We use a linear model since fixed effects probit estimates are inconsistent in short panels (Nickell 1982) and the relevant panel in this case is the micro-panel, where *T* is the number of children per mother. Standard errors are robust to arbitrary forms of heteroskedasticity and clustered at the district-level since the regressor of interest, *women*, is entered as a district-level share. This allows for correlations at any time and across time within district (e.g. Bertrand et al. 2004). It also allows for correlation of the standard errors across siblings because, by construction, siblings are all in the same district.

*In sum*, the estimated specification delivers an estimate of the difference in mortality risk across children of the same mother born at different times and under different political regimes, with the change in political regime constructed as a quasi-random assignment of the gender of politicians in the district of birth. We investigate the identifying assumptions

that underlie the use of close elections. In view of our use of mother fixed effects, we check on the further identifying assumption that there are no omitted mother-specific trends the effects of which may load onto the variable of interest. We conduct a range of sensitivity checks including three sorts of falsification exercises and investigate heterogeneity in impact by characteristics of the child/mother and political variables including competition and decentralisation.

We further estimate equations that replace *mortality* in equation (1) with a vector of village public goods on the one hand and a vector of individual antenatal and early postnatal health-seeking behaviours on the other. We investigate fertility and the sex of births as further outcomes. The identification strategies we use to analyse these investments and behaviours are similar and are detailed below, alongside the results.

## **V. Results**

### ***A. Survival Outcomes***

Estimates of equations (1) and (2) are in Table 2. Consider effect sizes in the richest specification, which appears in column (6). Raising the share of seats held by women at the district level (*women*) by one standard deviation (0.0740) is estimated to result in declines in neonatal [infant] mortality of 1.5 [1.4] percentage points. This is 24[14]% of the mean rate in the sample, which is 6.3[9.8]% and it corresponds to a 0.063[0.048] change in the s.d. of neonatal [infant] mortality. Since the distribution of *women* is skewed (Figure A1 in the Appendix), we consider the survival improvements brought about by electing one additional woman. The median district has nine constituencies and so nine seats. As it happens, close to 10% of districts in the sample have exactly nine seats and in this set the mean (0.034) and median (0) share of seats held by women is very similar to the sample mean (0.036) and median (0). If one additional woman wins, the median number of seats goes from 0 to 1 and the share from 0 to  $1/9=0.11$ . If we plug this in to the estimated equations, we get even larger numbers. The predicted decline in neonatal [infant] mortality is

now 2.3 [2.1] percentage points. Later we show that the effects of politician gender on infant mortality are entirely determined by its effects on neonatal mortality (section 5.3).

[Table 2 here]

Table 2 describes how the estimates for neonatal mortality evolve as the specification is enriched to improve identification. The OLS coefficient is small and insignificant. There is no significant change in it when the sample is restricted to mothers with at least two births in column (2), the sample that will be used in the final specification. Controlling for mother fixed effects in the OLS specification results in a larger negative coefficient but it remains poorly determined. The 2SLS coefficients are larger and, conditional upon mother fixed effects, statistically significant. The IV-fixed effects coefficient is not significantly different from the IV coefficient (col. 4 v col. 5). Introducing the household and district level controls listed in Notes to Table 2 makes no significant difference although the coefficient is slightly larger.<sup>12</sup>

The OLS coefficient may be biased by the endogeneity of voter preferences. Our finding that it is biased downwards is consistent with women being more likely to be elected in districts with poorer baseline health. The inclusion of mother fixed effects implies that results will not reflect fixed differences between districts, but it may be that female politicians are elected in districts with declining trends in health. The difference between the IV and OLS coefficients may also reflect a difference between the local average identified off close races and the average effect. How the local and the average impacts compare depends upon the underlying theoretical model, and our findings appear consistent with a citizen candidate model. If the cost of a close race is higher than of a non-close race then a candidate needs a higher payoff to play, suggesting that only candidates with far apart preferences will play (

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<sup>12</sup> Table 2 reports only the coefficient on the regressor of interest. Appendix Table A6 reports the full set of coefficients. Given that the effect we find for infant mortality is mainly capturing the effect on neonatal mortality, the extensions below will be for neonatal mortality. Parallel results for infant mortality are available on request.

Rehavi 2007). Another possibility is that women who run in close races against men are positively selected on characteristics, having overcome greater social barriers to compete on a male-dominated terrain.

In general IV coefficients will be larger when the variable of interest is measured with error but we expect that measurement error in counting women winners in Indian elections is small. Indeed if measurement error were important then, as this is exacerbated by differencing, we would expect attenuation of the coefficient upon introduction of mother fixed effects. Instead the coefficient is larger still, possibly indicating that parents reinforce endowments in allocations across siblings, consistent with other evidence that points in this same direction (Almond and Mazumder 2013).

The first stage of the 2SLS estimates is also reported in Table 2. The instrument is powerful and the share of women who win across all elections varies proportionately with the share of women who win in close elections.<sup>13</sup> With the share of women winners in close elections constant, the share of all women winners is negatively associated with the total share of close elections within a district since this now varies only with the share of male winners. The coefficient in the reduced form of the equation system (1), (2) is -0.1847 with a standard error of 0.070.

### ***B. Robustness Checks***

We first outline checks on the close elections assumption (Appendix Tables A1-A5) and then discuss robustness of the baseline specification to specification checks (Table 3, Appendix Tables A7-A9).

**The close elections assumption:** Following the initial use of the close election identification strategy in the context of close elections between Democrats and Republicans in the US (Lee 2008), concerns have been raised that the outcomes of close elections may be biased in favour of incumbents or

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<sup>13</sup> The coefficient on the fraction of constituencies won by women in close elections against men in Panel A is 0.89, which is not significantly different from 1. But some female politicians win in elections that are not close elections against men, some win against a man by a large margin and others may win against another woman.

in favour of candidates from parties that have strategic control over the election process (Snyder 2005, Caughey and Sekhon 2011, Grimmer et al. 2011). We investigate incumbency and party affiliation in addition to other observables but find no evidence of this problem for our sample of close elections between men and women in India. A simple regression shows that the proportion of women in a district who won in a close election against a man is independent of the history of close elections in the district, the history of women being elected, the share of seats that each of eight parties contested in close elections and a host of district-time varying demographic variables (Table A1 in the Appendix). Candidate and constituency characteristics are also independent of the gender of the winner in close elections. This includes the number of other female candidates in the constituency, whether the winner was the incumbent, the number of close elections in the past, the votes received by the winner and turnout (Table A2 in the Appendix). These results are consistent with the gender of the winner in close elections between men and women being random.

Our estimates rely upon districts that had at least one close election between a woman and a man and the share of all births in the sample that occur in these districts is 7.45%. In Tables A3 and A4 in the Appendix we investigate whether districts that had a close election are matched on observables to districts that did not. The observables include district demographics, the share of villages with hospitals and schools and the share of seats won by each gender in elections that were not close (Table A3). The mean differences are small but some are significantly different from zero, which underlines the importance of our controlling for the fraction of seats in the district that had close elections between women and men. The share of constituencies won by each of the main political parties in India is not significantly different between close and non-close elections (Table A4). We nevertheless control in the regressions for the district-level seat shares of each of seven parties for the reason stated in the next section.

**Electoral variables:** Since regression discontinuity underlies the close election IV approach, we investigate parameterisation of the vote margin between the winner and runner up. The coefficient of interest is not significantly altered by dropping these controls or restricting the order of the polynomial (Table 3, columns 1-3). The premise that the gender of the leader is random in a close election only holds in a small window around the point where the vote margin between the winner and the runner up is zero, but there is an efficiency-robustness trade off in deciding the width of the window. We therefore investigate the chosen threshold for identification of close elections, varying this downwards and upwards from the selected 3.5% to include the range 2 to 4%. There is no significant change in the coefficient on *women* (columns 4-7). We restricted the sample to district-year observations (41% of the baseline sample) in which there was at least one election between a man and a woman as these are the observations upon which identification rests. This makes no significant difference (col. 8).<sup>14</sup>

[Table 3 here]

To control for the possibility that certain political parties strategically appoint women candidates more than others (Rai 2002), we condition upon the district share of seats occupied by each of seven parties.<sup>15</sup> We also control for the district share of state assembly seats reserved for low caste candidates to allow for the possibility that women and men compete differently for these seats. These changes do not alter the coefficient of interest (Table 3, panel B, column 1). Using the outer fence of the interquartile range, we removed outliers in women's representation above the 75<sup>th</sup> percentile. This

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<sup>14</sup> In principle we could implement a sharp regression discontinuity using districts in which there was exactly one election between a man and a woman but in practice we do not have enough within-mother variation in this.

<sup>15</sup> Following Besley and Burgess (2002), seven party groups are constructed: Congress, Hard Left, Soft Left, Janata, Hindu, Regional parties, and Independents coupled with other small parties (these are the parties in A4 where Independent and other are separated; also see the data appendix).

amounts to removing the 1.45% of observations with *women* greater than 28%. The estimates are robust to this (Table 3, panel B, column 2).

**Within-mother trends:** In a model with mother fixed effects, identification of calendar time effects can only come from mothers having children at different moments in time, and there is necessarily a trend in birth order and the mother's age at birth between siblings. This makes it important to isolate the effects of time-variation in women's political representation from the effects of birth-order and the age of the mother at each birth. We already control flexibly for birth order, using dummies for each order. We now introduce age at birth dummies and more flexible controls for mother's age at birth, namely, dummies for mother's cohort\*child cohort (columns 3-4, Table 3: panel B). For a given year of birth of the mother, the year of birth of the child implies the age of the mother at birth. Since the estimated equations already include fixed effects for the child's birth year, this is equivalent to controlling for fixed effects in the age of the mother at birth and allowing their impact to vary by the birth cohort of the mother. Controls for mother's age at birth are also important in correcting for the fact that retrospective fertility (and mortality) data grow increasingly unrepresentative of mother's age as we go further back in time (Rindfuss et al. 1982). See the data appendix.

**Other unobserved trends:** We investigated controls for unobservable trends in general. The estimates in Table 2 are conditional upon state linear trends. We now include, in sequence, state-specific quadratic trends, state\*year dummies, district-specific trends and district\*mother's cohort dummies (columns 5-8, panel B). The last is important because a potential problem with using only district fixed effects is that different cohorts of mothers in different districts, having been exposed to different income and health conditions, are more or less likely to have children with lower survival

chances.<sup>16</sup> Again, as we condition upon fixed effects for the child's birth year, this effectively allows for district-varying trends in the age of the mother at birth. The estimates are robust to these fairly powerful controls.

**Lags:** Since interventions motivated by politicians may operate with a lag, we investigated lags as far back as the third year before birth; see Table 4. Some persistence is built in by virtue of politicians typically holding office for 5 years, as then politicians elected 2 years before a child is born will on average be in power the year before and the year the child is born.<sup>17</sup> We find that the gender composition of politicians in the two years before birth and in the birth year ( $t-2$ ,  $t-1$ ,  $t$ ) has a similar sized significant coefficient but the share of women in power more than two years before the index birth has no impact. In India most women have a birth within two years of marriage, so this particular pattern of fading is consistent with politicians targeting women at risk of conception (newly married women) with interventions such as prenatal care information, immunization or nutrition and this showing an impact on offspring born one or two years later.

[Table 4 here]

**Placebos:** To further mitigate concerns that we are capturing omitted trends, we implemented placebos. First, we modelled the neonatal mortality risk of cohorts born in  $t$  as a function of *future* politician gender ( $t+1$  to  $t+9$ ); these coefficients are all essentially zero (Table 4). Second, we used the sample of mothers that moved to the current location a year or more after the birth of their youngest child so their births are effectively matched to politicians that they were never exposed to. Again, the coefficient on *women* is small and insignificant (Table A8 column 1). However, there is a problem with this test which is that we do not know whether the previous location was

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<sup>16</sup> See Bhalotra (2010) on the importance of mother's birth year conditions for offspring health and see Kudamatsu (2012) for a related discussion in the context of mother fixed effects estimation of the impact of democratization on mortality.

<sup>17</sup> Politicians in power at  $t-5$  will tend to exit before or in the birth year  $t$ , politicians in power in  $t-4$  will tend to exit before or in  $t+1$ , and so on.



in the same district or in a different one, and this adds noise to the estimates. A third placebo was run using as an outcome, the height of mothers. Since adult height is largely determined in early childhood (Bozzoli et al. 2009), we expect politicians in the birth year of the child to have no impact on the height of the mother, and this is what we see (Table A9 in the Appendix).

**Migration:** The estimates in Table 2 are on the sample of mothers that have had all their births in their current place of residence. This is a strong restriction. We relax it to include only children who were born in the current place of residence of the mother. In order to investigate the extent of any selectivity in migration, we also estimate the model with no restrictions. The coefficient of interest grows smaller as we successively lift restrictions on the sample but it remains significant and within two standard errors of the baseline coefficient (Table A8 in the Appendix).<sup>18</sup> This puts aside any concern that our selection of a non-migrant mother sample is either creating a bias or producing results that may be altogether unrepresentative.

### *C. Extensions*

The results discussed here are displayed in Tables 5 and 6.

**Birth order:** If one of the ways that women politicians influence health outcomes is through information campaigns then we may expect exposure to women politicians to have a larger impact on outcomes for first births. This is what we find (col. 1-2, Table 5). Aizer and Currie (2004) adduce a similar distinction between first and higher order births in attempting to isolate the role of information diffusion in women's social networks. In principle, facilities constructed around the birth of a first child will also tend to persist within mother so that they are available to her further births even if the share of women in government has changed. However, there is widespread absenteeism of medical practitioners from public health facilities

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<sup>18</sup> The smaller coefficients are consistent with migrant women and children being matched in the data to politicians who had no influence on them because at least some of the migrants (those who moved district) had moved away from their domain.

(Muralidharan et al. 2011) so, in contrast to information, the *effective* availability of facilities will tend to vary within mother.

[Table 5 here]

**Heterogeneity in household status:** Since child mortality is concentrated amongst the poor, our finding of impacts of politician gender on mortality establish the progressiveness of women in state assemblies, in contradiction with debates that currently block the institution of quotas for women at this level (see section 2). We investigated this further by allowing the effects of politician gender to differ by the caste, wealth and rural/urban location of the family (Table 5). The coefficients are larger for lower castes, consistent with their greater “treatability”. They exhibit an inverted U-shape in wealth, with the smallest effects apparent for the poorest 20% of households. A likely explanation of this is that the productivity of public intervention is contingent upon private inputs and that women in the poorest households are harder to “treat” because they have lower stocks of health before birth or because they are less likely to take up publicly provided prenatal care. It may also be that wealth is correlated with education which, in turn, is correlated with preferences for health and with the ability to access and process relevant information.<sup>19</sup> We also investigated differences by the caste of politicians. We find no difference in impact: the coefficients differ by about half a standard error (Table A8).<sup>20</sup>

**Gender-specific impacts:** Earlier studies suggest that as women’s power in decision making within the household increases marginal gains for girls exceed those for boys (e.g. Duflo 2003, Thomas 1990). So, if politician

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<sup>19</sup> The rates of of neonatal and infant mortality in the higher status groups are not small. In the high caste group, the means are close to 5.2 and 8.1 per cent and in the top-40% of the wealth distribution they are 4.9 and 7.6 per cent.

<sup>20</sup> This result stands in contrast to Clots-Figueras (2011) who finds that re-distributive policy choices are driven by low caste women (see section 1 above). The difference may arise from our focus upon reproductive and child health (which women of all social groups may be motivated to improve- refer Campbell 2004 cited in section 1) but our results are not strictly comparable since we analyse individual outcomes while the earlier study analyses state-level aggregates.

identity influences policy outcomes then it is plausible that women politicians make a larger difference for girls, for example, by campaigning against son preference. So we investigated differences by child gender (Table 5). Using the baseline specification, we find slightly larger point estimates for boys, but the difference is not statistically significant. We probed this result by including in the model proxies for son preference. First, we include a dummy for the sex of the first-born child as previous research shows that this is quasi-random, and has a discernible impact on the tendency for parents to exercise son preference (e.g. Bhalotra and Cochrane 2010). The coefficient for girls is now much larger than that for boys and statistically significant, in favour of our motivating hypothesis. However it remains insignificantly different from the coefficient for boys. We also display results using controls for the sex composition of the entire sibship surviving to the birth of the index child (for instance, one boy, one girl, a girl and a boy, and so on). The results are similar to those for the more parsimonious indicator. Note that these estimates are for the sample of children of second or higher birth order. Since controlling for the exercise of son preference does raise the point estimate for girls, we re-ran the baseline model which pools gender (Table 2 column 6) with these controls. The point estimate increases to -0.23 and -0.24 respectively which is insignificantly different from the baseline -0.21.

**Pre/post devolution:** The 1993 Constitutional Amendment described in section 2 devolved control over the implementation of development programmes to the *panchayat* level and this included aspects of health provision such as drinking water and immunization delivered through the Integrated Child Development Services programme. We therefore allow for a weakening of the influence of politicians in state assemblies post-devolution. The estimated effect is 10% larger (in absolute value) after 1993 but not

significantly different (col. 1, Table 6).<sup>21</sup> So women's political representation at the state level mattered through the three decades in our sample and their power was not substantially eroded upon the empowerment of village councils.

[Table 6 here]

**Heterogeneity by political variables:** Previous work has shown that political competition disciplines politicians, making them more responsive to the needs of citizens (Besley and Burgess 2002) but there is no evidence of whether there is any effective interaction between political competition and political identity. If male and female politicians were perfect substitutes then we would expect to observe better policy outcomes in areas with strong political competition but not as a function of the gender of the leader. However if men and women having "separate spheres" of influence driven by their different preferences, competition may enhance domain-specific gender differences in policy outcomes.<sup>22</sup> To investigate this, we interact party competition with the share of women politicians. As in Besley and Burgess 2002, we also interact turnout and newspaper circulation with the share of women politicians. The interaction terms are poorly determined but the point estimates are sizeable (Table 6). They indicate that turnout and news circulation (especially in the regional language- as also found in Besley and Burgess 2002) tend to enhance the policy influence of women, while party competition tends to lower it.

We also investigated the relevance of the gender and party of the chief minister to allow, for instance, that legislators are more likely to receive funds

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<sup>21</sup> Following Iyer et al. (2012), we created a post-reservation dummy that varies at the state\*year level reflecting the state-specific implementation of the village level reform. Interactions with this dummy are similarly insignificant.

<sup>22</sup> There are other possibilities. If women's preferences for health are a "luxury", they may be overridden in competitive conditions. If women are marginal voters, women's preferences may hold greater sway under competition. If, as lay evidence suggests, women politicians receive more newspaper coverage than their male counterparts then the discipline flowing from increased newspaper circulation may result in greater effectiveness amongst women relative to men in government.

for their constituencies or have their voice heard if their preferences are more closely aligned with those of the chief minister. We find that belonging to the chief minister's party does not alter the impact of women legislators. However when the chief minister is a woman, there is a large and significant increase in the effectiveness of women legislators (Table 6). We conducted a further interaction (not shown) with political parties classified as left and right leaning and other, but could identify no significant difference in the effectiveness of women by their party affiliation. These results underline that we identify gender rather than party effects.

**Post-neonatal mortality:** We estimated the model for post-neonatal mortality and found no effects of politician gender (Table A10 in the Appendix).<sup>23</sup> The most likely explanation of this is that the post-neonatal mortality estimates are biased by endogenous selection into post-neonatal exposure to mortality risk- children only enter this sample if they have survived the neonatal period, and we have seen that neonatal survival is dependent upon politician gender. However the estimates for neonatal and infant mortality are unbiased (the selection issue does not arise since they take all children exposed from birth). The fact that the neonatal and infant mortality estimates are of similar size suggests that reductions in neonatal mortality account for almost all of the infant mortality reduction flowing from increases in women's representation.

This is consistent with the tighter identification of neonatal mortality with public health investments in women (maternal health, breastfeeding,

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<sup>23</sup> Mortality in the first month of life is neonatal and between the second and twelfth month is post-neonatal. Infant mortality is the sum of these. Mortality risk declines sharply with age, and neonatal mortality contributes two-thirds of infant deaths. Amongst children who died before the age of 5 in our sample, 50% died before the age of one month (neonatal), 29% between one month and one year (post-neonatal infant) and the remaining 21% of deaths occurred between the ages of one and five years. The risk of death for children who survive to the age of five is considerably smaller. Since the presence of women politicians has a mortality-deterring effect, survivors at the margin are negatively selected (see for instance, Bozzoli et al. (2009) on the relative roles of selection and "treatment").

place of delivery). As children age into the post-neonatal period, they are exposed to a more diverse set of household-level influences including income, nutrition and aspects of parenting behaviour like hygiene. It becomes harder to attribute improvements in post-neonatal health to public provision and this presumably weakens politician incentives. Also, there is growing evidence that health interventions are more productive the earlier in childhood they are made (Almond and Currie 2012, Bhalotra and Venkataramani 2013): this too rationalizes investments that improve health in the first month (neonatal) as more productive than investments that improve health after (post-neonatal).<sup>24</sup>

**Fertility:** By virtue of using mother fixed effects we have already accounted for endogenous variation in the composition of births. We now investigate the impact of politician gender on the probability that a child is born, and the sex of the child conditional upon birth, instrumenting with the outcome of close elections as before. As we have the entire fertility history of each mother, we expanded the data so that for every woman there is an observation for every year in which she is exposed to the risk of birth (starting with her age at marriage and stopping at age 49 or the date of interview). We find no significant change in the probability of a birth.

Having established that the probability of a further birth (conditional upon the last observed) is not endogenous, we can proceed to estimate child sex conditional upon birth. We find that the probability that a birth is female is increasing in the share of women politicians (Appendix Table A12). A ready explanation of this would be that women politicians have been more committed to controlling female feticide, although we were unable to nail this down.<sup>25</sup> Note that to the extent that women politicians lower female feticide

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<sup>24</sup> We also investigated under-5 mortality and mortality rates up until age 14 (Table A11). The coefficients are consistently negative but statistically insignificant. The selection and attribution arguments above apply *a fortiori* here.

<sup>25</sup> To test this hypothesis, we divided the sample at 1985, because ultrasound scanners started to be imported in 1987 and thereafter widely used for prenatal sex detection and it is only after the mid-80s that there is a significant trend departure of the sex ratio at birth from the biologically normal level (Bhalotra and Cochrane 2010). Our expectation

this will create a mechanical increase in fertility which may be balancing out a decline in fertility flowing from the decline in neonatal mortality induced by raising women's representation.

## VI. Channels

In this section we investigate the mechanisms underlying our finding that neonatal mortality responds to female political representation. We analyse the influence of women's representation on a range of indicators of antenatal and postnatal care at the mother level and health infrastructure at the village level.

### *A. Individual Indicators of Antenatal and Postnatal Care*

We investigated prenatal and early postnatal health-seeking behaviours that are known to impact upon early life mortality.<sup>26</sup> These data are available for births that occur in the four years preceding the survey date (1994/5-1998/9). There are siblings in this sample (for 15.3% of mothers) but they are characterised by short intervening birth intervals and given the evidence that a short preceding birth interval compromises the health of a child (e.g. Bhalotra and van Soest 2008), this is a selected sample. So, while we use the close elections instrument as before, instead of mother fixed effects, we use fixed effects for finely defined mother-groups, an approach similar to that taken in Kudamatsu (2012). We construct fixed effects for groups defined on a multidimensional space of mother traits, namely height (above/below mean), education (above/below primary), caste (SC, ST, OBC or high caste), urban/rural location and the birth year of the mother<sup>27</sup> and include these along

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was that we would see a positive impact of female political representation on the female-male ratio at birth after and only after 1985. However the post-85 coefficients are imprecise and we cannot reject the hypothesis that they are equal before and after 1985.

<sup>26</sup> Since the samples we use for estimation of equations for survival and for health inputs are constrained to be different, we cannot conduct an accounting exercise measuring the contributions of the inputs to neonatal mortality but the inputs we consider exhibit the expected correlations with neonatal and infant survival.

<sup>27</sup> This generates about 3000 fixed effects in a sample of just under 19000 mothers.

with district fixed effects. The estimated equation system otherwise looks similar to (1) and (2).<sup>28</sup>

The results are in Table 7 and the mean and s.d. of each dependent variable is in Table 1A. As in Table 2, OLS estimates (not shown) are smaller than the IV estimates and tend to be insignificant. The table displays 2SLS estimates with and without the quasi- mother fixed effects. Here we discuss the richer specification and present effects flowing from a 10 percentage point increase in women's representation which, we explained above, is the approximate change in share associated with moving the number of women politicians elected in the median district from 0 to 1.

[Table 7 here]

**Antenatal care, iron and tetanus:** The Indian government defines a complete course of antenatal care as including at least three antenatal visits, a three -month supply of iron tablets and two tetanus injections. Only 37% of women in our sample achieve a complete course. We find that women's political representation increases the number of antenatal visits a woman makes by 0.28, which is 12% of the mean. It raises the probability that iron supplements are taken by about 10 percentage points, relative to a mean of 57%. This is important given the high prevalence of anemia amongst Indian mothers (NFHS data) and the fact that it is a risk factor for low birth weight offspring and child survival (Bhalotra and Rawlings (2011)). The probability that a woman has at least one tetanus injection during pregnancy (which is important in lowering the risk of neonatal tetanus) is increasing in *women* but not significantly.

**Breastfeeding:** Women's political representation raises the probability of early breastfeeding by 13.4 percentage points, which is a 28% increase. Breastfeeding has been shown to influence the cognitive and emotional development of children (Del Bono and Rabe 2012). Although the vast

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<sup>28</sup> The samples used in this and the following section are different from the samples used to analyse mortality. We have checked that the first-stage is still very powerful; F-statistics and reduced form equations available on request.



majority of Indian women breastfeed and 77% of children are being breastfed at 20-23 months (UNICEF indicators), there are cultural inhibitions to breastfeeding immediately after birth and only 48% of Indian women in our sample initiate breastfeeding in the first 24 hours. This is when the benefits of breastfeeding are greatest, lowering the risk of contracting infectious diseases such as diarrhoea and pneumonia and, thereby, lowering neonatal mortality risk (Jones et al. 2003).

**Place of birth:** Some 69% of births to Indian women in our sample occur at home, while the remaining 31% are distributed equally between public and private facilities.<sup>29</sup> Since home births are typically unattended by professionals, they generate neonatal survival risks that can be averted in the presence of birth attendants trained in the management of sepsis (septicaemia, meningitis, pneumonia); see Bang et al. (1999). We find that a 10 percentage point increase in women's political representation lowers home deliveries by 8.8 percentage points. Moreover, there is evidence that this represents a shift of births into government hospitals: the same increase in *women* raises the probability of giving birth in a government facility by 11 percentage points while there is no significant change in the probability of giving birth in a private facility. Contrast this with the impact of a 10% increase in state income: this similarly lowers home deliveries (by a more modest 1.1 percentage points) but this arises entirely from a shift towards private sector deliveries; there is no change in the share of births in government hospitals (Bhalotra 2010). This suggests that putting women at the helm leads to improved access to government hospitals, consistent with the results in the next section.

### ***B. Village Public Goods***

We now investigate how women's political representation influences tangible supply-side investments. For this we use the under-used village

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<sup>29</sup> The share of home births is higher in rural areas, at 78%, compared with 33% in urban areas. The share of the residual births in public and private facilities is similar in rural and urban areas.

facilities survey of the NFHS to look at rural infrastructure choices. *Facility* is an indicator for a specified facility in village  $v$  in district  $d$  in year  $t$ . Instrumenting for the district share of women as before, we estimate this linear equation system using 2SLS-

$$(3) \text{facility}_{vdt} = \alpha_d + \theta_t + \beta \text{women}_{dt-1} + \sum_{j=1}^N \partial_{1j} I_{jdt} * F(m_{jdt}) + \sum_{j=1}^N \partial_{2j} I_{jdt} + \lambda \text{totalclose}_{dt-1} + Z_{dt}\mu + \varepsilon_{imdt}$$

$$(4) \text{women}_{dt} = \chi_d + \omega_t + \tau \text{womenclose}_{dt} + \sum_{j=1}^N \partial_{1j} I_{jdt} * F(m_{jdt}) + \sum_{j=1}^N \partial_{2j} I_{jdt} + \psi \text{totalclose}_{dt} + Z_{dt}\pi + e_{imdt}$$

The estimation sample contains 1591 villages (in the 246 districts), 76% of which are surveyed in 1998 and the rest in 1997.<sup>30</sup> As in the baseline specification for mortality, *women* is measured in the year before birth. We find that village public facilities vary significantly with the share of women politicians in the district in which the village lies (Table 8). The estimates are sharp and consistent across categories. A 10 percentage point increase in *women* raises the probability that the village has a primary health centre, a community health centre, a government dispensary and a government hospital by 0.22, 0.19, 0.18 and 0.057 respectively.

[Table 8 here]

With a view to identifying the sorts of village infrastructure that male politicians are more likely to provide, we ran the regressions reported in the lower panel of Table 8. A 10 percentage point increase in the share of women politicians (which decreases the share of male politicians) significantly lowers the probability that the village has a bank. We also observe sizeable negative coefficients on indicators for post offices and telephone facilities, although

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<sup>30</sup> The reported results are conditional upon year and district fixed effects but the coefficient of interest is insensitive to dropping the district effects.

these are poorly determined.<sup>31</sup> In sum, the evidence suggests that women politicians are significantly more likely to invest in the public health infrastructure while men are more likely to invest in the financial and telecommunications infrastructure. This defends our initial hypothesis. While we control for voter (and mother) characteristics we are unable to control for politician characteristics. Therefore some part of the relationships that we identify may flow from women who win in close [political] races against men being positively selected. However our finding that women favour health *over* financial infrastructure suggests that preferences do play a role.

### *C. Discussion*

We have identified significant impacts of women's political representation on a range of indicators of prenatal care, immunization, place of birth and breastfeeding, and on different dimensions of the village-level health infrastructure. This indicates the sorts of things that politicians do towards improving public health services, suggesting the mechanisms underlying our finding that childhood mortality rates respond to women's representation.

We argue that the improvements in health seeking behaviours are consistent with both infrastructure and information improvements. An advantage of studying outcomes (does a child survive, does a mother receive iron supplements, etc.) rather than direct inputs such as infrastructure or expenditure is not only that they reflect the effectiveness of inputs but also that they capture complementary policy influences including information campaigns. Information is important in the health domain, for example, Miller (2008) finds that door to door campaigns advocating home hygiene played a critical role in the steep decline in infant mortality in American history, and the uptake of publicly provided health services in India appears to be limited

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<sup>31</sup> So as to investigate the impact of these policy choices on state income, we used state-level panel data, instrumenting with close elections. We find no impact of the share of women in state assemblies on state income, by which measure it would appear that raising the political representation of women is a cost-effective way of improving public health.

because people do not fully recognise the benefits (Banerjee et al. 2011). Information may act in either of two ways: to encourage the uptake of public services (such as antenatal care), or to encourage private inputs (such as breastfeeding). Since breastfeeding does not require public infrastructure, the identified impact of women's representation on breastfeeding directly indicates a role for information campaigns. Also, the survey data we use asked women what inhibited them in making antenatal care visits and as many as 60% said that they did not think it was necessary, which suggests the potential for information campaigns to generate large improvements in the use of antenatal care.<sup>32</sup> As discussed in section 5.3, our finding that the largest impact of women's representation is on the first birth of a mother is also consistent with a role for information.

Although it is useful to look at timing, we cannot discriminate between information and infrastructure mechanisms based upon timing alone. While building new facilities may take more than a year, it is often the case in rural India that an existing building is brought into use so that in principle infrastructure can be raised within the space of a year. Moreover, politicians can act fairly rapidly to influence staffing and supplies in facilities or to discourage absenteeism among health providers (see Chaudhury et al. 2006, Muralidharan et al. 2011) by improving monitoring and incentives. In these ways politicians may improve the effectiveness of facilities (and hence mortality rates) without any construction of infrastructure.<sup>33</sup>

It is difficult to specify the trade-off involved in (randomly) replacing a woman with a man in political office without having a full specification of the influence of each on the entire portfolio of policies, but we noted earlier

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<sup>32</sup> Access to publicly provided care is clearly relevant too since the next highest frequency, 16%, was recorded for it costing too much. Only 3.5% said it was too inconvenient or too far and 1.6% said it was because a health worker did not visit.

<sup>33</sup> There may be complementarities in the outcomes we look at and between provision of facilities and information, for example, mothers who give birth in a facility rather than at home are more likely to receive information and encouragement to initiate breastfeeding soon after birth.

that raising the share of women relative to men leads to more public health infrastructure and less rural banking infrastructure. The premise for favouring an increase in women's political participation must be that this balances the representation of women's interests against those of men.

## **VII. Conclusions**

We investigated whether women politicians invest more in antenatal care and early child health because women incur the costs of reproduction and replacement fertility and they make greater investments in the early life health of their offspring. We identify significant causal impacts of women's political representation on neonatal survival, uptake of reproductive and postnatal care, and the village level public health infrastructure. The estimated impacts are large, suggesting that raising women's political representation may be an effective and novel way of improving public health delivery in developing countries. Women's representation at the state level is a live issue in the Indian Parliament today and quotas for women in government have been established in more than 100 countries in the last 15 years (Dahlerup and Francisco 2005). Our findings are relevant to debates concerning political reservations, especially in view of recent evidence that women holding office paves the way for competitively determined participation of women in politics (Beaman et al. 2009, Bhavnani 2009, Deininger et al. 2011, Casas-Arce and Saiz 2011, Bhalotra, Clots and Iyer 2013).

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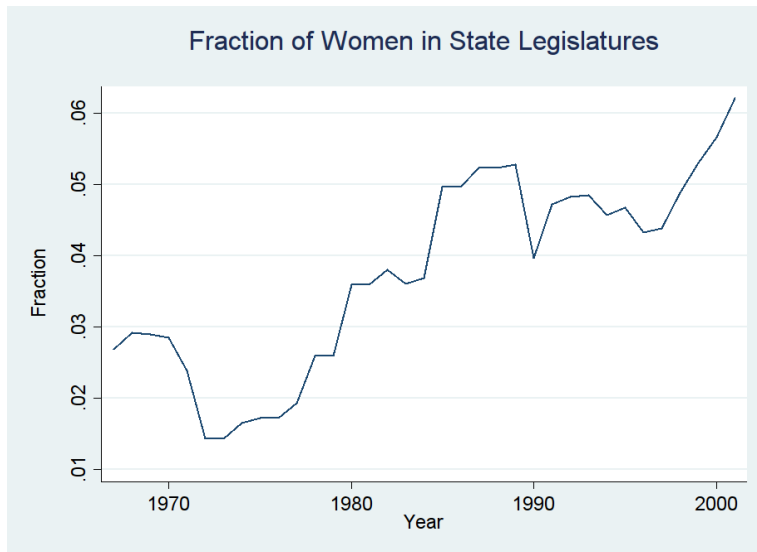
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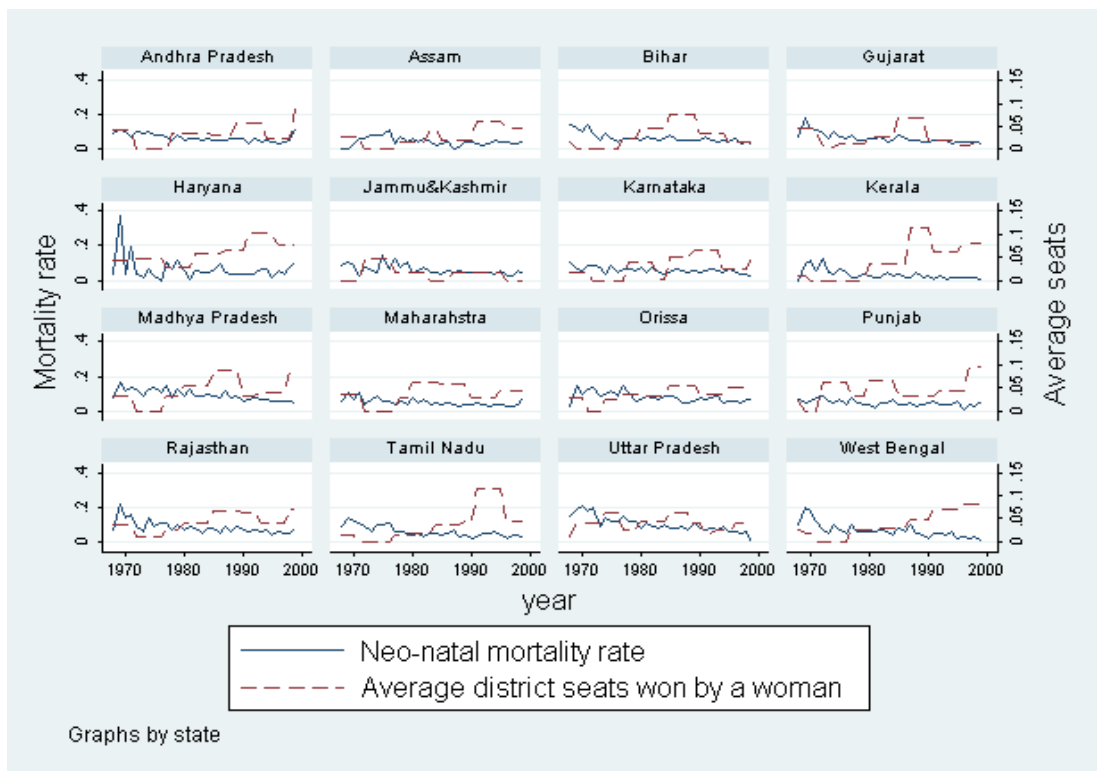
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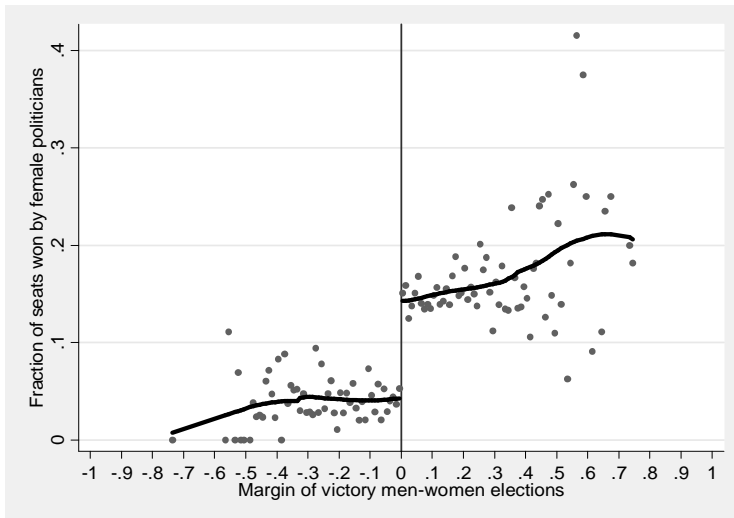
**Figure 1: Female Political Representation: All India**



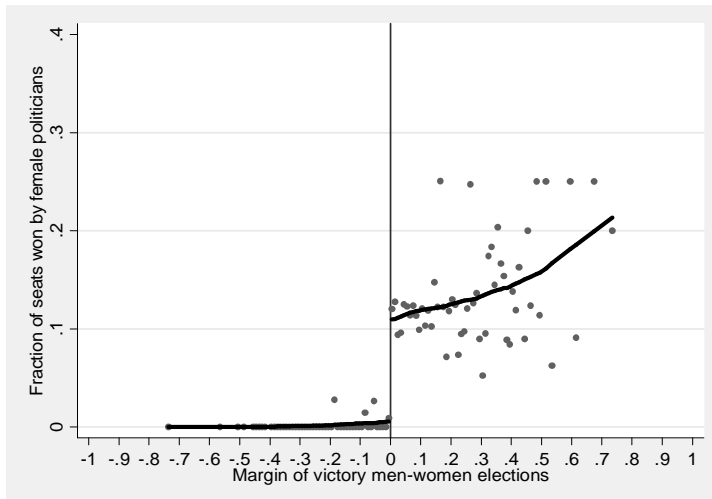
**Figure 2: Female Political Representation and Neonatal Mortality by State**



**Figure 3: First Stage Illustration.**



**Figure 3A**



**Figure 3B**

We plot the fraction of seats won by female politicians against the margin of victory in elections between women and men. Figure 3A uses all the elections between a male and a female politicians, while Figure 3B restricts the sample to those districts that had a single election between a male and a female politicians.



**TABLE 1A: DESCRIPTIVES***Unit of Observation: Individual (child), cohorts 1968-1999 in 246 districts*

	mean	sd
Neonatal mortality	0.0627	0.2423
Infant mortality	0.0980	0.2973
Birth order	2.7561	1.8099
Multiple birth	0.0138	0.1166
Girl	0.4770	0.4995
Scheduled Caste (SC)	0.1959	0.3969
Scheduled Tribe (ST)	0.1080	0.3104
Muslim	0.1070	0.3092
Christian	0.0113	0.1057
Other religion	0.0332	0.1791
Rural	0.7816	0.4131
Age of the mother at birth	22.7412	5.0420
Observations	75339	
<i>Unit of Observation: Individual (child) in the four years before the survey date (1998/9)</i>	mean	sd
Number of antenatal visits	2.3190	2.4403
Delivery at home	0.6920	0.4617
Delivery in a government institution	0.1618	0.3683
Delivery in a private institution	0.1462	0.3533
Breastfed in the first 24 hours	0.4816	0.4997
Took iron and folic acid supplements during pregnancy	0.5714	0.4949
Received a tetanus injection during pregnancy	0.7361	0.4408
Observations	8853	

Notes: National Family Health Survey, Round 2 (1998/0), summary statistics for estimation samples. The sample used to estimate childhood mortality outcomes consists of 75339 children born in 1968-1999 across 246 districts. The sample used to estimate health-seeking behaviours consists of 8853 children born in the four years before the survey date. "s.d." is standard deviation.

**TABLE 1B: DESCRIPTIVES***Unit of Observation: District in an electoral year , elections between 1967-1998 in 246 districts*

	mean	sd
Proportion of seats won by women ( <i>women</i> )	0.0364	0.0740
Proportion of districts that had at least one woman politician	0.2640	0.4409
Proportion of seats that had close elections between women and men, close: 3.5% margin ( <i>totalclose</i> )	0.0084	0.0333
Proportion of districts that had at least one election between a woman and a man	0.4025	0.4905
Proportion of districts that had at least one close election between a woman and a man (3.5%)	0.0745	0.2626
Proportion of seats won by women in close elections against men (3.5%) ( <i>womenclose</i> )	0.0041	0.0227
Proportion of districts that had at least one woman politician who won in a close election against a man (3.5%)	0.0383	0.1921
Dummy for the first election between a woman and a man ( <i>I<sub>j</sub></i> )*	0.4025	0.4905
Dummy for the second election between a woman and a man	0.1292	0.3356
Dummy for the third election between a woman and a man	0.0389	0.1934
Dummy for the fourth election between a woman and a man	0.0071	0.0841
Dummy for the fifth election between a woman and a man	0.0044	0.0661
Dummy for the sixth election between a woman and a man	0.0011	0.0331
Vote margin for the first election between a woman and a man ( <i>m<sub>j</sub></i> )**	0.0142	0.1820
Vote margin for the second election between a woman and a man	0.0307	0.2499
Vote margin for the third election between a woman and a man	0.0024	0.3002
Vote margin for the fourth election between a woman and a man	0.0115	0.3105
Vote margin for the fifth election between a woman and a man	-0.0241	0.3827
Vote margin for the sixth election between a woman and a man	0.0092	0.4649
Proportion of seats reserved for SC/ST candidates	0.2366	0.1785
Proportion of seats won by Congress parties	0.4519	0.3317
Proportion of seats won by Hard Left parties	0.0589	0.1512
Proportion of seats won by Soft Left parties	0.0231	0.0901
Proportion of seats won by Janata parties	0.1661	0.2727
Proportion of seats won by Hindu parties	0.1280	0.2290
Proportion of seats won by Regional parties	0.0772	0.2027
Female literacy rate	0.2844	0.1720
Male literacy rate	0.5352	0.1559
Proportion of SC/STs	0.2524	0.1344
Proportion of urban population	0.2041	0.1409
Proportion of female population	0.4811	0.0167
Women belonging to the same party of the Chief Minister	0.0146	0.0505
Women belonging to other parties	0.0218	0.0570
Political Competition Index (Besley & Burgess)	-0.4446	0.2105
Turnout rate	59.1858	9.5442
Female turnout rate	53.6500	11.5700
Female to male turnout ratio	0.8387	0.0876
Newspaper circulation (Besley & Burgess)	0.0599	0.0400
Newspaper circulation (Besley & Burgess) English	0.0085	0.0141
Newspaper circulation (Besley & Burgess) Hindi	0.0234	0.0251
Newspaper circulation (Besley & Burgess) Regional languages	0.0519	0.0342
Whether state Chief Minister is a woman	0.0356	0.1353
Observations	1826	

Notes: Electoral data from the Election Commission of India obtained at the constituency level and aggregated to the district level. Summary statistics for the estimation sample.

Individual health outcomes are matched to politician gender in their district of birth in the year before their birth, hence the one year difference in the year-range between the health and electoral sample.

\*The first, second, etc. election refer to the number of constituencies within a district that have a man-woman election.

So 40% of districts have at least one constituency (seat) that was won in a man-woman election, 13% of districts have at least two, and so on.

\*\*The vote margins are ordered by absolute value in districts with more than one election and now first, second, etc. margin refer to the first, second etc constituency.

**TABLE 1C DESCRIPTIVES***Unit of observation: village*

	mean	sd
Primary health subcentre	0.3517	0.4777
Primary health center	0.1418	0.3490
Community health centre	0.0815	0.2737
Government dispensary	0.1252	0.3310
Government hospital	0.0353	0.1846
Electricity	2.1797	0.6976
Post office	0.3755	0.4844
Telegraph	0.2015	0.4013
STD (telephone) booth	0.1958	0.3969
Bank	0.2831	0.4506
Observations	1558	

Notes: Data from the village file of the National Family Health Survey, 1988/9, containing information on facilities at the time of the survey.

Summary statistics for the estimation sample for the infrastructure equations.

**TABLE 2: BASELINE RESULTS**

PANEL A	Neonatal mortality					
	OLS 1	OLS 2	OLS 3	2SLS 4	2SLS 5	2SLS 6
Fraction of seats in district won by female politicians ( <i>women</i> )	-0.0050 (0.016)	-0.0148 (0.018)	-0.0303 (0.023)	-0.1176 (0.074)	-0.1942** (0.079)	-0.2061*** (0.078)
District FE	x	x		x		
Cohort FE	x	x	x	x	x	x
Mother FE			x		x	x
Controls						x
Vote Margins				3rd order polynomial	3rd order polynomial	3rd order polynomial
Observations	95,016	71,498	71,498	71,498	71,498	71,498
Number of mothers					18,754	18,754

**First Stage Regressions. Dependent variable: Fraction of seats in the district won by female politicians**

	1	2	3
Fraction of seats in district won by female politicians in close elections against male politicians ( <i>women</i> close)	0.8939*** (0.0714)	0.9011*** (0.0785)	0.8964*** (0.0789)
Fraction of seats in district that had close elections between female and male candidates ( <i>total</i> close)	-0.4303*** (0.0528)	-0.4514*** (0.0559)	-0.4495*** (0.0559)
F First Stage	156.27	131.6	129.1

PANEL B	Infant mortality					
	OLS 1	OLS 2	OLS 3	2SLS 4	2SLS 5	2SLS 6
Fraction of seats in district won by female politicians ( <i>women</i> )	0.0081 (0.019)	0.0029 (0.022)	-0.0143 (0.024)	-0.0540 (0.098)	-0.1645 (0.107)	-0.1916* (0.104)
District FE	x	x		x		
Cohort FE	x	x	x	x	x	x
Mother FE			x		x	x
Controls						x
Vote Margins				3rd order polynomial	3rd order polynomial	3rd order polynomial
Observations	91,169	68,665	68,665	68,665	68,665	68,665
Number of mothers					18,003	18,003

**First Stage Regressions. Dependent variable: Fraction of seats in the district won by female politicians**

	1	2	3
Fraction of seats in district won by female politicians in close elections against male politicians ( <i>women</i> close)	0.9202*** (0.0671)	0.9302*** (0.0744)	0.9255*** (0.0752)
Fraction of seats in district that had close elections between female and male candidates ( <i>total</i> close)	-0.4575*** (0.0507)	-0.4818*** (0.0537)	-0.4797*** (0.0540)
F First Stage	187.6	156.3	151.3

Robust standard errors clustered at the district level are reported between parentheses. \* Significant at the 10%, \*\* significant at the 5%, \*\*\* significant at the 1%. The first columns (1 and 2) show estimates with the full sample while subsequent columns show estimates for mothers with at least two births. Close elections are defined as elections in which the winner won against the runner up by less than 3.5% of votes. All regressions include district and cohort fixed effects (FE); the third and fourth columns replace district FE with mother FE. Note that mother FE include or imply district FE since mothers do not migrate across districts between births (by construction: this is investigated in section 5.2). All equations include state specific linear trends, district specific trends are included in Table 3. Controls at the district level include the fraction of the district population that is female, urban and SC/ST and male and female literacy rates. Controls at the individual level include dummy variables for gender, caste, religion, birth order, rural/urban location and for whether the birth is one of a multiple birth. All regressions shown also include as a control the fraction of constituencies in the district that had close elections between women and men (*total*close).

**TABLE 3: ROBUSTNESS CHECKS**

<b>PANEL A</b>	<b>Neonatal mortality</b>							
	1	2	3	4	5	6	7	8
Fraction of seats in district won by female politicians	-0.1473** (0.070)	-0.1902*** (0.073)	-0.1984*** (0.073)	-0.1969** (0.091)	-0.2616*** (0.088)	-0.2611*** (0.077)	-0.1422** (0.072)	-0.2289** (0.091)
Cohort FE	x	x	x	x	x	x	x	x
Mother FE	x	x	x	x	x	x	x	x
Controls	x	x	x	x	x	x	x	x
Margins	no polynomial	1st ord pol	2nd ord pol.	3rd ord pol.	3rd ord pol.	3rd ord pol.	3rd ord pol.	3rd ord pol.
<i>Robustness checks</i>	<i>victory margin</i>	<i>victory margin</i>	<i>victory margin</i>	<i>2%</i>	<i>2.50%</i>	<i>3%</i>	<i>4%</i>	<i>at least one w-m ele</i>
Observations	71,498	71,498	71,498	71,498	71,498	71,498	71,498	29,979
Number of mothers	18,754	18,754	18,754	18,754	18,754	18,754	18,754	9,989
<b>PANEL B</b>	<b>Neonatal mortality</b>							
	1	2	3	4	5	6	7	8
Fraction of seats in district won by female politicians	-0.1956** (0.078)	-0.2212** (0.086)	-0.2091*** (0.079)	-0.1886** (0.076)	-0.2001*** (0.076)	-0.2079** (0.082)	-0.1658** (0.082)	-0.1670** (0.080)
Cohort FE	x	x	x	x	x	x	x	x
Mother FE	x	x	x	x	x	x	x	x
Controls	x	x	x	x	x	x	x	x
Margins	3rd ord pol.	3rd ord pol.	3rd ord pol.	3rd ord pol.	3rd ord pol.	3rd ord pol.	3rd ord pol.	3rd ord pol.
<i>Robustness checks</i>	<i>political controls</i>	<i>outliers</i>	<i>age mother FE</i>	<i>mother cohort* child birth yr FE</i>	<i>quad trends</i>	<i>state*birth year FE</i>	<i>district trends</i>	<i>district*mother cohort FE</i>
Observations	71,498	70,581	71,498	71,498	71,498	71,498	71,498	71,498
Number of mothers	18,754	18,591	18,754	18,754	18,754	18,754	18,754	

See Notes to Table 2. Here we show robustness checks on the specification in column 6 of Table 2. Columns 1-3 in Panel A vary the functional form for the constituency-level victory margins, columns 4-7 vary the definition of close elections upwards and downwards from the baseline definition of a (constituency-level) vote margin that is less than 3.5%. Column 8 estimates the model on the sub-sample of districts that had at least one election between a woman and a man. Column 1 in Panel B includes controls for the share of votes won by each of seven political parties (see the data appendix) and the share of seats reserved for the low caste (SC/ST) population. Column 2 removes outliers in *women* (see the text). Columns 3 and 4 include increasingly flexible controls for mother's age at birth. Column 5 replaces linear with quadratic state trends, column 6 with state\*year dummies and column 7 adds district-specific cohort trends. Column 8 adds district-mother's birth year fixed effects.

**Table 4: Timing of the effects and Falsification Exercise**

<b>Coefficient: Fraction of seats in the district won by female politicians</b>					
Child's age when politician was in power	Neonatal Mortality		Child's age when politician was in power	Neonatal Mortality	
	coeff	se		coeff	se
-3	-0.0983	0.0900	4	-0.0114	0.0760
-2	-0.2257**	0.0990	5	0.0676	0.0700
-1	-0.2061***	0.0780	6	0.0247	0.0860
0	-0.1975**	0.0840	7	0.0200	0.0860
1	-0.0417	0.0830	8	0.0410	0.0950
2	-0.0338	0.0820	9	-0.0801	0.0940
3	-0.0121	0.0810			

See Notes to Table 2. The baseline specification in column 6 of Table 2 uses the first lag (-1) of the gender mix of politicians at the district level because neonatal mortality is largely determined by the foetal environment and birth conditions. That equation is re-estimated here without a lag to investigate the existence of contemporaneous effects and with two additional lags. We also investigate leads (1 to 9) of politician gender as a placebo (see the text).

**Table 5: Heterogeneity in Impact: Individual Variables. Neonatal Mortality.**

<i>PANEL A</i>	1	2	3	4	5	6	7		
Fraction of seats in district won by female politicians ( <i>women</i> )	-0.2528* (0.152)	-0.0760 (0.081)	-0.1432* (0.079)	-0.5677** (0.253)	-0.2648* (0.143)	-0.1715 (0.160)	-0.1911 (0.142)		
<i>Sample</i>	<i>first child</i>	<i>subsequent children</i>	<i>rural</i>	<i>urban</i>	<i>SC/ST</i>	<i>OBC</i>	<i>higher caste</i>		
Observations	18,562	52,936	56,273	15,225	21,827	23,134	26,537		
Number of mothers			14,401	4,353	5,416	6,085	7,253		
<i>PANEL B</i>	1	2	3	4	5	6	7	8	9
Fraction of seats in district won by female politicians ( <i>women</i> )	-0.1469 (0.138)	-0.2199** (0.112)	-0.3228** (0.153)	-0.2050 (0.156)	-0.3366** (0.159)	-0.2458* (0.143)	-0.0703 (0.179)	-0.3494*** (0.131)	-0.1219 (0.170)
<i>Sample</i>	<i>girls</i>	<i>boys</i>	<i>girls</i>	<i>boys</i>	<i>girls</i>	<i>boys</i>	<i>wealth bottom</i>	<i>wealth mid</i>	<i>wealth top</i>
<i>Control</i>			<i>gender first child</i>	<i>gender first child</i>	<i>gender siblings</i>	<i>gender siblings</i>	20%	40%	40%
Observations	27,886	31,871	17,025	18,699	17,025	18,699	19,217	30,823	20,558
Number of mothers	9,845	11,797	6,278	7,199	6,278	7,199	4,728	7,967	5,862

See notes to Table 2. The controls are those in the baseline model, col. 6 in Table 2. Columns in the top panel separate first and higher birth orders, rural and urban births, SC/ST, OBC, and higher caste births. In the bottom panel the columns separate girls and boys, and births in households in the bottom 20% of the wealth distribution, the mid 40% and the top 40%. We do not use mother fixed effects in the first two columns since a mother has only one first born child by definition.

**Table 6: Heterogeneity in Impact: Political Variables. Neonatal Mortality.**

<i>Political characteristics</i>	1	2	3	4	5	6	7	8	9
Fraction of seats in district won by female politicians ( <i>Women</i> )	-0.2043** (0.083)	-0.2241*** (0.079)	-0.2347** (0.091)		-0.1943** (0.079)	-0.1895** (0.080)	-0.2014** (0.081)	-0.2116** (0.091)	-0.2837*** (0.109)
Women interacted with a dummy for years after 1993	-0.0118 (0.119)								
Women interacted with political competition index		0.0597 (0.060)							
Women interacted with dummy for female Chief Minister			-0.8425*** (0.205)						
Women belonging to the same party as the Chief Minister				-0.2109** (0.092)					
Women belonging to a different party from the Chief Minister				-0.1995** (0.085)					
Women interacted with turnout					-0.0313 (0.066)				
Women interacted with female turnout						-0.0355 (0.074)			
Women interacted with female/male turnout ratio							-0.0265 (0.080)		
Women interacted with newspaper circulation								-0.1310 (0.205)	
Women interacted with newspaper circulation English									0.0199 (0.119)
Women interacted with newspaper circulation Hindi									0.1390 (0.214)
Women interacted with newspaper circulation Regional									-0.1547 (0.291)
P value joint significance of Women and interacted terms	0.0255	0.0162	0.000		0.0468	0.0095	0.0378	0.0568	0.0731
Observations	71,498	71,498	52,474	71,498	71,498	69,306	69,306	69,507	52,644
Number of mothers	18,754	18,754	15,664	18,754	18,754	18,589	18,589	18,258	15,639

See notes to Table 2. The controls are those in the baseline model, col. 6 in Table 2. In column 1 we interact a post-1993 dummy with the variable of interest (*women*). In column 2 we interact the variable of interest with the Besley & Burgess (2002) index of political competition. In column 3 we interact the variable of interest with a dummy for whether the Chief Minister was a woman. In column 4 we create variables denoting the share of female politicians that do and do not belong to the same political party as the Chief Minister. In columns 5, 6 and 7 we interact with turnout, female turnout and the ratio of male to female turnout. In column 8 we interact with newspaper circulation (Besley and Burgess 2002) and in column 9 we interact with newspaper circulation by language (Besley and Burgess 2002). Political competition, turnout and newspaper circulation are converted to z-scores before they are included in the model.



**TABLE 7: Health Seeking Behaviours**

Panel A	1	2	3	4	5	6	7	8
	Number of antenatal visits Iron & folic supplements				Tetanus injection in preg.		Breastfed first 24 hrs	
Fraction of seats in district won by female politicians ( <i>women</i> )	3.6524** (1.783)	2.7745* (1.662)	1.0663*** (0.364)	0.9518** (0.429)	0.1798 (0.320)	0.3113 (0.376)	1.2758*** (0.462)	1.3406** (0.535)
District FE	x	x	x	x	x	x	x	x
Cohort FE	x	x	x	x	x	x	x	x
Mother Group FE		x		x		x		x
F first stage	155.001	124.482	164.951	131.19	164.401	130.846	167.418	132.767
Observations	7,448	7,448	8,429	8,429	8,383	8,383	8,062	8,062

Panel B	1	2	3	4	5	6
	Delivery at home		Delivery gov. institution		Delivery priv. institution	
Fraction of seats in district won by female politicians ( <i>women</i> )	-0.8698** (0.433)	-0.8768* (0.522)	1.0344*** (0.346)	1.0904*** (0.394)	-0.1646 (0.344)	-0.2136 (0.350)
District FE	x	x	x	x	x	x
Cohort FE	x	x	x	x	x	x
Mother Group FE		x		x		x
F first stage	164.382	130.572	164.382	130.572	164.382	130.572
Observations	8,408	8,408	8,408	8,408	8,408	8,408

See Notes to Table 2. As in Table 2, *women* is instrumented with *womenclase*. Means of the dependent variables are in Table 1. *Mother group FE* are defined in the text, they are FE defined over multiple dimensions of mother-level observables.

**TABLE 8: Village Public Goods**

<b>PANEL A</b>					
	(1)	(2)	(3)	(4)	(5)
VARIABLES	Primary Health Subcentre	Primary Health Centre	Community Health Centre	Government Dispensary	Government Hospital
Fraction of seats in district won by female politicians ( <i>women</i> )	3.4769 (2.215)	2.2365** (1.120)	1.8871** (0.849)	1.7857** (0.900)	0.5682* (0.311)
Year dummies	x	x	x	x	x
District dummies	x	x	x	x	x
Observations	1,558	1,558	1,558	1,558	1,558
R squared	0.326	0.342	0.354	0.370	0.210
<b>PANEL B</b>					
	(1)	(2)	(3)	(4)	(5)
VARIABLES	Electricity	Bank	Post Office	Telegraph	STD (Telephone) Booth
Fraction of seats in district won by female politicians ( <i>women</i> )	0.9865 (1.225)	-2.4616* (1.494)	-2.1704 (2.218)	0.6037 (1.241)	-0.9293 (1.589)
Year dummies	x	x	x	x	x
District dummies	x	x	x	x	x
Observations	1,558	1,558	1,558	1,558	1,558
R squared	0.455	0.280	0.306	0.231	0.238

See Notes to Table 2. As in Table 2, *women* is instrumented with *women<sub>close</sub>*. All regressions include districts and year dummies and population controls. The means of the dependent variables are in Table 1. The F-statistic of the first stage regressions is 45.973.