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Type of Work Matters: Women's Labor Force Participation and the Child Sex Ratio in Turkey

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Summary. — Province-level census data for 1985 and 1990 from Turkey are used to examine the effect of the economic value of women on the 0–9 cohort population sex ratio. The sex ratio is interpreted as an indicator of the relative life chances of girls, reflecting their relative neglect in the household. We test two hypotheses: women's labor force participation lowers the sex ratio, and women's "gainful work outside" is more effective than the not-directly remunerative work in an integrated family production system in lowering the sex ratio. We find an inverse relationship between women's labor force participation and the sex ratio, but this effect is present only where women engage in unpaid family labor. © 2000 Elsevier Science Ltd. All rights reserved.

Key words — Middle East, Turkey, sex ratio, women's labor force participation, unpaid family labor, spatial econometrics

1. INTRODUCTION

Social scientists have paid considerable attention to the shortfall of women in the population and have often traced it to higher mortality rates among women in comparison with men (Visaria, 1969). This outcome is attributed to discriminatory practices against young women and girls in intrahousehold distribution of resources, female infanticide, or sex-selective abortion (D'Souza & Chen, 1980; Chen, Huq & D'Souza, 1981; Das Gupta, 1987; Patel, 1989; Yi *et al.*, 1993; Park & Cho, 1995). Significant crossregional variations in either the mortality or population sex ratios (number of males per hundred females) have also compelled researchers to pursue comparative analysis to identify the factors that explain the relative shortage of women (Bardhan, 1974, 1988; Miller, 1981; Dyson & Moore, 1983; Sen, 1989, 1990). Many studies have focused on the lower economic value of women's labor as an explanation for these gender discriminatory practices. The objective of this study is to examine the effect of crossprovince variations in the economic value of women, as reflected in the extent and type of women's labor force participation, on the relative life chances of girls in Turkey.

The overall sex ratio in Turkey does not reflect a significant deficit of women in the population, especially in comparison with the alarming figures from India and China. The 1985 and 1990 Population Censuses report population sex ratios of 102.72 (SIS, 1991) and 102.66 (SIS, 1993). [1] In the 0–9 age group, however, which is the age cohort most sensitive to gender discriminatory practices in intrahousehold allocation of resources, there is a shortfall of girls. In 1985, for instance, the sex ratio of this cohort was 105.17, higher than the expected ratio of 103.71 derived from the "East" model life table (Coale & Demeny, 1983), which is the standard reference of Turkish demography. [2] More significantly, this figure conceals wide variations across provinces and a distinct geographical pattern. As observed in Figure 1, the sex ratios are lower in the economically more developed West/Northwest and higher in the underdeveloped East/Southeast provinces, ranging from 101.10 (in 1985) in the Western Black Sea coast to 111.80 in the Southeastern corner. This study

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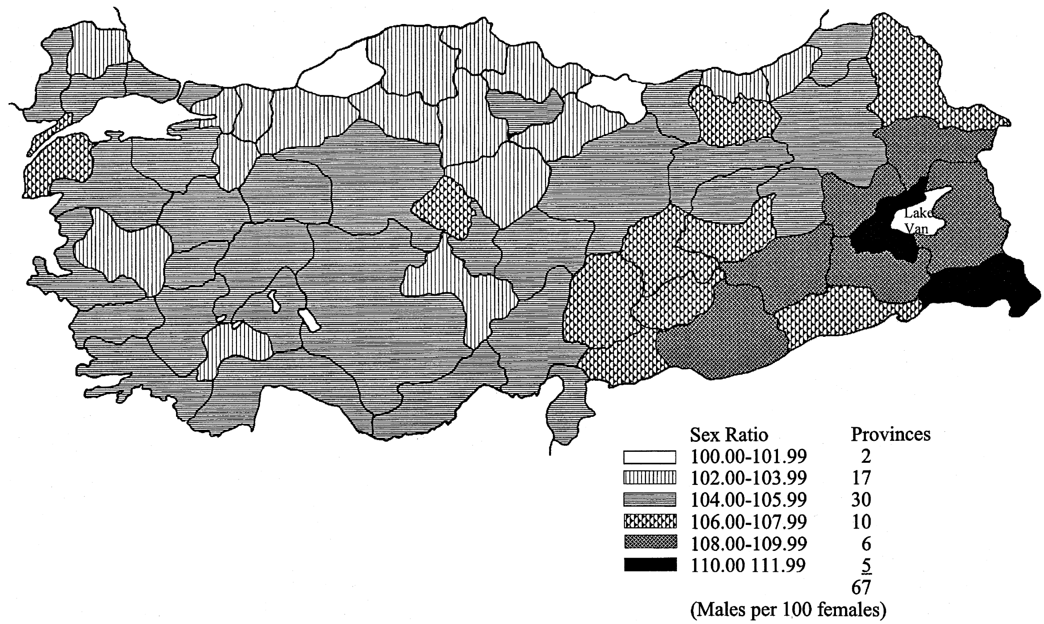

Figure 1. *0–9 age cohort population sex ratio by provinces (1985). Source: SIS (1989).*

examines the socioeconomic determinants of this geographic variability in the child sex ratio.

We follow the macro methodology pursued in a number of studies on this issue, which infers the extent of relative neglect of girls from the cross-sectional population or mortality rate sex ratios and investigates the relationship between these ratios and socioeconomic factors. The alternative methodology, which is increasingly applied in studies on gender differences in well-being, is the microeconomic analysis of intrahousehold resource allocation based on data from large-scale household surveys. [3] At present, data restrictions prevent us from pursuing the latter approach. We view the macro methodology as complementary to the micro studies in generating hypotheses for future household-level research.

This study has three main contributions: first, we differentiate among types of women's work and examine the impact of each on the sex ratio; second, we explicitly model the spatial dependence pattern among neighboring provinces observed in Figure 1; third, we draw attention to Turkey, a country that has received little attention in this literature. Our empirical analysis shows that the sex ratio is inversely related to women's labor force participation rate, which is consistent with the results of the few such studies (Rosenzweig & Schultz, 1982; Humphries, 1991; Kishor, 1993; Murthi, Guio & Dreze, 1995). The present study also reveals the unexpected result that this effect is due to women's unpaid family labor, which is the characteristic form of women's work under smallholder agriculture.

The paper is organized as follows. Section 2 presents a review of the literature on the relationship between women's economic value and well-being. Section 3 discusses the nature and limitations of Turkish data on gender differences in children's well-being. In Sections 4 and 5 the empirical model is presented and the results of crossprovince regression analysis of the association between women's economic value and the child sex ratio are reported. The final section discusses the conclusions and implications of the empirical analysis.

2. ECONOMIC VALUE AND WELL-BEING OF WOMEN

In explaining the trends and variations in the gender differential in child mortality or population sex ratio, much discussion has focused on the economic value of women measured by either their labor force participation or earnings. A wealth of studies trace excess female mortality to a general preference for sons (cf.D'Souza & Chen, 1980; Kynch & Sen, 1983; Das Gupta, 1987; Muhuri & Preston, 1991), which in turn is attributed to either higher expected returns to the labor of male over female children or anticipated old-age support from sons within the patrilineal kinship system (Miller, 1981; Das Gupta, 1987).

In an early exploration of the puzzling regional variations in life expectancy and mortality in India, Bardhan (1974) has linked the variations in relative survival chances of girls to the level of female participation in the cultivation of major crops. This view is consistent with Boserup's earlier argument (Boserup, 1970) that relates the low economic contribution of women in agriculture to their low social status and the prevalence of dowry. Subsequently, Bardhan (1988) has reformulated this hypothesis to incorporate Miller's (1981) research that links the regional variations in the cost of marrying daughters to their systematic neglect in the household: dowry payments arise so as to compensate for the lower economic value of women, and the prevalence of dowry (as a cost to fathers) is associated with the relative neglect of girls. [4]

The association between women's economic value and gender differentials in mortality has been confirmed by a number of studies (Rosenzweig & Schultz, 1982; Humphries, 1991; Kishor, 1993; Murthi *et al*., 1995). These studies have all estimated reduced form equations that were not equipped to explain the mechanisms whereby greater labor force participation of women leads to a more gender-equitable mortality outcome. Nor have they differentiated among different household models. Nonetheless, since the observed well-being outcomes can be consistent with more than one model, there has been debate over the intrahousehold process involved. In the case of gender differences in mortality among children, Rosenzweig and Schultz (1982) use the unitary model of the household to show that the well-being of girls improves where women have greater earning opportunities. They attribute this outcome to the economic rationality (i.e., investment concerns) of parents who allocate a larger share of household resources to those children who are expected to contribute more (in pecuniary or non-pecuniary terms) to the household. Folbre (1984) provides an alternative explanation for this finding that posits a gender difference in parental preferences. Accordingly, women whose employment opportunities increase may use their enhanced bargaining power to allocate more resources to their daughters, because they stand to benefit relatively more from daughters' (rather than sons') assistance in household tasks.

In an alternative framework, Sen (1989, 1990) argues that what enhances women's share of household resources is the type of work they perform, not their high level of participation in the labor force. Sen (1990) emphasizes the role of power differentials in shaping intrahousehold resource allocation, and specifies both the process whereby bargaining power would increase and the key variables that are likely to initiate this process. He proposes that access to "gainful work outside" will strengthen women's relative bargaining power in the household and that this could happen via three mechanisms: strengthening women's fall-back position; increasing their own as well as others' awareness of the value of women's work; or increasing women's awareness of the importance of their own well-being.

In Sen's discussion women's participation in "gainful work outside" is the "wedge" that promises to disrupt the perpetuation of gender inequalities. He argues that "gainful work outside" provides women easier access to income, social respect, legal protection, or exposure to the outside world, and these, in turn, enhance women's bargaining power in the household (Sen, 1989, pp. 22–23). By contrast, women's work within an "integrated system" of production (whether it is housework, subsistence or market-oriented work), where individual contributions may not be identifiable, is unlikely to be recognized as valuable "regardless of the amount of time and effort expended" (Sen, 1990, pp. 136, 139–140). Sen's argument implies that women's labor in family production (characteristic of smallholder agriculture across much of the Third World) is not gainful work, because it is commonly viewed as an integral part of their household chores. [5] Indeed, the farm is considered part of the territory of "home." This view is shared by others in this literature who presume that work in agriculture, often carried out in the context of kinship relations, allows the least autonomy and self-worth for women (cf. Preston, 1976; Kumar, 1989).

Sen also points out that increasing awareness of women's economic contributions and recognition of their entitlement to a larger share of household resources will have a favorable intergenerational effect on girls' share of resources *vis-à-vis* boys. While he does not elaborate on the parental decision-making process, one can speculate on the implications of his approach for the well-being of girls *vis-à-vis* boys. One scenario could be a case of differing preferences of fathers and mothers toward their sons and daughters as proposed by Folbre (1982). In this case, girls' share of household resources is expected to be higher where women, whose bargaining power increases due to gainful work outside, will allocate more household resources to their daughters. Evidence for gender differences in parental preferences toward children is sparse, however. [6] Under a second scenario, where there is no gender difference in parental preferences toward boys versus girls, daughters' share of family resources will increase as a result of growing awareness of women's economic contributions brought on by the availability of gainful work outside. Given the macro methodology pursued in this paper, we will not be able to analyze the mechanisms by which gainful work affects the relative life chances of girls. We will treat the issue of whether type of work matters as an empirical question and estimate whether it has any effect on the sex ratio. This brings us to the question of how to operationalize Sen's hypothesis on type of work.

In his discussion of empirical evidence, Sen fails to draw a clear distinction between gainful work outside and other kinds of work, which leaves his hypothesis on type of work untested. His contrasts between bargaining power of women lacemakers and beedirollers in India, for example, do not seem to be related to type of work, which in both cases is homebased production (Sen, 1990, pp. 144–145). Likewise, in drawing attention to broad regional contrasts in relative life expectancy of women, Sen uses the activity rate (i.e., labor force participation rate) as a proxy for gainful work outside (Sen, 1990, pp. 145–146). The measured labor force is, however, a heterogeneous category that encompasses both gainful work outside and other kinds of work. As revealed by the status in employment breakdown in population censuses, the measured labor force comprises of not only employees, employers and self-employed workers, but also unpaid family workers, who are overwhelmingly the female relatives of male employers or self-employed workers. These women have well-defined and regular stations in market-oriented family production but do not receive payment for their labor. These employment statuses are dissimilar in terms of the kinds of work relations, forms of remuneration, and their potential for social recognition as economically valuable. Variations in census practices notwithstanding, in many countries unpaid family workers account for a substantial proportion of female economically active population (ILO, 1995). Thus, Sen's hypothesis about gainful work outside cannot be tested by the activity rate (unless a particular country's census practice excludes unpaid family workers from the measured labor force).

In this study, we use the status in employment categories in population censuses to operationalize Sen's hypothesis that only gainful work outside or "outside work" improves the relative survival chances of girls. Following Sen's definitions, outside work is interpreted here as individuated, directly remunerative activities carried out outside the home, often in the context of non-kinship relations. We will call its obverse "inside" work, and along with Sen, expect that it will have either a smaller or no effect on the child sex ratio. In addition, we test the general hypothesis that the greater economic contribution of women (measured by their labor force participation rate) improves the relative survival chances of girls.

3. GENDER BIAS IN CHILDREN'S WELL-BEING IN TURKEY

Studies that pursue the macro methodology use mortality data broken down by sex as an indirect indicator of relative neglect of girls. [7] Unfortunately, there are no reliable province-level mortality data for Turkey, which leads us to use the population sex ratio as an indicator of female disadvantage in mortality. While Turkish Censuses collect data on children ever born and children surviving to each mother, they do not do so by sex of the child, which precludes separate evaluation of male and female mortality rates by province. Death registration statistics compiled by the State Institute of Statistics (SIS) are also unusable due to incomplete coverage, underestimation, and missing demographic information (Shorter & Macura, 1982; Bulut, Gokcay, Neyzi & Shorter, 1992). While National Demographic Surveys of 1988 and 1993 provide potentially better quality data on mortality, not only are these data unavailable at the province level, but survey samples are also too small to be broken down by provinces. [8]

On the other hand, there is ample evidence for son preference in Turkey. Survey data indicate that women have a stronger desire for having sons over daughters, although the intensity of son preference has declined over time. [9] Both women and men prefer having sons for economic reasons, but men also express a strong desire for having the family name continue (Kagitcibasi, 1982). Women's status in the family is closely related to bearing male children, peaking in extended families after the marriage of sons (Aksit, 1990). There is also evidence that son preference in rural Turkey has a strong positive effect on fertility rates. Rural women whose first live birth is a daughter have higher number of live births than women whose first live birth is a son (Ulusoy, 1986), and the number of dead male children has a positive effect on fertility rates (HIPS, 1989). Furthermore, a large proportion of women report that they have more children than they desired. In 1988, this proportion was a high of 61% in the 45–49 age group, but even among women under the age 20, 23% had more children than they desired (HIPS, 1989, p. 70). Thus, on the basis of this evidence one can form a presumption of relative neglect of daughters in families with a large number of children, especially those female children who are the unintended outcome of a fertility strategy of having surviving sons.

More direct evidence for this presumption, however, is not available. There are no studies on intrahousehold distribution of food and general care that examine differential treatment by age or sex, and results of earlier studies on medical care-seeking behavior are inconclusive (Aksit, 1989, 1990; Cerit & Unalan, 1988). The data on relative child mortality, morbidity and physical growth in the Turkish Demographic and Health Survey (TDHS) of 1993 are likewise inconclusive on the question of gender bias. First, while both post-neonatal and child mortality rates of girls are higher than that of the boys, the difference is not significant (Hancioglu, 1994, p. 84). Hancioglu suggests that the gender difference may be due to childcare practices that favor boys. Second, inquiry into provisioning of health care shows an advantage for girls, with the exception of treatment for diarrhea (Biliker, Haznedaroglu & Emiroglu, 1994). One possible explanation for the inconsistency between the mortality and health care/treatment data is that girls were at a disadvantage in receiving treatment earlier, but that the gender gap was closed over the decade. Third, anthropometric evidence indicates that there is little difference between boys and girls under the age of five in terms of nourishment. According to the TDHS data, 19.1% of boys and 18.7% of girls are chronically undernourished (i.e., stunted); 9.3% of boys and 9.8% of girls are underweight. Girls are slightly more likely to be "severely" stunted and underweight (among girls 6.3% and 1.9%, respectively, versus 5.4% and 1.7% among boys) (Tuncbilek, 1994). While these authors report significant variations in mortality, morbidity and nutritional status in terms of variables such as education level of parents, region, birth order and interval, they do not present the gender breakdown of these variations.

Notwithstanding the dearth of micro-level analyses on gender bias in the care of children and the inconclusive nature of the existing micro evidence, the variation in province-level sex ratios illustrated by Figure 1 still poses a challenging puzzle that warrants an investigation of its determinants.

4. THE EMPIRICAL MODEL

The dependent variable in our model is the 0–9 cohort sex ratio. The sex ratios for this cohort obtained from the 1985 and 1990 Censuses and the 1988 and 1993 Demographic Surveys are very close (105.17, 105.32, 105.52, 105.56, respectively) and, therefore, we have greater confidence in the reliability of the Census data for this age group. The focus on this age group helps to filter out the effect of sex-selective interprovince migration as well. The likely pattern of migration is that the father moves first, and the mother and dependent children join him later. Thus, for the older age groups, one may expect the number of men in regions with greater work opportunities to be larger relative to women, and, by the same token, the population sex ratio to be lower in the outmigration provinces. This bias is avoided by choosing the sex ratio among children as the dependent variable. [10] The 0–9 aggregation also circumvents the age misstatement problem of the Turkish Censuses wherein the size of the 0–4 age group is understated and that of the 5–9 group is overstated for both girls and boys (SIS, 1995, p. 97). [11]

The reliability of the data could also be questioned in light of the notion that girls tend to be underreported in the census data in the less-developed Eastern provinces, where parents are said to regard only sons as children. There is no documentation of this claim. If true, however, this could affect the interpretation of the reported regression coefficients and, therefore, requires further attention.

This potential underreporting problem is not fatal for our analysis because our objective here is not to obtain an accurate count of the shortfall of girls, but to detect the relationships between the shortfall of girls and the socioeconomic characteristics of provinces. We estimate the marginal impact of a host of explanatory variables on the level of the sex ratio, rather than try to obtain an accurate measure of the sex ratio itself. Our empirical methodology controls for and isolates several possible kinds of measurement errors as follows. Should the measurement errors in the sex ratio take the forms of either *random* or *uniform* underreporting of girls across provinces, then the intercept and the error term of the regression equation in the former case and the intercept in the latter case would absorb them. If the underreporting of girls in Turkey is a *regional* problem, primarily concentrated in the Eastern provinces, then we will attempt to control for this effect, alternatively, by employing a dummy variable or deploying appropriate econometric techniques to model spatial dependency. In all three types of measurement error, the sex ratio can serve as the dependent variable in estimating the marginal impact of explanatory variables on the excess mortality of girls.

If, however, underreporting is systematically linked with either some other explanatory variable in the empirical model (e.g., if it is more widespread where women's labor force participation is lower) or an omitted variable that is correlated with the explanatory variables, then the interpretation of the estimated coefficient would change. In this case, the estimated regression coefficient measures the impact of the relevant variable on the shortfall of girls, which may be due to underreporting, excess mortality or a combination of the two. Then, all of the "missing" girls cannot be proclaimed dead; some may have been deemed not worth counting. Undoubtedly, not being counted as children is an enormously less grave predicament than being dead. As Hull (1990, p. 76) points out in the case of China, however, it is still symptomatic of discrimination against girls, having serious consequences in terms of conditions of life, access to schooling and general care.

Lacking province-level earnings data broken down by sex, we use women's labor force participation rate (LFPR) as the measure of their economic value. Women's LFPR is defined as the ratio of the number of women in the labor force to the number of women in the 12–64 age group. We presume that the greater the actual or expected economic contribution of daughters, the larger is their share of household resources, and hence, the greater is their relative chance of survival. [12] Obviously, this is a rather crude measure since it does not indicate the extent or intensity of work of those in the labor force.

In order to operationalize the outside/inside work distinction, we disaggregated the female labor force into its unpaid and paid components, on the basis of status in employment information in the Censuses. We define "unpaid labor" as unpaid family labor, while "paid labor" comprises of employees, employers, and self-employed. Table 1 shows the breakdown of female labor force by status in employment in 1985 and 1990. In 1985 79% and in 1990 73% of women in the labor force were engaged in unpaid labor, and most of these were family workers in agriculture. [13] Women's paid labor is more common in the urban industry and service sectors and its relative importance has increased in the 1980s due to rapid internal migration and urbanization. [14] Besides testing Sen's hypothesis concerning the effect of type of women's work, given the high degree of overlap between unpaid and agricultural work, our work categories also provide an indirect test of the widely-held presumption that agricultural work/rural life is inimical to the well-being of women and girls. As an alternative proxy for the outside/inside work distinction, we also used wage and nonwage labor categories, where wage-work corresponds to the employees (salary- and wage-earner), and nonwage work

Table 1. *Distribution of women labor force by type of work (percentage of the women labor force)*[a]

	1985	1990
Unpaid labor (unpaid family labor)	79.22%	72.66%
Paid labor	20.78	27.34
Employee (wage or salary earner)	14.02	17.21
Employer	0.14	0.22
Self-employed	4.59	7.08
Unemployed	2.02	2.83

[a] *Sources*: SIS (1989, 1993).

comprises of the self-employed, employers, and unpaid family workers. Correlation coefficients between paid and wage labor force, and unpaid and nonwage labor force participation rates are in excess of 0.99 and in regression analysis the two classifications yield similar results. [15]

The conditioning set includes variables that measure the level of economic development, educational attainment, availability of health services, and household structure. The choice and definition of variables are informed by the available literature but inevitably circumscribed by data availability at the province level.

We use GDP per capita as an indicator of a province's level of economic development. If discrimination against daughters lessens in the course of economic development because of removal of resource constraints, as proposed by the early research on the sex ratio (cf. Preston, 1976), then this variable is expected to be negatively related to the child sex ratio. Apart from the well-known weaknesses of GDP per capita as a measure of economic development, however, there is considerable cross-sectional and time-series evidence that challenges this optimistic view of the effect of economic development. These studies find a positive relationship between capitalist development, wealth and income on the one hand and the sex ratio on the other (Kynch & Sen, 1983; Das Gupta, 1987; Sen, 1990; Humphries, 1991; Murthi *et al.*, 1995).

The general literacy rate (the number of literate divided by the population above age six) is an alternative measure of development that goes further than GDP per capita to capture the effectiveness of development interventions (i.e., access to/effectiveness of schooling). We expect the relative survival prospects of girls to improve with rising literacy rates. In contrast to much of the literature, we do not distinguish between the male and female educational attainment levels because province-level male and female literacy rates in Turkey are highly correlated ($r = 0.99$) and the data cannot distinguish between the separate effects of sex-specific literacy rates on the sex ratio.

The availability of health services is expected to be negatively related to the sex ratio. Here, availability of health services is interpreted as an indicator of their greater accessibility (lower cost) to parents and is measured by the number of hospital beds per 1,000 population. [16] Availability of medical facilities may not, however, reduce gender bias in health-seeking behavior if parents are less likely to use such facilities for girls as is shown by Bardhan (1988) and Murthi *et al.* (1995).

Finally, average household size is used as a proxy for the number of children in the household and tests two alternative hypotheses. The first is that the larger the number of children in the household the greater the preferential treatment of sons over daughters in access to household resources, especially where several daughters are the outcome of a fertility strategy driven by son preference. Micro studies indicating selective discrimination against higher birth-order girls inform this hypothesis (Das Gupta, 1987; Muhuri & Preston, 1991). One could also argue, however, that larger household size increases the available resources disproportionately through specialization. A higher level of resources, in turn, may make more of them available to daughters. The net impact on the sex ratio will depend on the relative importance of these two factors.

Basic statistics of variables used in the regression analysis are reported in Appendix A.

5. ESTIMATION

(a) *Regression model*

The unit of analysis in this study is the province, which is an administrative unit whose boundaries are more or less arbitrarily drawn. Since provinces are not behavioral units but aggregates of behavioral units (such as households or individuals), it is inappropriate to use crossprovince regression estimates to make individual-level inferences. Consequently, caution against ecological fallacy is warranted. The present analysis is valuable in detecting associations between the variations in the intensity of discrimination (i.e., the sex ratio) and the socioeconomic environment, and identifying an aggregate causal structure that may underlie individual-level analysis.

Figure 1 indicates that not only does the sex ratio rise as one moves from the North/Northwest to the East/Southeast, but there is also a clustering of the sex ratio such that a high sex ratio province tends to have similar neighboring provinces. Introduction of the dummy variable for the East/Southeast provinces may capture, at least partially, this pattern. [17] Alternatively, the observed spatial dependence may be modeled directly, in one of two forms. Measurement problems associated with the arbitrariness of geographic delineation and aggregation of provinces, spatial spillover effects, or omission of spatially correlated explanatory variables may cause spatial correlation in errors. This type of spatial dependence, called *spatial error autocorrelation*, makes the ordinary least squares (OLS) estimates inefficient. It can be corrected by adding an autoregressive process whereby the error term depends on its spatially lagged value and a random error term. The second type of spatial dependence, called *mixed regressive spatial autoregressive model*, is a more substantive problem that indicates the presence of a pattern of interaction whereby neighboring provinces influence each other's "behavior." This spatial lag in the dependent variable renders the OLS estimates biased and inconsistent (Anselin, 1988, pp. 8, 58–59; Case, 1991).[18] This functional relationship between the dependent variables across provinces is modeled by including the spatially-lagged sex ratio (i.e., the weighted average of the values of the sex ratio in neighboring provinces) among the right-hand-side variables of the regression equation. Since we do not have theoretical priors on the type of dependence that applies in the present context, the model is estimated under both specifications.[19]

The empirical model is:

$$\begin{aligned}
&\ln(\text{Sex ratio})\\
&= \alpha_0 + \alpha_1 \ln(\text{Per capita GDP})\\
&\quad + \alpha_2 \ln(\text{Literacy rate}) + \alpha_3 \ln(\text{Household size})\\
&\quad + \alpha_4 \ln(\text{Hospital beds per 1000})\\
&\quad + \alpha_5 \ln(\text{Women's LFPR variable})\\
&\quad + \alpha_6 \text{East/Southeast dummy}\\
&\quad + \rho \boldsymbol{W} \ln(\text{Sex ratio}) + \lambda \boldsymbol{W} u + v
\end{aligned}$$

where $\boldsymbol{W}$ is the spatial weights matrix whose elements reflect potential spatial interaction between two observations on the basis of contiguity of provinces; ρ and λ are spatial autoregressive coefficients for the spatially-lagged dependent variable and the error term, respectively; and u and v are spatial and non-spatial province-specific error terms, respectively. All variables are in natural logarithms because the dependent variable is a ratio and therefore asymmetric around the reference value. With logarithmic transformation, a deviation from a reference point becomes equidistant in either direction (Fossett and Kiecolt, 1991, p. 945). The regression coefficients are then elasticities of the sex ratio with respect to the explanatory variables. We estimate this equation under alternative assumptions of $\rho = \lambda = 0$, $\alpha_6 = \lambda = 0$ and $\alpha_6 = \rho = 0$ in order to evaluate the stability of the estimates. Data come from 1985 and 1990 Censuses, with 67 and 73 province observations, respectively.[20]

(b) *Estimation results*

The regression results are reported in Tables 2–4. In Table 2, women's LFPR is measured in the aggregate without distinguishing between inside and outside work. In Tables 3 and 4, women's LFPR is disaggregated into its unpaid and paid labor components in order to estimate their individual effects on the sex ratio. The first three columns of Table 2 report estimates based on the 1985 data. The OLS regression (column 1) ignores spatial dependence. The second and third regressions assume spatial-lag and spatial error, respectively, and both are estimated by the maximum likelihood (ML) method.[21] The remaining three columns repeat estimations with the 1990 data.

In the case of the OLS estimation based on the 1985 data, the adjusted coefficient of determination indicates reasonable statistical fit. Consistent with expectations, the estimates reported in column 1 of Table 2 indicate that girls' relative survival chances are poorer in provinces that have lower literacy rates, lower women's LFPR, and that are located in the East/Southeast region ($p < 0.01$). The sex ratio is also inversely related to GDP per capita and household size but these have only marginal levels of statistical significance. The 1990 OLS estimates (column 4) are similar to the 1985 results, except that the magnitudes of the estimated coefficients and t-statistics decline overall. In fact, GDP per capita and household size are no longer statistically significant at the conventional levels, and the number of hospital beds is marginally significant. The relationship between the sex ratio and the other three variables remain robust, however. The sex ratio declines with rising literacy rate and women's LFPR, and is higher in the East/Southeast provinces than in the rest of the country.

These estimates, however, are not free from problems. Even after controlling for the regional effect with the East/Southeast dummy, the Lagrange Multiplier (LM) tests indicate the presence of spatial dependence. In fact, the observed heteroskedasticity may be a symptom of this problem. Comparisons of the

Table 2. *Regression model for 0–9 age cohort sex ratio (women's unpaid and paid LFPR not separated)*[a]

	1985			1990		
	No spatial-dependence ($\lambda = \rho = 0$) (OLS) (1)	Spatial Lag ($\alpha_6 = \lambda = 0$) (ML) (2)	Spatial Error ($\alpha_6 = \rho = 0$) (ML) (3)	No spatial-dependence ($\lambda = \rho = 0$) (OLS) (4)	Spatial Lag ($\alpha_6 = \lambda = 0$) (ML) (5)	Spatial error ($\alpha_6 = \rho = 0$) (ML) (6)
ln(GDP per Capita)	−0.0126	−0.0114	−0.0091	0.0051	0.0002	0.0029
	(1.70)*	(1.82)*	(1.42)	(1.24)	(0.05)	(0.72)
ln(Literacy rate)	−0.0569	−0.0427	−0.0669	−0.0349	−0.0184	−0.0371
	(2.86)***	(2.60)***	(2.99)***	(2.13)**	(1.25)	(2.03)**
ln(Household size)	−0.0312	−0.0215	−0.0019	0.0193	0.0173	0.0403
	(1.75)*	(1.43)	(0.09)	(1.21)	(1.21)	(2.56)**
ln(Hospital beds	−0.0075	−0.0037	−0.0022	−0.0053	−0.0040	−0.0022
per 1000)	(2.12)**	(1.21)	(0.67)	(1.68)*	(1.43)	(0.76)
ln(Women's LFPR)	−0.0267	−0.0201	−0.0234	−0.0194	−0.0169	−0.0182
	(3.47)***	(2.99)***	(3.10)***	(3.19)***	(3.13)***	(3.19)***
E/SE Dummy	0.0174			0.0185		
	(3.22)***			(4.21)***		
Spatially lagged		0.5834			0.5962	
ln(Sex ratio)		(5.28)***			(5.72)***	
Spatially lagged			0.6125			0.6161
error			(5.37)***			(5.46)***
Intercept	0.1982	0.1522	0.1131	−0.0783	−0.0266	−0.0789
	(2.00)**	(1.81)*	(1.28)	(1.12)	(0.42)	(1.16)
R^2	0.69	0.61	0.43	0.73	0.63	0.45
Adj-R^2	0.66			0.71		
Breusch-Pagan	10.72*	8.57	9.25*	22.57***	16.12***	12.95**
LM (lag)	14.74***		3.10*	13.52***		3.98**
LM (error)	3.26*	2.42		10.86***	0.12	
Log-likelihood	−206.64	−210.87	−207.56	−235.42	−237.38	−234.52
AIC	−399.85	−407.75	−403.13	−456.84	−460.75	−457.05

[a] *t*-values (regressions 1 and 4) and *z*-values (regressions 2, 3, 5 and 6) are in parentheses. Since standard errors calculated by the ML method are asymptotic, hypothesis tests are based on the normal distribution, and *z*-values (rather than *t*-values) are reported. Statistical insignificance of the Breusch–Pagan test denotes that the null-hypothesis of heteroskedasticity is rejected. R^2s of the ML estimations is the "pseudo-R^2," defined as the ratio of the variance of the predicted values to the variance of the observed values and they are not strictly comparable to those of the OLS estimations.

* Significant at the 10% level.

** Significant at the 5% level.

*** Significant at the 1% level.

significance levels of LM(lag) and LM(error) tests reported in columns 1 and 4 suggest that spatial dependence is pertinent in both data sets and warrants spatial regression modeling. We estimated spatial lag (columns 2 and 5) and error (columns 3 and 6) models alternatively to confirm the presence of spatial effects. The highly significant positive signs of spatial coefficients indicate that the sex ratio of a province that borders provinces with high sex ratios is likely to be high as well. On the basis of regression diagnostics, we conclude that the observed spatial dependence is attributable more to a functional relationship among the sex ratios across provinces than to autocorrelated spatial errors, and we estimate the spatial lag models. [22]

In the case of 1985 data, introduction of the spatial lag variable (column 2) lowers parameter estimates and their statistical significance, most notably those of the household size and the number of hospital beds. The literacy rate ($p = 0.01$) and women's labor force participation rate ($p = 0.002$), however, remain as the critical variables, both having negative effects on the sex ratio, and the stability of their coefficients and standard errors attests to the robustness of these results. In the case of the 1990 data, consistent with the earlier findings, women's LFPR again has a strong negative impact on the sex ratio. The results from the 1990 Census are otherwise not as satisfactory. The heteroskedasticity problem persists, and none of the other variables is significantly different from zero. The most conspicuous among the latter is the literacy rate, although it still has the negative sign. This result may be the outcome of the somewhat high degree of collinearity between the GDP per capita, literacy rate and household size variables. We investigate whether the estimates are confounded by collinearity in the more detailed regressions reported in Tables 3 and 4.

In the regressions reported in Tables 3 and 4, we separate women's LFPR into its unpaid and paid components. This adds more information to the emerging picture and underscores the importance of disaggregating the female labor force. Tables 3 and 4 report the 1985 and 1990 estimates, respectively. Columns 1 to 4 of each table are the OLS estimates of four different specifications, and the spatial lag model estimates are reported in columns 5–8. We will discuss the 1985 results first. According to column 1 of Table 3, the literacy rate and the E/SE dummy coefficients are again highly significant with the expected signs. More importantly, the negative relationship between the sex ratio and women's LFPR is shown to be driven by women's unpaid work. The coefficient of this variable is significantly different from zero ($p = 0.02$). Women's paid work, on the other hand, has no impact on the sex ratio. The coefficients of GDP per capita and household size have the expected signs but are not statistically significant. One possible reason for these outcomes is the confounding effects of collinearity noted above. The correlation coefficients between paid work, GDP per capita, literacy rate, and household size variables are in the 0.6–0.9 range (in absolute values). In order to check the possible degradation of the estimates by collinearity, we estimated regressions in which the last three regressors are deleted alternatively, and report the results in columns 2–4. [23] In all of these regressions, the findings related to the unpaid and paid work variables remain unchanged. Thus, the statistical insignificance of the paid work variable is not attributable to the collinearity among the regressors. [24]

Findings concerning the nature of spatial dependence again indicate significant spatial lag dependence, which columns 5–8 of Table 3 correct for. In the spatial lag model, the magnitudes of the coefficients (in absolute values) and their levels of statistical significance decline. The primary results, however, do not change substantially. The sex ratio is explained by women's unpaid LFPR and the literacy rate. Women's unpaid LFPR is inversely related to the sex ratio and statistically significant, and paid LFPR does not have any effect on the sex ratio. In addition to unpaid labor, literacy rate is also observed to have a significant negative impact on the sex ratio. Similar to the OLS estimations, this result holds across specifications.

The 1990 results reported in Table 4 are not as robust as the 1985 results. Heteroskedasticity remains a problem, making estimates less efficient. In the absence of strong priors on the nature of heteroskedasticity, we chose not to correct this problem. The two main conclusions of the previous analysis, however, hold in Table 4 as well: women's unpaid LFPR has a strong inverse effect on the sex ratio, while paid LFPR has no impact. The magnitude of the impact of the unpaid work on the sex ratio is slightly lower but comparable to that found in the 1985 sample. The instability of the standard errors of literacy rate and household size variables across the specifications also suggests that collinearity may be more of a problem in the 1990 sample.

Table 3. *Regression model for 0–9 age cohort sex ratio—1985 census women's unpaid and paid LFPR separated)*[a]

	No spatial-dependence ($\lambda = \rho = 0$) (OLS)				Spatial lag ($\alpha_6 = \lambda = 0$) (ML)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln(GDP per capita)	−0.0077		−0.0123	−0.0028	−0.0079		−0.0113	−0.0045
	(1.06)		(1.63)	(0.45)	(1.30)		(1.87)*	(0.88)
ln(Literacy rate)	−0.0603	−0.0651		−0.0547	−0.0439	−0.0491		−0.0389
	(2.92)***	(3.23)***		(2.70)**	(2.63)***	(2.95)***		(2.42)**
ln(Household size)	−0.0267	−0.0144	−0.0131		−0.0178	−0.0054	−0.0037	
	(1.25)	(0.80)	(0.59)		(1.02)	(0.36)	(0.22)	
ln(Hospital beds	−0.0065	−0.0066	−0.0097	−0.0063	−0.0029	−0.0030	−0.0044	−0.0028
per 1000)	(1.80)*	(1.81)*	(2.62)**	(1.73)*	(0.96)	(0.97)	(1.50)	(0.92)
ln(Women's	−0.0110	−0.0092	−0.0126	−0.0097	−0.0082	−0.0066	−0.0090	−0.0073
unpaid LFPR)	(2.46)**	(2.23)**	(2.69)***	(2.22)***	(2.22)**	(1.85)*	(2.42)**	(2.01)**
ln(Women's	−0.0001	−0.0002	−0.0002	0.0033	0.0000	−0.0000	0.0002	0.0024
paid LFPR)	(0.03)	(0.03)	(0.03)	(0.68)	(0.01)	(0.00)	(0.06)	(0.58)
E/SE Dummy	0.0178	0.0175	0.0259	0.0162				
	(3.14)***	(3.10)***	(4.95)***	(2.92)**				
Spatially lagged					0.6050	0.5968	0.7244	0.5956
ln(Sex ratio)					(5.60)***	(5.48)***	(8.30)***	(5.57)***
Intercept	0.1460	0.0433	0.1873	0.0617	0.1138	0.0093	0.1336	0.0571
	(1.46)	(1.86)*	(1.79)*	(0.83)	(1.38)	(0.50)	(1.60)	(0.93)
R^2	0.67	0.66	0.62	0.66	0.57	0.58	0.38	0.59
Adj-R^2	0.63	0.63	0.58	0.63				
Breusch-Pagan	10.75	10.07	11.44*	8.75	8.02	7.20	8.40	7.10
LM (error)	3.03*	3.54*	3.45*	4.40***	3.47*	2.10	3.57*	1.08
LM (lag)	13.06***	12.42***	17.44***	12.02**				
Log-likelihood	−204.43	−203.80	−199.90	−203.56	−209.45	−208.61	−206.57	−208.93
AIC	−392.86	−393.61	−385.80	−393.12	−402.89	−403.22	−399.14	−403.85

[a] t-values (regressions 1–4) and z-values (regressions 5–8) are in parentheses. See also notes to Table 2.

* Significant at the 10% level.

** Significant at the 5% level.

*** Significant at the 1% level.

Table 4. *Regression model for 0–9 age cohort sex ratio—1990 census (women's unpaid and paid LFPR separated)*[a]

	No spatial-dependence ($\lambda = \rho = 0$) (OLS)				Spatial lag ($\alpha_6 = \lambda = 0$) (ML)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln(GDP per capita)	0.0065		0.0066	0.0043	0.0017		0.0014	−0.0005
	(1.61)		(1.60)	(1.12)	(0.47)		(0.40)	(0.16)
ln(Literacy rate)	−0.0331	−0.0335		−0.0486	−0.0170	−0.0165		−0.0290
	(1.99)*	(1.99)*		(3.44)***	(1.14)	(1.11)		(2.27)**
ln(Household size)	0.0295	0.0207	0.0484		0.0247	0.0219	0.0328	
	(1.70)*	(1.24)	(3.26)***		(1.61)	(1.54)	(2.53)**	
ln(Hospital beds	−0.0059	−0.0061	−0.0062	−0.0065	−0.0043	−0.0043	−0.0043	−0.0047
per 1000)	(1.77)*	(1.83)*	(1.84)*	(1.94)*	(1.48)	(1.48)	(1.49)	(1.62)
ln(Women's	−0.0089	−0.0109	−0.0085	−0.0097	−0.0076	−0.0081	−0.0073	−0.0083
Unpaid LFPR)	(2.75)***	(3.66)***	(2.57)**	(2.99)***	(2.68)***	(3.07)***	(2.59)***	(2.92)***
ln(Women's	0.0039	0.0053	0.0037	−0.0004	0.0017	0.0020	0.0015	−0.0020
paid LFPR)	(0.60)	(0.84)	(0.57)	(0.07)	(0.31)	(0.37)	(0.26)	(0.39)
E/SE Dummy	0.0184	0.0179	0.0204	0.0207				
	(4.10)***	(3.95)***	(4.57)***	(4.74)***				
Spatially lagged					0.5903	0.5995	0.6331	0.6385
ln(Sex ratio)					(5.61)***	(5.87)***	(6.55)***	(6.59)***
Intercept	−0.093	0.0067	−0.1241	−0.0353	−0.0504	−0.0222	−0.0582	0.0069
	(1.43)	(0.30)	(1.77)*	(0.59)	(0.80)	(1.22)	(0.94)	(0.13)
R^2	0.73	0.72	0.71	0.71	0.61	0.59	0.56	0.53
Adj-R^2	0.70	0.69	0.68	0.69				
Breusch-Pagan	15.04**	7.91	11.76*	10.57	12.50*	9.84*	10.79*	15.26***
LM (error)	7.72***	7.37***	9.21***	7.66***	0.18	0.29	0.50	0.98
LM (lag)	12.20***	13.80***	15.61***	14.50***				
Log-likelihood	−234.70	−233.27	−232.53	−233.10	−236.38	−236.26	−235.76	−235.08
AIC	−453.39	−452.53	−451.07	−452.21	−456.75	−458.53	−457.52	−456.16

[a] t-values (regressions 1–4) and z-values (regressions 5–8) are in parentheses. See also notes to Table 2.

* Significant at the 10% level.

** Significant at the 5% level.

*** Significant at the 1% level.

Especially noteworthy is that the literacy rate becomes statistically significant once the household size is dropped from the equation. It is not clear, however, why the coefficient of the household size variable is positive in the 1990 regressions.

The negative relationship between women's LFPR and the child sex ratio in Turkey indicates that the availability of work opportunities for women improves the relative survival chances of girls. This result is consistent with the district and county-level findings of Rosenzweig and Schultz (1982), Humphries (1991) and Murthi *et al.* (1995). Type of work matters as well, although not in the way predicted by A. K. Sen. We find that only women's participation in unpaid employment is associated with a lower sex ratio, whereas paid work has no impact. This result also warrants caution against reflexive association of agriculture/rural areas with greater discrimination against women. The explanation for this result may lie in the indispensability of women's labor for the survival of the family enterprise in smallholder agriculture. Case studies done in the early 1980s indicate that, in Turkey, the number of female workers in the household is a key determinant of household income level, ensuring either farm viability or wealth accumulation among smallholder farmers (Morvaridi, 1992; Sirman, 1988). Small farms rely on the unpaid labor of female family members, and in some cases, the network of female kin and neighbors, whom they recruit for work parties in peak periods through reciprocal, nonmonetized exchanges. By relying on family labor, small farms minimize labor costs, recruit labor without competing for wage workers, and are able to coexist with large landowners who use wage workers. If paid labor were used, such farms would not be economically viable. By contrast, women's paid work, which is still limited in Turkey, may still be perceived as a supplementary source of income. An interesting parallel may also be drawn with this result and Humphries' (1991) finding for 19th century England. In England, women's greater mortality was directly related to the rise of capitalist agriculture and the ensuing loss of self-employed farm work for women. More than a century later, in Turkey, we observe a similar coincidence of a relatively less hostile environment for girls with the prevalence of agricultural, unpaid family work.

The second variable that shows up as an important determinant of the sex ratio, especially in the 1985 sample, is the literacy rate. Regression results indicate that girls fare better where the literacy rate is higher. It is interesting that another index of development, GDP per capita, does not have any impact on the sex ratio. Our data do not permit utilization of sex-specific literacy rates as explanatory variables and therefore our results are not comparable with those of Rosenzweig and Schultz (1982), Humphries (1991) and Murthi *et al.* (1995).

Another implication of these results is that the province sex ratio is the outcome of two factors that work in opposite directions. In Turkey, higher levels of literacy are associated with lower rates of women's unpaid labor force participation ($r = -0.5$). Since a province with a high literacy rate tends to have a lower incidence of women's unpaid family work, controlling for other variables, the positive impact of the former variable on the girls' relative life chances appears to be counteracted by the negative effect of the latter.

6. SUMMARY AND CONCLUSION

In this paper we investigated the relationship between the extent and type of women's labor force participation and the child population sex ratio in Turkey. Economic environment and discrimination against women interact via the intermediation of social and cultural institutions, perceptions and customs. The methodology adopted here does not describe or analyze the processes through which discrimination intensifies or weakens. We are able to detect, however, the net impact of a number of variables on the child sex ratio. Our findings indicate that in the 0–9 age group the sex ratio is higher where women's LFPR, women's unpaid LFPR, and the literacy rate are lower. In estimation, we controlled for the geographic pattern of the province sex ratios observed in Turkey. Our analysis diagnosed a strong positive spatial dependence, which suggests the presence of functional interaction of sex ratios between the neighboring provinces. The investigation of this pattern is a critical topic for further research.

The main objective of this study was to operationalize and test A. K. Sen's prediction that women's and girls' well-being is conditional upon paid work outside of the kinship-based, integrated production system. We tested this hypothesis at the province-level with Population Census data from Turkey, where three-quarters of women in the labor force are engaged in unpaid family labor, and found that the prevalence of agricultural, unpaid family labor coincides with a relatively less hostile environment for girls. While our findings on type of work contradict Sen's prediction, they do not invalidate, or indeed test, his argument concerning the importance of either perceptions or women's increased bargaining power in shaping intrahousehold resource allocation. Rather, we infer that inside work is not necessarily ineffective in setting in motion the changes that lower the sex ratio.

This result raises the question of the implications of higher female labor force participation beyond relative survival chances of girls. We suggest that appreciation of the economic value of daughters in smallholder agriculture and their consequent higher rate of survival may accompany life-sapping, but not life-threatening, living and working conditions. Daughters may enjoy more equitable health care and nutrition as they come to be seen as valuable in the household, but if these improvements are accompanied by increasing workload and exploitation, then discrimination against women is transformed but not eliminated. This interpretation sets us apart from some other authors. Notably, Murthi *et al.* (1995) find that rising *overall* female labor force participation improves the relative life chances of girls and attribute this outcome to "women's agency and empowerment." But if inside work is driving this result then there is reason to be skeptical about this conclusion.

Support for our interpretation comes from case studies that show that the use of unpaid female labor to lower costs has resulted in labor intensification. For Igdir (Eastern Turkey), Morvaridi (1992) reports rise in women's age at marriage (attributable in part to the desire of fathers to hold onto their daughters), use of force against women in order to maintain labor productivity, and further intensification of women's labor when small farmers resort to unsound farming practices (such as excessive fertilizer use, overirrigation). If women and girls lack control over their working conditions and the fruit of their labor, and face more arduous working days, then "ill-being" rather than "well-being" may accompany the improvement in their life chances. This argument may have relevance beyond rural Turkey for cases of family-based agricultural and industrial production for the market, which is arguably becoming more widespread on a global scale. [25]

Obviously this study's findings are tentative until they are corroborated by micro-level studies in Turkey. Evidence from other countries is also needed to determine whether these findings are specific to the Turkish experience. In the meantime, the reported findings serve as a cautionary note. If a "better" sex ratio accompanies a lower quality of life, then our empirical results do not unequivocally support policies to preserve and extend family-based-agriculture (e.g., via credit and pricing policies) and thereby women's work opportunities in its midst. Improving the life chances of girls is necessary but not sufficient for enhancing their quality of life, and policy measures that turn women into beasts of burden are ill-conceived even if they result in a more favorable sex ratio.

NOTES

1. The first Population Census of the Republic of Turkey was undertaken in 1927. According to the historical census figures, Turkey experienced a shortage of males until 1945, primarily due to the effects of WW I and the War of Independence. The overall population sex ratio increased steadily from 92.65 in 1927 to 101.10 in 1945 and 104.22 by 1960. It fluctuated between 100.32 and 105.82 during the following two decades and stabilized around 103 after 1980.

2. The expected ratio is based on the following assumptions: in the absence of sex-selective abortion, the sex ratio at birth is 106, but it declines with age because infant boys are more susceptible to diseases; women's life expectancy at birth is 60; and the gross reproduction rate is 2.00. These figures approximate the mortality and fertility conditions preceding the 1985 Census.

3. See Haddad, Hoddinott and Alderman (1997) for a recent survey of this literature.

4. According to another line of research, kinship patterns and cultural differences underlie the variation in the economic undervaluation of women (Dyson & Moore, 1983). This has led some to conceptualize the factors that affect women's well-being in terms of a cultural-economic dichotomy. For example, Kishor (1993) distinguishes between women's cultural worth (measured by the incidence of patrilocal exogamy) and

women's economic worth (measured by their labor force participation). We do not find such dichotomous interpretations persuasive because the economic and the cultural cannot be delineated unambiguously. See Sen (1989) and Murthi *et al.* (1995, p. 753) for critiques. While we refer to the "economic value" of women we do not counterpose it to their "cultural value."

5. Sen (1989, p. 26; 1990, p. 146) states this argument explicitly in his discussion of China's reforms. He suggests that the decline in women's life expectancy in China in the 1980s is partly attributable to the conversion from collective farming to family-based farming for the market. In collective farms, where production was team-based, women's labor was remunerated by work points and their individual contributions were visible. Under the new household responsibility system women's contribution is no longer identifiable and, therefore, lacks recognition (see also Aslanbeigui & Summerfield, 1989).

6. Thomas (1990) shows that mothers' access to income improves the health of daughters to a greater extent than sons' health, while fathers' income improves only sons' health. Klasen's historical analysis shows that the number of children in the family increases the mortality rate of mothers by a greater amount than the fathers' mortality, but the number of boys in the family increases fathers' mortality while the number of girls in the family increases mothers' mortality (Klasen, 1997). This suggests fathers are more willing to reduce their share of household resources in favor of sons while mothers do the same for daughters.

7. The exception is Rosenzweig and Schultz (1982), who deploy model life tables to derive the mortality figures from the age-specific population data in the Indian Census.

8. Some scholars also question the reliability of mortality data in these surveys, with suspicion centering on incomplete recording of female infant deaths especially in rural areas (Akadli & Cerit, 1988, pp. 14–15; Aksit & Aksit, 1989, p. 574). For international evidence on underreporting of female deaths see D'Souza (1978). D'Souza suggests that in an environment of strong son preference the lower reported mortality of females relative to males may be illusory because of underreporting of female births in the first place.

9. In a 1975 national sample survey, in response to a forced-choice question, son preference was found to be strong among both men and women, but stronger among men (93% of men vs. 75% of women) (Kagitcibasi, 1982). Comparable survey data on women's preferences indicate that the proportion of women who desired having a son declined from 42% to 39% during 1978–88 with a corresponding rise in the proportion of women who preferred to have a daughter from 26% to 29% (HIPS, 1989, p. 64).

10. Rosenzweig and Schultz (1982) and Bardhan (1988) also cite this as the reason for focusing on the sex ratio of the 0–9 age group.

11. The number of children in the 5–9 age group in the 1990 census is greater than the number of children in the 0–4 age group in the 1985 census. Demographers who examined the causes of this puzzle observe that this pattern is present in earlier censuses as well. They rule out both international immigration and better enumeration in successive censuses as causes and conclude that it is due to age misstatement (SIS, 1995). We examined whether there is a gender difference in age misstatement (overall and at the province level) by calculating the percentages of excess number of 5–9 year old females and males in the 1990 Census over their counterparts in the 0–4 age group in the 1985 census. The correlation coefficient of 0.93 between the percentage "surpluses" of boys and girls across provinces suggests that there is no gender bias in age misstatement.

12. In this paper the causality runs from women's LFPR to the sex ratio. It is possible, however, that having more daughters than sons in the household permits women to increase their labor force participation because of greater help from daughters in household chores. This two-way causality introduces simultaneity bias and confounds statistical analysis. By focusing on younger children, we expect to minimize this problem. The appropriate testing of simultaneity hypothesis also requires household-level data.

13. In 1985, unpaid family labor in agriculture constituted 80.2% and 92.7% of the total and agricultural female labor force, respectively. Eighty-seven percent of the total female labor force was employed in agriculture (SIS, 1989).

14. Well-known enumeration problems in censuses may result in some underestimation of women's paid labor. It is probable that many women who engage in paid work in the growing urban informal sector are misclassified as out of the labor force. There is no reliable statistical information on either the extent of women's participation in the informal sector or its interprovince variation. Therefore, it is not possible to assess the extent of underreporting of women's paid labor in the census. Thus, our empirical results on paid labor must be interpreted with caution. In rural areas, on the other hand, there is a tendency to classify all

women who work on land as unpaid family workers, regardless of their additional participation in nonagricultural activities as wage workers. Even if it may include some wage-workers, however, our unpaid labor category still provides an accurate proxy for inside work, because such wage-work still takes place within the kinship relations. In the case of rural carpet weaving, for example, production is organized, sales mediated, and earnings collected by male members of the family. Even weavers who work on a piece-wage basis have very limited control over their earnings (Berik, 1987).

15. The results of regressions that use the wage/nonwage work distinction are available from the authors upon request.

16. Urbanization rate (the ratio of the urban dwellers to the province population) may also be a proxy for the availability of health and educational facilities (see Murthi *et al.*, 1995). We omitted this variable on theoretical and empirical grounds. First, urbanization rate may also be a proxy for other, possibly contradictory influences (e.g., breakdown or strengthening of kinship or community ties, impoverishment or rising living standards), and it is difficult to determine which effect the variable captures. Thus, we opted for using more direct measures of availability of services. Second, urbanization rate and women's LFPR are negatively highly correlated (–0.90), and inclusion of both as regressors results in high standard errors. When either one of the variables is excluded, the remaining one captures the urbanization/labor force effect robustly. Since urbanization rate does not seem to have an effect independent of the LFPR, we excluded this variable. While this choice is empirically arbitrary, it can be justified by the ambiguity of the relationship between urbanization and the sex ratio in contrast to the clear theoretical link between women's economic value and the sex ratio. The negative correlation between female labor force participation and the urbanization rate reflects the withdrawal of women from the labor force with rural–urban migration. This may be due to both migrants' attempt to emulate urban middle class gender norms and the limited urban employment opportunities for women.

17. We adopt the SIS (1991) classification of the East/Southeast provinces.

18. The spatially lagged sex ratio of a province is the weighted average of the values of the variable in the neighboring provinces. In order to find the spatially lagged sex ratio for 1985 we used Figure 1 to create a 67×67 matrix, which corresponds to the first order contiguities (e.g., if provinces 1 and 2 are contiguous, then the corresponding cells of the matrix—x_{12} and x_{21} — are equal to one, otherwise they are zero), and row-standardized it by dividing the elements of each row by the corresponding row sum. The product of this matrix ($\boldsymbol{W}$) and the sex ratio vector yields values of the spatially lagged sex ratio.

19. Kishor (1993) and Murthi *et al.* (1995) also detect strong spatial dependence and assume that it is the spatial-error type.

20. Between the two censuses certain province borders were redrawn and six new provinces created. This affects the spatial dimension of the analysis, and the incompatibility of the two data sets precludes a panel regression based on pooling the 1985 and 1990 data.

21. *SpaceStat* software package was used in the estimations (Anselin, 1995). The spatial lag model was also estimated by the instrumental variables (IV) technique. In comparison with the ML estimates, IV estimations yielded slightly lower z-values.

22. Our choice of spatial lag over error model is based on three grounds. First, Anselin (1988) suggests comparing the LM tests for spatial error and spatial lag obtained from the OLS estimation, and choosing the spatial model that has the higher level of statistical significance. As reported in columns 1 and 4, LM(lag) has a higher level of significance than LM(error) in the OLS estimates. Second, in regressions 2 and 5, the LM(error) tests show that the lag model eliminates any residual spatial error ($p = 0.32$ and $p = 0.83$, respectively), while the error model of columns 3 and 6 LM(lag) tests indicate that spatial lag dependence still persists ($p = 0.08$ and $p = 0.04$, respectively). Third, log-likelihood and Akaike Information Criterion (AIC) values of the lag specification are somewhat higher than those of the OLS indicating some improvement in the fit of the regression, whereas the fit of the spatial error model does not indicate as much improvement over the OLS estimate. These findings hold for alternative specifications of the regression equation as well. Henceforth, estimates of the spatial error specification are not reported. They are available from the authors upon request.

23. We also estimated regressions where all three variables were deleted. In these regressions results did not change but the explanatory power declined significantly. They are not reported here but are available from the authors.

24. Estimating new regressions by separating women's paid labor force participation into its components (employees, employers, self-employed labor) did not

alter these results either: in these regressions, each of the components of paid labor turned out to be statistically insignificant. These results are not reported here but are available from the authors upon request.

25. See, for example, Hsiung's (1996) study of the crucial role of unpaid family labor in the success of export-oriented, small-scale industrial workshops on Taiwan.

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APPENDIX A

Descriptive statistics[a,b]

Variable	1985		1990	
	Mean	(Standard deviation)	Mean	(Standard deviation)
Male-female sex ratio (0–9 age group)	105.15	(2.13)	105.27	(1.99)
Per-capita GDP (in Turkish Liras)	43,599	(24,665)	1,158,575	(579,494)
Literacy rate (%)	73.99	(10.75)	76.62	(10.79)
Household size	5.75	(1.18)	5.54	(1.18)
Number of hospital beds per 1000	1.55	(0.94)	1.82	(0.93)
Women's labor force participation rate (%)	51.96	(11.48)	50.79	(10.64)
Women's unpaid labor force participation rate (%)	46.55	(13.70)	40.95	(12.20)
Women's paid labor force participation rate (%)	5.40	(3.14)	9.84	(3.49)

[a] Means are unweighted. The large difference between the per capita GDP of 1985 and 1990 is due to the fact that they are calculated with different base years.

[b] *Sources*: Per capita GDP data are from Ozotun (1988); other data are from SIS (1989, 1993).