J Popul Econ (1998) 11:435–452

Journal of Population Economics

The demand for children in Arab countries: Evidence from panel and count data models

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Received: 30 May 1995 / Accepted: 19 February 1998

Abstract. This paper provides empirical evidence on fertility determinants in Arab countries. Adopting a macro and micro framework and exploiting panel and count data models the paper estimates the impact of cultural and economic factors on the demand for children. The results obtained strongly support the hypothesis that cross-country heterogeneity buttresses differentiated fertility and that female education mitigates high fertility. Child mortality and parent's preferences for sons positively affect fertility. By and large, demand for children is price and income inelastic.

JEL classification: J13, C25, C33

Key words: Fertility, panel data, negative binomial, pro-natal policies

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Introduction

The single most remarkable demographic aspect of the Arab region is the high fertility level-the average level of childbearing is 6 children per woman. Analysis and policy-makers commonly list young age at marriage, lagging female education, son preference, and religious beliefs as causes of the persistence of high fertility in the Arab countries. With the exception of a few studies (Aly and Shields 1991; Caldwell 1977; Obermeyer 1992; Weeks 1988), one notes a relative scarcity of empirical work that seeks to

I would like to thank two anonymous referees and A. Cigno for helpful comments on an earlier draft of this paper. Remaining errors are those of the author. *Responsible editor:* Alessandro Cigno.

test fertility-related hypotheses. Most identification rests on surprisingly thin evidence and the origins and ramifications of differences in fertility at the micro and macro levels are far from fully understood. Without denying the role of culture, this paper poses the question whether, over and above it, there are economic forces at work that influence fertility decisions. Second, the study inquires if observed economic and non-economic heterogeneity across and within countries contribute to fertility variations. Third, the paper investigates the extent to which improvements in female education modify the high fertility pattern. Relying on time series and cross section data, the paper investigates the impact of economic forces and cultural norms on fertility variations across and within Arab countries. At the aggregate level the paper reviews cross-country fertility trends and applies a random-effect model to panel data covering eleven Arab countries and derives estimates of cross-country fertility determinants. At the micro-level, the paper applies the Poisson and Negative binomial models to micro data of four Arab countries and ferrets out the causal impact of demographic, cultural and economic factors on the demand for children. In the process, the paper tests a set of common hypotheses that the literature has posed. Among the common hypotheses is that an exogenous reduction in child mortality rates, assuming the demand for children is price inelastic, is associated with a decline in the demand for births (Olsen 1980; Schultz 1997). On the other hand, it is widely believed that increased education of women raises the cost of childbearing and reduces fertility (Rosenzweig and Schultz 1989). The paper is organized in four sections. The second presents the method, variable description and empirical findings of the panel data analysis. The third section introduces our approach to micro data analysis. It describes country family-level data, and discusses the results of applying the Poisson and negative binomial models to the country data. The last section offers a summary of main findings and discusses their policy implications. The remainder of this section reviews salient trends in Arab fertility.

1. Fertility trends

Whereas fertility levels are high in the region, disparities exist across countries. In the mid-1994 while Kuwait, Oman, Saudi Arabia, the West Bank, Gaza and Yemen maintained a TFR of between 6 and 7.5 births per female national, TFR was less than 4 in Egypt, and less than 3 in Lebanon (Escwa 1996). Differences are also noted within the same country. In Algeria for instance, women still gave birth to more than six children in the southern part of the country but less than four in the north. While urban Yemen had a TFR of 5.6 in 1992, its rural areas had a higher rate, 8.2. Lebanon despite its small size harbors strong regional contrasts: from Beirut (2.3 children per woman) to the north (4.3 children per woman). In Egypt the average family numbers only 3.6 children in Port Said but 8.2 in Fayoum (Escwa 1996; Fargues 1994). In contrast to its lagging fertility, the region has experienced a very rapid decline in mortality rates at every age. In the mid-1994, the crude death rate for the Arab countries was under 7 per 1000, which is less than the World's 9.3 crude death rate. Current levels resulted from remarkable improvements in health standards. For instance, Egypt's infant mortality rate was reduced from 170 per 1000 births in 1960 to 110 in 1981 and to 62 in 1992. In the Gulf region the improvement has been substantial: Saudi Arabia's 1992 rate of 31 is much lower than the 190 rate that existed in 1960. Other countries recorded improvements also but the current levels of IMR remain high particularly in Sudan (100), Yemen (107).

Arab countries can be grouped into three broad categories based on fertility trends. First is the group of countries with persistent high fertility rates and declining mortality. This group includes Jordan, Oman, Syria, Yemen, the West Bank and Gaza where the per capita income level is low to moderate. The birth rate among these countries was 44 births per 1000 population in 1990, well above the birth rate of 30 for all developing countries in that year. Infant mortality rate (IMR) dropped sharply from 196 per 1000 live birth in 1950 to 65 deaths per 1000 in 1990. Life expectancy increased from 30 to 60 years for males and from 40 to 63 for females. In the second group fertility is declining at rates that are faster than the rate of decline in mortality rates leading to a deceleration in the natural growth rate. The group includes Morocco, Egypt and Lebanon whose socio-economic development is at an intermediate level. The third group is the Gulf countries that are characterized by high fertility and rapidly declining mortality rates. These include Bahrain, Iraq, Kuwait, Qatar, Saudi Arabia, and the United Arab Emirates. In the early 1950's the birth rate for this group of countries was 49 births per 1000, the death rate was 23 deaths per 1000. During the next few decades, the harvesting of oil caused the influx of waves of immigrants and contributed to the rapid improvement in health standards and socioeconomic development in general. Average birthrate for these countries (which includes substantial numbers of immigrants) dropped to 36 by 1990, while the death rate plummeted to 6 per 1000 (Omran and Roudi 1993).

Heterogeneity of Arab countries is also gleaned from the role of government policy. In some countries (e.g., Tunisia, Egypt) governments are more active than others in the area of family planning programs; ensuring health care assistance (e.g., Kuwait, Oman) especially for the poor and the disadvantaged women and children. However, Arab governments' commitment to reducing fertility is generally low. United Nations records indicate that, between 1978 and 1987, most Arab governments considered that their fertility levels were satisfactory despite that fact that they are the highest in the world (Caldwell and Larson 1989). This general pattern notwithstanding, a few have succeeded in pursuing ambitious family planning programs. For instance, the government of Tunisia has made modern contraceptives and sterilization and abortion procedures available to a large segment of the Tunisian population. In response, the total fertility rate has dropped from 7 births per woman in the 1960s to 4.4 and 3.8 births per woman by 1988 and 1992 respectively.

2. Panel data analysis

The paper investigates fertility at both the micro and macro level of analysis. At the aggregate level, it applies the fixed effect and random error models to a set of balanced panel data covering eleven Arab countries. The countries are Algeria, Bahrain, Egypt, Kuwait, Morocco, Oman, Saudi Arabia, Sudan, Syria, Tunisia, and the United Arab Emirates. The data are for the years 1971, 1976, 1981, 1986, 1991 and 1996. The data sources are the World Tables (1995), Escwa (1993, 1996) and official country sources. The use of panel data allows us to control for unobservable country-specific characteristics that may be correlated with fertility behavior. Some variables may be country invariant or time invariant that may affect fertility. Other variables may be hard to obtain and their omission leads to bias in the resulting estimates. Panel data are able to control for these country and time-invariant variables whereas a time series or cross section study can not. In general, failure to condition on these unobservables will result in inconsistent estimates of the coefficients of these variables (Baltagi 1995). Examples of factors that are unobservable in our data but could induce country heterogeneity include country differences in the distribution of income among households and according to geographic location. Similarly, the extent of availability and accessibility of birth control methods may vary over time and from one country to another. As noted in the introduction, government policy can also influence overall fertility. Some governments follow a pro-natal population policy (e.g. Saudi-Arabia, Oman) while others discourage high fertility (e.g. Tunisia). Country heterogeneity also covers differences in government food subsidy policies. A few countries (e.g., Egypt, UAE, Saudi Arabia, and Yemen) have food subsidy programs that vary in coverage, targets and success. Food and other commodity subsidies reduce the cost of having and rearing children. Our model deals with these and other sources of country heterogeneity-correlation between the explanatory variables and omitted country attributes – in an appropriate manner. We apply panel data estimators to the following fertility equation:

$$F_{it} = a_0 + X_{it}\beta + a_i + \varepsilon_{it} \quad i = 1, ..., N;\ t = 1...T\,, \tag{2.1}$$

where F_{it} is the total fertility rate, X_{it} contains four observable explanatory variables of country i in period t. The variables are real per capita GDP, rate of urbanization, infant mortality rate and female education. Enrollment of females at the second level of education, lagged 5 years, is introduced as a measure of female education. The lag captures the duration required for school accreditation at the second level as well as the time lapse between schooling and marriage. In order to control for the potential cross impact between education and infant mortality, the analysis includes a variable gauging the cross impact of school enrollment and infant mortality rate.[1] The a_i are fixed effects that may vary by country and reflect unobservable country-specific characteristics that may be correlated with X_{it}. The ε_{it} are typical disturbances terms, assumed to be iid with a zero mean and a constant variance σ_ε^2. Two approaches to estimate equation (1) are taken. The first is to ignore the individual country effect as a distinct component of the error term. Alternatively, we employ fixed effects approach that amounts to estimating equation (2) in deviations from means. That is:

$$(F_{it} - \bar{F}_i) = (X_{it} - \bar{X}_i)\beta + (\varepsilon_{it} - \bar{\varepsilon}_i) \tag{2.2}$$

Table 1. Results from estimation of robust, random and fixed effect models (standard erros in parentheses)

Variable	White	Panel 1[c]	Panel 2[c]	Elasticity[d]	Variable mean[a]
Per capita GDP	0.000063* (0.0000255)	0.0000254 (0.000038)	0.00007** (0.000039)	0.0825	8.08528
IMR	–0.00357 (0.00738)	0.017981* (0.007197)	0.019177* (0.00707)	0.1684	3.96337
Second enroll. lag 5	–0.02199** (0.012412)	–0.01924** (0.010779)	–0.01975* (0.00981)	–0.1016	
Enroll. Lag·IMR	–0.0000538* (0.0002518)	–0.006664* (0.000256)	–0.00043** (0.000263)		3.40041
Urban rate	–0.023087* (0.007756)	–0.0198*** (0.013551)	–0.05361* (0.01966)	–0.4398	3.81568
Intercept	8.470941* (0.743023)	7.44144* (1.02685)	9.44874* (1.4186)		1
χ^2 (or F statistic)[b]	13.95	80.55	22.08		
R^2	0.5337	0.6838	0.7390		
N	55	55	55	55	

Panels 1 and 2 are estimated using the random and fixed effect models respectively.
[a] Variables are in logarithmic scale.
[b] Values of χ^2 are for the random effect model. F values are for the fixed effect model and for White's robust regression.
[c] Average values of individual estimates based on the fixed effect model.
[d] The value of the Hausman specification test is 23.88 for the random effect model.
* Indicates that the estimate is significant at the 5%.
** Indicates that the estimate is significant at the 10%.
*** Indicates that the estimate is significant at the 15%.

The random-effects estimator is a weighted average of the estimates produced by the between and within estimators. In particular, the random-effects estimator is:

$$(F_{it} - \theta \bar{F}_i) = (1 - \theta)a_0 + (X_{it} - \theta \bar{X}_i)\beta + [(1 - \theta)v_i + (\varepsilon_{it} - \theta \bar{\varepsilon}_I)] \qquad (2.3)$$

where θ is a function of σ_v^2 and σ_ε^2.

In addition to the random and fixed effects models, we report estimates that obtain when White's (1980) heteroscedastic variance covariance matrix is applied. Finally, we apply the Hausman specification test in order to check the appropriateness of the random effect model for our data.

Table 1 summarizes our findings. First, note that, for all specifications, the signs of the variables are according to theoretical expectations. Second, the results indicate that unobserved country heterogeneity is statistically important in our sample. In the model where country effects are controlled for, we obtain estimated effects for per capita GDP, infant mortality and education parameters, that are more consistent than those obtained when country heterogeneity is ignored. This is corroborated by the Hausman specification test that leads us to conclude that heterogeneity is statistically important in our sample and accept estimators that condition on country effects. The Hausman specification tests if the difference between the coefficients of the fixed and random effects is not systematic. The value of the test statistic, which is asymptotically distributed as χ^2 (4), is 23.88 for the random-effects model, much larger than the theoretical value. This leads us to reject the hypothesis of equivalence between fixed and random effect coefficients in favor of heterogeneity in our data.

At this level of aggregation, our findings suggest that Arab countries that have higher per capita income maintain higher fertility rate, implying a positive albeit small income elasticity of demand for children. All else constant, countries with higher urban structure have lower fertility and education of females reduces fertility. The results also indicate that infant mortality has a positive influence on fertility which corroborates earlier work that documented that declines in infant mortality in the context of an Arab country, Sudan, reduced the need for high fertility (Maglad 1994). The negative sign of the education variable implies that increasing female secondary school enrollment exerts a negative impact on fertility. The strong influence of female education overtakes the positive effect of infant mortality. That is, the cross product of education and infant mortality reduces fertility.

Column 5 of the table displays estimates of the elasticity with respect to the explanatory variables. According to the fixed-effects model, elasticity of fertility with respect to per capita income is in the vicinity of 0.08. A 10% increase in the rate of female secondary school enrollment is accompanied by a 1.6% reduction in fertility.

Reducing infant mortality by 10% could induce 1.68% reduction in fertility while a corresponding increase in urbanization rate is associated with a 4.4% reduction in women's total fertility. Finally, we note that country heterogeneity is also reflected in the obtained country-specific elasticities. In the oil producing countries of Saudi Arabia, Bahrain and Kuwait the elasticity of fertility with respect to infant mortality is in the range of 0.05 to 0.11, smaller than the corresponding elasticity for Tunisia, Egypt, Syria, Algeria and Morocco where the range falls between 0.12 and 0.31. The fertility elasticity with respect to female's education varies from a low and negative value of 0.03 in Oman and Sudan to 0.14 and 0.18 in Egypt, UAE and Tunisia (data are not shown). Further, the size of the elasticity estimates tends to decline in absolute value over time. This suggests that significant fertility reductions will obtain only when marked changes occur in the time-dependent behavior of child mortality, women's education and urbanization.

3. Micro-level analysis

In order to corroborate findings from aggregate panel data, we apply econometric techniques to micro unit (family) level data sets of four Arab countries. Many variables bearing on fertility are more accurately measured at the micro level and biases resulting from aggregation over households and countries are eliminated. We utilize a simple demand function for children whose general form is as follows (Cigno 1991):

$$D = f(A, F, C, W) , \tag{3.1}$$

where *D* stands for the demand for children, measured in our data by the number of births ever born to a mother (CEB), *A* is the demographic and personal characteristics of the individual woman including age, education, and age at first marriage. *F* stands for family characteristics including marriage duration, child mortality rate and preferences for male offspring. *C* stands for community location characteristics (urban versus rural) and *W* stands for economic factors including wife's wage, proxies for standards of living and measured family income. Because childbearing is a highly time-intensive activity, economic theory predicts that a rise in women's wage would lead to a decrease in the demand for children. Further, children are expected to be a normal good with positive income elasticity. Since it is possible that the impact of wages (price of women's time) may stem directly from female education in the fertility equation, we test the hypothesis that wages are endogenous. We assume that the number of children ever born (CEB) is a good indicator of desired fertility given the socioeconomic, health and cultural environments which parents encounter. Married women's fertility is assumed to be endogenous; that is, the number of children born to a family is a choice variable. Following the neoclassical theory, the family is assumed to maximize a utility function that includes the number of children and all other goods. The number of children born to married women depends on age cohorts 15–19, 20–24, 25–29,... 45–54, female's age at first marriage, marriage duration, education, and labor market earnings of working females, household income and religion. The variables age and age at first marriage control for the biological productive supply of the women. The literature indicates that the expected impact of education on fertility is negative. It is effectively a correlate with the desire for increased child quality as argued by Becker (1981), and improved knowledge about availability and effectiveness of contraceptive methods (Rosenzweig and Schultz 1989). Further, education reduces fertility presumably through its impact on (i) increasing women's age at marriage; (ii) increasing the value of her time; and (iii) increasing her desire for self-esteem and (family) improvement. The analysis includes the sex of children (the ratio of the surviving female to male offspring) in order to capture the potential impact of son preference. Parents may have preference for boys because the net economic productivity of boys may exceed that of girls. Remittances to parents may be smaller from daughters than sons and, accordingly, boys may be better old age insurance value for parents than girls or because they maintain family line (Schultz 1997).

To test the hypothesis that child mortality leads to higher fertility, the regression analysis below incorporates child mortality which enters as a dichotomous variable taking a value of 1 if the family had experienced a child loss and zero otherwise. This formulation reduces the extent of positive association between the dependent variable (number of births) and child mortality.[2] We note that our concern here is with the probable impact of mortality on fertility. This abstracts away from the reverse causality, i.e., from fertility to infant mortality. A number of studies have shown that fertility regulation can reduce child mortality. High fertility can affect the prospects of child survival through termination of breast-feeding, sibling competition for limited family resources, and through the negative impact on mother's physical and psychological health. Other contributing factors include the impact of the risk of poor child health that may be attributable to short conception intervals and birth order (Miller et al. 1992; Koeing et al. 1990).

Our strategy is to employ the Poisson and the Negative binomial distribution models that are the simplest and perhaps the most common count models in order to estimate equation (1) above. For a sample of N individuals, we observe y_i, $i=1...N$, number of births that a married woman gives during her productive cycle. We suppose that y_i has a Poisson distribution with parameter ξ>0. Then

$$\Pr(\mathrm{y}_{I=\mathrm{r}}) = \exp(-\xi_i)\xi_i/\mathrm{r}! = 0, 1, 2, ..., \tag{3.2}$$

ξ_i can be expressed as a function of observable individual characteristics and family attributes which are included in the k dimensional vector $\boldsymbol{x}$. It is customary to let $\xi_i = \exp(\boldsymbol{x}'\beta)$. Substituting this into (1) yields the log likelihood function for β that is globally concave:

$$\mathrm{L}_i(\xi_i) = -\xi_i + \mathrm{y}_i \log \xi_i - \log \mathrm{y}_i\,! \,. \tag{3.3}$$

The number of births per period or the conditional mean is given by:

$$\mathrm{E}[\mathrm{y}_i|\boldsymbol{x_i}] = \xi_i = \xi(\boldsymbol{x_i}, \beta)\ , \tag{3.4}$$

where ξ(.) is a specified function, $\boldsymbol{x_i}$ is a vector of exogenous regressors that includes a constant term, and β is a $\boldsymbol{k}\times\mathbf{1}$ parameter vector. For this model, the conditional variance is restricted to equal the conditional mean:

$$\mathrm{V}ar(\mathrm{y}_I|\mathrm{x}_i) = \xi_{i..}\ . \tag{3.5}$$

The likelihood equations are:

$$\partial \log \xi_I / \partial \log \beta = \sum_{i=1}^{\mathrm{n}} (\mathrm{y}_i - \xi_i)\mathrm{x}_i = 0\ . \tag{3.6}$$

The Hessian is:

$$\partial^2 \log \xi_i / \partial\beta\partial\beta' = -\sum_{i=1} \xi_i \mathrm{x}_I \boldsymbol{x_{i'}}\ . \tag{3.7}$$

Cameron and Johansson (1997) and Winkelmann and Zimmermann (1995) have observed that for a number of reasons the Poisson model might be inadequate. The equality between the conditional mean and conditional variance is often rejected in econometric applications. Most empirical data omit information on unobservable characteristics for the set of explanatory variables, x. Under these circumstances, it is more appropriate to think of ξ_i as stochastic and to characterize the interpersonal heterogeneity in a mathematically convenient way. When such heterogeneity is a feature of the data we expect the count to exhibit overdispersion, i.e. greater variance than is consistent with the Poisson model. Estimation under the Poisson assumption and neglect of overdispersion will generally lead to inefficient estimates. Cameron and Travedi (1986) suggest that the negative binomial distribution becomes appropriate if statistical tests indicate the existence of overdispersion. The negative binomial distribution can be regarded as an extension to the Poisson model with a heterogeneity term gamma. More specifically, let ξ_i be gamma distributed with parameters $\gamma_i > 0$ and $\delta > 0$, so

$$\mathrm{f}\,(\xi_i = \mathrm{z}) = 1/\Gamma(\gamma_i) \cdot \delta_i^{\gamma} \exp(-\delta \mathrm{z}) \mathrm{z}^{\gamma i-1} \,, \tag{3.8}$$

with $E[\xi_i] = \gamma_i/\delta$ and $\mathrm{var}(\xi_i) = \gamma_i/\delta^2$. Since γ_i is Poisson (ξ_i), we have

$$\Pr(\mathrm{y}i = \mathrm{r}) = \int \Pr(\mathrm{y}i|\xi_i = \mathrm{z})\mathrm{f}\,(\xi_i = \mathrm{z})\mathrm{dz} \tag{3.9}$$

$$= \Gamma(\gamma_I + \mathrm{r})/\Gamma(\mathrm{r}+1)\Gamma(\gamma_I)(\delta/1+\delta)_I^{\gamma}(1/1+\delta)' \,,\ \mathrm{r} = 0, 1, 2, \ldots \tag{3.10}$$

which is negative binomial distribution with $E[y_i] = \gamma_i/\delta$ and $\mathrm{var}(yi) = \gamma_i(1+\delta)/\delta^2$. The model accounts for overdispersion in the data since the variance is greater than the mean. Since we employ data sets that are diverse in terms of coverage, size and main purpose, we apply both the Poisson and Negative binomial models to each data set. We then use the appropriate log likelihood ratio in order to test the null hypothesis that the Poisson model is appropriate for each data set.

3.1 Micro-level data

The paper utilizes a variety of official micro-level data sets that encompass two samples of population censuses, a demographic survey, and a multipurpose household survey. For two countries, Yemen and Oman, the paper accesses large samples drawn from their population censuses. The data files provide diverse information on individual and family characteristics that affect fertility. They also differ in size, coverage and year of enumeration. Yemen's sample is randomly drawn from its 1994-population census and represents 2% of all enumerated individuals including more than 30 000 eligible mothers. Oman's sample ratio is larger representing about 20% of the country's population during the 1993 census year. The sample of Yemen contains information on urban location, sex of children ever born (CEB) to married women, and proxies for standards of living. These are in terms of housing characteristics and its main source of cooking fuel and lighting source (kerosene vs. electricity). Oman's data covers more than 60 000 eligible mothers and contains information on their demographic and education characteristics, urban location and family ownership of car and home appliances. Moreover, the sample contains information on female age at marriage and duration of marital life.

The West Bank and Gaza demographic survey, conducted by the Palestinian Central Bureau of Statistics (PCBS) in 1995, collected information on fertility of women aged 15 to 54. Information is gleaned on more than 100 000 individuals that are members of 15 000 households. Collected information covers age and gender of surviving children, mother's age, religion and education and surrogates for family resources. For Jordan, our study utilizes the 1987 Health, Nutrition, Manpower and Poverty survey that covered about 40 000 individuals who are members of 5500 households. Jordan's data set is unique in that it collected information on wages and family income. Information covering age, education, religion, number of births, age at marriage and marriage duration is also readily available from the data. Income from all sources, including wages, is reported for the family and each employed member. The paper utilizes this survey in order to estimate Jordan's fertility determinants and explicity gauge the impact of wages of working women and measured family resources on fertility variations.

3.2 Empirical findings

Tables 2 through 4 report variable means and estimation results. Regressions for the Poisson and negative binomial models are reported separately. The log likelihood ratio test displayed along with estimates of the negative binomial model tests for the possibility that the data are over-dispersed. A low value of the likelihood ratio test statistic leads to accepting the null hypothesis that the data follow a Poisson distribution. Of the four country data sets, Oman's data is the most overdispersed and hence the negative binomial model fits Oman's data better than the Poisson model at the 5% significance level. For the other countries, the test leads us to accept the hypothesis that their data are Poisson. This is also gleaned from the minute value of the estimate of the overdispersion parameter, i.e. close to zero. Further, parameter estimates of the two models are of the same magnitude in the case of Jordan. West Bank, Gaza and Yemen but are different in the case of Oman. All the variables have the expected sign and the overall fit of the model is quite satisfactory for cross-section regressions. The findings indicate that demographic characteristics clearly influence fertility behavior of Arab women. The expected number of births procreated by women who have completed their reproductive cycle is 8.1 in Oman, 8.4 in the West Bank and Gaza, 7.8 in Yemen and 8.2 in Jordan. Second, while clearly high, women's fertility is mitigated by her education as shown by the statistically significant sign of the education estimate in all country regressions. In general, university-educated women who have completed their reproductive cycle have nearly one-half the number of births of illiterate women.

Third, duration of marriage life increases women's potential fertility and women who enter into their first marriage at a young age have a higher expected fertility rate than women who marry later in life.[3] In Jordan (1987) and the West Bank and Gaza (1995), the expected number of births is about 10 and 9.5 respectively if the mother (45–54 years old) had entered into marriage at the age of 15. Her expected number of births declines to 7.4 and 7.8 in the two countries respectively if her age at first marriage was 20 and to about 6 and 6.3 births respectively if she married at the age

Table 2. Variable means of count data regression models

Variable	Jordan 1987	Oman 1993	Yemen 1994	West Bank and Gaza 1995
Age 20–24	0.1211010	0.1724603	0.1445763	0.1721688
Age 25–29	0.1659531	0.158957	0.1852689	0.1892672
Age 30–34	0.1480122	0.144823	0.1680025	0.1728349
Age 35–44	0.300509	0.25315	0.2684853	0.2485566
Age 45–54	0.231600	0.1826665	0.1624135	0.1665433
Age at marriage	19.1979	16.55766	–	18.87694
Dead child = 1	0.2879252	0.3639856	0.2964302	0.2252759
Duration of marriage	16.39653	16.64431	–	15.76489
Education	5.48297	1.306405	0.6776465	6.521271
Ed* religion	5.216514	–	–	7.505961
Ed* son preference	5.242281	0.6633481	0.4489145	5.952154
Family income [a]	137.6166	–	–	–
Moslem = 1	0.971735	–	–	0.9779423
Son preference	1.04289	1.037562	1.038222	1.057251
Urban = 1	–	0.36109	0.1619397	0.4123612
Woman's wage	3.966800	–	–	–
Car = 1	0.2757342	0.6178954	–	0.2884045
Charcoal = 1	–	–	0.5998988	–
Freezer = 1	0.0324152	0.5620643	–	–
Kerosene = 1	–	–	0.5124266	–
Oven = 1	0.9531051	0.8937802	–	0.6563658
Own home = 1	0.248122	–	0.8917714	0.8267378
Phone = 1	0.3070012	0.5643841	–	0.188292
Refrigerator = 1	0.8167285	0.7708318	–	0.847351
TV = 1	0.9095419	0.8605811	–	0.894479
VCR = 1	0.1355751	0.2386434	–	0.1648759
Washer = 1	0.8428385	0.7565545	–	0.7295295

[a] Family monthly income is net of wages of working women.
Source: Computed from the micro data of the country samples.

of 25. The association between age at first marriage and decline in maternal fertility does not necessarily imply that late marriage itself is the cause of an early decline in marital fertility, or that early marriage is the direct cause of the lack of control of fertility within marriage. Rather, the social and cultural conditions that influence the age of entry into marriage also can be themselves barriers to voluntary controls of maternal fertility.[4]

However, in all countries under study younger cohorts do not marry as young as did their older cohorts and their expected fertility is lower and hence a demographic transition is taking place across generations. Several social and cultural forces influence this transition including improved female education and developments in health standards and facilities. These developments mitigate child mortality within marriage and reduce the need for replacement. The positive sign and high significance of the variable capturing son preference indicates that cultural forces exert a powerful influence on fertility: on average, mothers are likely to experience more births the larger the sex (female/male) ratio of their surviving children. Because parent's demand is not births per se, but surviving children, child mortality is associated with a positive influence on fertility as parents at-

Table 3. Poisson model of fertility determinants (standard errors in parentheses)

Variable	Jordan	Oman	Yemen	West Bank and Gaza
Age 20–24	0.4895835*	0.9599325*	0.4013652*	0.5161059*
	(0.0951946)	(0.0161098)	(0.031324)	(0.0351593)
Age 25–29	0.8104637*	1.415004*	0.7908809*	0.9353404*
	(0.0944554)	(0.0166321)	(0.0301714)	(0.0354984)
Age 30–34	1.004201*	1.652053*	1.070635*	1.183882*
	(0.0967993)	(0.018014)	(0.0299969)	(0.0374353)
Age 35–44	1.083397*	1.750422*	1.267232*	1.33225*
	(0.102909)	(0.0206047)	(0.0297534)	(0.0416908)
Age 45–54	0.9849592*	1.582924*	1.264845*	1.194097*
	(0.1139867)	(0.025639)	(0.0299244)	(0.0505794)
Age at marriage	–0.0126329*	–0.017094*	–	–0.0290473*
	(0.0025422)	(0.0008285)	–	(0.0016106)
Marriage duration	0.0214508*	0.0069011*	–	0.0191844*
	(0.0022566)	(0.0006677)	–	(0.0012444)
Education	–0.0143408*	–0.0308182*	–0.0320565*	–0.0116752*
	(0.0014397)	(0.000937)	(0.0019176)	(0.0073359)
Son preference	0.0356945*	0.0617187*	0.0512862*	0.050082**
	(0.0048505)	(0.001273)	(0.0021645)	(0.0041624)
Moslem = 1	0.289392*	–	–	0.3854342*
	(0.079012)			(0.0743107)
Ed* religion	0.007831	–	–	0.001042
	(0.008651)			(0.0073123)
Ed* son preference	0.006645*	0.0212892*	0.0089958*	0.0034682*
	(0.003215)	(0.0000785)	(0.0015905)	(0.0006568)
Dead child = 1	0.183194*	0.2128097*	0.3337513*	0.1995913*
	(0.0118836)	(0.003577)	(0.0048735)	(0.0085401)
Urban = 1	–	–0.0039201*	–0.022778*	–0.007335*
		(0.0005892)	(0.0076912)	(0.0051066)
Own home = 1	0.032471*	–	0.0252498*	0.0715402*
	(0.013658)		(0.0093484)	(0.0106825)
TV = 1	0.0595536*	0.0181032*	–	0.0049467
	(0.0232275)	(0.0062087)		(0.0129905)
VCR = 1	0.0280512**	0.0078608*	–	0.0372743*
	(0.0177661)	(0.0039583)		(0.0109484)
Oven = 1	0.0399306	0.0034831	–	0.0100973*
	(0.0283233)	(0.005719)		(0.0082546)
Refrigerator = 1	0.0150615	0.0140824*	–	0.0366751*
	(0.017583)	(0.0044883)		(0.0115086)
Freezer = 1	0.054315*	0.0242791*	–	–
	(0.0330467)	(0.0035837)		
Washer = 1	0.0493317*	0.0641006*	–	0.0420607*
	(0.0193946)	(0.0049344)		(0.0096652)
Phone = 1	0.01414073	0.0244945*	–	0.0520346*
	(0.014291)	(0.0038359)		(0.0105756)
Car = 1	0.0320032*	0.0368075*	–	0.0101438*
	(0.0137377)	(0.0035265)		(0.0089409)
Charcoal = 1	–	–	–0.0186649*	–
			(0.0058895)	
Kerosene = 1	–	–	–0.0299876*	–
			(0.0056257)	
Women's wage	–0.0009524	–	–	–
	(0.000351)			

Table 3 (continued)

Variable	Jordan	Oman	Yemen	West Bank and Gaza
Family income	0.0000891*	–	–	–
	(0.0000434)			
Constant	0.7645966*	0.2319302*	0.4881385*	0.3704014*
	(0.1050043)	(0.0211339)	(0.0308708)	(0.0874539)
χ^2	5340.676	91163.188	23438.688	19045.086
Pseudo R^2	0.2131	0.2132	0.1480	0.2632
No. of observation	5418	63643	30781	13359
Log likelihood	–9858.273	–168217.906	–67450.016	–26661.168

Note: Pseudo R^2 is a measure of goodness of fit. Fundamentally, it is obtained by subtracting from one the ratio of two log likelihood functions: the numerator is a constant only MLE regression function. The denominator is a MLE regression with all explanatory variables (or the perfect fit). The measure is bounded between 0 and 1 (Greene 1997).
* Indicates the variable is significant at the 5%.

tempt to replace deceased children. Surrogates of standards of living indicate that improvements in the economic wellbeing are generally associated with increased fertility. The sign of home ownership variable is positive and tests significant in all country regressions implying that, ceteris paribus, families that own their homes tend to have higher fertility levels. Indicators of main source of lighting and cooking suggest that Yemeni families that rely on cheap sources of energy (charcoal and kerosene) have lower fertility than families relying on modern energy sources. In general, the signs and significance of asset ownership variables indicates a positive impact on fertility.

Jordan's regression offers an opportunity to test the impact of woman's wage and family income resources. The negative sign of the wage coefficient indicates that fertility and wage earnings compete for women's time and that a rise in women's wage lowers fertility. As indicated in Table 5, the wife's wage elasticity is –0.32; thus the demand for children is inelastic. However, Hausman's test failed to support the hypothesis that women's wages are exogenous. (The Hausman test rejected the hypothesis that women's wages are exogenous. The value of the specification test is 1420, which is much higher than the theoretical χ^2 value.) While not truly exogenous, the inclusion of both wages and education in the fertility equation improves the efficiency of the variable gauging female education (Macunovich 1996). Similar to our aggregate findings from the panel model above, children appear to be a normal good. The sign and significance of family income variable in the case of Jordan are in the expected direction but the demand for children has a minute income elasticity of about 0.029.

Urban households have lower fertility than rural households. In Yemen, for example, the expected number of births for rural women in the age group (45–54) is 7.6 relative to 5.9 for women in urban areas and the corresponding numbers in Oman are 8.4 and 7.6 respectively. The coefficient of the variable controlling for religion has a positive and statistically significant sign, i.e. Moslem parents demand more children than Christian par-

Table 4. Negative binomial model of fertility determinants (standard errors in parentheses)

Variable	Jordan 1987	Oman 1993	Yemen 1994	West Bank and Gaza 1995
Age 20–24	0.489578* (0.0952725)	0.940301* (0.017424)	0.401364* (0.03207)	0.516108* (0.034682)
Age 25–29	0.8104565* (0.0947883)	1.37759* (0.0186331)	0.79088* (0.0309665)	0.9353425* (0.035046)
Age 30–34	1.004194* (0.0971681)	1.599765* (0.0209963)	1.07063* (0.0307935)	1.183884* (0.0369286)
Age 35–44	1.083388* (0.1034665)	1.679821* (0.0250896)	1.267231* (0.0305568)	1.332253* (0.0412821)
Age 45–54	0.9849486* (0.1146289)	1.481492* (0.0326934)	1.264844* (0.0307246)	1.1941* (0.0502915)
Age at marriage	–0.0236328* (0.0025711)	–0.0168161* (0.0011204)	– –	–0.029047* 80.0015989)
Marriage duration	0.0214509* (0.0022646)	0.0100982* (0.0009218)	– –	0.019185* (0.001241)
Education	–0.0143409* (0.0014693)	–0.0314924* (0.0011113)	–0.0426086* (0.001919)	–0.011673* (0.006978)
Son preference	0.0356946* (0.0049345)	0.0758754* (0.0018768)	0.0512863* (0.0021644)	0.050079* 80.004162)
Moslem = 1	0.388965* (0.0747210)	–	–	0.3854502* (0.0752406)
Ed.* religion	–0.000828 (0.008201)	–	–	0.0010399 (0.0069353)
Ed.* son preference	0.004679* (0.001039)	0.0238892* (0.000875)	0.0098803* (0.0016037)	0.0034683* (0.0006593)
Dead child = 1	0.1831946* (0.0119645)	0.2163094* 80.0048276)	0.3337512* (0.0048721)	0.199589* (0.008564)
Urban = 1	–	–0.0096452* (0.004091)	–0.0227779* (0.0076903)	–0.007366** (0.004159)
Own home = 1	0.034271* (0.01391128)	–	0.02525* (0.0093664)	0.071542* (0.010704)
TV = 1	0.595554* (0.0233384)	0.0192753* (0.0082526)	–	0.0049431* (0.012935)
VCR = 1	0.028051** (0.0178262)	0.0127286* (0.0053047)	–	0.037274* (0.01091)
Oven = 1	0.039934 (0.0279454)	0.0102327* (0.0076216)	–	0.010097* (0.008321)
Refrigerator = 1	0.0150633 (0.0174441)	0.017749* (0.006035)	–	0.0366756* (0.0116755)
Freezer = 1	0.0543159** (0.00320551)	0.0246492* (0.0047957)	–	–
Washer = 1	0.049332* (0.029749)	0.0705512* (0.0065862)	–	0.0520365* (0.0097507)
Phone = 1	0.041436* (0.014093)	0.0303881* (0.005146)	–	0.0520352* (0.0105859)
Car = 1	0.0320051* (0.013705)	0.0372805* (0.0047294)	–	0.010143* (0.0089405)
Charcoal = 1	–	–	–0.0186647* (0.0058858)	–
Kerosene = 1	–	–	–0.0299878* (0.0056239)	–
Women's wage	–0.0009524* (0.0003912)	–	–	–

Table 4 (continued)

Variable	Jordan 1987	Oman 1993	Yemen 1994	West Bank and Gaza 1995
Family income	0.0000892*	**	**	**
	(0.0004335)			
Constant	0.764028*	0.2101281*	0.4881395*	0.3703877*
	(0.1049705)	(0.0256588)	(0.0316461)	(0.0879903)
Log dispersion	–16.6625	–2.842076*	–17.7951	–19.77571
parameter[a]	(–18.9379)	(–0.019343)	(–11.3376)	(–18.3687)
χ^2	4199.48	53349.70	18867.13	14682.58
Pseudo R^2	0.1756	0.1513	0.1225	0.2159
No. of observations	5418	63643	30849	13359
Log likelihood	–9858.2782	–149664.5788	–67584.5655	–26661.1736

[a] *Note:* Overdispersion parameter is the exponent of the value of the coefficient. The parameter estimate is not statistically different from zero in all but Oman's regression. The negative binomial regression fits Oman's data at the 5% level of significance.

* Indicates that the estimate is significant at the 5% level.

** Indicates that information is not available to estimate parameter.

Table 5. Fertility elasticity in Arab countries

Elasticity	Jordan	Oman	West Bank and Gaza	Yemen
Education	–0.0651	–0.03416	–0.07615	–0.01845
Age at marriage	–0.2555	–0.028809	–0.54827	–
Marriage duration	0.45081	0.222241	0.30243	–
Son preference	0.04011	0.07694	0.00579	0.05337
Wife's wage	–0.3221	–	–	–
Family income	0.02862	–	–	–

Source: Computed based on individual observations from the Poisson and negative binomial regressions.

Note: Elasticity parameters are computed as the average over all observations for each country. For Oman the elasticity parameters are derived from the negative binomial model, for the other countries estimates are based on the Poisson model.

ents, all else equal. We find that the interaction between religion and women's education does not adversely influence fertility. Further, all regressions do not yield any significant impact of education on the strong cultural preference for male children. The fertility mitigating effect of education, while significant, does not seem to be strong enough to change cultural preferences for male offspring or reverse pronatal religious beliefs.

Empirical evidence point to inter-country differences with respect to the impact of cultural, economic and education forces on fertility. The findings therefore corroborate the fixed-effects results obtained in Sect. 2 above. For example, son preference seems to be strongest in Oman and Yemen that have highest rural population. This suggests that boys are favored to girls because of the differential economic role they play as farm workers, wage earners and providers of social security to retiring (especially rural) parents and because they maintain family line. Second, the negative fertility impact of education appears strongest in Jordan and the West Bank and Gaza. This may be due to the fact that the overall level of education in Yemen and Oman is significantly lower than Jordanian and Palestinian level of education. The population censuses in both Yemen and Oman showed that an overwhelming majority of women have received little or no education and are illiterate. By contrast, women in Jordan and the West Bank and Gaza have 5.5 and 6.5 years of education respectively (Table 2). Fertility elasticity with respect to education varies between 0.018 in Yemen to 0.076 in the West Bank and Gaza (Table 5). Accordingly, increasing women's level of education from, say 4 to 8 years of schooling, would induce a fertility reduction of 7.6% in the West Bank and Gaza compared with 1.8% in Yemen. Fertility resulting from replacement of child death appears strongest in Yemen and Oman, the two countries with the largest rural population 80% and 64% respectively and highest infant mortality rates. Therefore, policy intervention aimed at reducing infant and child mortality and improving overall health standards can be effective in lowering fertility rates. Finally, some empirical evidence suggests that the fertility impact of improvements in living standards is conditional on the overall country level of economic and social development. In Oman where standards of living are much higher but education levels are lower than in Jordan, the West Bank and Gaza, improvements in standards of living are more strongly associated with elevated demand for children. Yemeni families with equally low education but markedly lower standards of living also tend to increase their demand for children as economic opportunities improve.

4. Conclusions and policy implications

Research conducted in this paper has relied on a variety of panel and micro level data in order to glean cross and within country variations in the pattern of Arab fertility. By and large our findings indicate that country heterogeneity is an important factor to take into consideration. Country invariant and time invariant patterns tested significant utilizing a panel data covering 11 countries during 1971–1996. Micro unit (family level) data for Jordan, Oman, West Bank and Gaza and Yemen have corroborated panel data findings and provided a microscopic treatment of the determinants at the family level. Our findings indicate that preference for male offspring and religious beliefs positively influence the number of births. One of our findings, robust at both the aggregate and family levels of analysis, pertain to the powerful negative impact of female education on fertility. The finding has implications for social resource allocation. Because of cultural and economic factors, the private allocation of resources (expenditure on education and health of boys and girls) may not be efficient from the prespective of society at large. However, women's education produces externalities (improving children's health, longevity, and scholastic achievement of children). It also influences family size. If the argument that large families and high population growth rates produce negative social externalities is correct (Chomitz and Birdsall 1991), then subsidizing female education could promote a socially more efficient pattern of investment that could maximize economic growth and decelerate population growth. Wages of working women negatively impact fertility. Increasing women's employment opportunities and 'exogenous' increases in their wage increase the opportunity cost of having children and lower fertility. Although further investigation is needed, our results indicate that the demand for children is price-inelastic. This implies that incentive-based government policies that attempt to reduce family size will require large penalties (large disincentive) in order to reduce family size and may even be regressive from the perspective of social welfare. Existing government food subsidies and pro-natal policies reduce the cost of having children and reinforce private (family) decisions for high fertility. To the extent that fertility is under control, child mortality influences birth as parents attempt to replace deceased children. There is a differential in rural-urban fertility that largely results from a combination of higher illiteracy rate and lower standards of living and higher infant mortality rate in rural areas. Policy intervention to reduce infant and child mortality and improve overall health standards should therefore especially target disadvantaged (mostly rural) households.

Endnotes

[1] We included other cross terms including education times urban rate. The results lowered the significance of the model fit.

[2] One referee pointed out that the number of dead children is correlated with the total number of births. Two options were tested. The first removes the number of dead children from the dependent variable and keeps it in the set of explanatory variables. The other keeps it but gauges mortality as a dichotomous variable. Total births (including those who died) is a better measure of fertility and it is the measure that is employed here. Regressions (not shown) that included the actual number of dead children were similar to those reported here although the size of the estimated coefficient was slightly smaller than in the categorical variable case.

[3] In addition to the impact of increased girl's education, nuclearization of the families and rising urbanization trends shift the age at marriage upwards and reduce fertility.

[4] Our findings here are in general agreement with the empirical findings elsewhere (see, for instance, Coale 1991).

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