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Age at motherhood in Japan*

John Ermisch[1] **and Naohiro Ogawa**[2]

[1] ESRC Research Centre on Micro-Social Change, University of Essex, Wivenhoe Park, Colchester CO4 35Q, UK
[2] Population Research Institute, Nihon University, 3-2 Misaki-Cho 1-Chome, Chiyoda-Ku, Tokyo 101, Japan

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Abstract. The paper analyzes factors influencing the age of motherhood in Japan, using both cross-sectional and time-series data. Both hazard rate and time series analyses support the hypothesis that better women's earning opportunities, as indicated by their educational attainments and relative pay, encourage Japanese women to marry and become mothers later in their lives. But both these analyses indicate that the trend toward later marriage and motherhood in Japan cannot be fully accounted for by improvements in women's educational attainments and earning opportunities, and the hazard analysis indicates that the strength of the trend increases with a woman's educational attainment.

Introduction

Japan was the first non-Western country to accomplish a demographic transition. Immediately after World War II, Japan's total fertility rate (TFR) was more than four children per woman. Subsequent to this short baby boom (1947–1949), the TFR fell precipitously at an unprecedented rate. By 1957, it had declined to 2.04 children per woman. Except for the "year of the fire horse" in 1966,[1] it remained virtually unchanged at the replacement level until the early 1970s, after which it declined continuously, reaching 1.54 in 1990.[2] Such rapid fertility declines,

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[1] A long-lasting Japanese superstition says that a woman born in this particular year is destined to an unhappy life and will kill her husband if she marries (Hodge and Ogawa 1991). In 1966, the TFR dropped to 1.58 per woman.

[2] Japan's current TFR is not as low as that for Italy (1.31), but is considerably lower than that for other Western industrialized nations such as France (1.78), England and Wales (1.84), and Sweden (2.14).

together with remarkable mortality improvements, have contributed to the rapid aging of the Japanese population, and it is expected that Japan will be the world's oldest known human population in the early part of the next century (Ogawa 1989).

In demographic terms, the sources of Japan's postwar fertility decline have changed over time (Atoh 1988; Ogawa and Retherford 1993). In the 1950s and 1960s, it was lower marital fertility that played the predominant role. This fall in marital fertility was facilitated by widespread abortion and a growing use of contraception (Hodge and Ogawa 1991). Following the enactment of the Eugenic Protection Law in 1948 (Oakley 1978), the number of reported abortion cases was more than one million per year between 1953 and 1961 (compared to 1.7 million births per annum), but the use of contraception became increasingly important since the early 1960s. The proportion of married women currently practicing contraception increased from 42% in 1961 to 58% in 1990.

Since the early 1970s, however, later marriage has been the primary contributor to the decline in fertility. For the age group 25–29, the proportion of women married decreased from 81.9% in 1970 to 59.6% in 1990 (Statistic Bureau 1992). The corresponding rise in the singulate mean age at first marriage for females was from 24.7 to 26.9 during this period (Ogawa and Retherford 1993).

These developments have been accompanied by rapid urbanization and a secular rise in educational enrollment, particularly for women. During the period 1950–1990, the proportion of people residing in urban areas increased from 37 to 77% (Ogawa 1986; Statistics Bureau 1992). Women's enrollment in junior colleges and universities has grown remarkably over the past few decades. In 1955, only 5% of women of the relevant age enrolled in either junior college or university, as opposed to 15% for men. In 1990, the former exceeded the latter by a slight margin, 37 and 35%, respectively (Ministry of Education 1991).[3]

The diminishing sex differential in education has contributed to a rapidly growing proportion of women working as paid employees (Ogawa and Ermisch 1993). Shimada and Higuchi (1985) claim that the growth of women's paid employment has been the most rapid in the recorded experience of advanced economies. The ratio of women's to men's hourly pay has also increased dramatically for recent generations of entrants to the labour force. Among persons aged less than 30 in full-time paid employment, it rose from 0.70 in 1970 to 0.83 in 1990 (Ministry of Labour, various years).[4]

Recent studies using aggregate data on age-specific fertility (Ogawa and Mason 1986; Osawa 1988) suggest that these changes in women's labor force participation and gender differentials in earnings have been largely responsible for the fertility decline. The present study uses data which is more closely associated with individual behavior: women's marital and birth histories and aggregate birth rates specific to both birth order and age.

Unlike Western industrialized nations, multigenerational households are still fairly common in contemporary Japan, although their proportion has been

[3] University enrollment remains lower for women: 15% compared to 33% for men; but the women's percentage is more than double its level in 1970 (7%).

[4] For all persons in full-time paid employment, the corresponding ratio increased from 0.54 in 1970 to 0.59 in 1990.

declining (Preston and Kono 1988).[5] According to data gathered from the 1990 round of the National Survey on Family Planning (Population Problems Research Council 1990),[6] approximately 40% of married Japanese women of childbearing age resided with their own or their husband's parents and/or grandparents immediately after their marriages. More than 80% of the coresident cases are the "patrilocal" type: a couple takes up residence with the husband's parents.

Another characteristic of traditional Japanese family life, arranged marriages, has, however, declined more markedly (Hodge and Ogawa 1991). The proportion of married women aged below 50 whose marriage was arranged decreased from 50% in 1981 to 40% in 1990; only 15% of recent marriage cohorts had arranged marriages (Population Problems Research Council 1990).

Although family size goals have become increasingly homogeneous in Japan over the last few decades, converging on two children (Hodge and Ogawa 1991), shifts in the timing of the first birth may have major effects on the fertility pattern. For this reason, there has been increased interest in the analysis of the age at motherhood in Japan (Feeney 1986; Ishikawa 1990; Otani 1988).

A few studies during the past decade have analyzed the associations between the age at motherhood and women's education, coresidence with parents, marriage form, urban residence and paid employment by women (Morgan et al. 1984; Otani 1989a, b). These studies used data from the 1974, 1982 and 1987 National Fertility Surveys. In contrast to the findings obtained by Morgan et al. (1984), the two most recent analyses, using the 1982 and 1987 surveys, revealed that women's educational achievements and pre-marital participation in paid employment play an important role in influencing the age at motherhood.[7]

In the present paper, we analyze the factors influencing the age at motherhood, using both cross-sectional and time-series data. The next section uses a sample of women from the 20th round of the National Survey on Family Planning conducted by the Mainichi Newspapers of Japan in June 1990. In the third section, the first birth probability is examined using time-series data compiled from vital statistics. Both types of data support the hypothesis that higher women's earning power relative to men's encourages postponement of motherhood, but changes in women's educational attainments and relative pay are only one factor in the rise in Japanese women's age of motherhood during the past 25 years.

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5 The proportion of three-generational households decreased from 16% in 1970 to 12% in 1990, while the proportion of nuclear households rose from 57 to 60% during the same time period (Statistics Bureau 1992). Changes in childbearing, mortality conditions and accompanying changes in age structure have placed considerable strain on Japan's traditional household structure. Moreover, equally rapid socioeconomic development and urbanization have occurred in parallel with demographic change. All of these factors have contributed to changes in the characteristics of Japanese households (Mason et al. 1992).

6 The 1990 round of the National Survey on Family Planning is the source of the micro-level data used in this paper. Further discussion of this data follows.

7 Otani (1989b) has also analyzed factors affecting the age at first marriage. His results show that the factors influencing the timing of the first birth basically account for the timing of first marriage.

Analysis of variation in age at motherhood among Japanese women

Since 1950, a series of nationwide sample surveys dealing with fertility and family planning have been undertaken almost every other year by the Mainichi Newspapers.[8] As in previous rounds, the 20th round used a stratified, multi-stage sampling procedure (Population Problems Research Council 1990). Although the target population of the earlier rounds was currently married women of childbearing age (less than 50 years old), the 20th round gathered information from all women, both married and single, aged 16–49. A total of 5,720 questionnaires were distributed and 3,769 completed, yielding a response rate of 72%.

Because educational attainment is an important explanatory variable and some young women may not have completed their education at the time of the survey, we restrict our analysis to women aged 22 and above in 1990. Thus, the analysis includes the generations born between 1941 and 1968. The age at first birth is measured in months since the time a woman reached the age of 17 years, 6 months (the youngest age at first birth in the sample was 17 years, 7 months).

The potential explanatory variables that could plausibly be considered as predetermined for the marriage and first birth decisions are limited in these data. They are a woman's education, her childhood residence, the number of siblings she has, and her birth cohort, Descriptive statistics from them are given in Table 1.

A number of theoretical perspectives would suggest that these variables might influence the timing of marriage and motherhood. For instance, economic theories associate a woman's education with her human capital, which is directly related to the opportunity costs of children, and Michael (1973) discusses other channels of influence of education on fertility, from an economist's perspective. Decisions about family size and the timing of births are affected by a woman's human capital (see Cigno and Ermisch 1989; Happell et al. 1984). The costs of bringing up children are also likely to differ between urban and rural residence.

There are 2,991 women in our sample, of whom 2,271 have had a first birth, and the remaining 720 are still childless at the time of the survey. The statistical methods that we employ use the information on these "censored" cases as well as the information on women who gave birth to identify systematic differences between women in the timing of motherhood.

Statistical models

In order to be more confident that our results do not reflect relatively arbitrary statistical assumptions, we consider a number of models which relate the age at first birth to the explanatory variables discussed above. A model of the *hazard rate* for the first birth is explicit or implicit in each of these models. Loosely speaking, the hazard rate is the probability of having a first birth in a short period of time conditional on surviving childless up to that time.

For demographers, the so called *proportional hazards* model is probably most familiar. In the proportional hazards model, the hazard rate for the first birth at age $t+17.5$ years is given by

$$h(t,X) = \lambda(t) \exp(X\Pi) \ , \quad \text{or} \quad \log h(t,X) = X\Pi + \log \lambda(t) \ , \tag{1}$$

[8] A detailed description of these national sample surveys is available elsewhere (see, for example, Hodge and Ogawa 1991).

Table 1. Means and standard deviations of the variables

	Mean	S.D.
Hazard rate analysis		
Woman's educational attainment:		
Less than senior high school[a]	0.123	
High school	0.531	
Junior college	0.264	
University	0.082	
Childhood residence: Urban	0.599	
Rural[a]	0.401	
Number of siblings	2.690	1.820
Age in 1990:		
22 – 24	0.076	
25 – 29	0.158	
30 – 34	0.157	
35 – 39	0.198	
40 – 44[a]	0.227	
45 – 49	0.184	
$N = 2{,}991$		
Aggregate time series analysis (1964 – 90)		
ln (RW [20 – 24])	– 0.495	0.071
ln (RW [25 – 29])	– 0.582	0.078
ln (MW [25 – 29])	0.599	0.242
ln (MW [30 – 34])	0.784	0.247
TRADB	0.185	0.071
logit (FBR [20 – 24])	– 2.741	0.331
logit (FBR [25 – 29])	– 1.472	0.234
ED [20 – 24]	23.09	10.51
ED [25 – 29]	17.87	11.10

[a] Reference category.

where $\lambda(t)$ is a function expressing the relationship between the woman's age and the hazard (that need not be specified); Π is a vector of parameters to be estimated, relating the explanatory variables to the hazard of a first birth. It is called the *proportional* hazards model because the shape of the hazard function with age is the same for everyone, but its level varies proportionally with the variables in X.

Because we do not need to specify $\lambda(t)$, this would appear to be a fairly flexible formulation, its main restriction being the proportionality assumption.[9] It proves, however, difficult to introduce unobserved woman-specific heterogeneity into this approach (see Lancaster 1991, pp. 263 – 272), and this is an important consideration.

In the model described by Eq. (1), we have assumed that all systematic differences between women in their first birth timing are accounted for by the set of explanatory variables, X. There may, however, be persistent differences in the first birth birth hazard between women that are not accounted for by the measured variables (e.g., differences in fecundity). In general, failure to account for such "unobserved heterogeneity" biases estimates of the pattern of age variation in the first hazard in the direction of a falling hazard with age (e.g., because a larger pro-

[9] Cox's (1972) partial maximum likelihood method is used to estimate the parameters in (1).

portion of the remaining childless women are less fecund), and it also produces biased estimates of β, even if the unobserved characteristics of women are uncorrelated with the variables in X (e.g., see Lancaster 1991, Chapts. 4 and 8).

To facilitate the introduction of such unobserved heterogeneity into our analysis, we also estimate so called *accelerated failure time models.* Before introducing such residual heterogeneity, they take the following form:

$$\log D = \alpha + X\beta + \sigma \log d_0 \ , \quad \text{or} \quad D = \exp(\alpha + X\beta) d_0^{\sigma} \ , \tag{2}$$

where D is the time to first birth (in months); X is a vector of explanatory variables; d_0 is a first birth time from a "baseline" distribution; α, β and σ are parameters to be estimated (α and σ are often called the intercept and scale parameters). The parameters are estimated by maximum likelihood under different assumptions about the distribution of d_0 (e.g., see Kalbfleisch and Prentice 1980). The three distributions considered are Weibull, Log-Logistic and Log-Normal.[10]

A model of the hazard rate is implicit in each of these failure time models. If an explanatory variable X_i has a positive effect on the time to first birth (i.e., its coefficient β_i is positive), then it has a negative effect on the hazard rate. The baseline distribution assumptions allow the hazard rate for first births to vary with age. The Weibull assumption implies a monotonic relationship between the hazard and age, while the Log-Logistic and Log-Normal distributions are more flexible, allowing for an increase and then a decrease with age. The pattern of variation of the hazard with age depends on the estimate of σ. Among the failure time models, only the Weibull is also a proportional hazards model.[11]

For our purposes, the main advantage of these failure time models is the relative ease with which we can introduce unobserved heterogeneity into the model. In order to explore the sensitivity of our estimates to the presence of such unobserved differences between women, we estimate two types of model that allow for unobserved heterogeneity. The first, suggested by Greene (1989), introduces a woman-specific variable, v, as a factor that scales the survival function, where v has a Gamma distribution with a unit mean and variance Θ.[12] We obtain the observed hazard function by integrating out v using the Gamma distribution:

$$\begin{gathered} h(t, X, \Theta) = \mu p (\mu t)^{p-1} / [1 + \Theta (\mu t)^p] \ , \\ \text{where} \quad p = 1/\sigma \quad \text{and} \quad \mu = \exp(-\alpha - X\beta) \end{gathered} \tag{3}$$

and if $\Theta = 0$, then this hazard function is identical to the Weibull hazard. Thus, we shall call this model a Weibull model with unobserved heterogeneity.

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10 These correspond to distributions of $\log(d_0)$ of Extreme Value (minimum), Logistic and Normal respectively (see Kalbfleisch and Prentice 1980).

11 Solving for the Weibull hazard rate from (2), $h(t,X) = \mu p (\mu t)^{p-1}$, where $t = D$, $p = 1/\sigma$ and $\mu = \exp(-\alpha - X\beta)$; or in terms of the logarithm of the first birth hazard: $\log h(t,X) = -X\beta p + [\log p - p\alpha] + (p-1)\log t$. In the Weibull specification, therefore, $\lambda(t) = p[\exp(-p\alpha)]t^{p-1}$ and $\Pi = -\beta p$.

12 Greene (1989) specifies a conditional survivor function $S(t|v) = v[\exp(-\mu t)]^p$, where $\mu = \exp(-\alpha - X\beta)$ and $p = 1/\sigma$. After integrating out v, the unconditional survivor function is $S(t) = [1 + \Theta(\mu t)^p]^{-1/\Theta}$.

The other model assumes that the unobserved heterogeneity takes the simple form of having two types of women. One type never will become mothers, and the probability of being this type is 1-*P*. Thus, the models in (2) only apply to women of the other type, who are observed with probability *P*. We specify $P = 1 - \Phi(\gamma)$, where $\Phi(\cdot)$ is the standard normal distribution function, and we estimate γ along with the other parameters.[13] This is often called "mover-stayer heterogeneity", and Heckman and Walker (1990) find that it performed best in modelling Swedish fertility dynamics. It clearly also addresses explicitly the issue of permament childlessness.[14]

Childlessness is relatively rare among Japanese women. In the 1941–45 birth cohort, only 9.8% of women remained childless. Each of the above models other than the last implies that, if women were fecund over their entire life, none would remain childless. The last model allows explicitly for childlessness, but we also find that some of the estimated models which do not make such an allowance entail childlessness at the time of menopause that is not grossly inconsistent with the observed incidence of childlessness.

Parameter estimates

The Kaplan-Meier estimate of the median age at first birth in our sample is 26.6. Systematic variation in age at motherhood among women is indicated by estimates of the parameters β, and these are shown for the two "best fitting" (in terms of log-likelihood) accelerated failure time models in the first two columns of Table 2.[15] The estimated β's are similar in the two models.

A powerful influence on the timing of motherhood is a woman's education. Better educated women become mothers significantly later. For instance, in the Log-Logistic model with mover-stayer heterogeneity, women completing senior high school have a median age at motherhood that is 9 months longer than those with less education; those having two years additional education to the junior college level wait about 15 months longer than women who complete high school, and women who study for an additional 2 years to obtain a university degree start childbearing about 2 years later than junior college graduates.

Women who resided in an urban area during their childhood become mothers about 5 months later than women from a rural background. In addition, the dichotomous age group variables indicate a tendency for later motherhood among women from the generations born since 1960, but these trend coefficients are not estimated very precisely, particularly in the Weibull model. The number of brothers and sisters that a woman has does not affect her first birth timing. Indeed, it is never statistically significant in all of the analyses that follow.

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13 The observed survivor function $S(t)$ is, therefore, $S^*(t)\cdot P + 1 - P$, where $S^*(t)$ is the survivor function for the women who are at risk of having children. The corresponding observed hazard function is $h(t) = h^*(t)\cdot P\cdot S^*(t)/S(t)$.

14 This model is sometimes also called a "split population" model (Schmidt and Witte 1989). It is possible to allow P to vary with observed variables, and to allow also for unobserved heterogeneity among the women who are at risk of having children.

15 We do not show the results for the Log-Normal model because we could not obtain a maximum for the log-likelihood when we allowed for "mover-stayer heterogeneity". When we ignored unobserved heterogeneity, the Log-Normal model consistently produced a lower log-likelihood than the models in Table 2, although the inferences that we would draw from the estimates of β are similar.

Table 2. Models of time to first birth. Women aged 22 or more in 1990 ($N = 2{,}991$)[a]

	Logistic[b]	Weibull[c]	Logistic-Interaction[b]	
Constant	4.402	4.386	4.397	
	(129.66)	(100.62)	(129.66)	
Woman's education				Education × age in 1990
High school	0.085	0.077	0.351	−0.00585
	(3.46)	(2.59)	(2.65)	(1.81)
Junior college	0.223	0.238	0.901	−0.0171
	(7.69)	(6.90)	(5.80)	(4.32)
University	0.414	0.461	1.274	−0.0221
	(9.27)	(9.29)	(5.17)	(3.30)
Childhood residence	0.072	0.083	0.073	
Urban	(4.19)	(4.04)	(4.31)	
Number of siblings	−0.004	−0.006	−0.003	
	(0.86)	(0.95)	(0.69)	
Age in 1990				
22−24	0.097	0.085	−0.075	
	(1.92)	(1.62)	(1.02)	
25−29	0.071	0.039	−0.092	
	(2.48)	(1.18)	(1.80)	
30−34	0.041	0.007	−0.060	
	(1.47)	(0.20)	(1.52)	
35−39	−0.094	−0.044	−0.060	
	(0.78)	(1.46)	(2.18)	
45−49	0.016	0.002	−0.045	
	(0.65)	(0.05)	(1.63)	
σ	0.222	0.214	0.222	
	(66.72)	(48.33)	(66.57)	
Θ		2.155		
		(23.60)		
γ	−1.293		−1.290	
	(32.54)		(32.75)	
P	0.902		0.902	
$\log L$	−2086.13	−2238.21	−2071.4	

[a] Absolute value of ratio of parameter estimate to its standard error in parentheses.
[b] "Mover-stayer" model with the probability of remaining childless being $1\text{-}P = \Phi(\gamma)$, where $\Theta(\cdot)$ is the normal distribution function.
[c] With unobserved heterogeneity according to a Gamma distribution with unit mean and variance Θ.

The estimates of the models in Table 2 indicate that the relationship between the hazard rate and age is peaked and skewed. As Fig. 1 illustrates for women born during 1946−50 in an urban area and obtaining a high school education, both the Weibull and Log-Logistic models have a peak *observed* hazard at about 27 years of age, although in other respects the age pattern of the hazard is quite different.

In the case of the Weibull model (with unobserved heterogeneity), the decline in the *observed* hazard reflects persistent unobserved differences between women (e.g., fecundity). Note the stastistically significant variance for the unobservable variable, v. The estimates of this variance (Θ) and σ indicates that while the

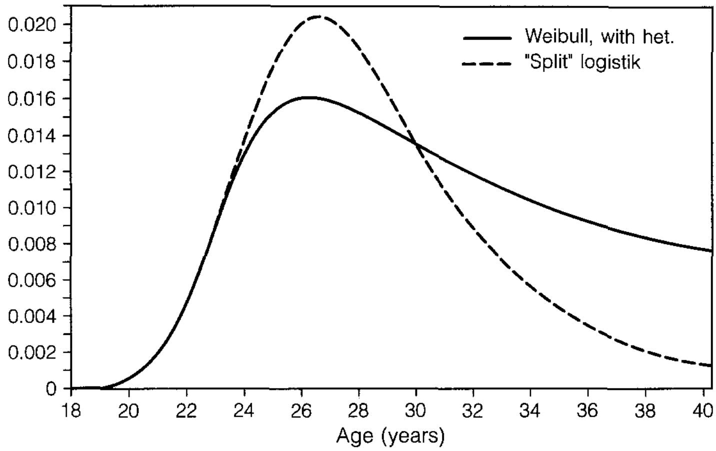

Fig. 1. First birth hazard rate

hazard increases with age for a given woman, the observed hazard in the population first rises and then falls.[16]

The estimated Log-Logistic model (with mover-stayer heterogeneity) exhibits, a much more concentrated hazard rate. The steep fall in the hazard supposedly reflects duration dependence of the hazard for women at risk of childbearing, but it could, in part, also reflect unobserved heterogeneity among these women.[17]

As the estimate of P indicates, the Log-Logistic model implies that 9.8% of the population are in the childless group of women, so that 10.6% are childless at the age of 40. In the Weibull model, which makes no explicit allowance for childlessness, 6.9% are still childless at 40 (see Fig. 2).

Table 3 shows estimates of the parameters of Cox's proportional hazards model. These again show powerful effects of a woman's education above high school, childhood residence and a trend toward later motherhood among recent generations. Women with higher educational attainment beyond high school have a lower first birth hazard, women completing junior college having about a 26% lower hazard and women completing university about a 53% lower hazard than women with high school education. Women who spent their childhood in urban areas have a 20% lower first birth hazard, and those born since 1960 have much lower hazards. These results support the qualitative conclusions from Table 2.

Thus, there are a number of characteristics of women that are systematically related to the age at which they become mothers. These results contrast with those of Morgan et al. (1984), particularly with their finding of no statistically significant associations between educational attainment and birth timing. This contrast may arise because of the more recent generations that we have analyzed. Morgan

[16] Heckman and Walker (1990) find that their best model of Swedish first births has an increasing Weibull hazard and "mover-stayer" heterogeneity. We also attempted to estimate a Weibull model with mover-stayer heterogeneity, but we were unable to find a maximum for the log-likelihood.

[17] We tried to estimate a Weibull model with mover-stayer heterogeneity *and* the woman-specific variable v (distributed as gamma with unit mean and variance Θ) for women at risk at having a birth, but we could not obtain a maximum for the log-likelihood.

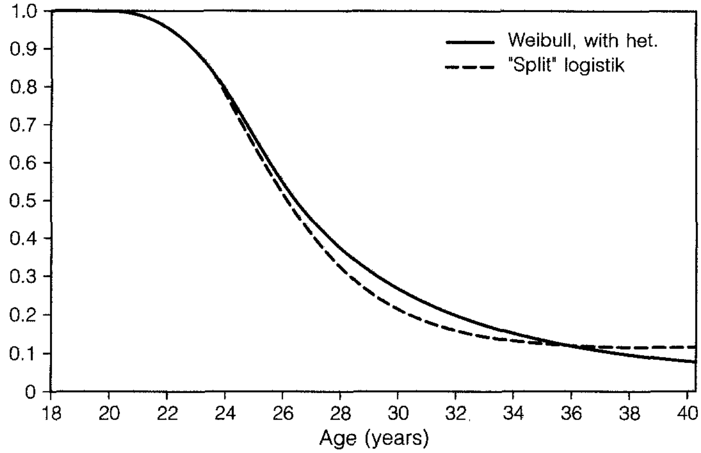

Fig. 2. Proportion childless

Table 3. First marriage and first birth proportional hazards models (Cox Model). (Women aged 22 or more in 1990)[a]

	Age at motherhood	Age at first marriage	First birth interval
Woman's education			
High school	0.066	−0.117	0.104
	(1.004)	(1.75)	(1.51)
Junior college	−0.231	−0.399	0.050
	(3.00)	(5.13)	(0.63)
University	−0.687	−0.786	−0.262
	(6.31)	(7.31)	(2.34)
Childhood residence	−0.217	−0.238	−0.177
Urban	(4.96)	(5.40)	(3.89)
Number of siblings	0.019	0.012	0.005
	(1.53)	(0.92)	(0.40)
Age in 1990			
22–24	−0.408	−1.067	0.337
	(2.00)	(5.46)	(1.50)
25–29	−0.209	−0.386	0.091
	(2.59)	(4.95)	(1.10)
30–34	0.015	−0.094	0.148
	(0.22)	(1.38)	(2.14)
35–39	0.112	0.024	0.156
	(1.84)	(0.40)	(2.49)
45–49	0.022	0.031	−0.006
	(0.35)	(0.50)	(0.10)
Log L	−16330	−16051	−14676
N	2991	2742	2254

[a] Absolute value of ratio of parameter estimate to its standard error in parentheses.

Table 4. Estimated median age at motherhood and percentage remaining childless (Log-Logistic model, Table 2)

	Median Age				Percentage childless	
	1945 Cohort		1968 Cohort		1945 Cohort	1968 Cohort
Woman's education						
Less than high school	26 years	2 months	25 years	3 months	10.5	10.2
High school	26 years	9 months	26 years	11 months	10.8	10.9
Junior college	27 years	3 months	30 years		11.0	13.8
University	28 years	7 months	33 years	7 months	12.1	21.6

Estimated median age at marriage and percentage never marrying (Log-Logistic model, Table 5)

	Median age				Percentage never	
	1945 Cohort		1968 Cohort		1945 Cohort	1968 Cohort
Woman's education						
Less than high school	22 years	2 months	23 years	11 months	4.0	3.9
High school	24 years	11 months	26 years	3 months	4.2	4.9
Junior college	25 years	3 months	29 years	2 months	4.4	8.4
University	25 years	10 months	33 years		4.6	19.5

et al. analyzed the 1929–1938 birth cohorts using data from the 1974 Japan Fertility Survey.

As a test of this hypothesis, we interact woman's age in 1990 with the educational variables, and include these variables in the model. If the discriminatory power of educational attainment for first birth timing has increased for more recent generations, then the coefficients on these interaction variables should be negative, and the size of the coefficients should increase with the level of education. As the final column of Table 2 shows, they are indeed negative and statistically significant, both individually and as a group, and increase in size with education level.[18]

Furthermore, the difference in the impact of education on age of motherhood between recent and older generations of women is large. To illustrate, Table 4 calculates the median ages of motherhood from the parameter estimates for the Log-Logistic interaction model in Table 2.[19] The effects on median age from moving between the consecutive levels of education are 7, 6 and 16 months for women born in 1945, but these rise to 20, 37 and 43 months for women born in

[18] The chi-square test of their joint significance yields a value of 29.4 (3 degrees of freedom). Similar results are obtained in the Weibull ($\chi^2 = 25.6$) and proportional hazards ($\chi^2 = 20$) models, although in the latter the interaction coefficient for high school education is not individually significant.

[19] Note that, in this interaction model, the parameter β_j associated with educational level j varies with birth cohort: $\beta_j = \beta_{oj} + \beta_{aj} \cdot \mathrm{AGE}$, where AGE is the age of the woman in 1990, and β_{oj} *and* β_{aj} are, respectively, the estimated coefficients of educational level j and its interaction with age of the woman. At the mean age of the woman in our sample, 36.5 (born in 1953–54), these are 0.14, 0.27 and 0.47, which are, as we would expect, similar to the estimates of the effects of education in the first column of Table 2.

1968. The cohort trends in the interaction model vary with educational level. While they are, on average, similar to those discussed earlier, Table 4 illustrates that they are steeper for higher education levels.

Thus, systematic variance in age at motherhood associated with education has become increasingly important over time. It is not surprising that it was not apparent among the generations studied by Morgan et al. (1984).

This increasing impact of education on the timing of the first birth may reflect the rapid rise in the participation of women in paid employment noted above. Women with higher educational attainment have higher earning potential (Ogawa and Ermisch 1993), entailing a higher opportunity cost of childbearing for more educated women who work outside the home. As women in more recent generations expect to be employed for more of their life, the opportunity cost element of education increases in importance, and this may explain the emergence of a strong positive gradient between educational attainment and age at motherhood.

It is very rare for Japanese women to become mothers before marrying. In our sample, only 1.4% of the women aged 22 and over who had given birth, gave birth before marriage. Later motherhood could reflect postponement of marriage or a longer interval between marriage and motherhood. In order to study which is more important, we estimate similar models for age at first marriage and for the timing of first birth within marriage.

Age at first marriage

Marriage has been almost universal in Japan. Only 2% of women born during 1941–45 remained unmarried in 1990. The time to first marriage is measured in terms of months since a woman's 16th birthday. Again we restrict our analysis to women aged 22 or above in June 1990, and after rejecting women who were remarried, divorced or widowed, or who did not state whether or not they had remarried, there are 2,742 women in the sample, of whom 2,288 had married by the time of the survey.[20] The median age at first marriage in the sample is 25 years.

Again, the Log-Logistic model with mover-stayer heterogeneity produces a higher log-likelihood than the Weibull model with unobserved heterogeneity (first two columns of Table 5). It indicates that 96.5% of women will eventually marry. The estimates of the impacts of the explanatory variables are, however, similar across the two model specifications, and similar to those from the proportional hazard model (Table 3, column 2).[21]

They all indicate that, not surprisingly, women who remain in education longer marry later, but the size of the effects are interesting. For instance, using the same base case as in the previous analysis, the results from the Log-Logistic model suggest that 2 years of junior college beyond high school increases the median

[20] The date of first marriage was not available for remarried, divorced or widowed women, of whom there were 153 (about 5%) in our sample. Another 32 women did not state whether or not they were remarried, and for 62 women the recorded date of marriage was after the survey date. The youngest age at marriage in the sample is 16 years, 5 months.

[21] Both estimated models have a median age at marriage of 24 for the base case, and the first marriage hazard reaches a maximum at the age of 25 years 4 months in both models. As in Fig. 1, the hazard in the Log-Logistic model declines faster after the peak.

Table 5. Models of time to first marriage. Women aged 22 or more in 1990 ($N = 2{,}742$)[a]

	Logistic[b]	Weibull[c]	Logistic-interaction[b]	
Constant	4.403	4.371	4.393	
	(146.12)	(125.50)	(129.66)	
Woman's education				Education × age in 1990
High school	0.085	0.093	0.456	−0.00825
	(3.87)	(3.84)	(3.58)	(2.69)
Junior college	0.211	0.227	0.928	−0.0178
	(8.16)	(8.07)	(6.50)	(4.93)
University	0.350	0.383	1.389	−0.0267
	(9.23)	(9.78)	(6.65)	(4.80)
Childhood residence	0.065	0.071	0.067	
Urban	(4.30)	(4.26)	(4.44)	
Number of silblings	−0.003	−0.004	−0.002	
	(0.66)	(0.70)	(0.43)	
Age in 1990				
22−24	0.233	0.224	0.016	
	(5.58)	(5.39)	(0.24)	
25−29	0.093	0.084	−0.099	
	(3.85)	(3.21)	(2.04)	
30−34	0.036	0.033	−0.086	
	(1.49)	(1.27)	(2.34)	
35−39	−0.001	−0.005	−0.050	
	(0.03)	(0.22)	(2.04)	
45−49	0.012	0.006	0.052	
	(0.52)	(0.25)	(2.08)	
σ	0.200	0.191	0.199	
	(59.20)	(38.70)	(59.05)	
Θ		1.516		
		(19.26)		
γ	−1.806		−1.798	
	(29.40)		(29.76)	
P	0.965		0.964	
$\log L$	−1409.5	−1465.1	−1389.3	

[a] Absolute value of ratio of parameter estimate to its standard error in parentheses.
[b] "Mover-stayer" model with the probability of remaining single being $1\text{-}P = \Phi(\gamma)$, where $\Phi(\cdot)$ is the normal distribution function.
[c] With unobserved heterogeneity according to a Gamma distribution with unit mean and variance Θ.

age of marriage by 13 months, and 4 years of university delays marriage by 29 months, compared to women who stop their education at high school graduation. Thus, an additional year of education postpones marriage by much less than a year. Women who grew up in urban areas marry about 6 months later than those who were children in rural areas.

There is also evidence of a trend toward later marriage among women born since 1960, which is independent of the upward trend in educational attainments in successive generations. The estimates from the Log-Logistic model suggest that the median age of marriage for the latest cohort (born 1966−68) will be 26 years, or 25 months later than the 1946−50 cohort.

The final column of Table 5 shows estimates of the education-birth cohort interaction model. As with age at motherhood, education becomes a better predictor of age at marriage for more recent generations. For instance, as Table 4 illustrates, the marginal effects of consecutively higher levels of education on median age at marriage change from 9, 4 and 7 months for women born in 1945 to 28, 35 and 46 months for women born in 1968.[22]

First birth interval within marriage

Marriage and parenthood are almost synonomous in Japan. Among married women born during 1945–45, only 4% remained childless. Of the 2,254 women in our sample who were in their first marriage and did not have a child before marriage, 2,136 had a child before the time of the survey.[23] The median duration in the sample is about 15 months. As this suggests, many Japanese women have very short first birth intervals: in this survey, 12% have a birth within 8 months of marriage and 20% within 9 months. The Ninth Japanese Fertility Survey also indicates that many Japanese women are pregnant at marriage: about a quarter of those married during 1980–84 (Institute of Population Problems 1988; Otani 1990, 1991).

This suggests difficulties for interpretation of hazard models of the interval between marriage and first birth, because marriage and childbearing tend to be joint decisions in Japan. From an analytical point of view, it is, therefore, odd to condition the study of first birth timing in Japan on marriage. Nevertheless, such hazard models have descriptive value, and Table 6 shows that the two accelerated failure time models have similar "explanatory" power. Both, and the Cox proportional hazard model in column 3 of Table 3, point to the same conclusions.

First, while there is less systematic variation in the first birth interval than in age at first marriage, there is a significant amount. In particular, university educated women become mothers later in marriage, their median interval being about 4 months longer. Women who grew up in urban areas also wait about a month longer after marriage to have their first child than women who were children in rural areas.[24] Thus, university educated women and those brought up in urban areas both marry later and have their children later within marriage.

There is also a trend toward more rapid childbearing within marriage among women born since 1950, which appears to have accelerated among women born since 1960. According to the estimates of the Weibull model in column 2, the latest cohort (born during 1966–68) will have a median first birth interval 6 months shorter than the 1946–50 cohort. This is also evident from Table 4, which shows that the gap between median ages of motherhood and marriage is much smaller for the 1968 cohort than for women born in 1945. These results are consis-

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22 Again, similar results are obtained from the Weibull and proportional hazards models. The chi-square statistics for the joint significance of the interaction terms (3 degrees of freedom) are 40.4, 35.2 and 30 for the Log-Logistic, Weibull and proportional hazards models respectively.

23 Thirty-four women had a birth before marriage, and 454 had never married.

24 Whether a woman had an urban or rural residence at marriage makes no significant difference to the first birth interval; neither does the number of siblings of the husband. These variables were excluded from the model, and this is statistically acceptable.

Table 6. Models of first birth interval. Women aged 22 or more in 1990 ($N = 2{,}254$)[a]

	Logistic[b]	Weibull[c]	Weibull-interaction[c]	
Constant	2.666	2.466	2.466	
	(46.64)	(40.55)	(40.49)	
Woman's education				Education × age in 1990
High school	−0.011	0.009	−0.146	0.004
	(0.27)	(0.20)	(0.61)	(0.68)
Junior college	0.002	0.024	0.496	−0.0102
	(0.05)	(0.48)	(1.48)	(1.46)
University	0.259	0.246	0.484	−0.0064
	(3.87)	(3.74)	(1.26)	(0.65)
Childhood residence	0.100	0.084	0.086	
Urban	(3.47)	(2.97)	(3.06)	
Number of siblings	−0.001	0.000	0.001	
	(0.07)	(0.05)	(0.12)	
Age in 1990				
22–24	−0.327	−0.465	−0.429	
	(2.87)	(3.93)	(2.76)	
25–29	−0.099	−0.216	−0.231	
	(2.08)	(4.83)	(2.59)	
30–34	−0.066	−0.111	−0.125	
	(1.54)	(2.66)	(1.87)	
35–39	−0.082	−0.083	−0.084	
	(2.04)	(2.10)	(1.83)	
45–49	0.015	0.022	0.014	
	(0.36)	(0.49)	(0.28)	
σ	0.361	0.227	0.227	
	(52.97)	(32.29)	(32.19)	
Θ		3.118	3.131	
		(20.59)	(20.58)	
γ	−1.834			
	(32.71)			
P	0.967			
$\log L$	−2433.0	−2412.4	−2408.1	

[a] Absolute value of ratio of parameter estimate to its standard error in parentheses.
[b] "Mover-stayer" model with the probability of remaining childless being $1 - P = \Phi(\gamma)$, where $\Phi(\cdot)$ is the normal distribution function.
[c] With unobserved heterogeneity according to a Gamma distribution with unit mean and variance Φ.

tent with the upward trend in the proportion of women who are pregnant at marriage (Institute of Population Problems 1988; Otani 1990, 1991).

While the individual interaction coefficients in Table 6 are not statistically significant,[25] there is evidence of changes in the impact of education over time in Table 4. University educated women born in 1945 had the largest difference between median ages of motherhood and marriage (33 months), but university

[25] The chi-square (3 degrees of freedom) value of 8.6 (in both the Weibull and Log-Logistic models) indicates joint statistical significance of the interaction coefficients at the 0.05 level, but not at the 0.01 level.

educated women born in 1968 had the smallest (7 months). The gap declined much less for women who did not complete high school.

Thus, while these recent generations are marrying later, they are having their children earlier in marriage.[26] Our analysis clearly shows that the later marriage trend dominates in determining the trend toward later motherhood among these recent cohorts.

"Traditional values" and the timing of motherhood

In the time series analysis of the next section, we examine the effect of two measures of "traditional values" on first birth rates. These variables can also be used in the analysis of variation in the timing of first birth among women. But as these are only measured for married women, we must restrict our sample to married women. This can create a selection bias problem: for instance, women who wish to have children earlier probably marry earlier, thereby disproportionately representing young married mothers in the sample. To reduce this problem, we restrict the sample to married women aged 30 or more in 1990, and as the results are similar with other models, we only report the results for the Weibull model with unobserved heterogeneity, which was the preferred model in terms of log-likelihood.

The traditional value variables are dichotomous and composite. The first (called TRADA later) is unity if the women had an arranged marriage *and* resided with the husband's parents after marriage *and* had an education up to junior high school or less. It has an effect on the timing of motherhood on the margin of statistical significance (t-ratio = 1.67): women who are "traditional" in this sense (only 3% of the sample) have their first birth about 9 months later. The second traditional values variable (later called TRADB) equals unity if the woman had an arranged marriage *and* resided with the husband's parents after marriage. It has a smaller effect, which is never statistically significant. The individual components of these variables are, however, statistically significant and interesting, because they affect the timing of motherhood in different directions, as we now show.

The first column of Table 7 shows the estimates of the parameters of the age at motherhood model. The age group (cohort) variables were not statistically significant as a group or individually; thus they have been dropped from the model. Having an arranged marriage is associated with a median age at motherhood about 9 months later, while living in a "patrilocal household" (i.e., with the husband's parents) is associated with bringing motherhood forward by about 7 months.

Caution should be exercised in interpretation both of these results because of potential endogeneity. The second and third columns show that the delaying impact of arranged marriages almost solely operates through delay in marrying, and it may be the case that it is because a woman is still single at an "advanced" age that a marriage is arranged (Otani 1991).

Living in a patrilocal household after marriage is associated with both earlier marriage and quicker motherhood after marriage, but the latter impact is not statistically significant at levels below 0.10: in terms of the medians, it is associated

[26] When age at marriage is entered as a regressor in the models for the first birth interval it is not, however, statistically significant.

Table 7. Models of first marriage and first birth. Married Women aged 30 or more in 1990[a]

	Age at motherhood	Age at marriage	First birth interval
Constant	4.415	4.472	2.537
	(110.30)	(137.57)	(33.65)
Woman's education			
High school	0.057	0.065	0.002
	(2.26)	(3.27)	(0.05)
Junior college	0.165	0.164	−0.010
	(5.45)	(6.72)	(0.19)
University	0.330	0.247	0.224
	(6.63)	(6.06)	(3.36)
Childhood residence Urban	0.087	0.065	0.117
	(3.80)	(3.51)	(3.46)
Residence at marriage Urban	−0.004	0.000	−0.054
	(0.17)	(0.02)	(1.37)
Number of siblings			
Wife	0.003	0.002	0.012
	(0.60)	(0.59)	(1.48)
Husband	−0.0166	−0.0154	−0.0015
	(3.66)	(4.21)	(0.20)
Patrilocal household	−0.072	−0.043	−0.045
	(3.87)	(2.89)	(1.53)
Arranged marriage	0.097	0.101	0.059
	(5.19)	(6.70)	(1.87)
Age at marriage			−0.0014
			(3.68)
σ	0.189	0.199	0.205
	(33.30)	(35.16)	(31.53)
Θ	1.499	0.718	3.427
	(17.65)	(11.48)	(20.58)
Log L	−1033.16	−569.69	−2040.77

[a] Absolute value of ratio of parameter estimate to its standard error in parentheses; $N = 1{,}962$.

with marriage 4 months earlier and a first birth interval less than a month shorter. But it may be that women who wish to have their children sooner choose to live with their husband's parents after marriage to obtain help with child care. Questions in the survey concerning reasons for coresidence with parents or parents-in-law suggest that the main motives are cultural: only 13% said they were coresiding so that their parents (-in-law) could help with housekeeping and child care.[27] This reduces concern about the endogeneity of living in a patrilocal household, but its effect is still relatively small.

Education and childhood residence have similar impacts on this truncated sample of married women to those discussed above, and in this sample we also find that women who married husbands with more brothers and siters married earlier, leading them to become mothers earlier. There is also some evidence that

[27] Up to two reasons could be given. The reason chosen most often (by 60% of respondents) is "obligations as eldest son-daughter, and 23% said they we coresiding because of the parents' (-in-law) wishes".

women who marry later have their children sooner within marriage, but the effect is very small.

Analysis of variation in first birth rates over time

In econometric analyses of changes in birth rates over relatively short periods of time (such as the 27 years in this analysis), it is natural to focus on economic variables, because these are more likely to be measured at regular intervals, and they are also likely to exhibit more short term (high frequency) variation. We draw, therefore, primarily on economic theories of fertility in selecting explanatory variables.

These suggest that men's and women's wages should have an important effect on fertility. Higher real wages mean higher real income, thus a couple can afford more children and to have them sooner. The latter would be particularly important if there are constraints on borrowing against future income. But they also mean more income lost by those caring for additional children. Since most childrearing is done by the mother, higher women's wages raise the cost of a child by increasing earnings foregone. Thirdly, the higher income that higher real wages entail means that more can be spent in raising each child (higher "child quality"), but this means that each child costs more. Thus, with mothers doing most of the childrearing, higher men's wages mainly affect childbearing through their effect on a couple's income, while higher women's wages have their main effect through their impact on the opportunity cost of children.

The age at which women decide to become mothers reflects the influences on both the timing and number of children. For instance, in Cigno and Ermisch's (1989) model, higher men's earnings are predicted to encourage women to start childbearing earlier, but the effect on completed family size is ambiguous, because they also raise the demand for child quality. Higher earning capacity by the wife at marriage is predicted to encourage earlier childbearing but smaller families, while a higher general level of women's wages is predicted to discourage early childbearing.

Their model assumes that couples are able to borrow if necessary,[28] but in the absence of sufficient assets and access to capital markets, there is a consumption-smoothing motivation in the timing of childbearing. A couple would like to synchronize the costs of children with a period in which the husband's earnings are relatively high. By raising the cost of children, a wife's higher earning capacity would be associated with a later start to childbearing (see, for example, Happell et al. 1984).

This discussion has focused on childbearing decisions by a couple, but age at motherhood also reflects age at marriage. If is often contended, however, that probably the most important source of benefits from marriage in Japan is the raising of children, with the relationship with one's spouse given much less importance (Coleman 1983; Morgan et al. 1984). Thus, factors which encourage people to become parents earlier (later) in their lives also encourage earlier (later) marriage, particularly because childbearing outside marriage is rare in Japan.

[28] It also assumes that child quality depends on the age of the mother at birth and the amount of resources expended on a child, and it takes account of the interaction between the time patterns of childbearing and maternal earnings.

The theories discussed do not always give unambiguous predictions. For example, while a higher earning capacity of the wife at marriage may encourage earlier childbearing for a given desired family size, it also reduces the latter, thereby tending to reduce birth rates throughout the woman's life and possibly producing a higher age at motherhood. The strong effect of women's education on the age at motherhood estimated in the previous section suggests, however, that higher women's real wages encourage later marriage and motherhood, and the theories above suggest that higher men's wages have the opposite effect.

Data

In the analysis of aggregate data over the period 1964–1990, we model the age-specific first birth rate, defined relative to the population of childless women of a given age. Changes in the "average" first birth rates for women aged 20–24 and 25–29 are shown in Fig. 3.[29]

We measure the general level of women's and men's age-specific wages by gross and average monthly contractual earnings of full-time employees for five-year age-groups of each sex (Ministry of Labour, various years), deflated by the consumer price index. As women marry men somewhat older, the men's earnings relevant to a specific five-year age group of women are assumed to be those of the next oldest age group of men. These variables change annually. When they do so, they not only affect current resources and resources foregone, but also expectations about the future values of these variables.[30]

Figure 3 also shows the full-time monthly pay of women relative to men (five years older) for the two age groups of women. The negative correlation between the first birth rates and relative pay in the figure supports the predictions of the economic theories.

We have not, however, ignored non-economic variables. As Japan has been undergoing relatively rapid social change, variables measuring changes in *societal values* may also play a part in explaining marriage and fertility developments (Retherford et al. 1994). We construct two measures of "traditional values" in Japan based on responses to bi-annual surveys by The Mainichi Newspapers (like that used in the previous section): the proportion of women married in a particular year who had an arranged marriage *and* resided with the husband's parents after marriage *and* who had education up to junior high school or less, TRADA, or, alternatively, who satisfy only the first two criteria, TRADB. Because of sampling variation in the survey, these indicators were smoothed by exponential smoothing. Figure 4 shows that these measures of traditional values fall over time. As the exogeneity of the traditional value indicators may be questionable, we estimate models with and without them.

In order to enhance the comparability of the aggregate analysis to the hazard rate analysis above, we also construct a cohort-specific educational attainment variable: the percentage of women in a birth cohort who enrolled in junior college

29 These are computed by taking the simple average of the "logit" of the first birth rate over the five ages in the group in each year and transforming this average back to its natural value. The "logit" is the natural logarithm of the birth rate divided by its complement.

30 We also considered unemployment rates, the inflation rate, real housing rents, and pension benefits as time-varying influences on first birth rates, but these did not make a significant contribution to the explanation of changes in first birth rates over time.

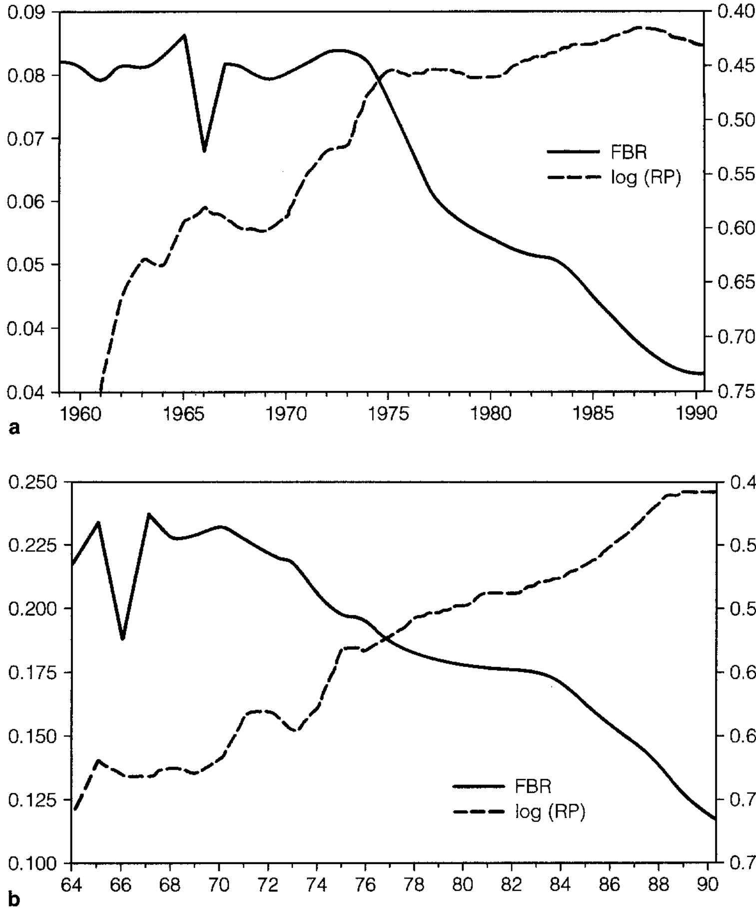

Fig. 3. a First birth rate and relative pay (women aged 20–24), **b** First birth rate and relative pay (women aged 25–29)

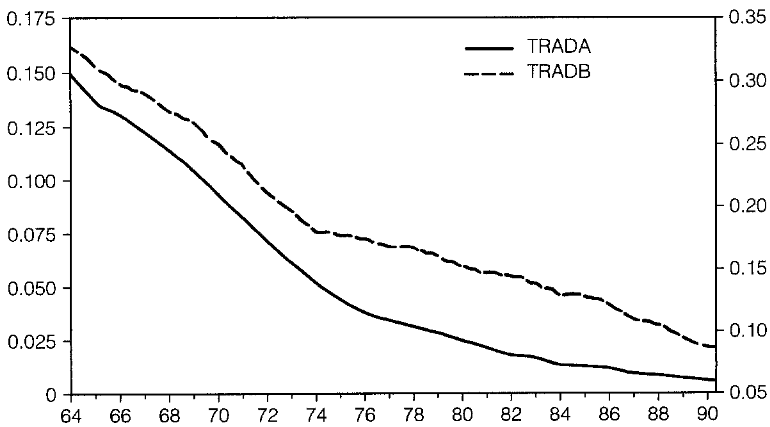

Fig. 4. Measures of "traditional values"

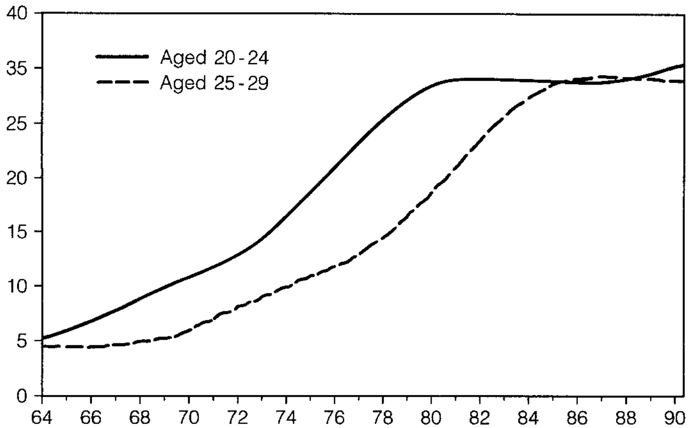

Fig. 5. Percentage of women who enrolled in University of junior college

or university.[31] It is averaged for the five cohorts represented in an age group in a given year, and is shown in Fig. 5. It could be considered as a lifetime measure of women's earning power, which is not subject to the year-to-year fluctuations which characterize women's annual pay.

Figure 3 illustrates the effect of the "year of the fire horse" on first birth rates in 1966 and the years preceding and following. In our estimation we try to control for this phenomenon by including a variable in the birth rate equations that takes the value 1 in 1966 (the year of the "fire horse"), -0.5 in 1965 and 1967 and zero elsewhere. The values for 1965 and 1967 are meant to reflect any tendency to shift births out of 1966 into the adjoining years in order to avoid the year of the fire horse.

Estimates of first birth rate equations

Our theories tell us little about how responses to economic variables are distributed over time. They only suggest which variables may be important in a steady-state environment. The dynamic responses of age of motherhood to these must, therefore, be inferred from the data. One bit of dynamic information is, however, available. Allowing for biological lags, it takes at least one year for the time-varying economic variables to affect the birth rate.

Thus, the foundation of the analysis is a set of equilibrium relationships relating the level of the first birth rate at each woman's age to the levels of the economic and social variables. To ensure that birth rates remain between zero and one, the birth rate is transformed by taking the natural logarithm of the rate divided by its complement (its "logit"):

$$y_{ka} = A_a + \beta_{1a} X_{at-1} + \beta_{2a}\,\mathrm{TRAD}_t + \beta_{3a}\,\mathrm{ED}_k \ , \tag{4}$$

[31] It should be noted that in Japan virtually all people who enroll at a given stage of education complete that stage.

where k, a and t designate birth cohort of the women, age and year respectively ($a = t - k$), and

y is the "logit" of the first birth rate (FBR),
X_a is a vector of time-varying economic variables, $\ln(\mathrm{RW})_a$ and
$\ln(\mathrm{MW})_{a+5}$ ($\ln(\cdot)$ denotes the natural logarithm).

RW_a is the ratio of full-time monthly earnings of women aged a to the monthly earnings of men aged $a+5$, MW_{a+5}; TRAD is one of the indicators of traditional values discussed above; and ED is the percentage of women in a cohort who enrolled in junior college or university. The vector β_{1a}, β_{2a}, β_{3a} and A_a are parameters to be estimated, and these vary with age. In the estimation, women are grouped into age groups defined over five years, with responses assumed constant within the group. Thus, an age group is designated by a and all variables subscripted by k are simple averages over the cohorts in the particular age group in year t.

Figures 3, 4 and 5 indicate that the first birth rate, women's relative pay, women's education and traditional value series are trended, and this is also true of men's real wages. It is now well known that there is a danger of spurious regressions and that the distributions of the parameter estimates are non-standard if such trended variables are non-stationary (e.g., Phillips 1986). Examination of the correlograms (the set of sample correlation coefficients between the contemporary value of a variable and its lagged values) for each of our variables suggests that while the first-order autocorrelation coefficient is high, the series may be stationary; in each case the correlogram "dies out".[32] Because of likely bounds on the birth rate, relative pay, women's education and traditional value series, our a priori belief in stationarity is also strong for these variables, but less so for men's real wages. In any case, we treat all series as stationary in the analysis which follows.

We begin by estimating some simple, "static" regressions corresponding to Eq. (4), and these are shown in Table 8. As we consider women's relative pay and their education as alternative measures of women's earning power, only one of the two is included in a regression. Of the equations with relative pay, those which omit the potentially endogenous traditional value variables show strong negative effects of women's pay and positive effects of men's pay, as our theories and Fig. 3 had suggested. Of the two traditional value measures, the second (TRADB) performs better, and we concentrate on it. When it is included, the explanatory power of women's relative pay for women aged 20 – 24 diminishes substantially, but for women aged 25 – 29 their relative wage retains considerable explanatory power, although its coefficient is halved. When women's education substitutes for their relative pay, it also exhibits a strong negative effect on the probability of a first birth, but the inclusion of the traditional value variable substantially reduces the estimated impact of education, and indeed eliminates its effect for women aged 25 – 29.

The low values of the Durbin-Watson statistic in Table 8 confirm that we need to pay attention to the dynamics of responses. We focus on the dynamic specifica-

[32] A battery of tests for a "unit root" (non-stationarity) have been developed recently. It is well known that these lack power to reject the hypothesis of a unit root when the series is stationary but exhibiting high autocorrelation. This is particularly the case in our small sample.

Table 8. Static regressions, 1964–1990. First birth rates of Japanese women in their twenties (Dependent Variable: ln $[FBR_{at}/(1\text{-}FBR_{at})]$)

Age group	Women aged 20–24		Women aged 25–29	
	(1)	(2)	(1)	(2)
Constant	−5.867	−6.539	−3.964	−4.262
	(6.99)	(15.46)	(15.47)	(16.92)
DumFireH	−0.094	−0.101	−0.123	−0.138
	(0.63)	(1.37)	(2.42)	(3.04)
$\ln(RW_a)_{t-1}$	−5.601	−0.418	−3.803	−1.898
	(4.36)	(0.47)	(3.44)	(2.50)
$\ln(MW_{a+5})_{t-1}$	0.561	2.403	0.324	1.065
	(1.55)	(8.52)	(3.44)	(3.73)
$TRADB_t$		12.046		4.680
		(8.47)		(2.71)
R^2	0.737	0.938	0.938	0.953
S.E.	0.1803	0.0893	0.0620	0.0549
D.W.	0.293	0.884	0.953	0.781
Constant	−2.135	−5.593	−1.142	−4.573
	(29.50)	(13.05)	(19.41)	(7.27)
DumFireH	−0.143	−0.110	−0.185	−0.152
	(1.15)	(1.72)	(2.44)	(3.00)
ED_k	−0.043	−0.014	−0.020	0.002
	(6.04)	(2.72)	(6.95)	(0.48)
$\ln(MW_{a+5})_{t-1}$	0.685	2.299	0.027	1.753
	(2.39)	(9.29)	(0.23)	(5.39)
$TRADB_t$		10.124		9.428
		(8.10)		(5.47)
R^2	0.815	0.953	0.860	0.941
S.E.	0.1515	0.0776	0.0929	0.0619
D.W.	0.110	0.918	0.238	0.840

Note: Absolute value of "t-ratios" in parentheses.

tions with women's relative pay rather than their education. Education was never statistically significant in the dynamic equations for women aged 25–29. While it was significant over the period 1964–84 for women aged 20–24, its coefficient was not stable, and it was not significant over the full sample period, 1964–90. Comparison of Figs. 3 and 5 suggests why education is not significant: educational attainments level off during the 1980s while first birth rates fall dramatically.

Our analysis initially allows for a relatively general distributed lag formulation, which includes two lags of the wage variables and one of the dependent variable (the smoothed traditional value variable is not lagged). If any distributed lags in response are to refer to roughly the same women, then any lags of y should be taken *within a cohort* of women, and we do this. The distributed lag formulation could be related to the formation of expectations about economic variables in an uncertain world, to behavioral inertia, or to biological factors delaying conception.

We simplify the general formulation by dropping statistically insignificant variables and modelling the disturbance term u_t where appropriate. The parameter estimates that we show below are for the period 1964–90, but as the true test of any econometric model is how well it does outside the sample period within which it was estimated, we also searched for the best dynamic specification over the period 1964–84 and performed tests of parameter stability for 1985–90.

When omitting the potentially endogenous traditional values variable, the most satisfactory specifications for women aged 20–24 and 25–29 respectively are:[33]

$$\begin{aligned} y_{kt} = & \underset{(0.62)}{-0.292} + \underset{(7.17)}{0.846} \cdot y_{kt-1} - \underset{(2.18)}{1.022} \cdot \ln(\mathrm{RW}_a)_{t-2} - \underset{(9.97)}{0.236} \cdot \mathrm{DumFireH} \\ & - \underset{(1.05)}{0.310} \cdot \ln(\mathrm{MW}_{a+5})_{t-2} - \underset{(1.70)}{0.522} \cdot \Delta \ln(\mathrm{MW}_{a+5})_{t-1} + u_t \end{aligned} \tag{5}$$

$$\text{and } u_t = \underset{(6.35)}{0.782} \cdot u_{t-1} + \varepsilon_t$$

$$R^2 = 0.990,\ \mathrm{SE} = 0.0378,\ \mathrm{DW} = 1.98,\ \chi^2(6) = 109.2$$

$$\begin{aligned} y_{kt} = & \underset{(2.65)}{-1.618} + \underset{(2.95)}{0.475} \cdot y_{kt-1} - \underset{(2.19)}{1.421} \cdot \ln(\mathrm{RW}_a)_{t-2} - \underset{(7.53)}{0.232} \cdot \mathrm{DumFireH} \\ & - \underset{(1.37)}{0.774} \cdot \Delta \ln(\mathrm{RW}_a)_{t-1} + u_t \end{aligned} \tag{6}$$

$$\text{and } u_t = \underset{(2.63)}{0.596} \cdot u_{t-1} + \varepsilon_t$$

$$R^2 = 0.981,\ \mathrm{SE} = 0.0366,\ \mathrm{DW} = 1.66,\ \chi^2(6) = 226.2.$$

The forecast errors for the period 1985–90 are high relative to the standard error of the equation, indicating parameter instability.[34] The forecasts understate the fall in the first birth rate during this recent period. This could reflect an increasing impact of women's relative pay on first birth rates over time, similar to the increasing effect of education that we found in the hazard rate analysis. These results may also reflect the absence from the model of important influences on fertility.

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[33] The t-statistics of the coefficients are given in parentheses. The summary statistics are as follows: R^2 is the coefficient of determination; SE is the standard error of the equation; DW is the Durbin-Watson statistic; χ^2 (6) is a post-sample forecast test statistic (over the period 1985–90) based on forecasts from the best dynamic model estimated over 1964–84, and it is asymptotically distributed as χ^2 variate with 6 degrees of freedrom (its 0.01 critical value is 16.8). The variable ε_t is a white noise random variable.

[34] These forecasts are not fully dynamic: lagged dependent variables are not predicted within the forecast, but pre-forecast residuals continue to exert an influence in the specifications with autoregressive or moving average disturbances. For women aged 20–24, the forecast error varied from three to five times the equation standard error in this period, and for women aged 25–29 from three to nine times.

When the traditional value variable, TRADB, is included in the modelling exercise, the best dynamic equations that emerge for women aged 20–24 and 25–29 during 1964–90 are:

$$\begin{aligned} y_{kt} = {} & \underset{(2.72)}{-1.642} + \underset{(6.85)}{0.777} \cdot y_{kt-1} - \underset{(1.79)}{0.774} \cdot \ln(\mathrm{RW}_a)_{t-2} - \underset{(9.27)}{0.219} \cdot \mathrm{DumFireH} \\ & + \underset{(2.95)}{0.691} \cdot \ln(\mathrm{MW}_{a+5})_{t-2} + \underset{(3.49)}{3.625} \cdot \mathrm{TRADB}_t + u_t \end{aligned} \tag{7}$$

$$\text{and } u_t = \underset{(2.62)}{0.585} \cdot u_{t-1} + \varepsilon_t$$

$$R^2 = 0.993,\ \mathrm{SE} = 0.0321,\ \mathrm{DW} = 1.92,\ \chi^2(6) = 42.8$$

$$y_{kt} = \underset{(5.42)}{-1.778} + \underset{(6.02)}{0.231} \cdot y_{kt-1} + \underset{(4.45)}{4.030} \cdot \mathrm{TRADB}_t - \underset{(6.25)}{0.231} \cdot \mathrm{DumFireH} + u_t \tag{8}$$

$$\text{and } u_t = \underset{(11.96)}{0.827} \cdot u_{t-1} + \varepsilon_t + \underset{(1.87)}{0.627} \cdot \varepsilon_{t-1}$$

$$R^2 = 0.990,\ \mathrm{SE} = 0.0265,\ \mathrm{DW} = 2.01,\ \chi^2(6) = 45.0\ .$$

These equations forecast better during 1985–90 than the earlier equations, but not well enough to accept the hypothesis of parameter stability.[35] They suggest that the fall in first birth rates was "caused" by the decline in traditional values. Given the relatively small effect of the indicators of traditional values in the hazard rate analysis, it is, however, difficult to believe that a decline in traditional values (as measured here) had such a large effect.

It is possible that changes in our measure of traditional values may have been responding to the same underlying factors as first birth rates; thus, their decline may not "explain" the fall in fertility. In other words, there are doubts about the (weak) exogeneity of TRADB_t in the estimation of these equations. We tested for this using a Wu-Hausman test (see Hausman 1978), with two lagged values of ln (MW) and ln (RW) for both age groups and the fire-horse dummy variable as instruments for TRADB. For neither age group could we reject the hypothesis of the exogeneity of TRADB_t: the respective test statistics are $t = -0.28$ for women aged 20–24 and $t = 0.73$ for women aged 25–29. Nevertheless, these tests may be biased if the model is mis-specified, which, in light of the poor forecasting ability of the model, is likely to be the case here.

The case for the exogeneity of women's relative pay is stronger, and we have seen that when we exclude the traditional value variable, it appears to be a signifi-

[35] The ratio of the forecast error to the equation standard error varies from 1 to 4 for women aged 25–29, with the ratio remaining below 1.7 during 1985–87; for women aged 20–24, the ratio consistently remained between 2 and 2.9 for the entire forecast period.

cant determinant of first birth rates.[36] This is consistent with substantial and growing impact of education on the age at motherhood in the hazard rate analysis. Thus, the rise in Japanese women's pay relative to men since the mid-1960s appears to play an important role in the fall in their first birth rates. But it cannot explain fully the recent steep decline in first birth rates, and it is likely that factors other than relative pay are also at work during the period 1964–84.

Conclusions

Our hazard rate analysis shows that Japanese women with higher educational attainment begin childbearing later in their lives, with this impact of education on the age of motherhood increasing for more recent generations. In other words, the higher the educational level, the stronger the trend toward later childbearing. For more recent generations, this relationship between age at motherhood and education comes about entirely because more educated women marry later. Our analysis of aggregate first birth rates to women in their twenties is consistent with these findings. It suggests that higher women's pay delays motherhood while higher men's real wages encourage earlier motherhood.

The econometric analysis of the time series data suggests that while the upward trend in women's pay relative to men's encouraged postponement of childbearing, other factors, including a decline in traditional values, were also working in the same direction. The hazard rate analysis also shows trends toward later marriage in cohorts of women born since 1960 which cannot be accounted for by trends in their educational attainment. The postponement of marriage among these recent generations is fully responsible for their higher age at motherhood, as our hazard rate analysis shows that these recent generations are having their children earlier in marriage.

Thus, both types of analysis support the hypothesis that better women's earning opportunities (as indicated by their educational attainment and relative pay) encourage Japanese women to marry and become mothers later in their lives. But both these analyses indicate that the trend toward later marriage and motherhood in Japan cannot be fully accounted for by improvments in women's education and earning opportunities.

Finally, we can use our hazard model to explore how trends in education may affect future ages of marriage and motherhood. It suggests that, if the impact of education remains the same as that for the 1968 cohort, then the generation born in 1975 will have a median age at marriage of 27 years, 11 months, and their median age at motherhood will be 28 years, 7 months, with 6.4% never marrying and 12.1% remaining childless.[37] If Japanese women go on to achieve the same percentage attaining a university degree as men, while maintaining their present advantage over men in the percent remaining in education beyond high school (43.4 compared with 38.5 in 1993) these median ages are about 6 months higher.

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[36] In both age groups, the year of the "fire horse" produced about a 20% lower first birth rate, while the first birth rates were about 10% higher than would have otherwise been the case in the years preceding and following 1966.

[37] That is, we assume no further trends in the ages of marriage and motherhood other than those caused by the changes in the distribution of women's educational attainment. The Log-Logistic model is used in the calculations.

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