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Association of late childbearing with healthy longevity among the oldest-old in China

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Statistical analysis of a large and unique longitudinal data-set demonstrates that childbearing after age 35 or 40 is associated with survival and healthy survival among very old Chinese women and men. The association is stronger for women than for men. The estimates are adjusted for a variety of confounding factors: demographic characteristics, family support, social connections, health practices, and health conditions. Further analysis based on an extension of the Fixed-Attributes Dynamics method shows that late childbearing is positively associated with long-term survival and healthy survival from ages 80–85 to 90–95 and 100–105. This association exists among oldest-old women and men, but, again, the effects are substantially stronger for women than for men. We discuss four possible factors that may explain why late childbearing affects healthy longevity at advanced ages: (1) social factors; (2) biological changes caused by late pregnancy and delivery; (3) genetic and other biological characteristics; and (4) selection.

Keywords: late childbearing; healthy longevity; China; oldest-old

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Introduction

This paper focuses on the association between late childbearing and healthy longevity. Perls et al. (1997) and Doblhammer (2000) investigate the association between late childbearing and longevity for women. Smith et al. (2003) explore the association for men as well as women, as we do here. In contrast to previous studies, we analyse healthy longevity as well as longevity itself. Furthermore, we use an unprecedentedly large follow-up sample of nearly 9,000 elderly people, aged 80–105, including 2,400 centenarians. This sample is from China, a less developed country where the chance of surviving from birth to advanced old age was very low compared with the cohorts of Americans and Europeans studied by Perls et al. (1997), Doblhammer (2000), and Smith et al. (2003).

This paper contributes to a literature on how events early in life—in the uterus, in childhood, in adult ages before 50 or so—and genetically predetermined characteristics affect health and survival at older ages. Nearly all this literature focuses on ages younger than 80. Notable exceptions are a study by Preston et al. (1998) of African American cohorts born at the beginning of the twentieth century, a study by Snowdon (2001) of elderly nuns, and a study by Doblhammer and Vaupel (2001) of how month of birth influences longevity. Our paper also contributes to the literature on the determinants of healthy longevity, especially the determinants of exceptional longevity and the determinants of health among the oldest-old, i.e., those aged 80 or above (e.g., Christensen and Vaupel 1996; Hitt et al. 1999; Perls et al. 2000; Kerber et al. 2001; Puca et al. 2001).

We apply logistic and survival-analysis models to health data and follow-up mortality data collected from 1998 to 2000 on nearly 9,000 Chinese elderly aged 80 and above. In our analyses, we adjust for a variety of confounding factors. We then use an extended Fixed-Attributes Dynamics (FAD) method and the 1998 baseline survey data to investigate the association of late childbearing with healthy longevity from age 80–85 to 90–95 and 100–105.

Data

Data used in this paper are from the Chinese Longitudinal Survey on Healthy Longevity, which is the largest longitudinal study of very old people ever undertaken in a less developed country; it is also the

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largest longitudinal study of centenarians, nonagenarians, and octogenarians conducted anywhere. The baseline survey and the first follow-up survey were conducted in 1998 and 2000. They were carried out in 631 and 777 counties/cities in 1998 and 2000, respectively, randomly selected from about half of the counties and cities in 22 Chinese provinces. The increase in numbers of survey units between the two surveys was partly because some towns that belonged to a county when the baseline survey was conducted had become cities by the time of the later wave, and partly because some selected counties/cities had no centenarians in 1998 but did have some by 2000 (see discussion below on centenarians). The 22 surveyed provinces are Liaoning, Jilin, Heilongjiang, Hebei, Beijing, Tianjing, Shanxi, Shaanxi, Shanghai, Jiangsu, Zhejiang, Anhui, Fujian, Jiangxi, Shangdong, Henan, Hubei, Hunan, Guangdong, Guangxi, Sichuan, Chongqing (there are 31 provinces in total in China). These provinces have a population of 985 million, constituting 85 per cent of the total population of China.

For each wave of the survey, an interview and a basic health examination were conducted at the interviewee's home. The topics on which data were collected include the following: family structure; living arrangements; number, age, and proximity of children; activities of daily living (ADL); self-rated health; self-evaluation of life satisfaction; cognitive function; medical care; social and religious activities; diet; smoking and alcohol consumption; psychological characteristics; economic resources; caregivers and family support.

The survey team tried to interview all centenarians who had voluntarily agreed to participate in the study. For each centenarian, the team also tried to interview one nearby octogenarian (aged 80–89) and one nearby nonagenarian (aged 90–99) of a predesignated age and sex. The term 'nearby' is loosely defined to encompass someone living in the same village or the same street or, if applicable, the same town, county, or city. Predefinition of the age and sex of interviewees was used to ensure that approximately equal numbers of male and female nonagenarians and octogenarians were randomly selected, using the code numbers of the centenarians. The aim was to have comparable numbers of more or less randomly selected male and female octogenarians and nonagenarians at each age from 80 to 99 (see endnote 4 in Zeng et al. 2002 for details). The total size of the sample used for the baseline survey was 8,805 elders aged 80–105, including 458 male and 1,806 female centenarians, 1,298 male and 1,714 female nonagenarians, and 1,787 male and 1,741 female octogenarians.

Of the participants in the 1998 baseline survey, 4,691 (53.3 per cent) were alive and re-interviewed in the follow-up survey in 2000. A total of 3,264 (37.1 per cent) had died and their date and cause of death and health status before dying were collected by interviewing one of their close family members. Survival time for decedents or censoring time for survivors was calculated in days between the date of the baseline interview and the date of death or second interview. In the follow-up survey we were unable to obtain details of 850 members (9.6 per cent) of the baseline sample (usually owing to change of address) and they were excluded from the study.

Coale and Li (1991) concluded that age reporting for very old persons in these provinces of China, where Han Chinese constitute the majority, is as reliable as it is in developed countries. But in other regions (such as Xingjiang), where the majority or a significant proportion of the population belongs to ethnic groups other than the Han, age reporting may be inaccurate. This is the main reason why we restrict our survey to the 22 provinces where Han Chinese are the overwhelming majority. The other nine provinces (Xinjiang, Qinghai, Ningxia, Inner Mongolia, Tibet, Gansu, Yunnan, Guizhou, and Hainan) all have a high proportion of inhabitants belonging to ethnic minorities. In the case of Xinjiang, we are sure that the quality of age reporting is very poor, but the other eight were excluded because we were not sure about the quality of age reporting among their ethnic groups and we did not have the detailed age-specific and ethnicity-specific data required to evaluate its quality.

The results of a study of the validity of the age reporting of Chinese Han centenarians using a comparison of demographic indices with Sweden, Japan, France, and Italy were consistent with Coale and Li's findings (Wang et al. 1998). The age distribution of the centenarians interviewed in our 1998 survey was found to be very similar to that of Swedish centenarians (see Figure 1 in Zeng et al. 2001). These findings lead us to believe that age reporting in the Chinese Healthy Longevity Survey is generally acceptable up to age 105. A careful evaluation (based on reliability coefficients, factor analysis, the rates of logically inconsistent answers, etc.) showed that the data quality of the Chinese survey on healthy longevity was generally good (Zeng et al. 2001, 2002). The reliability coefficients of the ten categories of variables were found to be reasonable (e.g., see Table 3 in Zeng et al. 2001). For example, the ADL reliability coefficients were 0.88 and 0.87 in the 1998 and 2000 surveys, respectively, figures comparable to the coefficients of 0.87 obtained in the Duke Older Americans Resources and Services Program survey (Fillenbaum 1988) and 0.89 in the Canadian 1991–92 elderly survey (Penning and Strain 1994). A factor analysis demonstrated that interviewees' answers to questions within the same category but concerning different issues were generally consistent. The rates of logically inconsistent answers and 'missing' data were found to be reasonably low (see Appendix B in Zeng et al. 2001).

Measures

The frequency distributions of the covariates used in this paper and based on the 1998 baseline survey are listed in Table 1. These variables were chosen because of their potential association with survival or health status. Following similar studies (e.g., Landerman and Fillenbaum 1997; Palmore and Burchett 1997; Smith and Kington 1997; Strawbridge et al. 1997; Clark et al. 1998; Koenig et al. 1999), all variables were dichotomized, except age, number of births after age 35 or 40, age at first marriage, and number of surviving children. The variables are described below. More detailed descriptions are given in the Appendix.

Late childbearing was measured by the number of births after age 35 and age 40.
Demographic variables include age (80–89, 90–99, 100–105), sex (male, female), ethnicity (Han, minorities), education (0 years of schooling, ≥1 year schooling), age at first marriage, and residence (rural, urban).
Physical and mental health variables include ADL, cognitive function measured by the Mini Mental Status Exam (MMSE), self-rated health, and depression symptoms.
Family support includes marital status, number of surviving children at the time of the 1998 survey, proximity to children, and attendance at religious events.
The social connection/support index was based on questions about playing cards or mah-jong regularly, watching TV or listening to the radio regularly, retirement wages, adequate medical service, and caregivers.
Health practice variables include cigarette smoking, alcohol consumption, diet, and exercise.

Since the percentage of missing values for all baseline variables (excluding 'not able to answer') used in our models is less than 2 per cent except for age at first marriage (5 per cent), we imputed the mean value for missing values of all variables (Landerman et al. 1997) except age at first marriage for which multiple imputation was used (Allison 2002).

Methods

Statistical modelling

We performed three kinds of statistical analysis. First, we used logistic regression to analyse how late childbearing is associated with health status. In such models, the dependent variables are ADL scores, MMSE scores, self-rated health, and depression symptoms. The independent variables are late childbearing after age 35 or 40, plus other covariates measuring demographic characteristics, family support, social connections, and health practice (to be described later).

Second, we used a Cox proportional hazards regression model (Cox 1972) to examine the association between late childbearing and survival, controlling for confounding factors. The proportionality hypothesis was tested graphically and with the normal score test (Grambsch and Therneau 1994), and confirmed. Survival time from the interview in 1998 to death or re-interview in 2000 was measured in days.

Third, we used ordinal logistic regression (Harrell 2001, p. 331; Meurer et al. 2001) to investigate the association of late childbearing with healthy survival, adjusting for confounding factors (the proportional odds assumption required for ordinal logistic regression was tested and met for both males and females). The dependent variable was defined as follows. The individuals who survived with good health in the second wave (2000) were coded as 1; those who survived with poor health were coded as 2; those who died between the two waves (1998–2000) were coded as 3. 'Good health' is defined as no impairment in ADL, normal cognitive function, self-reported good health, and no depression symptoms (measurement issues will be discussed later). Persons who are not in 'good health' are considered to be in 'poor health'.

We employed a sequential modelling strategy (e.g., Strawbridge et al. 1997; Koenig et al. 1999) in our Cox survival analysis and ordinal logistic regression analysis. In our sequential models, Model I adjusts for demographic variables of gender, age, rural/urban residence, education, and ethnicity. Model II further adjusts for family support and social connection variables pertaining to the number of surviving children, marital status, living arrangement, proximity to children, religious attendance, social contacts, retirement wage, adequate medical care, and caregiver, in addition to the demographic variables. Model III adds health practice variables pertaining to smoking, alcohol drinking, diet, and exercise. Model IV further controls for health status as measured by ADL

Table 1 Frequency distributions of variables used for respondents aged 80+ in the 1998 baseline survey used for studying the association between late childbearing and healthy longevity in China

Variables	Number		%	
	F	M	F	M
Late childbearing variables				
Birth after age 35				
0	2,434	1,033	46.3	29.2
1	1,256	688	23.9	19.4
2	849	651	16.1	18.4
3+	723	1,171	13.7	33.1
Birth after age 40				
0	3,687	1,634	70.1	46.1
1	988	792	18.8	22.4
2	398	527	7.6	14.9
3+	189	590	3.6	16.7
Demographic variables				
Age				
80–89	1,741	1,787	33.1	50.4
90–99	1,715	1,298	32.6	36.6
100–105	1,806	458	34.3	13.0
Residence				
Urban	1,912	1,418	36.3	40.0
Rural	3,350	2,125	63.7	60.0
Education				
0 year schooling	4,562	1,310	86.7	37.0
≥1 year schooling	700	2,233	13.3	63.0
Ethnicity				
Han	4,878	3,307	92.7	93.3
Minority	384	236	7.3	6.7
Age at first marriage				
≤18	2,772	891	52.7	25.1
19–20	1,351	691	25.7	19.5
21–24	866	884	16.5	24.9
≥25	273	1,077	5.2	30.4
Family support and social connection				
Married				
No	5,000	2,373	95.0	67.0
Yes	262	1,170	5.0	33.0
Living alone				
No	4,702	3,216	89.4	90.8
Yes	560	327	10.6	9.2

Variables	Number		%	
	F	M	F	M
Family support and social connection (continued)				
Proximity to children				
Low	1,310	813	24.9	22.9
High	3,952	2,730	75.1	77.1
Social connection index				
Low	4,031	1,844	76.6	52.0
High	1,231	1,699	23.4	48.0
Number of surviving children				
0	783	428	14.9	12.1
1–2	1,845	908	35.1	25.6
3–4	1,667	1,148	31.7	32.4
5+	967	1,059	18.4	29.9
Religious activities				
No	4,220	3,144	80.2	88.7
Yes	1,042	399	19.8	11.3
Health practice				
Currently smoking				
No	4,876	2,435	92.7	68.7
Yes	386	1,108	7.3	31.3
Strong alcohol drinker				
No	5,087	3,060	96.7	86.4
Yes	175	483	3.3	13.6
Healthy diet				
No	4,121	2,564	78.3	72.4
Yes	1,141	979	21.7	27.6
Currently exercising				
No	4,245	2,176	80.7	61:4
Yes	1,017	1,367	19.3	38.6
Physical and mental health				
ADLs impaired				
No	2,982	2,581	56.7	72.8
Yes	2,280	962	43.3	27.2
Cognitive function impaired				
No	2,295	2,497	43.6	70.5
Yes	2,967	1,046	56.4	29.5
Self-reported health				
Good/very good	2,615	2,082	49.7	58.8
Fair/poor	2,647	1,461	50.3	41.2
Depression symptoms				
No	4,437	3,279	84.3	92.5
Yes	825	264	15.7	7.5

Notes: F = females; M = males.
Source: Longitudinal Healthy Longevity Survey, China: baseline survey 1998.

limitations, MMSE score, self-reported health, and depression symptoms. All of the values of the covariates of Models I, II, III, and IV in the analyses presented in Tables 3 and 4 are obtained from the 1998 baseline survey. The sequential models allow us to examine how the estimated association of late childbearing with survival and healthy survival is affected when confounding factors are added to the model step by step. The likelihood ratio tests show that the goodness of fit of our models is improved significantly as additional factors are added (see Table 3).

Fixed-Attributes Dynamics (FAD) method

The models and data described above can be used to explore the effects of late childbearing on survival and healthy survival from age x to $x + 2$. Because the longitudinal data are available for only 2 years, these analyses cannot be used to investigate the association of late childbearing with healthy survival over long intervals. Hence, we apply a FAD method (described below) to study the association of late childbearing with healthy longevity from ages 80–85 to 90–95 and 100–105.

The idea of the FAD method, also known as the Survival Attribute Assay, was initially proposed by Vaupel (1992, formulas 9.12 and 9.17) and developed by him and Anatoli Yashin in a series of papers (Yashin et al. 1998, 1999, 2000; Gerdes et al. 2000). The method is based on the fundamental demographic insight that the prevalence of a fixed attribute in a population can change with age even though no individual can change his or her variant of the attribute, and that therefore much can be learned about the impact of the attribute on survival. Fixed attributes that affect survival at older ages include genetic factors and various non-genetic factors that are fixed in earlier life. Such factors could be those fixed at birth (e.g., genetic make-up, birth weight, and birth order) but any factor that is fixed before the onset of old age (e.g., diseases suffered earlier in life, education, occupation, and pattern of childbearing) can also be used in a study of survival at older ages.

The basic FAD method for use in a study of longevity can be formulated as follows. Let $N(x)$ denote the number of persons aged x; $p_1(x)$ the proportion of individuals who are x years old and have the fixed attribute; $s_1(x + n)$ the conditional survival probability from age x to $x + n$ for those who have the fixed attribute; $s_0(x + n)$ the conditional survival probability from age x to $x + n$ for those who do not have the fixed attribute; and $S(x + n)$ the conditional survival probability from age x to $x + n$ for all members of the cohort.

On the one hand, because $N(x)p_1(x)s_1(x + n) = N(x)S(x + n)p_1(x + n)$, it follows that

$$s_1(x+n) = S(x+n)\left(\frac{p_1(x+n)}{p_1(x)}\right). \tag{1}$$

On the other hand, it similarly follows that

$$s_0(x+n) = S(x+n)\left(\frac{1-p_1(x+n)}{1-p_1(x)}\right). \tag{2}$$

Dividing (1) by (2) gives the ratio of survivorship (RS) of those with the fixed attribute to those without the attribute:

$$RS = \frac{s_1(x+n)}{s_0(x+n)} = \frac{(1-p_1(x))p_1(x+n)}{p_1(x)(1-p_1(x+n))}. \tag{3}$$

If RS is greater than one, the attribute is positively associated with longevity. Statistical tests based on the method developed by Mantel and Haenszel (1959) can be performed to test whether there is a statistically significant difference of survivorship between those with the fixed attribute and those without it (i.e., to test the null hypothesis that the ratio of survivorship equals one).

The research by Vaupel, Yashin, and colleagues using the FAD method is based on the use of formulas (1) and (2) to derive estimates of the risk of mortality for people with some attribute relative to the risk for those without the attribute (Yashin et al. 1998, 1999, 2000; Gerdes et al. 2000). In this paper, however, we propose and emphasize the use of formula (3). Formula (3) does not require estimates of cohort survivorship ($S(x)$), which were not available for this study and may not be available for many other studies, especially those applied to less developed countries.

We extend the FAD method expressed in formula (3) to study the association between a fixed attribute and the combined status of survival and good health. Good health is defined by some measure, such as 'ADL independence', the ability to perform the basic activities of daily living independently. Let $l_1(x + n)$ denote the probability that individuals who are x years old and who have some fixed attribute will survive to age $x + n$ and be in good health at age $x + n$. Note that these individuals do not have to be in good health at age x. Similarly, define $l_0(x + n)$ as the probability that individuals who are x years old and who do not have the fixed attribute will survive to age $x + n$ and be in good health at age $x + n$. Let $h(x + n)$ be the proportion of persons who are in good health at age $x + n$. Finally, let $\pi(x + n)$ be the proportion of persons having the fixed attribute among those who are healthy at age $x + n$.

On the one hand, because $N(x)p_1(x)l_1(x + n) = N(x)S(x + n)h(x + n)\pi(x + n)$, it follows that

$$l_1(x+n) = \frac{S(x+n)h(x+n)\pi(x+n)}{p_1(x)} \tag{4}$$

where $S(x)$ and $p_1(x)$ are defined as before. On the other hand, it is also the case that

$$l_0(x+n) = \frac{S(x+n)h(x+n)(1-\pi(x+n))}{(1-p_1(x))}. \tag{5}$$

The ratio of the probability of healthy survival of those with the fixed attribute to the probability for cohort members without it, which we call the ratio of healthy survivorship (*RHS*), is estimated by dividing formula (4) by formula (5):

$$RHS = \frac{l_1(x+n)}{l_0(x+n)} = \frac{(1-p_1(x))\pi(x+n)}{p_1(x)(1-\pi(x+n))}. \qquad (6)$$

If *RHS* is greater than one, the fixed attribute is positively associated with healthy longevity.

Formula (6) is an extension of formula (3). Note that formula (3) can be expressed in terms of the number of people with and without some attribute at two ages:

$$RS = \frac{N_0(x)N_1(x+n)}{N_1(x)N_0(x+n)}.$$

This is a formula used by Hill (1999). So the method expressed in formula (6) can be considered to be an extension both of the FAD method and of Hill's method.

One of the main strengths of the FAD method is that it can be used to analyse the association between fixed attributes and longevity or healthy longevity based on two or more independent cross-sectional samples, rather than on a longitudinal survey. The ratios of survivorship and healthy survivorship using the FAD method can also be estimated using one cross-sectional dataset. This can be done if it is reasonable to assume that the initial distribution of the fixed attribute and its effects on survival do not differ substantially between older and younger cohorts. In this case, the older and younger cohorts are connected in one hypothetical cohort. As will be discussed later, we assume that the initial distributions of late childbearing—and the effect of this fixed attribute on healthy survival—did not differ significantly among the cohorts aged 80–105 in 1998.

The FAD method assumes that the number of migrants is small or, alternatively, that migrants do not differ significantly from non-migrants with respect to the fixed attribute and survival. The magnitude of international and internal migration before the 1980s in China was small owing to slow economic development and policies restricting residential movement. Mobility greatly increased after 1980, but not among the oldest-old.

A major limitation of the standard approaches to survival analysis is the need for prospective tracking of study participants to ascertain the occurrence of exit events. If the transition of interest requires many years to occur, the required follow-up period often becomes prohibitively expensive and time-consuming (Hill 1999, pp. 497–8). The main strength of the FAD method is that it circumvents this major limitation. In the present study, one of the objectives is to investigate the association of late childbearing with healthy longevity from age 80–85 to 90–95 and 100–105. To do so, and using the standard approach of survival analysis, we would need 10–20 years of follow-up data, which are not available. Furthermore, almost all of the oldest-old aged 80–105 would die within the 10–20 year follow-up period even if such a study were practicable in other respects. This is the main reason why we use the FAD method based on the cross-sectional baseline survey data, in addition to the survival analysis using the baseline and the 2-year follow-up data.

The strength of the FAD method, however, is bought at a price because it has a major limitation. In addition to its crucial assumptions, discussed earlier, the FAD method, as currently developed, permits a univariate analysis to ascertain the 'de facto' association between the fixed attribute and healthy survival, while controlling for age and sex. But an analysis based on the FAD method does not ascertain whether the association of the fixed attribute with healthy survival is due to the attribute itself or is caused or mediated by other factors. That is why we also use logistic analysis and hazards models to investigate the association of late childbearing with health status at the baseline and with survival and healthy survival during the 1998–2000 follow-up period, while controlling for various confounding factors.

Findings

Results based on the statistical models and the 1998–2000 survey data

The association of late childbearing with health status in the1998 baseline survey. Table 2 shows the estimates of the odds ratios (OR) of being ADL-disabled, MMSE-abnormal, having self-reported poor health, and having depression symptoms (we are satisfied with the goodness of fit of the logistic regression models presented in Table 2, based on the χ^2-tests suggested by Hosmer and Lemeshow 1989). These estimates allow comparison of health status in the 1998 baseline survey between the oldest-old who had 1, 2, or 3+ births after age 35 or 40 with those who did not have late births. The estimates of the OR are adjusted for the other 15 confounding factors of demographic characteristics, family support, social connections, and health practice listed in Table 1. Of

Table 2 Effects (odds ratios) of late childbearing on health status of the oldest-old in China, based on multivariate logistic regression applied to data collected at 1998 baseline survey, adjusted for covariates of demographic characteristics, family support, social connections, and health practice

	ADL disabled	MMSE impaired	Self-reported bad health	Depression symptoms
Birth after age 35				
Women (0 birth after age 35)				
1 birth after age 35	0.890	0.949	0.923	0.854
2 births after age 35	0.939	0.835#	0.906	0.827
3+ births after age 35	0.765*	0.772*	1.031	0.626**
Hosmer–Lemeshow χ^2, df, and significance level of the model (goodness-of-fit test)	10.5 (df = 8) (p = 0.235)	9.4 (df = 8) (p = 0.307)	10.1 (df = 8) (p = 0.256)	12.8 (df = 8) (p = 0.118)
Men (0 birth after age 35)				
1 birth after age 35	0.954	1.111	1.002	0.846
2 births after age 35	0.879	0.937	0.858	0.757
3+ births after age 35	0.880	0.791	0.939	0.711
Hosmer–Lemeshow χ^2, df, and significance level of the model (goodness-of-fit test)	11.2 (df = 8) (p = 0.193)	6.8 (df = 8) (p = 0.562)	12.6 (df = 8) (p = 0.126)	14.8 (df = 8) (p = 0.063)
Birth after age 40				
Women (0 birth after age 40)				
1 birth after age 40	0.909	0.890	1.022	0.810#
2 births after age 40	0.987	0.979	1.153	0.572**
3+ births after age 40	0.481***	0.765	0.837	0.947
Hosmer–Lemeshow χ^2, df, and significance level of the model (goodness-of-fit test)	9.9 (df = 8) (p = 0.269)	9.1 (df = 8) (p = 0.333)	6.2 (df = 8) (p = 0.622)	8.5 (df = 8) (p = 0.390)
Men (0 birth after age 40)				
1 birth after age 40	0.952	0.978	0.884	0.616*
2 births after age 40	0.707*	0.865	0.878	0.812
3+ births after age 40	0.803	0.828	0.936	0.562*
Hosmer–Lemeshow χ^2, df, and significance level of the model (goodness-of-fit test)	8.9 (df = 8) (p = 0.348)	9.4 (df = 8) (p = 0.311)	10.4 (df = 8) (p = 0.241)	3.0 (df = 8) (p = 0.937)

Notes: (1) The category in parentheses is the reference group in each case. (2) The goodness-of-fit tests were performed following the approach of Hosmer and Lemeshow (1989). (3) df = degree of freedom. (4) #$p < 0.10$; *$p < 0.05$; **$p < 0.01$; ***$p \leq 0.001$.
Source: As for Table 1.

the 48 estimates of the OR of being in poor health presented in Table 2, 43 are less than one and 5 are greater than one but not statistically significant. Among the OR that are less than one, ten are statistically significant at different levels. For instance, the OR of being ADL-disabled for very old women who had 3+ births after age 35 or 40 is 0.77 or 0.48; these estimates are statistically significant at a level of $p < 0.05$ or $p < 0.001$. This means that, controlling for the other 15 confounding factors of demographic characteristics, family support, social connections, and health practice, the likelihood of being ADL-disabled for the women who had 3+ births after age 35 or age 40 is 23 or 52 per cent lower than that of those who did not have such late births. While controlling for the confounding factors, the oldest-old women and men who had 1, 2, or 3+ births after age 35 or 40 tend to be somewhat less likely to be ADL-disabled, MMSE-abnormal, to have self-reported poor health, and to have depression symptoms. The general pattern is clear: late childbearing may be associated with better health at advanced ages; and in most cases, the larger the number of births after age 35 or 40, the stronger the association. Most of the estimates are not statistically significant, however. Thus, we must be cautious in drawing conclusions about the strength of the association between late childbearing and health status among the oldest-old.

The association of late childbearing with survival. The second and third columns in Table 3 present the relative mortality risk (RMR) between 1998 and 2000 for the oldest-old who had 1, 2, or 3+ births after age 35 or 40 vs. those who did not have any late births. In all Models (I, II, III, and IV), the mortality risk of very old females who had 1 birth after age 35 or age 40 did not differ significantly from those who did not have such late births. Adjusted for age, residence, education, age at first marriage, and ethnicity (Model I), the mortality risk of women who had 3+ births after age 35 or had 2 or 3+ births after age 40 was reduced by 28, 21, or 43 per cent, respectively, of those who did not have any late births. These estimates are statistically significant. After the covariates of family support, social connection, and health practice were added to the model (II and III), the impact of late childbearing in reducing the relative mortality risk of women remains almost the same as that estimated in Model I, which adjusted for demographic variables only.

Model IV, which further adjusts health status, shows that the mortality risk of oldest-old women who had 3+ births after age 35 or had 2 or 3+ births after age 40 remains significantly lower than that of those who had no late births. More specifically, controlling for the other 19 confounding factors—including ADL, MMSE, self-reported health, depression symptoms, health practice, family support, social connections, and demographic characteristics—the mortality risk of very old females who had 2 or 3+ births after age 40 is 19 and 38 per cent lower than that of those who had no late births. The mortality risk of women who had 3+ births after age 35 is 23 per cent lower than that of those who had no late births.

Remarkably enough, we also found late childbearing to be associated with lower mortality risk among very old males. Controlling for the demographic characteristics of age, residence, education, age at first marriage, and ethnicity, having 2 or 3+ offspring after age 40 reduces the 2-year period mortality risk of the oldest-old men by 20 or 33 per cent. (By men's late birth(s), i.e., births after age 35 or 40, we mean of course that men's wives had given birth after the men had reached these ages.) After the covariates of family support (including the number of surviving children), social connections, and health practice are added to the model, having 2 or 3+ offspring after age 40 still reduces the 2-year period mortality risk of the oldest-old men by 20–26 per cent. Even after ADL, MMSE, self-reported health, and depression symptoms are added to Model IV, in addition to demographic characteristics, family support, social connections, and health practice as control variables, very old males who had 2 or 3+ offspring after age 40 still had a lower risk of death of 16–22 per cent. These estimates are all statistically significant (see Table 3).

The estimates show that the addition of the covariates of family support (including the number of surviving children) and social connections as control variables to Model II slightly weakens the association between late childbearing and survival among the oldest-old men, compared with Model I. Additions of the covariates of health practice and health conditions to the models for males further weaken the association. The survival analysis also demonstrates that for all Models I, II, III, and IV, the larger the number of births after age 35 or 40, the stronger the association between late childbearing and survival among both male and female oldest-old (see Table 3).

The association of late childbearing with healthy survival. The last two columns of Table 3 present the ordinal logistic regression estimates of the association of late childbearing with healthy survival, sequentially adjusted for the other 19 confounding factors of demographic characteristics, family support, social connections, health practice, and health conditions observed in the 1998 baseline survey. As described earlier, healthy survival is defined as survival plus good health in the second wave of the survey in 2000. 'Good health' is measured by 'no impairment' in all of the four main aspects of health—active in daily living, normal cognitive function, self-reported good health, and no depression symptoms.

The ordinal logistic regression analysis shows that the OR of death or unhealthy survival among oldest-old women who had 2 or 3+ births after age 35 or 40 were 15–44 per cent lower in Models I, II, III, and IV than amongst women who did not have such late births. These estimates for females are all statistically significant, with one exception. The risk of death or unhealthy survival among very old males was found to be statistically significant among those who had 3+ births after age 35 or had 2 or 3+ births after age 40 in Model I compared with those who did not have late births. The estimates also show that the association of having 2 or 3+ births after age 35 or 40 and

Table 3 Effects of late childbearing (relative risks and odds ratios) on survival and healthy survival of the oldest-old in China between 1998 and 2000 based on multivariate Cox proportional hazards and ordinal logistic regression models

Models	Relative risk of Cox hazards models survival analysis		Odds ratios of ordinal logistic regression healthy survival analysis	
	Women	Men	Women	Men
Gave birth after age 35				
Model I (0 birth after age 35)				
1 birth after age 35	0.941	0.910	0.914	1.103
2 births after age 35	0.850*	1.077	0.758***	0.930
3+ births after age 35	0.722***	0.747***	0.697***	0.769**
Model II (0 birth after age 35)				
1 birth after age 35	0.941	0.919	0.977	1.209#
2 births after age 35	0.858*	1.100	0.847#	1.045
3+ births after age 35	0.715***	0.800*	0.766*	0.895
Model III (0 birth after age 35)				
1 birth after age 35	0.949	0.919	0.967	1.157
2 births after age 35	0.861*	1.116	0.828*	1.066
3+ births after age 35	0.720***	0.821#	0.747**	0.889
Model IV (0 birth after age 35)				
1 birth after age 35	0.977	0.947	1.003	1.169
2 births after age 35	0.902	1.181	0.854	1.111
3+ births after age 35	0.767**	0.889	0.791*	0.944
–2 log Likelihood (–2LL), Model I	32,138.0***	18,981.1***	8,577.6***	6,360.1***
Improvement of –2LL, Model II vs. Model I	79.5***	47.6***	96.8***	50.3***
Improvement of –2LL, Model III vs. Model II	65.9***	71.9***	70.4***	98.5***
Improvement of –2LL, Model IV vs. Model III	231.4***	233.4***	286.1***	304.9***
Gave birth after age 40				
Model I (0 birth after age 40)				
1 birth after age 40	0.925	0.908	0.945	0.978
2 births after age 40	0.785**	0.798**	0.739**	0.762**
3+ births after age 40	0.569***	0.673***	0.560***	0.719**
Model II (0 birth after age 40)				
1 birth after age 40	0.947	0.912	1.018	1.043
2 births after age 40	0.791**	0.792*	0.812#	0.818#
3+ births after age 40	0.568***	0.717**	0.584*	0.817#
Model III (0 birth after age 40)				
1 birth after age 40	0.942	0.935	1.000	1.076
2 births after age 40	0.792*	0.798*	0.810#	0.810#
3+ births after age 40	0.565***	0.744**	0.568*	0.842
Model IV (0 birth after age 40)				
1 birth after age 40	0.962	0.979	1.019	1.131
2 births after age 40	0.811*	0.842#	0.792*	0.850
3+ births after age 40	0.620***	0.782*	0.627**	0.873
–2 log Likelihood (–2LL), Model I	32,135.0***	18,981.2***	8,578.2***	6,359.1***
Improvement of –2LL, Model II vs. Model I	82.6***	46.2***	104.6***	48.3***
Improvement of –2LL, Model III vs. Model II	66.5***	71.0***	70.6***	98.6***
Improvement of –2LL, Model IV vs. Model III	230.3***	231.9***	285.0***	305.6***

Notes: (1) The category in parentheses is the reference group in each case. (2) Covariates in Model I are late childbearing plus demographic variables of age, gender, residence, education, and ethnicity. Model II is Model I plus covariates of family support and social connection. Model III is Model II plus covariates of health practices. Model IV is Model III plus covariates of health conditions. (3) The number of degrees of freedom for Models I–IV is 9, 17, 21, 25, respectively. (4) #$p < 0.10$; *$p < 0.05$; **$p < 0.01$; ***$p < 0.001$.
Source: Longitudinal Healthy Longevity Survey, China: baseline survey 1998 and follow-up survey 2000.

healthy survival is much stronger among the oldest-old women than the oldest-old men.

A comparison of estimates between Model II and Model I shows that the addition of the covariates of family support (including the number of surviving children) and social connections as control variables to the model moderately weakens the association between late childbearing and healthy survival. The addition of the covariates of health conditions to the models further weakens the association. These observations hold for both oldest-old women and men.

Results based on the 1998 baseline survey data and the FAD method

The 1998 survey shows that the fertility level among the three cohorts, aged 100–105, 90–95, and 80–85 in 1998, was similar. We also believe that the impact of giving birth at later ages on longevity and healthy longevity did not change significantly across the cohorts of our subjects. Thus, it seems reasonable to investigate the association between late childbearing and healthy longevity using the FAD method based solely on the 1998 survey data. Because the FAD method is a univariate method of analysis, when we compare those with and those without the fixed attribute, we compare: those who had one or more (1+) late births with those who had no late birth; those who had two or more (2+) late births with those who had fewer than two or no late birth; and those who had three or more (3+) late births with those who had fewer than three or no late birth. Estimates of the ratio of survivorship show that women who had 1+, 2+, 3+ births after age 35 would increase the likelihood of survival from age 80–85 to 100–105 by 44, 59, and 127 per cent, respectively, compared with those women who did not have such late births. The increase in the likelihood of survival from age 90–95 to 100–105 due to having 1+, 2+, 3+ births after age 35 is 32–53 per cent. These estimates are highly significant. The impact of having 1+, 2+, 3+ births after age 40 on survival from age 80–85 to 100–105 and from 90–95 to 100–105 is even more dramatic—the increased likelihood of survival ranges from 102 to 322 per cent and from 41 to 120 per cent, respectively (see Table 4a).

The estimates also indicate that later childbearing

Table 4a Ratio of survivorship (*RS*) of elders who had one or more births after age 35 or 40 to those who did not have such late births

	Ages 100–105 vs. 80–85				Ages 100–105 vs. 90–95				Ages 90–95 vs. 80–85			
$P_1(x)$	$P_1(80–85)$	$P_1(100–105)$	*RS*	*p*	$P_1(90–95)$	$P_1(100–105)$	*RS*	*p*	$P_1(80–85)$	$P_1(90–95)$	*RS*	*p*
Birth(s) after age 35												
Men												
1+ births	69.0	71.0	1.10	0.449	72.3	71.0	0.93	0.596	69.0	72.3	1.17	0.106
2+ births	50.1	54.4	1.19	0.125	49.9	54.4	1.20	0.121	50.1	49.9	0.99	0.913
3+ births	30.2	35.6	1.28	0.036	32.7	35.6	1.14	0.281	30.2	32.7	1.12	0.233
Women												
1+ births	49.8	58.8	1.44	0.000	51.3	58.8	1.35	0.000	49.8	51.3	1.06	0.489
2+ births	24.6	34.2	1.59	0.000	28.2	34.2	1.32	0.001	24.6	28.2	1.20	0.059
3+ births	8.9	18.1	2.27	0.000	12.6	18.1	1.53	0.000	8.9	12.6	1.48	0.005
Birth(s) after age 40												
Men												
1+ births	48.9	57.9	1.43	0.001	57.6	57.9	1.01	0.926	48.9	57.6	1.42	0.002
2+ births	25.2	37.6	1.79	0.000	33.7	37.6	1.18	0.158	25.2	33.7	1.51	0.000
3+ births	10.8	21.2	2.21	0.000	17.1	21.2	1.30	0.070	10.8	17.1	1.70	0.000
Women												
1+ births	21.9	36.2	2.02	0.000	28.7	36.2	1.41	0.000	21.9	28.7	1.43	0.000
2+ births	5.6	15.3	3.02	0.000	10.8	15.3	1.49	0.000	5.6	10.8	2.03	0.000
3+ births	1.3	5.4	4.22	0.000	2.5	5.4	2.20	0.000	1.3	2.5	1.91	0.047

Notes: (1) $P_1(x)$ = percentage of elders who gave birth after age 35 or after age 40 among those aged x. (2) p = significance level of χ^2–tests based on the method of Mantel and Haenszel (1959) to test whether there is a statistically significant difference of survivorship between those with the fixed attribute and those without it.
Source: As for Table 1.

is positively associated with the healthy longevity of women. The estimates of the ratio of healthy survivorship presented in Table 4b show that giving birth after age 40 or 35 is positively associated with the likelihood of survival plus ADL independence. Most of the estimates are statistically significant; the association between giving birth after age 40 and healthy longevity is very dramatic and highly significant. The chance of surviving with ADL independence for women who had 1+, 2+, 3+ births after age 40 increased by 133–542 per cent from age 80–85 to 100–105, and increased by 62–235 per cent from age 90–95 to 100–105, compared with women who did not have such late births.

All of the estimates of the ratio of survivorship and ratio of healthy survivorship among the oldest-old men who had one birth or more after age 40 (i.e., the men whose wives had one or more births after the men reached age 40) are greater than one; most estimates are statistically significant. The likelihood of surviving with ADL independence from age 80–85 to 100–105 for men who had 1+, 2+, and 3+ births after age 40 was 58–143 per cent higher than for men who did not have such a late birth. The increases in the likelihood of healthy survival from age 90–95 to 100–105 for the men who had 1+, 2+, and 3+ births after age 40 were 11–43 per cent. The likelihood of healthy survival from age 80–85 to 90–95 among the men who had 1+, 2+, or 3+ births after age 40 were 40–69 per cent higher than for those who did not have such a late birth.

Discussion

A study using logistic regression, controlling for a few major socio-economic factors and based on two large datasets—one from Great Britain the other from Austria—finds a positive relationship between giving birth after age 40 and longevity among women (Doblhammer 2000). These results are consistent with two previous studies (Voland and Engel 1986; Perls et al. 1997). Perls et al. (1997) found that among 92 centenarians born in 1896, 19 per cent had borne children after age 40, whereas only 6 per cent had done so among a control group of 49 women born in the same year who had died at age 73. One possible explanation is that factors conducive to bearing a

Table 4b Ratio of healthy survivorship (*RHS*) of elders who had one or more births after age 35 or after age 40 to those who did not have late births

	Ages 100–105 vs. 80–85				Ages 100–105 vs. 90–95				Ages 90–95 vs. 80–85			
$P_1(x)$, $\pi(x+n)$	$P_1(80–85)$	$\pi(100–105)$	*RHS*	*p*	$P_1(90–95)$	$\pi(100–105)$	*RHS*	*p*	$P_1(80–85)$	$\pi(90–95)$	*RHS*	*p*
Birth(s) after age 35												
Men												
1+ births	69.0	71.4	1.12	0.506	72.3	71.4	0.95	0.778	69.0	72.3	1.17	0.151
2+ births	50.1	56.3	1.28	0.103	49.9	56.3	1.29	0.097	50.1	49.2	0.96	0.709
3+ births	30.2	39.3	1.50	0.010	32.7	39.3	1.34	0.069	30.2	31.5	1.07	0.556
Women												
1+ births	49.8	60.5	1.54	0.000	51.3	60.5	1.45	0.000	49.8	51.9	1.09	0.891
2+ births	24.6	36.7	1.78	0.000	28.2	36.7	1.48	0.000	24.6	29.0	1.25	0.045
3+ births	8.9	20.7	2.68	0.000	12.6	20.7	1.81	0.000	8.9	13.9	1.67	0.001
Birth(s) after age 40												
Men												
1+ births	48.9	60.2	1.58	0.003	57.6	60.2	1.11	0.496	48.9	57.3	1.40	0.001
2+ births	25.2	39.8	1.97	0.000	33.7	39.8	1.30	0.097	25.2	33.6	1.51	0.000
3+ births	10.8	22.8	2.43	0.000	17.1	22.8	1.43	0.057	10.8	17.0	1.69	0.000
Women												
1+ births	21.9	39.5	2.33	0.000	28.7	39.5	1.62	0.000	21.9	29.7	1.50	0.000
2+ births	5.6	18.4	3.77	0.000	10.8	18.4	1.85	0.000	5.6	11.7	2.21	0.000
3+ births	1.3	8.0	6.42	0.000	2.5	8.0	3.35	0.000	1.3	3.4	2.59	0.005

Notes: (1) $P_1(x)$ = percentage of elders who gave birth after age 35 or age 40 among those aged x. (2) $\pi(x + n)$ = percentage of elders who gave birth after age 35 or age 40 among those age $x + n$ and ADL independent. (3) p = significance level of χ^2-tests based on the method of Mantel and Haenszel (1959) to test whether there is a statistically significant difference of healthy survivorship between those with the fixed attribute and those without it.
Source: As for Table 1.

child at a later age might also be conducive to having aged at a slower pace and living longer. It is possible that later menopause is an additional marker.

Snowdon et al. (1989) and Snowdon (1990) show that the mortality risk for women who had their natural menopause before age 40 is nearly twice as high as that of women who experienced menopause at age 50–54. It may be that the positive impact of late menopause results from the extended period of endogenous oestrogen production. A positive relationship between age at natural menopause and breast cancer (Heck and Pamuck 1997), and an increase in the risk of breast cancer due to oestrogen therapies (Law et al. 1996; Paganini-Hill 1996) have, however, also been found. The delay of menopause itself is perhaps unlikely to be the cause of increased longevity (Perls et al. 1997; Perls and Fretts 2001).

In addition to the effects of an extended period of endogenous oestrogen production, later pregnancy, childbirth, and breastfeeding may also stimulate women's biological systems and positively affect survival and health. Furthermore, some social factors might also contribute to the positive relationship between later childbearing and longevity. The contribution of social factors can be investigated by analysing the fertility of men and longevity data. Doblhammer (2000) could not address this issue owing to a lack of information on the fertility of males in her datasets. With the exception of one recent study, none of the previous studies investigated the association between late childbearing and longevity among men (Doblhammer 2000). Based on the Utah genealogical database, Smith et al. (2003) found that late age at last birth is associated with greater post-reproductive longevity among women; husbands of these women experienced weaker longevity benefits associated with late fertility.

The 1998 baseline and 2000 follow-up surveys in China on the determinants of healthy longevity collected data on fertility, health status, and mortality/survival from oldest-old men as well as women. This enables us to explore possible social factors, such as whether the advantage in health longevity enjoyed by those very old people who had births at late ages is partly a consequence of having younger children to take better care of them. Statistical analysis based on the 1998–2000 follow-up data demonstrates that having 3+ births after age 35 or 2 or 3+ births after age 40 is significantly associated with survival and healthy survival during the 2-year period among the oldest-old Chinese women and men. The association of late childbearing with healthy longevity is stronger in oldest-old women than in their male counterparts. All of these statistical model estimates are adjusted for a variety of confounding factors of demographic characteristics, family support, social connections, health practice, and health conditions.

The FAD analysis based on the 1998 baseline survey shows that late childbearing, especially having 3+ births after age 35 or 2 or 3+ births after age 40, is positively associated with long-term survival and healthy survival from ages 80–85 to 90–95, and 100–105. This association exists among oldest-old women and men, but, again, the effects are substantially stronger in women than in men.

An extended period of endogenous oestrogen production, later pregnancy, birth delivery, and breastfeeding may stimulate biological systems and positively affect survival and health for women. But the positive effect of later childbearing on survival and healthy survival for oldest-old fathers cannot be explained by biological mechanisms related to late pregnancy and delivery. Thus the effect on men indicates that various social factors may also be important. One possible explanation is that the oldest-old who had children after age 35 or 40 might receive better care from their relatively younger offspring. Other hypotheses are that the family environment of the middle aged and the elderly might be more enjoyable and longevity-enhancing if it includes babies and younger children. Further, people who produce late offspring might tend to take better care of their health because they have the responsibility of raising children and want to survive to see their children get married and produce grandchildren; such effects may continue from middle age to old age.

From the ordinal logistic regression analysis, we found that the oldest-old who had 2+ surviving children born after age 35 or 40 and had at least one such late-born daughter living nearby had statistically significant advantages in healthy survival over those who also had 2+ surviving children born after age 35 or 40 but not a nearby late-born daughter. This indicates that young daughters may provide better care to their very old parents than do sons. This is consistent with the literature on the gender dimension of family care and intergenerational relations (Gerstel and Gallagher 1993; Grundy et al. 1999; Légaré and Martel 2000; Zeng and George 2000; Zhi 2001; Grundy 2002).

Enlightened by the discussions in the literature (e.g., Hutchinson and Rose 1991; Kirkwood and Rose 1991; Perls 1997; Perls and Fretts 2001; Perls et al. 2002), we also hypothesize that women who have 3+ births after age 35 or 2 or 3+ births after age 40 and who remain healthy up to the oldest-old ages of 80–105 may have specific longevity gene(s) or other unobserved biological characteristics. The possession of such genetic or biological characteristics, or both, may be associated with higher fecundity after ages 35–40 and healthy longevity among women. The genetic hypothesis may be less relevant among oldest-old men who have 3+ offspring after age 35 or 2 or 3+ after age 40. This is because the decline of fecundity with the increase in age among men is much slower than it is among women. Most men are able to provide fertile sperm after age 40, 45, or even 50, and thus produce late-born offspring if the wife or female partner is able to conceive (Check et al. 1989; Schill 2001). In our sample, 16, 8, and 7 per cent of the oldest-old men had 1, 2, or 3+ offspring, respectively, after age 45, and 9, 3, and 2 per cent of the oldest-old men had 1, 2, and 3+ offspring, respectively, after age 50. The statistical analysis of the 1998–2000 follow-up data shows that, among oldest-old men, the association of late childbearing with healthy longevity after age 45 or 50 is similar to that of women giving birth after age 40.

In addition to possible biological and social factors, discussed above, selection may also play an important role. It may be that physically robust and healthy persons are more likely to have children after age 40 and also enjoy healthy longevity. Rich and healthy men might have married much younger women, and thus had higher chances of having children at later ages.

It seems clear that late childbearing is significantly associated with health longevity at ages 80–105. The magnitude of the effects of late childbearing on maternal mortality and healthy survival at middle age and the early stages of elderly life in China is, however, uncertain. We had no information for the early stages of the life course. It is possible that, among the cohort members born 80–105 years ago, some women who gave birth after age 35 or 40 have died owing to difficulties at childbirth or the extra burden and hardship of raising more children. The husbands of these unfortunate and perhaps genetically or biologically weak women might have died, too, as a result of sadness about the wife's death and the burden of raising small children as a widowed father with poor resources. As Chen and Jones (1989, p. 73) noted: in high-mortality populations, the aged are those who have survived the dangers of being born, the risks of infancy and childhood, and the sicknesses and accidents of middle age. The sicknesses and accidents of middle age might include delivery difficulties resulting from late childbearing and maternal mortality and its associated psychological and physical losses.

In general, the oldest-old observed in the 1998–2000 follow-up surveys are a select group who survived the possible hardships of their middle ages and are generally robust in health. Also some of them may be genetically strong. Perhaps robust women who gave birth after age 35 or 40 were able to overcome the difficulties caused by late childbearing during and after the deliveries, and, therefore, could survive to age 80–105. The weaker women who happened to give birth after age 35 or 40 might have died owing to problems during and after the deliveries. Thus the selected, (because more robust) oldest-old women with late childbearing experience may have advantages in healthy survival. Research on the extraordinarily selected population of Chinese oldest-old (including women who gave birth after age 35 or 40 and are likely to be more robust) may provide important insights into why some people survive to advanced old age in good health while others do not (Zeng et al. 2001). Nevertheless, it is not appropriate to draw firm conclusions from this study about the association of later childbearing with healthy survival at middle age and at the early stage of elderly life among the Chinese, although European studies have indicated that such an association exists (Doblhammer 2000).

In sum, we believe that four factors may possibly explain why late childbearing may positively affect healthy longevity at advanced ages: (1) social factors; (2) biological changes caused by late pregnancy and delivery; (3) genetic or other biological characteristics, or a combination of genetic and other biological factors, associated with healthy longevity and higher fecundity; and (4) selection. The biologic hypothesis, (2), is probably applicable to women only. The genetic/biological characteristics hypothesis, (3), may be more applicable to women than to men, but it is not clear how much of an effect genes may have and how large the sex differentials are. The social and selection hypotheses may be applicable to both women and men. Although relatively weaker than the effects for the oldest-old women, the remarkable positive association of late childbearing with healthy longevity among the oldest-old men has led us to conjecture that social factors and selection are important in addition to biological and genetic factors.

Appendix

Physical and mental health variables

The ADL functional statuses of eating, dressing, transferring, using the toilet, bathing, and continence are used to measure the elderly person's status of independence in daily living, based on the international standard of Katz's ADL index (e.g., Katz et al. 1963) adapted to the Chinese cultural/social context and carefully tested by pilot studies/interviews. In this paper, if none of the six ADL activities is impaired, the elder is classified as 'ADL active'; if one or more activities are impaired, the elder is classified as 'ADL impaired'.

The mental state of the Chinese oldest-old was screened using the Chinese version of the MMSE, which was culturally adapted and translated into the Chinese language. It was based on the international standard MMSE questionnaire (Folstein et al. 1975), and carefully tested in our pilot survey interviews. The Chinese version of the MMSE is appropriate for the cultural and socio-economic conditions among oldest-old persons in China. The questions are easily understandable and can be readily answered if the subject's cognitive function is normal (see Zeng 2002, for details). A perfect score on the MMSE is 30; the methods for determining the score based on items referring to orientation, registration, attention, calculation, recall, and language are the same as for the international standard version. We also use the same cut-offs as the MMSE international standard to define a score of 24+ as 'Normal cognitive function', and <24 as 'Impaired cognitive function' (see, e.g., Osterweil et al. 1994; Deb and Braganza 1999).

Each interviewee was asked the question 'How do you rate your health at present?', with possible responses being 'excellent', 'good', 'fair', 'poor', or 'not able to answer'; no proxy responses were allowed. Responses were dichotomized into very good/good vs. fair/poor. Those oldest-old who were not able to answer the question on self-rated health owing to health problems are included in the fair/poor category, which is a reasonable solution because these interviewees were generally not in good health.

The depression symptoms measure was based on answers to three questions about whether the subject 'looked on the bright side of things', 'did not feel lonely', and 'felt as happy now as when young'. If an elder answered 'no' to all of these three questions, he or she was classified as 'no depression symptoms'. Otherwise, he or she was classified as 'having depression symptoms'.

Family support and social connections

Marital status. Subjects who were widowed, divorced, separated, or never married in 1998 were compared with those who were married.

Surviving children. The number of surviving children at the time of the 1998 survey was coded as 0, 1–2, 3–4, and 5+.

Proximity to children. Subjects who lived with children or had at least one child close-by (in the same village or in the same neighborhood, i.e. within a small urban area belonging to the same neighborhood committee) were compared with those who neither had children living with them nor had children living close-by.

Religious activities. These (which were social activities and not manifestation of religiosity, such as personal prayer) were dichotomized into 'do not participate' (83.6 per cent) and 'participate almost everyday or sometimes' (16.4 per cent).

Social connection/support index. This was based on the following questions addressed to each interviewee. (1) Do you play cards or mah-jong regularly at the present time? (2) Do you watch TV or listen to the radio regularly at the present time? (3) Are retirement wages your first or second most important means of financial support? (4) Can you get adequate medical service when you are sick? (5) When you are sick, who usually takes care of you? The possible answers to questions (1)–(4) are 'yes' or 'no'. If the answer is 'yes', a score of 1 is given; otherwise, a score of '0'. The options for the answer to question (5) are (a) children and/or spouse; (b) other family members; (c) live-in caregiver; (d) friends; (e) social service; (f) nobody. Those who provided answers (a) and/or (b) and/or (c) were given a score of 1; answers (d) and (e) were given a score of 0.5; and answer (f) was given a score of 0. The sum of the scores covering these five questions was dichotomized at approximately the midpoint of the distribution of the sample into low and high social connection/support categories.

Health practice

Cigarette smoking. Smoking status was determined with the question 'Do you smoke regularly at the present time?' Response options were 'no' and 'yes'.

Alcohol consumption. Alcohol use was determined with the following three questions: (1) Do you drink alcohol at the present time? (2) If you drink alcohol at the present time, how much per day on average do you drink? The measurement unit employed in the answer was 'liang', which is approximately equal to 50 grams and is widely used as a unit for measuring alcohol consumption in people's daily conversations in China. (3) If you drink at the present time, what kind of alcohol do you mainly drink? The options for the answer were 'liquor', 'wine', or 'rice wine'. We coded subjects into 'strong alcohol drinker' if they drank more than 2 liang (100 grams) of liquor per day or more than 6 liang (300 grams) of wine per day or 8 liang (400 grams) of rice wine per day; otherwise, the subject was coded as 'not a strong alcohol drinker'.

Diet. If a respondent ate both meat and fresh vegetables often, he or she was considered to have a 'healthy diet'; otherwise, he or she was considered to have an unhealthy diet.

Exercise. If a subject exercised regularly, he or she was coded as 'yes'; otherwise, 'no'.

Some self-report variables such as adequacy of medical service, diet, exercise, and depression symptoms are inevitably vulnerable to the vagaries of subjective judgement but can be useful when objective data are impossible to obtain, as was the case in our survey. For example, many previous studies have demonstrated that self-assessed health status acts as a significant and independent predictor of the functioning and mortality of older people (e.g., Lee 2000).

Notes

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References

Allison, P. 2002. *Missing Data*. Thousand Oaks, CA: Sage.

Check, J. H., B. Shanis, A. Bollendorf, H. Adelson, and E. Breen. 1989. Semen characteristics and infertility in aging, *Archives of Andrology* 23: 275–277.

Chen, A. J. and G. Jones. 1989. *Aging in ASEAN: Its Socioeconomic Consequences*. Singapore: Institute of Southeast Asian Studies.

Christensen, K. and J. W. Vaupel. 1996. Determinants of longevity: genetic, environmental and medical factors, *Journal of Internal Medicine* 240(6): 333–341.

Clark, D. O., T. E. Stump, and S. L. Hui. 1998. Predictors of morbidity and basic ADL difficult among adults aged 70 years and older, *Journal of Aging and Health* 10(4): 422–440.

Coale, A. and S. Li. 1991. The effect of age misreporting in China on the calculation of mortality rates at very high ages, *Demography* 28(2): 293–301.

Cox, D. R. 1972. Regression models and life tables (with discussion), *Journal of the Royal Statistical Society* 34B: 187–220.

Deb, S. and J. Braganza. 1999. Comparison of rating scales for the diagnosis of dementia in adults with Down's syndrome, *Journal of Intellectual Disability Research* 43(5): 400–407.

Doblhammer, G. 2000. Reproductive history and mortality later in life: a comparative study of England & Wales and Austria, *Population Studies* 54(2): 169–176.

Doblhammer, G. and J. W. Vaupel. 2001. Lifespan depends on month of birth, *Proceedings of the National Academy of Sciences of the United States of America* 98(5): 2934–2939.

Fillenbaum, G. G. 1988. *Multidimensional Functional Assessment of Older Adults: the Duke Older Americans Resources and Services Procedures*. Hillsdale, NJ: Erlbaum Associates.

Folstein, M. F., S. E. Folstein, and P. R. McHugh. 1975. Mini-Mental State: A practical method for grading the cognitive state of patients for the clinician, *Journal of Psychiatric Research* 12(3): 189–198.

Gerdes, L. U., B. Jeune, K. Andersen-Ranberg, H. Nybo, and J. W. Vaupel. 2000. Estimation of Apolipoprotein E genotype-specific relative mortality risks from the distribution of genotypes in centenarians and middle-aged men: Apolipoprotein E gene is a 'frailty gene', not a 'longevity gene', *Genetic Epidemiology* 19: 202–210.

Gerstel, N. and S. Gallagher. 1993. Kin keeping and distress: gender, recipients of care, and work–family conflict, *Journal of Marriage and the Family* 55: 598–607.

Grambsch, P. M. and T. M. Therneau. 1994. Proportional hazards tests and diagnostics based on weighted residuals, *Biometrika* 81: 515–526.

Grundy, E., M. Murphy, and N. Shelton. 1999. Looking beyond the household: intergenerational perspectives on living kin and contacts with kin in Great Britain, *Population Trends* 97: 33–41.

Grundy, E. 2002. Gender and ageing. Paper presented at the IUSSP (International Union for Scientific Studies of Population) Seminar on Longer Life and Healthy Aging, Beijing, 22–24 October.

Harrell, F. E. 2001. *Regression Modeling Strategies: With Applications to Linear Models, Logistic Regression, and Survival Analysis*. New York: Springer-Verlag.

Heck, K. E. and E. R. Pamuck. 1997. Explaining the relation between education and postmenopausal breast cancer, *American Journal of Epidemiology* 145(4): 366–372.

Hill, M. E. 1999. Multivariate survivorship analysis using two cross-sectional samples, *Demography* 36: 497–503.

Hitt, R., Y. Young-Xu, and T. Perls. 1999. Centenarians: the older you get, the healthier you've been, *Lancet* 354: 652.

Hosmer Jr, D. W. and S. Lemeshow. 1989. *Applied Logistic Regression*. New York: Wiley.

Hutchinson, E. W. and M. R. Rose. 1991. Quantitative genetics of postponed aging in *Drosophila melanogaster*, I: analysis of outbred populations, *Genetics* 127: 719–727.

Katz, S., A. B. Ford, R. W. Moskowitz, B. A. Jackson, and M. W. Jaffe. 1963. Studies of illness in the aged—the index of ADL: a standardized measure of biological and psychosocial function, *Journal of the American Medical Association* 185(12): 914–919.

Kerber, R. A., E. O'Brien, K. R. Smith, and R. M. Cawthon. 2001. Familial excess longevity in Utah genealogies, *Journal of Gerontology: Biological Science and Medical Science* 56A(3): B130–B139.

Kirkwood, T. B. L. and M. R. Rose. 1991. Evolution of senescence: late survival sacrificed for reproduction, *Philosophical Transactions of the Royal Society of London* B332: 15–24.

Koenig, H. G., J. C. Hays, D. Larson, L. K. Geroge, et al. 1999. Does religious attendance prolong survival? A six-year follow-up study of 3968 older adults, *Journal of Gerontology: Medical Science* 54A(7): M370–M376.

Landerman, L. R. and G. G. Fillenbaum. 1997. Differential relationship of risk factors to alternative measures of disability, *Journal of Aging and Health* 9(2): 266–279.

Landerman, L. R., K. L. Land, and C. F. Pieper. 1997. An empirical evaluation of the predictive mean matching method for imputing missing values, *Sociological Methods and Research* 26(1): 3–33.

Law, T. M., P. J. Hesketh, K. A. Porter, L. Lawn-Tsao, R. McAnaw, and M. J. Lopez. 1996. Breast cancer in elderly women: presentation, survival, and treatment options, *Surgical Clinics of North America* 76(2): 289–308.

Lee, Y. 2000. The predictive value of self assessed general, physical, and mental health on functional decline and mortality in older adults, *Journal of Epidemiology and Community Health* 54(2): 123–129.

Légaré, J. and L. Martel. 2000. Living arrangements of older persons, in *Proceedings of the United Nations, Economic Commission for Europe, Population Activities Unit Conference on Status of the Older Population: Prelude to the 21st Century, Sion, Switzerland*. Geneva: ECE.

Mantel, N. and W. Haenszel. 1959. Statistical aspects of the analysis of data from retrospective studies of diseases, *Journal of the National Cancer Institute* 22: 719–748.

Meurer, L. N., P. M. Layde, and C. E. Fuse. 2001. Self-rated health status: a new vital sign for primary care?, *Wisconsin Medical Journal* 100(7): 35–39.

Osterweil, D., P. Mulford, K. Syndulko, and M. Martin. 1994. Cognitive function in old and very old residents of a residential facility: relationship to age, education, and dementia, *Journal of the American Geriatrics Society* 42(7): 766–773.

Paganini-Hill, A. 1996. Estrogen replacement therapy in the elderly, *Zentralblatt für Gynäkologie* 118(5): 255–261.

Palmore, E. B. and B. M. Burchett. 1997. Predictors of disability in the final year of life, *Journal of Aging and Health* 9(3): 283–297.

Penning, M. J. and L. A. Strain. 1994. Gender differences in disability, assistance and subjective well-being in later life, *Journal of Gerontology: Social Sciences* 49B: S202–S208.

Perls, T. 1997. Centenarians prove the compression of morbidity hypothesis, but what about the rest of us who are genetically less fortunate?, *Medical Hypotheses* 49(5): 405–407.

Perls, T., M. Shea-Drinkwater, J. Bowen-Flynn, S. B. Ridge, S. Kang, E. Joyce, M. Daly, S. J. Brewster, L. Kunkel, and A. A. Puca. 2000. Exceptional familial clustering for extreme longevity in humans, *Journal of American Geriatrics Society* 48: 1483–1485.

Perls, T., L. Kunkel, and A. Puca. 2002. The genetics of exceptional human longevity, *Journal of the American Geriatrics Society* 50(2): 359–368.

Perls, T. T., L. Alpert, and R. C. Fretts. 1997. Middle aged mothers live longer, *Nature* 389: 133.

Perls, T. T. and R. C. Fretts. 2001. The evolution of menopause and human life span, *Annals of Human Biology* 28(3): 237–245.

Preston, S. H., M. E. Hill, and D. L. Drevenstedt. 1998. Childhood conditions that predict survival to advanced ages among African-Americans, *Social Science and Medicine* 47: 1231–1246.

Puca, A. A., M. J. Daly, S. J. Brewster, T. C. Matise, J. Barrett, M. Shea-Drinkwater, S. Kang, E. Joyce, J. Nicoli, E. Benson, L. M. Kunkel, and T. Perls. 2001. A genome-wide scan for linkage to human exceptional longevity identifies a locus on chromosome 4, *Proceedings of National Academy of Sciences of the United States of America* 98(18): 10505–10508.

Schill, W. B. 2001. Fertility and sexual life of men after their forties and in older age, *Asian Journal of Andrology* 3: 1–7.

Smith, J. B. and R. Kington. 1997. Demographic and economic correlates of health in old age, *Demography* 34: 159–170.

Smith, K. R., G. P. Mineau, and L. L. Bean. 2003. Fertility and post-reproductive longevity, *Special Biodemography Issue of Social Biology.*

Snowdon, D. A., R. L. Kane, W. L. Beeson, G. L. Burke, J. M. Sprafka, J. Potter, H. Iso, D. R. Jacobs, and R. L. Phillips. 1989. Is early natural menopause a biological marker of health and aging?, *American Journal of Public Health* 79: 709–714.

Snowdon, D. A. 1990. Early natural menopause and the duration of postmenopausal life: findings from a mathematical model of life expectancy, *Journal of the American Geriatric Society* 38(4): 402–408.

Snowdon, D. A. 2001. *Aging with Grace.* New York: Bantam Books.

Strawbridge, W. J., R. D. Cohen, S. J. Shema, and G. A. Kaplan. 1997. Frequent attendance at religious services and mortality over 28 years, *American Journal of Public Health* 87: 957–961.

Vaupel, J. W. 1992. Analysis of population changes and differences. Paper presented at the annual meeting of the Population Association of America, April 1992.

Voland, E. and C. Engel. 1986. Ist das postmenopausale sterbealter variable einer fitness-maximierenden reproduktionsstrategie?, *Anthropologischer Anzeiger* 44(1): 10–34.

Wang, Z. L., Zeng Yi, B. Jeune, and J. W. Vaupel. 1998. Age validation of Han Chinese centenarians, *Genus* 54(1/2): 123–141.

Yashin, A. I., G. D. Benedictis, J. W. Vaupel, Q. Tan, K. F. Andreev, I. A. Iachine, M. Bonafe, M. D. Luca, S. Valensin, L. Carotenuto, and C. Franceschi. 1998. Combining genetic and demographic information in population studies of aging and longevity, *Journal of Epidemiology and Biostatistics* 3(3): 289–294.

Yashin, A. I., G. D. Benedictis, J. W. Vaupel, Q. Tan, K. F. Andreev, I. A. Iachine, M. Bonafe, M. D. Luca, S. Valensin, L. Carotenuto, and C. Franceschi. 1999. Gene, demography and life span: the contribution of demographic data in genetic studies of aging and longevity, *American Journal of Human Genetics* 65: 1178–1193.

Yashin, A. I., G. D. Benedictis, J. W. Vaupel, Q. Tan, K. F. Andreev, I. A. Iachine, M. Bonafe, M. D. Luca, S. Valensin, L. Carotenuto, and C. Franceschi. 2000. Genes and longevity: lessons from centenarian studies, *Journal of Gerontology: Biological Sciences* 55A(6): B1–B10.

Zeng, Yi and L. George. 2000. Family dynamics of 63 million (in 1990) to more than 330 million (in 2050) elders in China, *Demographic Research* 2(5).

Zeng, Yi, J. W. Vaupel, Z. Y. Xiao, C. Y. Zhang, and Y. Z. Liu. 2001. The healthy longevity survey and the active life expectancy of the oldest old in China, *Population: An English Selection* 13(1): 95–116.

Zeng, Yi, J. W. Vaupel, Z. Y. Xiao, C. Y. Zhang, and Y. Z. Liu. 2002. Sociodemographic and health profiles of the oldest old in China, *Population and Development Review* 28: 251–273.

Zhi, Yi. 2001. Study on intergenerational interaction of 1,303 elderly persons in urban areas of Zhejiang Province, *Journal of Chinese Gerontology* 3: 81–82 (in Chinese).