Observations on the History of Cohort Fertility in the United States

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This is a study of the completed fertility of cohorts of women in the United States, beginning with those born in 1867 and ending with those born in 1955. The term *cohort* is used to signify persons born in a specified period. The history of cohort fertility in the United States is well known in broad terms: a long-run decline on which there is superimposed a transitory reversal, peaking about a generation ago. An earlier study of the same time series focused on the contrast between cohort and period measures, essentially a question of the changing tempo of cohort reproduction (Ryder, 1980). In the present study, we examine some details of reproductive change from a parity-specific viewpoint. To further that end, we propose a model of the components of parity progression. The formulation provokes new questions for those concerned with the theory of fertility decline that seem to deserve attention despite the admittedly speculative character of some estimates used in the model.

Some technical preliminaries

The principal source of information is the set of fertility tables for birth cohorts of American women (Heuser, 1976). Those tables provide alternative representations of a set of basic data: birth rates for separate orders of birth, specific for woman's age and period of occurrence. The materials have two distinctive features. Although the numerator of each birth rate is restricted to births of a particular order, the denominator is a person-years count for all women of the specified age in the specified period. Furthermore, the tables are so arranged as to simulate the experience of birth cohorts—to be precise, the combinations of age and period values for which the difference between period and age is the same. In consequence, the resultant cohorts have birth dates centered on the beginning of each designated year of birth.

Birth rates of this kind have been produced for each year, beginning with 1917. Although the published volume concludes with data for 1973, supplements have been published annually for 1974–81 in the volume on Natality of the annual report of vital statistics (National Center for Health Statistics, 1985), and the author has been provided with the comparable unpublished information for 1982 and 1983. The rates are superior to those published elsewhere in reports of vital statistics because their denominators are corrected for misenumeration. Their quality is likely to be lower for recent years, however, because the presence of an illegal population of uncertain but probably substantial magnitude has thwarted efforts at confident correction.

The analysis of completed fertility can exploit a much longer time series than that encompassed by the records for the years 1971–83 because estimates have been produced of cumulative birth rates by order, as of the beginning of 1917, for all cohorts then in the reproductive ages. The principal source of those estimates is the report of parity for ever-married women in the 1910 census (Heuser, 1976, Appendix III). Accordingly the time series studied here begins with the birth cohort of 1867, which was exact age 50 at the beginning of 1917.

For different purposes, the information to be considered is presented in three alternative forms. The cumulative birth rates by birth order are convenient for computation but analytically unevocative. To study the determinants of reproductive behavior, the preferred form is the set of parity progression ratios, that is, the conditional probability of having an $i+1^{\text{th}}$ birth, given that an i^{th} birth has occurred. To analyze the consequences of reproductive behavior, there are some advantages to the parity distribution. A clarification of the relations among these forms of measure may make the subsequent account more intelligible.

The basic element in the cohort fertility tables is a birth rate for members of a female birth cohort in one or another reproductive age, with the numerator of the rate restricted to births of a particular order. The age-specific i^{th} order fertility rate for women of age x (that is, in the exact age interval $(x,x+1)$) is $F(i,x) = B(i,x)/Y(x)$, where $B(i,x)$ are births of order i to women of age x, and $Y(x)$ are the person-years of women in age x. (This cannot be called a parity-specific rate because the denominator is unidentified by parity.)

One can identify the total i^{th} order fertility rate, $F(i)$, as the sum of $F(i,x)$ over all reproductive ages x. (Here and elsewhere the summation is for the cohort in question.) $F(i)$ may be interpreted as the proportion of women in the birth cohort who have an i^{th} order birth at some time in their reproductive careers. (As is customary demographic practice, such measures ignore the implications for fertility of changes in the cohort population attributable to migration and mortality.)

A parity distribution is constructed from these rates by differencing. The proportion of women ending their reproduction in parity i, $P(i)$, is given by $P(i) = F(i) - F(i+1)$, where $F(0) = 1$. As an analytic alternative to the parity distribution, the author devised the concept of the parity progression ratio (Ryder, 1951). Subsequently and independently, the same measure, with the name ''probabilité d'aggrandisement,'' was developed by Louis Henry (Henry, 1953). The progression ratio for parity i, R_i, is defined by $R_i = F(i+1)/F(i)$.

Thus one may represent the (cohort) total fertility rate, F, in three alternative forms:

$$F = \sum_{i=1}^{\omega} F(i) = \sum_{i=1}^{\omega} i\,P(i) = \sum_{i=0}^{\omega-1} \prod_{j=0}^{i} R_j \quad ,$$

where ω is the highest birth order (and parity) attained.

The expression for the total fertility rate in terms of parity progression ratios is isomorphic with the expression for the expectation of life at birth, with the progression ratio substituted for the probability of surviving from one age to the next. Indeed, the system of fertility measures shown here can be characterized in life table terms, but with parity rather than age serving as the argument of the functions. Just as persons progress (survive) from age to age, so do they progress from parity to parity. The parity progression ratio is the analogue of the p_x value in the life table, the i^{th} order fertility rate is the analogue of the l_x value, and the proportion in parity i is the analogue of the d_x value.

In the following account, we have introduced into the parity progression sequence the probability of progressing from never-married to ever-married, calculated by estimating the proportion of each cohort of women who were ever-married at exact age 45, say M. That measure is used to partition the conventional progression ratio for parity zero, $R_0 = F(1)$, into M and $R_0' = F(1)/M$, and similarly to partition the (cohort) total fertility rate, F, into M and $F' = F/M$, henceforth called mean marital parity. This is admittedly not a completely legitimate calculation because some never-married members of a cohort contribute to its fertility. On the other hand, it represents an approximation, faute de mieux, of the analytic distinction between changes in fertility attributable to nuptiality and those attributable to marital fertility. Moreover, some such procedure is subsequently required because some data used in the model to be presented are descriptive of the reproductive behavior of ever-married women.

The values for M, the proportion ever-married by exact age 45, were based on the relevant census reports of women by age and marital status. Intra-cohort analysis of data from the census volumes on Age at Marriage, for the 1960 and 1970 censuses (US Bureau of the Census, 1966 and 1973), indicated the desirability of increasing the reported proportions ever-married. Two sources of undercount were identified: (1) the never-married are less likely to be enumerated than the ever-married; (2) some marriages occurring at earlier ages (in the earlier census) seem to have been expunged from the record at later ages (in the later census). Although it is probably true that underestimates of proportions ever-married are somewhat counterbalanced by the tendency of some never-married women who are fertile to report themselves as ever-married, this last source of error is considered inoffensive in light of the use to be made of M in the present account. The two sets of census data were used to produce a corrected set of M values for the cohorts born in the years 1900–25. (In the subsequent text, references to cohorts born over a group of years like this will be symbolized, as a convenient abbreviation, by cc.1900–25.) Similar adjustments were made to the values for prior and subsequent cohorts to preserve temporal comparability. Unless otherwise specified, all fertility measures in the following account refer to the reproductive behavior of birth cohorts of women.

Fertility projection for cohorts with incomplete reproduction

Notwithstanding the historical focus of the present account, there is also considerable interest in the most recent experience in the United States: fertility is now lower than ever before. Were we to be purist in confining our attention to those cohorts that have completed their childbearing, the most recent cohort available for the time series would be in its late 40s at the end of 1983, displaying the twentieth century peak in the series but moot on the sequel. Although some may be tempted to achieve contemporaneity by resorting to a period mode of temporal aggregation, the evidence is clear that this practice, as regrettable as it is commonplace, yields a view of what is happening that may be substantially distorted (Ryder, 1980). But, for the record, the period total fertility rate in the United States has been steady within about 3 percent of 1.80 over the years 1978–83, and advance reports of births in 1984 and 1985 indicate that it remains within that narrow range.

We think the preferable way to achieve contemporaneity is to make a fertility projection for cohorts currently in the childbearing ages, provided they have already accomplished a substantial part of their reproductive record. We have made a projection for birth cohorts through cc.1951–55, on the assumption that each incomplete cohort will subsequently experience the same fertility rates, specific for age and parity, observed in the comparable ages in 1983. The risk of error would not seem to be large: the least complete cohort projected has already had five-sixths of its first births and two-thirds of its second births; these two birth orders constitute more than three-quarters of its total fertility. The total fertility rate for cc.1951–55 is estimated as $F = 1.92$. It should not be surprising that this value is substantially higher than the 1.80 cited above for period total fertility currently. Whenever the age pattern of cohort reproduction is sliding upward—as it has been recently—the consequence is to distort downward the level of period fertility.

The recent history of cohort fertility for the United States shows a continuing decline from $F = 3.20$ for cc.1931–35 to $F = 1.92$ for cc.1951–55. The interesting question is whether the decline is likely to continue. A new source of data has recently become available that can help to answer this question. Every five years, the Current Population Survey collects birth histories for all women. (See, for example, US Bureau of the Census, 1982.) We have exploited the public use file for the June 1985 survey to calculate fertility rates for birth cohorts, of higher specificity than is feasible with data from the cohort fertility tables. Instead of controlling only for woman's age and parity, we are also able to partition age into the sum of age at entry into the parity, and the length of the open interval (the length of time already spent in that parity). At this level of detail, the reproductive records for cc.1955–59, as of mid-1985, are virtually the same as those for cc.1950–54, as of five years earlier.

For those cohorts with incomplete reproduction in 1985, the remainder of their fertility history was projected on the assumption that they would experience the same birth rates within each parity, age at entry into parity, and length of time in that parity, as their predecessors experienced in the 1980–85 period. The resultant total fertility rates for the quinquennial birth cohorts of 1935–39 through 1955–59 were projected to be 2.889, 2.469, 2.070, 1.935, and 1.915 respectively.

For purposes of comparison, the total fertility rates for cohorts 1936–40 through 1951–55, as projected from the cohort fertility data, were 2.950, 2.470, 2.108, and 1.920 respectively. The close correspondence between the two projected series gives some confidence in the survey of birth histories as a reliable source of up-to-date information on fertility.

Although the projections based on the birth histories of 1985 are more speculative than those based on the cohort fertility tables—in the sense that the record to date ends at average age 28 for the most recent cohorts in the former, but at average age 30 in the latter—the risk of error is attenuated by the circumstance that the control variables available for the former permit a firmer purchase on the time pattern of childbearing. The implication of the birth history projection is that the 20-year decline in the cohort total fertility rate has virtually ended. Furthermore, the concomitant rise in the mean age of fertility would also seem to have ended: its projected value is 26.7 for both cc.1950–54 and cc.1955–59. The best evidence currently available suggests, then, that, at least for the moment, cohort fertility has stabilized.

The time series of aggregate fertility

In Table 1, we show five summary measures of fertility for the cohorts of 1867–1955. In the first column, the cohort total fertility rate, F, follows the well-known pattern of decline from 4.000 for the initial cohort group to a local minimum of 2.286 for cc.1906–10. This is followed by a strong rise to a local maximum of 3.201 for cc.1931–35. Current experience is manifested in the historic low of 1.920 for cc.1951–55.

The second column shows the proportion of women ever-married by exact age 45, M. Although the variations are small, there was a steady rise in

TABLE 1 Fertility measures for US cohorts, 1867–1955

Cohort	Total fertility rate (*F*)	Proportion ever-married (*M*)	Mean marital parity ($F'=F/M$)	Coefficient of variation of marital parity (*c*)	Mean sibship size ($F'(1+c^2)$)
1867–70	4.000	0.922	4.338	0.827	7.304
1871–75	3.773	0.927	4.068	0.854	7.039
1876–80	3.532	0.927	3.809	0.880	6.758
1881–85	3.322	0.936	3.547	0.900	6.420
1886–90	3.137	0.939	3.342	0.906	6.085
1891–95	2.933	0.942	3.107	0.907	5.663
1896–1900	2.676	0.941	2.842	0.931	5.306
1901–05	2.442	0.948	2.574	0.965	4.971
1906–10	2.286	0.952	2.402	0.976	4.689
1911–15	2.354	0.962	2.446	0.919	4.512
1916–20	2.574	0.971	2.649	0.829	4.470
1921–25	2.857	0.973	2.936	0.745	4.568
1926–30	3.079	0.974	3.160	0.702	4.716
1931–35	3.201	0.976	3.280	0.642	4.630
1936–40	2.950	0.969	3.044	0.614	4.193
1941–45	2.470	0.952	2.594	0.611	3.564
1946–50	2.108	0.945	2.230	0.617	3.078
1951–55	1.920	0.917	2.094	0.600	2.848

M, to a maximum for cc. 1931–35, and since then a decline to the lowest level in this record. Because those variations are small, the value for mean marital parity, F', in the third column, tells much the same story as the total fertility rate. The decomposition of this index is the subject of what follows.

The fourth column of Table 1 presents a less familiar measure, the coefficient of variation of the parity distribution, *c*. This measure has interest in its own right, as a manifestation of what may be a general characteristic of fertility decline—the tendency for the parity distribution, as it evolves, to evince first an increase and later a decrease of reproductive heterogeneity in the population. This is what one would expect from a pattern of dissemination of a new low-fertility pattern among a progressively larger proportion of the population.

A further reason for displaying the coefficient of variation of the marital parity distribution is its relationship to the values of mean sibship size, shown in the last column of Table 1. Mean sibship size measures fertility from the standpoint of the children involved, just as mean parity measures fertility from the standpoint of the parents. If a woman bears *n* children during her lifetime, each of those *n* children is a member of a sibship of size *n*. Mean sibship size is the relevant measure to use in investigating hypotheses about the subsequent behavior of children as a function of the size of the child's family of orientation. Samuel Preston (1976) has provided a compelling account of the value of this measurement. Currently there is a lively debate on the role of sibship size in educational stratification (Blake, 1981, 1985, 1986; Mare and Chen, 1986a and b).

The formula for calculating mean sibship size for the children born to a cohort of women is

$$\Sigma\, i^2P(i) / \Sigma\, iP(i) \quad ,$$

where $P(i)$ is the proportion of the women in the cohort who end with parity i. Algebraically, this is readily transformed into $F' \ (1+c^2)$, where F' is the mean marital parity and c is the coefficient of variation of the marital parity distribution. During the period of substantial rise in mean marital parity, from cc.1906–10 to cc.1931–35, the decline in the coefficient of variation was sufficient to leave the mean sibship size for children of those women essentially unchanged.

The distinction between the parent-oriented and the child-oriented measures of fertility is important in evaluating a popular hypothesis concerning the time series of fertility. It has been widely noted that the low birth rates of the 1930s were followed, a generation later, by the high birth rates of the late 1950s; yet another generation later, birth rates are low again. This has provoked speculation that the sequence is more than coincidental. But if the chain of causation is presumed to run from the size of the family in which a child is raised to the subsequent reproductive behavior of that child, the appropriate measure of parental fertility is mean sibship size.

It seems to us that the task of ascertaining the parental characteristics of the members of birth cohorts has not heretofore been addressed squarely. Beyond the recognition that the children of more fertile parents will be more generously represented among the offspring than those of less fertile parents, it is also necessary to incorporate in the calculations the circumstance that the members of any birth cohort are recruited from a cross-section of those parental cohorts who are in the various ages of reproduction in the period in which the children are born.

We have made estimates of mean sibship size for the cohorts of children born in the quinquennial periods 1920–24 through 1955–59. They are shown in Table 2, together with the subsequent fertility of the cohorts concerned; for

TABLE 2 Mean sibship size for cohort members and their subsequent total fertility rate: Quinquennial cohorts, 1920–59, United States

Cohort	Mean sibship size	Percent change	Total fertility rate	Percent change
1920–24	5.336		2.800	
1925–29	5.175	−3	3.035	+ 8
1930–34	4.958	−4	3.177	+ 5
1935–39	4.633	−7	3.000	− 6
1940–44	4.348	−6	2.566	−14
1945–49	4.214	−3	2.180	−15
1950–54	4.302	+2	1.935	−11
1955–59	4.357	+1	1.915	− 1

the last four quinquennial cohorts, the fertility values come from the projections described previously. The basic data are birth rates by birth order and cohort, for the specified periods (Heuser, 1976, Table 2A). For each cohort of parents, we have the distribution of its eventual fertility by order. Births of a given order represent mean sibship sizes based on a distribution restricted to that order and all higher orders, giving an order-specific sibship size. Once the sibship sizes by order, for each cohort in each period, are obtained in this way, the results must then be weighted by the size of the parent cohort. (Cohort size is corrected for underenumeration, using the estimates presented in Whelpton and Campbell, 1960, Table A-6.) In this way we obtain a mean sibship size for the births of the period in question, that is, for the cohort originating in that period.

It is evident from a comparison of the intercohort changes in the two measures shown in Table 2 that there is no relationship between the average number of children in the cohort's families of orientation and the subsequent fertility of the cohort.

The parity distribution and parity progression

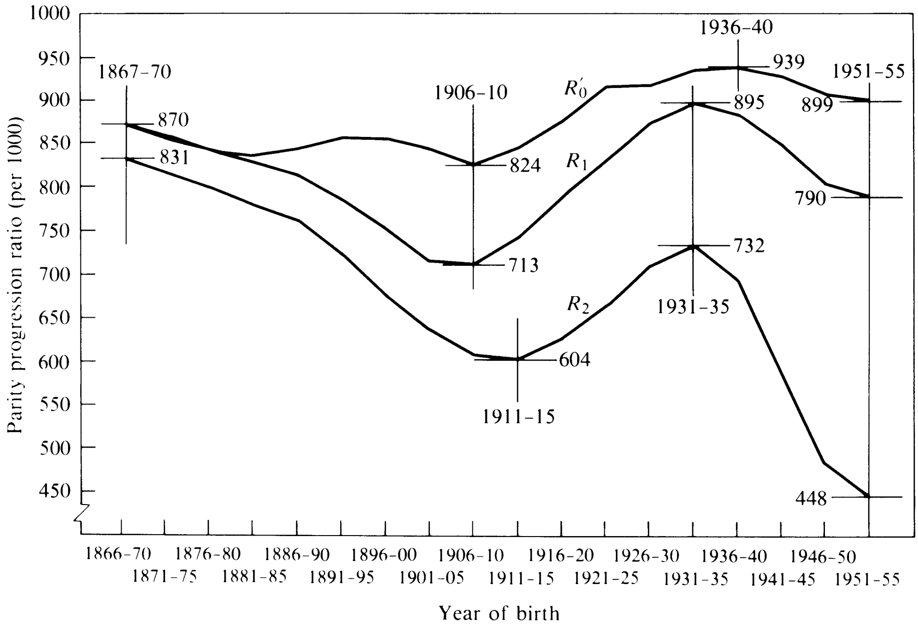
The trends in the mean and coefficient of variation of the parity distribution have been shown in Table 1. These are summary indexes of the parity distribution itself. We do not propose to present a table showing changes in the observed parity distribution, for several reasons. In the first place, from the standpoint of understanding the determinants of fertility, such changes are an epiphenomenon. The parity distribution shows the outcome of a series of parity progression ratios, each ratio having due effect on all subsequent parity distribution components. The distributional elements thus do not clearly reveal the consequence of changes in progression from one parity to the next. This was our principal reason for devising the progression ratio approach. For determinant analysis, the progression ratios are the desideratum, and we consider them below.

In the second place, although the parity distribution, as the outcome of the sequence of progression probabilities, would at least seem to be suited for consequential analysis (Ryder, 1971), it is clearly inferior there too, for reasons to do with the sibship distribution, as should be apparent from the preceding section.

In the third place, the parity distribution offers little guidance in distinguishing between what women intended to have happen, and what did happen willy-nilly. For that purpose, we propose a model of the components of parity progression, in the following section.

The time series of parity progression ratios are displayed in Figures 1 and 2. The evident justification for splitting these series into two groups is that the story told by the ratios for the higher parities is completely different from that for the lower parities. The sources of the major fluctuation in fertility in

FIGURE 1 Parity progression ratios (per thousand) for parities 0, 1, and 2: US cohorts 1867–1955



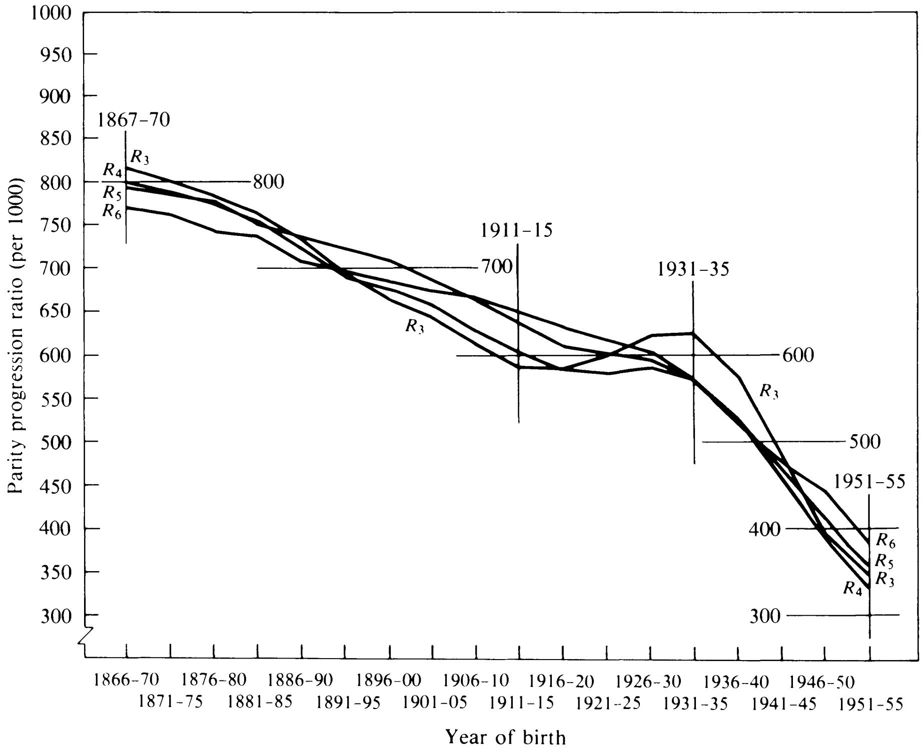
the twentieth century are clearly revealed in Figure 1. For perspective, note that, in the sequence of ratios, those for a lower parity are much more potent in their import for total fertility than those for any higher parity.

The dominant role in the fertility fluctuation has been played by R_1 and R_2, the progression ratios for parities one and two. The noteworthy characteristic of the time series for R'_0, the parity-zero progression ratio for ever-married women, is the systematic and rather abrupt shift from a lower to a higher level. Although its current value is less than it was a generation ago, it is much higher than the values for the first half of the series. There is a further detail with respect to R_1 and R_2: the low values of these ratios for the cohorts whose childbearing was concentrated in the 1930s were the culmination of a long downward slide in the progression probabilities. The phenomenon deserves more than an ad hoc reference to the tribulations of a particular depression decade; a more general explanation of the reproductive evolution is called for.

One interesting sidelight is that the time series for progression from parity two, R_2, is virtually identical in shape with the time series for mean marital parity itself (not shown). That stands as empirical justification for the conceptual framework of the Princeton Fertility Study, a longitudinal investigation of the progression from parity two to parity three (Westoff et al., 1963).

The time series for higher parity progressions are shown in Figure 2. The detail for each parity is effectively lost because the four curves are so

FIGURE 2 Parity progression ratios (per thousand) for parities 3, 4, 5, and 6: US cohorts 1867–1955



much alike. Otherwise said, there is little information contained therein that would not be conveyed simply by reporting a time series of their mean values. On average, the lowest of the four values for each cohort is only 6 percent lower than the highest. And even this is likely to be an overstatement of the range. Of the 18 cohort groups, the lowest value is that of R_4 in nine of the cases and that of R_6 in another six.

We think it unlikely that the circumstance that the low values are mainly for even parities is a chance occurrence. The progression ratios for even parities have odd-ordered births in the numerator and even-ordered births in the denominator; the opposite is true of ratios for odd parities. Accordingly, should there be any tendency to over-report even-numbered relative to odd-numbered births, there would be a double effect on the respective progressions. The tendency is less pronounced in the early part of the series, but those data mainly derive from enumeration rather than registration; any tendency for an even-number bias in the reported parity is attenuated by the circumstance that births by order are inferred from a cumulation of the parity distribution.

Yet we would be remiss in neglecting to mention an alternative possibility to misreporting. There may be a tendency for couples to be a little more likely to attempt to terminate childbearing after an even rather than after an odd number of children. The preference for two children rather than one or three is well established, and defended conversationally at great length. Perhaps there is also a preference for four rather than three or five, and so on. It may be difficult to distinguish these two possible sources of influence on the sequence of progression ratios.

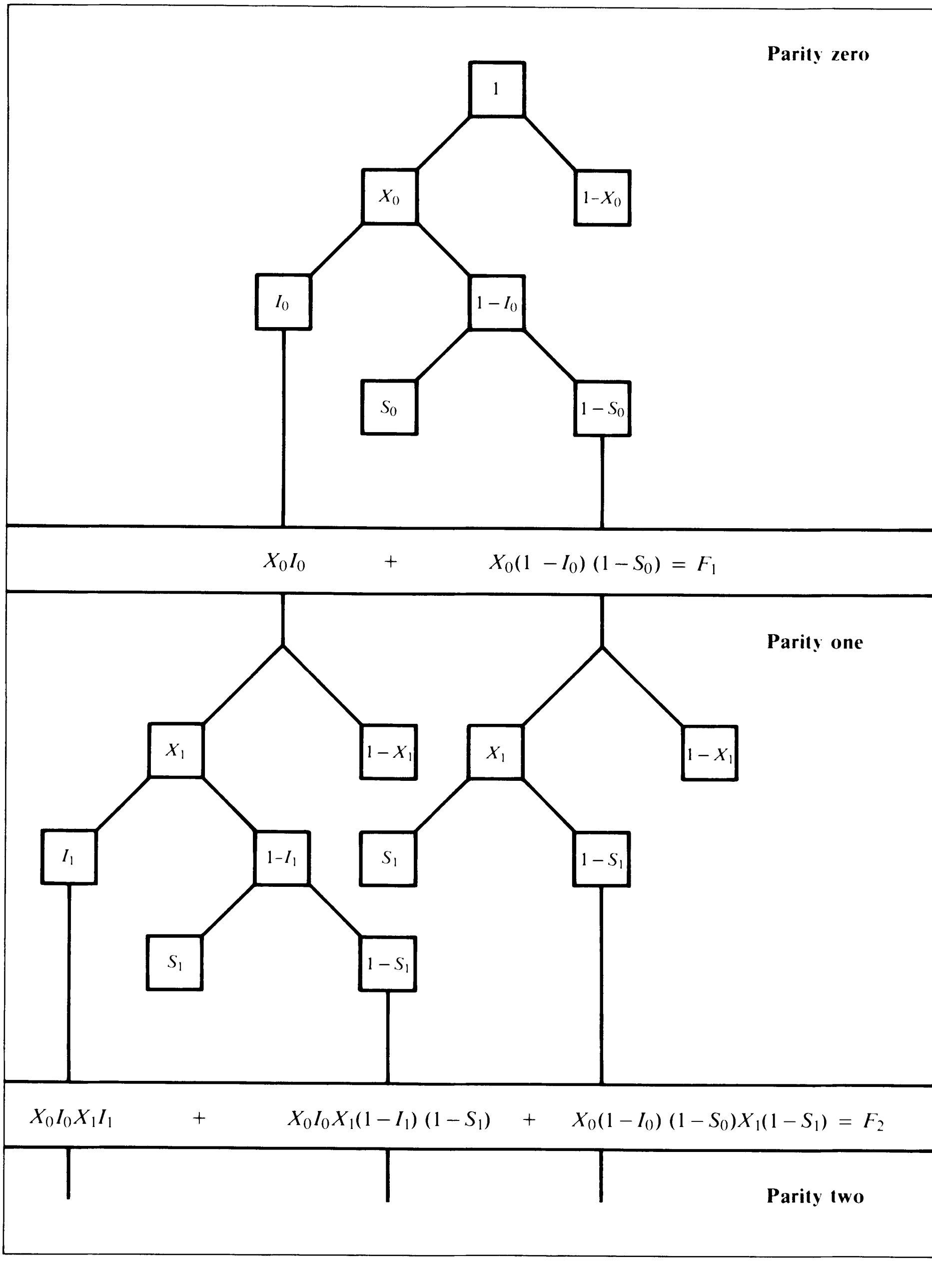
Two questions of considerable interest are provoked by the time series displayed in Figure 2. The first is why the values are so similar for any one cohort. In the abstract, one would anticipate that some kind of selection process is involved, with those who progress singled out by that fact as being different in relevant ways from those who do not. This question is addressed in the final section of this article. The second question concerns the virtually monotonic decline in higher-parity fertility. Most thinking about fertility evolution, explicit or implicit, postulates a growing dissemination of the two-child family. Yet that would be expected to produce a trough in the sequence of progression ratios, deepening over time, with progression at the higher parities much less affected. The phenomenon revealed in Figure 2 deserves the attention of those interested in the theory of fertility decline.

A parity progression model

It seems to us that more intensive analysis of the time series of progression, parity by parity, requires some recognition of the effect on progression ratios of likely changes in fecundity, and in the effectiveness of fertility regulation, as a function of parity (and implicitly of age as well). To this end, we propose that the process of progression from each parity to the next be partitioned into three components. The argument follows the diagram in Figure 3. In the first instance, the women who have attained a particular parity i have a probability, X_i, of being exposed to the risk of progressing to the next parity. There are two aspects to this exposure. It may terminate either with the disappearance of fecundity or with the cessation of copulation (most likely as a sequel to marital dissolution). The fecundity in question here differs from that ordinarily incorporated in fertility models, since the calculation is confined to those who have just entered a particular parity by virtue of having a birth. Note too that the model proceeds parity by parity rather than age by age; time is suppressed in the account. The issue is considered at greater length below.

In the upper panel of Figure 3, of all women of parity zero (and, in the present data set, ever-married), a proportion X_0 are exposed to the risk of a first birth. For those so exposed, the next question is whether they intend to have a first birth, with probability I_0, or not. We assume, in this model, that those who do intend to have a first birth, and are exposed to risk, do indeed have a first birth; it is considered an intended birth. For the proportion who do not intend to have any births, $1-I_0$, the next step is to consider whether they are successful in fulfilling that intention. The proportion of them who are

FIGURE 3 A model of parity progression



successful, S_0, terminate their childbearing at that point. The proportion of them who are unsuccessful, $1-S_0$, have a first birth.

The total number of first births consists of two parts, as indicated in Figure 3: those who intended (at least) a first birth, and those who intended no births, but were unsuccessful. Proceeding to parity one, the issue is raised again concerning the probability of being exposed to risk of (at least one more) birth, X_1. Again there will be a certain proportion of the exposed, say I_1, who intend to progress at least to parity two, and who do so. Their complement, $1-I_1$, who do not intend another birth, will be successful in that endeavor with probability S_1. Those who have already had an unintended first birth are presumed to have an unchanged intention. For them as well, the question concerns their probability of success in terminating childbearing, S_1.

For second births, then, the total is the sum of three probability products, representing progression by those who intend (at least) two, by those who intended one, and by those who intended none. The process continues parity by parity, with one additional element in each higher parity. It is apparent from this account why the straightforward progression ratio, R, would be an unreliable discriminator of parity-specific intentions.

The succession of equations, by straightforward algebra (see Appendix), yields the following algorithm:

$$F^*_{i+1} = F_i\left(X_i - \frac{X_i - R_i}{S_i}\right) \quad ,$$

where F^*_{i+1} is the number of intended $i+1^{\text{th}}$ births, and R_i is the conventional progression ratio for parity i, that is,

$$R_i = F_{i+1}/F_i \quad .$$

Using this algorithm, it becomes feasible, with a set of observed R_i (and thus F_i) values, and assumed values for X_i and S_i, to calculate the numbers of intended births by order, F^*_i. One can also use the values F^*_i to calculate the intended parity distribution,

$$P^*_i = F^*_i - F^*_{i+1} \quad .$$

Most desirably, one can calculate the parity-specific intention probabilities

$$I_i = (F^*_{i+1}/F^*_i)/X_i \quad .$$

How might such data be collected? Consider an interview with women who have completed the reproductive age span with at least n births. The first question concerns those who have not had an $n+1^{\text{th}}$ birth. The problem is to determine whether this was because they were unable or unwilling. Those who fall into the "unable" category are those who became sterile, or who were without a sexual partner, subsequent to the n^{th} birth. This would give the value of X_n. Those unwilling to have an $n+1^{\text{th}}$ birth are the successful terminators, a fraction $S_n(1-I_n)$. For those who did have an $n+1^{\text{th}}$ birth, the question is whether they intended to have that birth at any time in the future. If so, their proportion constitutes the value I_n. This corresponds closely with the procedures followed in the successive fertility surveys in the United States.

This is a highly simplified abstract of a more realistic model. For example, those who are classified as not progressing because they were not exposed to risk undoubtedly consist in part of those who were exposed to risk at least for a time, but subsequently became sterile or stopped copulation. For another example, the value of S, probability of success in fulfilling the intention to terminate, depends not only on the choice of method of fertility regulation, but also on whether the person who intends not to have another child actually employs some mode of fertility regulation to accomplish that objective. Moreover, some of the apparent success in terminating at the intended level is undoubtedly attributable to unrecognized nonexposure. Others may think there is nothing to decide, since they believe they are incapable of having a child. If they are surprised by a newcomer, they face the quandary of whether, had they known they were exposed to risk, they would have intended another child or not. All these are the familiar difficulties of interpreting such data from fertility surveys.

Rather than attempt to put an unreasonable burden on the somewhat unreliable yield from fertility surveys on such questions, we focus attention on the two considerations that create a discrepancy between actual progression and intended progression: change in exposure to risk from parity to parity (concerning which we assert confidently that X_i declines with increase in i) and change in the probability of success in terminating, from parity to parity (concerning which we assert equally confidently that S_i rises with increase in i). The basis for confidence in the last statement is that the higher the parity at which termination is intended, the shorter the length of exposure to risk of having an unintended birth.

Estimation of exposure to risk and of success in fertility regulation

The simple interpretation of the exposure variable, X_i, is that it would be the progression ratio for unregulated fertility. The population chosen to represent unregulated fertility is taken from the 1941 census of Canada: ever-married women, aged 45–54, Quebec residence, French mother tongue, Roman Catholic religion, rural, born on a farm, and with less than nine years of schooling (Charles, 1948). The mean marital parity for these women was 8.53. Their parity distribution permits progression ratios to be calculated directly. To arrive at a value for X_0, we take advantage of the fundamental inequality $R_0 \leqslant X_0 \leqslant 1$. Since R_0 reaches the value of 0.94 in the series for US cohorts, we choose to set $X_0 = 0.97$. It was approximately the case that $\ln X_i$ (the natural logarithm of X_i) for the French-Canadian data followed a quadratic curve (giving a reasonable concave shape to the function). The actual formula employed for the value of X_i was $\ln X_i = -0.0030459\ (i^2 + i + 10)$. This causes the values of X_i to decline by about one percent over each of the first several parities, and more rapidly thenceforth.

The mean parity implied by this set of X_i values is 7.20, some 16 percent less than that observed for the French-Canadians. That seems reasonable for the American experience, given the likelihood that a substantial proportion of intervals in the American birth histories would be lengthened by fertility regulation. Note further that the parity progression ratios reflect not only the growth of infecundity with parity (and inferentially with age), perhaps reflecting the accumulation of reproductive insults, but also experience with marital dissolution (net of remarriage). On the one hand, male mortality in the French-Canadian population was not negligible; on the other hand, divorce would be much more prevalent in the American cohorts.

We assume that the X_i values do not change from cohort to cohort. What is the likelihood that there has been a trend in exposure, downward or upward, across the range of cohorts considered here? Were male mortality the only mode of marital dissolution, one would surmise that there had been a gradual rise in exposure. In the United States, to the contrary, divorce has substantially supplanted widowhood (in the ages prior to the end of reproduction) as a mode of dissolution. Moreover, the incidence of remarriage is somewhat higher for the divorced than for the widowed. And there may be some exposure to risk subsequent to marital dissolution.

Aside from the effect of marital dissolution (net of remarriage) on the incidence of exposure, there is the question of fecundity itself. Although evidence is scarce, it would not seem unreasonable to anticipate some positive correlation between the level of fecundity and the general health of the population. Rather than attempt an adjudication among the various influences upward and downward on the trend in exposure, we are content to hold it constant, and propose to examine subsequently the consequences of error in this assumption.

With respect to the estimation of values for S_i, success in terminating reproduction, there are two questions: the form of the function with respect to parity and the level of success. With respect to the form, we settled on a conventional survival function, with parity rather than age as its argument. For the exponent of the function, we calculated from the French-Canadian parity distribution the expectation of additional births for each parity, say d_i, and thus, for those intending to terminate at a particular parity, the length of risk (measured by number of additional births to be otherwise expected).

The form of the equation is

$$S_i = (1 - Q)^{d_i} \quad ,$$

where Q is the probability of contraceptive failure per birth. For the French-Canadian data, expectation of additional births, as a function of parity, was approximately linear, with slope approximately -0.6. For application to a population like that of the United States, in which those regulators who failed to terminate would ordinarily succeed at least in lengthening the interval prior to failure, we set the slope at -0.7, with intercept 9. With this specification the formula is

$$S_i = (1-Q)^{9-0.7i}$$

The second question is, what is an appropriate value for contraceptive failure, Q, for the series of birth cohorts in the United States? On this question, some data were available. Drawing on the 1965 and 1970 National Fertility Studies, the author has estimated the mean number of unintended births per woman, for marriage cohorts (real and synthetic), for whites and nonwhites combined, where the limit on age at first marriage was between 25 and 30 (Ryder, 1978). Drawing on the 1970 and 1975 National Fertility Studies, the author has estimated the mean number of unintended conceptions per woman, for marriage cohorts (real and synthetic), for whites in intact first marriages, where the limit on age at first marriage was 25.0 (Ryder, 1981). To meet present purposes, it was necessary to engage in a series of estimated conversions, from synthetic to real cohorts, from marriage to birth cohorts, from conceptions to births, from white births to total births, and from one age-at-marriage limit to another. The outcome was a series of mean number of unintended births per woman, for the quinquennial birth cohorts of 1916–20 through 1941–45.

Using the algorithm for intended fertility F^*_{i+1} as a function of R_i, X_i, and S_i, with the observed R_i and the assumed X_i (as described above), we determined by trial and error the values of Q that would, for each cohort, yield the mean number of unintended births per woman, U, estimated from the National Fertility Studies. The results are shown in Table 3. Elsewhere we have argued that the numbers of unintended births per woman in the National Fertility Studies are underestimates (Ryder, 1976 a and b).

TABLE 3 Mean number of unintended births per woman (U) and probability of failure per birth (Q), based on a model of parity progression components, for six quinquennial birth cohorts, United States

Cohort	U	Q
1916–20	0.320	0.0453
1921–25	0.419	0.0593
1926–30	0.596	0.0807
1931–35	0.565	0.0793
1936–40	0.462	0.0653
1941–45	0.308	0.0433

Considerable speculation is involved in estimates of comparable Q values for cohorts prior to cc. 1916–20. Concerning the low number of unintended births per woman for cc. 1916–20 relative to subsequent cohorts, we elsewhere argued that this was explained by a combination of lower exposure to risk (because of a higher age at second birth) and greater diligence in use of (premodern) contraceptives, associated with the gravity of the consequences of failure in stringent circumstances (Ryder, 1982). Since both of these arguments would apply, a fortiori, to the predecessors of cc. 1916–20, that is, the cohorts of 1906–15, we assigned to them the Q value for 1916–20.

The question of an appropriate value for contraceptive failure, Q, for the cohorts of the nineteenth century hinges on whether there was an upward trend in regulatory efficacy. On the one hand, there may have been a slight improvement in contraceptive technology and some tendency to shift from less to more effective methods, perhaps associated with the increases in urbanization and education. On the other hand, it is not unlikely that the process of extension of the franchise of fertility regulation to ever larger proportions would have the consequence of reducing the efficacy of the average contraceptor. Rather than pretend to resolve the quandary, we selected the conservative assumption that the failure rate was unchanging for the nineteenth century cohorts, at a level corresponding to the average of the cohorts of 1916–40. (We excluded the cohorts of 1941–45, on whom we also had survey data, because they were already substantially participating in the contraceptive revolution.) To avoid introducing a gross discontinuity into the time series, we set the value of Q for the 1901–05 cohorts midway between that for 1896–1900 and that for 1906–10.

Finally there is the question of the effectiveness of fertility regulation for the two quinquennial cohort groups subsequent to cc. 1941–45. These cohorts were full participants in the contraceptive revolution. For instance, the 1940 cohort was age 30 in 1970. In that year, of contraceptive users under age 30, 49 percent were using the pill, whereas of users age 30 and over, only 21 percent were using the pill (Westoff and Ryder, 1977, Table II-3). We were also impressed by the evidence of greatly increased resort to sterilization, as well as to abortion. Our assumption, accordingly, was that the failure rate, Q, would be 25 percent lower for cc. 1946–50 than for cc. 1941–45, and another 25 percent lower for cc. 1951–55.

We hold no strong brief for these particular assumptions about appropriate values of Q and of X_i. After reporting on the results of applying the model, for which such estimates are required, we consider the robustness of the findings should the assumptions prove to be erroneous.

Intended and unintended births

The first results from the model are presented in Table 4. The first column repeats, for convenience, the time series of mean marital parity found in Table 1. The second column shows the mean parity that would have resulted had

TABLE 4 Intended and unintended fertility for US birth cohorts, 1867–1955

Cohort	Mean marital parity	Intended births	Unintended births	Intrinsic rate of natural increase	
				Observed	Intended[a]
1867–70	4.338	4.021	0.317	0.0106	0.0082
1871–75	4.068	3.721	0.347	0.0097	0.0068
1876–80	3.809	3.432	0.377	0.0086	0.0051
1881–85	3.547	3.139	0.408	0.0076	0.0039
1886–90	3.342	2.909	0.432	0.0066	0.0018
1891–95	3.107	2.646	0.461	0.0053	−0.0006
1896–1900	2.842	2.349	0.493	0.0028	−0.0042
1901–05	2.574	2.155	0.419	0.0001	−0.0065
1906–10	2.402	2.093	0.309	−0.0017	−0.0068
1911–15	2.446	2.140	0.307	0.0000	−0.0048
1916–20	2.649	2.329	0.320	0.0038	−0.0008
1921–25	2.936	2.517	0.419	0.0082	0.0027
1926–30	3.160	2.564	0.596	0.0119	0.0042
1931–35	3.280	2.715	0.565	0.0143	0.0072
1936–40	3.044	2.582	0.462	0.0120	0.0055
1941–45	2.594	2.286	0.308	0.0054	0.0003
1946–50	2.230	1.992	0.238	−0.0005	−0.0050
1951–55	2.094	1.917	0.177	−0.0038	−0.0073

[a] Based on intended births.

intentions to control fertility been successfully executed. The third column, mean number of unintended births per woman, is the difference between the first and second column values. (As explained above, the values for cc.1916–45 have been estimated from the National Fertility Studies.) The numbers of unintended births are a direct reflection of the assumptions made about the probability of failure, for cohorts prior to 1916 and subsequent to 1945. In brief, whatever usefulness the model may have lies elsewhere. It is worth noting, however, that the number of unintended births tends to vary inversely with the number of intended births, ceteris paribus. That is the reason for the rise in unintended births from cc.1867–70 to cc.1896–1900 (despite the underlying assumption that the probability of failure was fixed over that part of the series). The reason is that the greater the number of births intended, the shorter the length of exposure to risk of subsequent unintended births. (It also follows that the not uncommon index, proportion of births unwanted, is intrinsically unsatisfactory.)

In the fourth column of Table 4, we present the time series of intrinsic rates of observed natural increase for birth cohorts in the United States. (The survival component of that calculation is predicated on a linear rise in the expectation of life at birth for females, from age 50 years for the cohorts of 1867–70 to age 75 years for the cohorts of 1951–55. Even a considerable departure from that pattern of change would have relatively slight consequences for the calculation reported here.) To the best of the author's knowledge, no previous attempt has been made to estimate the intrinsic rate of natural increase for such a lengthy series of US data.

Contrary to the impression one might gain from consideration solely of fertility, the intrinsic rate of natural increase was lower in the first half of the series than in the second half—simply because of a rise in the survival component of net reproduction (from approximately 72 percent survival from birth to the average age of reproduction to approximately 98 percent).

Consider now the consequences for our understanding of the sources of population growth when the progression components model is applied. In the final column of Table 4, the calculation of the intrinsic rate of natural increase is repeated, but with observed fertility replaced by intended fertility. Understandably the values are systematically lower, but it is somewhat surprising to see that nearly one-half of them are negative. For the conventional intrinsic rate of natural increase, the average for the series is 5.6 per thousand per annum; for the "intended" intrinsic rate, the average is somewhat less than 0.5 per thousand per annum. Given the suspicion that levels of unintended fertility are understated, and given the conservative nature of the assumption that fertility regulation was as effective in the nineteenth century as it was in most of this century, it seems safe to assert that, were it not for unintended births, the intrinsic rate of natural increase would have been, on average, negative for the United States over the past 80 years.

To further illustrate the kind of results yielded by the progression components model, we show in Table 5 an abbreviated description of parity distributions, observed and intended, for four selected cohorts. Between the initial and final cohort groups, the two other cohorts are the ones at the pivots of the interim fluctuation in fertility. (The parameters for the omitted cohorts in each case follow approximately linear paths between the values shown in Table 5.) The chosen parameters are the proportions with fewer than two, two, and more than two births, together with the mean number of additional births for those with at least three.

TABLE 5 Observed and intended parity distributions for four selected quinquennial cohorts, United States

Cohort	Percent in parities			Mean additional births (for those with at least 3)
	0/1	2	3+	
	Observed			
1867–70	24	13	63	3.04
1901–05	35	22	43	1.94
1931–35	16	22	62	1.46
1951–55	29	39	32	0.52
	Intended			
1867–70	30	13	57	3.20
1901–05	53	19	28	2.29
1931–35	23	32	45	1.32
1951–55	37	43	20	0.48

It is immediately evident that a quite different message is purveyed in the set of "intended" data from that in the observed series. Thus the proportions with fewer than two are markedly higher for the "intended" than for the observed. The proportion with two births in the observed series shows a rise to a saddle, between cc.1901–05 and cc.1931–35, the so-called baby boom; in the "intended" series, on the contrary, there is not only an increase in the proportion intending two during that interval, but a larger increase than in the other pieces of the time series. With respect to the final column of the table, it is of considerable interest that intended fertility (in the higher reaches of the parity distribution) was higher than observed fertility in the first half of the series, but lower in the second half.

Earlier we were disparaging the analytic utility of the observed parity distribution. We would not find similar fault with the intended parity distribution. To the extent that one is willing to give credence to the assumptions underlying the model, the intended parity distribution is a meaningful expression of the structure of reproductive intentions (conditional upon continuation of exposure to risk). We feel that the data in the lower panel of Table 5 represent the appropriate explanandum for a theory of fertility decline in the United States.

Progression and parity-specific selection

The last question to be addressed concerns the pattern of progression ratios by parity, within cohorts. In truth, this was the starting point for the work reported herein. In a recent book (reviewed for this journal by the author), John Hobcraft observed that ". . . at least at the societal level, subsequent fertility behaviour does not depend very much on achieved parity beyond the second birth" (Hobcraft, 1985, pp. 78–79). In his conclusion he said: ". . . potentially the most important and far-reaching findings are those appearing in the more elaborate analyses of birth intervals. These analyses throw considerable doubt upon the parity-specific nature of controlled fertility behaviour. If these findings are confirmed, it will involve the profession in a major rethinking of its basic assumptions." His position corresponds with that taken by the authors of the most important of the analyses of birth intervals to which he refers (Rodriguez et al., 1984).

We agree with the empirical observation (see Figure 2 above, for example); we also agree with the bearing this would have on fertility theory, if observed parity progression ratios were trustworthy indicators of what the theorists seem to be talking about. The purpose in developing the progression components model was to reconsider the question from a vantage point somewhat closer to the structure of reproductive intentions than the observed progression ratios can bring us.

The theoretical proposition at the heart of the matter is that the driving force behind fertility decline is the adoption of a small-family norm by progressively larger segments of the population. This would be expected to manifest itself in two ways in the time series of the progression sequence. First, one would anticipate a growing trough in the neighborhood of parity two. Second, one would look for a rising probability of progression for successive parities higher than two. That should occur as a reflection of a selection process that would eliminate from membership in the higher parities those with low reproductive aspirations, so that the remainder would be selected for pronatalism. The observed data show nothing of the sort.

In Table 6, we show the values of the observed progression ratios, R_i, and the intention probabilities, I_i, developed by way of the progression components model. To repeat, I_i is the proportion of those exposed to risk in parity i who intend at least one more birth. The values are shown for the average of the first nine and last nine cohort groups.

In both the earlier and the later sequences of Rs, in Table 6, the similarity of values beyond parity two is evident. The situation is otherwise with respect to the sequence of intention probabilities, I_i: these clearly rise with advancing parity. The one small exception to this tendency, I_6 in the earlier sequence, may be a reflection of the previously noted tendency for even-numbered parities to be depressed by misreporting. To quantify the contrast between the slopes of R and I, we calculated the change from each parity to the next higher, over parities three and higher, in R and in I, for the 18 cohort groups in the time series. The median inter-parity change in R was -0.8 percent; the same for I was $+5.2$ percent. Thus the intention values display the pattern of change expected on theoretical grounds, even though the phenomenon is concealed in the observed progression ratios.

With respect to evidence for a trough in the progression sequence in the neighborhood of parity two or parity three, it is clear that the observed progression ratios are lower than their predecessors but not lower than their successors. In the sequence of I values, on the contrary, the lowest value in the earlier set is clearly that for parity two, and the lowest two values in the later set are those for parities two and three. Moreover, as would be expected on theoretical grounds, the depth of the trough is greater in the later sequence.

TABLE 6 Observed (*R*) and intended (*I*) parity progression ratios for cohorts 1867–1910 and 1911–55, United States

	Cohorts 1867–1910		Cohorts 1911–55	
Parity	***R***	***I***	***R***	***I***
0	0.8475	0.7552	0.9071	0.8969
1	0.7971	0.7994	0.8281	0.8194
2	0.7372	0.7414	0.6162	0.5283
3	0.7245	0.7945	0.5354	0.5221
4	0.7222	0.8227	0.5127	0.5759
5	0.7361	0.8729	0.5308	0.6652
6	0.7149	0.8515	0.5434	0.6856
7+	0.6733	0.9328	0.5555	0.7642

In summary, then, the probabilities of intention by parity resulting from application of the progression components model go a long way toward dispelling the mystery of parity-independence in the observed sequence of progression ratios. From the evidence of Table 6, there is no basis for rejecting the idea that the structure of intentions plays an important role in determining fertility as a function of parity.

With respect to the overall pattern of change over time, parity by parity, we note that the percentage declines in the intention probabilities from the earlier to the later set are 29, 34, 30, 24, 19, 18 for parities two, three, . . . , seven and more. Although there is a tendency for a larger relative decline in the lower parities in this sequence, it deserves emphasis that the process of decline is well dispersed among representatives of all parities. Accordingly, fertility theory is insufficient as an explanation of that process to the extent that it emphasizes the adoption of a small-family norm. Similarly, the recent review of World Fertility Survey findings suggests that fertility decline in developing countries is not confined solely to the behavior of particular parities (Cleland and Hobcraft, 1985).

Sensitivity of the model to assumptions about exposure and success

The credibility of the inferences based on consideration of the estimates of intended fertility reported above obviously depends on the robustness of the results in relation to assumptions about exposure (the values posited for X_i) and about success in regulating fertility (the values posited for S_i). To find out what would happen if we used a different assumption about exposure, we recalculated the results for the four cohorts examined in Table 5, assuming a value of 96 percent or 98 percent for X_0, rather than the 97 percent used above. Although that may seem a small change, there is in fact little leeway for the value of X_0. Moreover, the shape of the X function by parity guarantees increasing differences with advancing parity; any larger range for X_0 would have yielded impossible outcomes in the highest parities.

For the four cohort groups considered, the result is a 2 percent change in mean intended parity—lower when X_0 is 98 percent and higher when X_0 is 96 percent. The reason for what may appear to be an anomalous direction of effect is that unintended fertility varies directly with exposure, and thus intended fertility varies inversely with it. Of more interest is what happens to the parity pattern of intention probabilities, I, when exposure is modified. From parity to parity, the relative change is of the order of one percent. In brief, the argument of the previous section is virtually independent of the exposure assumption.

The assumptions concerning success in regulating fertility require careful scrutiny, because they are likely to have a larger effect on the outcome than assumptions about exposure. Two separate questions are involved. In the first place, our assumptions about the rate of failure for cohorts prior to and subsequent to those for which we have survey evidence, cc.1916–45, may be incorrect. To exemplify the problem, we postulate a failure rate, Q, of 0.0807 (the highest observed in the surveyed cohorts) for cc.1886–90, some 22 percent higher than that postulated for nineteenth century cohorts in the model. As would be expected, there is a substantial rise in the mean number of unintended births, from 0.43 to 0.57. As already intimated, the values reported for unintended births depend directly on the quality of what are no more than speculations about the level of failure for unsurveyed cohorts.

Of more concern is the possible impact of an erroneous assumption about the failure level on the parity sequence of intention probabilities, I. For cc.1886–90, at the higher failure rate, the value of I_0 is reduced by 5 percent; otherwise changes parity by parity are of the order of one percent. In brief, the reported pattern is essentially undisturbed.

We carried out a similar exercise for cc.1951–55, assuming that the failure rate for these cohorts was the same as that for cc.1941–45 ($Q = 0.0433$) rather than the posited $Q = 0.0244$. The consequences for the sequence of intention probabilities, by parity, are as follows (percent change): -2, -4, -30, -5, $+16$, $+12$, $+8$, and $+4$. The source of the very large difference for parity two (30 percent smaller) is the very large change in observed progression ratios, from $R_1 = 0.76$ to $R_2 = 0.41$. Evidently the results for the most recent cohorts are highly sensitive to the assumption about the failure rate; they should not be taken too seriously. In this situation there is, however, a consoling thought. Assuming that fertility surveys are continued in the United States, direct estimates of the failure rate can eventually be made for these cohorts, as for their predecessors.

Apart from the overall rate of failure assumed, there is the question of the pattern of failure by parity. To examine this, we chose cc.1921–25, for which the level of unintended fertility is soundly based on survey data. Two recalculations were undertaken. In the first place, we reduced the slope of the success function, vis-à-vis parity, to zero; in the second place, we steepened that slope to a like degree. In both instances, the mean number of unintended births was held fixed at the level indicated in survey data. The consequence for the sequence of intention probabilities by parity was very small. At their most extreme, the values for I_3 and I_4 were reduced by 5 percent with a zero slope, and raised by 5 percent with a steeper slope. The reported pattern of parity-specific intention is insensitive to the shape of the success function.

Our general conclusion from these tests is that the results are quite robust in response to plausible modifications of assumptions about exposure and regulatory success. The estimates of the proportions intending no children, for the nineteenth century cohorts, are rather untrustworthy, as are the numbers of unintended births per woman. With the sole exception of the most recent cohorts, the intention structure by parity seems to have been well delineated by the model.

Conclusion

This article has exploited a valuable set of data: completed parity distributions for the cohorts of 1867 through 1955, for women in the United States. We have attempted by various projection procedures to bring the series up to the present. The history is a familiar one: long-term decline, with a transitory reversal about a generation ago. As for the most recent experience, the cohort total fertility rate has declined from 3.20 for cc.1931–35 to 1.92 for cc.1951–55. Evidence from the birth histories collected in the June 1985 Current Population Survey gives a clear indication that fertility has for the present stabilized at the (low) level reported for cc.1951–55.

In considering the time series of aggregate fertility, we have also introduced a measure of mean sibship size in order to represent fertility from the standpoint of the children rather than, as is conventional, from the standpoint of the parents. Historically, there has been a much more substantial decline, both absolutely and relatively, in this measure than in the total fertility rate. Moreover, the pronounced local maximum in the latter (known as the baby boom) is reduced to no more than an extended hiatus in the tendency of mean sibship size to decline.

Understanding of the determinants of the time series of cohort total fertility is enhanced by a consideration of the time series of progression ratios for successive parities. From this standpoint, the sources of the transitory reversal of fertility a generation ago are clearly identified as fluctuations in the progression ratios for parities one and two, coincident with an upward shift in the level of progression for parity zero. Even the concept of a fluctuation is brought into question, since the low values characteristic of those cohorts whose childbearing was concentrated in the 1930s, for parities one and two, were merely the culmination of a lengthy downward drift in those measures.

Progression ratios beyond parity two show characteristics quite unlike those for the lower parities. For each cohort they are approximately the same, parity by parity, and they have experienced almost monotonic decline throughout the recorded time span, except for a temporary saddle a generation ago. This decline in higher-parity progression is a phenomenon on which fertility theorists have been silent.

In order to come to closer grips with theoretical issues, we have developed a model of the components of parity progression. The model is designed to elucidate the separate roles played by three factors in determining the observed progression sequence: exposure to risk (which declines parity by parity); success in terminating fertility (which rises parity by parity); and the structure of reproductive intentions.

Using evidence for a population with essentially unregulated fertility, and data from the series of National Fertility Studies concerning the incidence of unintended births, we have developed assumptions about exposure to risk and about the time series of success in terminating fertility that permit us to infer the structure of reproductive intentions. The results appear to be reasonably invariant to alternative assumptions.

The model is first used to contrast the values of the fundamental demographic parameter, the intrinsic rate of natural increase, for all births and then for intended births. Our conclusion is that population growth in the United States over the past eight decades has had three components: immigration, mortality decline, and unintended births.

The pattern of change in the intended parity distribution shows a systematic monotonic rise in the proportion intending two births from 13 percent for the earliest cohort to 43 percent for the latest cohort, without a transitory deviation. This transformation has been accompanied by an equally systematic decline in the mean number of intended additional births for those with at least three births, from 3.2 at the beginning of the series to 0.5 at the end of the series. While the growth in the intention to have two children is consonant with expectations from fertility theory, the decline in intended parity for those with families already large (at least by today's standards) is not.

The final application of the model is to test the proposition that progression is insensitive to parity, beyond parity two. While this is true for observed progression, it is not true for intended progression. There is clear evidence of the theoretically expected selection process, as well as of an evolving concentration on small families.

Although the speculative elements in our model of the components of progression merit skeptical scrutiny, we feel confident that the principles of its construction are sound, and that the pattern of findings derived from its application makes sense. Although theorists have been remiss in ignoring the pervasiveness of fertility decline over the entire range of parities, their emphasis on the relevance of the structure of reproductive intentions does not, by the present account, seem to be misplaced.

Appendix

Derivation of the algorithm $F^*_{i+1} = F_i\left(X_i - \frac{X_i - R_i}{S_i}\right)$,

where F^*_{i+1} = the number of intended $i+1^{\text{th}}$ births

F_i = the total number of i^{th} births

X_i = the proportion exposed to risk of an $i+1^{\text{th}}$ birth

S_i = the probability of success in terminating childbearing with the i^{th} birth

R_i = the i^{th} parity progression ratio, F_{i+1}/F_i .

From Figure 3 above

$$F_1 = X_0 I_0 + X_0 (1 - I_0)(1 - S_0)$$
$$= X_0 I_0 + X_0 - X_0 I_0 - X_0 S_0 + X_0 I_0 S_0$$
$$= X_0 I_0 S_0 + X_0 (1 - S_0) \quad .$$

But $F_1 = R_0$ and $X_0 I_0 = F^*_1$.

Thus $R_0 = F^*_1 S_0 + X_0 (1 - S_0)$

and $F^*_1 = X_0 - \frac{X_0 - R_0}{S_0}$. Note that $F_0 = 1$.

Again from Figure 3 above

$$F_2 = X_0 I_0 X_1 I_1 + X_0 I_0 X_1 (1-I_1)(1-S_1) + X_0 (1-I_0)(1-S_0) X_1 (1-S_1)$$
$$= X_0 I_0 X_1 I_1 + X_1(1-S_1)(X_0 I_0(1-I_1) + X_0(1-I_0)(1-S_0))$$
$$= X_0 I_0 X_1 I_1 + X_1(1-S_1)(X_0 I_0 + X_0(1-I_0)(1-S_0) - X_0 I_0 I_1)$$
$$= X_0 I_0 X_1 I_1(1-(1-S_1)) + X_1(1-S_1)(X_0 I_0 + X_0(1-I_0)(1-S_0))$$

But $X_0 I_0 X_1 I_1 = F^*_2$ and $X_0 I_0 + X_0(1-I_0)(1-S_0) = F_1$ and $F_2 = F_1 R_1$

Thus $F_1 R_1 = F^*_2 S_1 + X_1(1-S_1)F_1$

$$\text{and } F^*_2 = F_1\left(X_1 - \frac{X_1 - R_1}{S_I}\right)$$

The comparable elaboration of F_3 leads to $F^*_3 = F_2\left(X_2 - \frac{X_2 - R_2}{S_2}\right)$, and so forth.

Furthermore, since $F^*_i = \prod_{j=1}^{i}(X_{j-1} I_{j-1})$,

$$F^*_{i+1}/F^*_i = X_i I_i, \quad \text{from which} \quad I_i = (F^*_{i+1}/F^*_i)/X_i$$

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