MARRIAGE AND DIVORCE IN THE U.S.: 
TESTING FOR COHORT SIZE EFFECTS 

by 
Diane J. Macunovich 

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Marriage and Divorce in the U.S.:
Testing for Cohort Size Effects

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Abstract: Although the common wisdom — based largely on Vital Statistics available only for the period through 1988 — is that marriage is fast becoming a defunct institution, evidence derived from the annual March CPS (1964-95) is used to demonstrate that there have been substantial changes in the long-term trends in marriage and divorce within the 20-24 age group since 1988. For males and females working fulltime year round, and for those with less than a completed college education, age-specific marriage rates appear to have stabilized — and divorce rates have even reversed their direction. This change in a longstanding trend appears to be a cohort-related, rather than a period, phenomenon, since it clearly began with the youngest age groups and is gradually spreading to older groups over time. In addition, the patterns show a clear relationship with male relative income, as hypothesized by Richard Easterlin. However, the relationship between male relative income and relative cohort size appears to have weakened substantially in recent years, and a ‘marriage squeeze’ variable based directly on relative cohort size produces counterintuitive signs on a number of variables in a full model, and loses its significance when entered along with a relative income variable. Thus it is not clear that marriage and divorce rates can be forecast directly on the basis of relative cohort size: their relationship appears to be closest with male relative income.
A. Introduction

After more than a decade of strong increases, the cohort size of young people in their early twenties began to decline in the early 1980s, and their relative cohort size (their numbers relative to numbers in their parents' cohorts) began to decline in the mid-1980s. Thus the period since the mid 1980s should provide evidence -- if such there is -- of hypothesized cohort size effects on marriage and divorce rates. Both the 'marriage squeeze' hypothesis of Glick, Beresford and Heer (1963) and the 'relative income' hypothesis of Easterlin (1980), suggest that declining cohort size should reverse the trends which we have observed in these rates in the decades following 1970.

These hypotheses are difficult to test using data from Vital Statistics, however, since the latest data available in the series on marriage and divorce are the advance statistics for 1988: this is the only comprehensive source of U.S. data on age-specific (as opposed to crude) rates. It presents the patterns everyone has come to know: declining rates of marriage, and increasing incidence of divorce, as shown for the age group 20-24 in Figure 1. For women in this age group, the marriage rate declined from 234.2 per 1000 in 1970 to only 104.9 in 1987 -- and then rose slightly to 105.5 in 1988. For men the decline was even more precipitous: from 205.7 to only 73.6 in 1987, with again a slight rise to 73.7 in 1988. Not surprisingly, during this same period the median age of all brides increased steadily from 21.7 to 26.1 and the mean age increased from 25.1 to 28.8 (Appendix). For first-time brides, the spread between mean and median was not so large, but the steady increase in both was the same: median age increased from 20.6 to 23.7, and the mean increased from 21.6
unmarried

(only 1970 age-specific divorce rate available prior to 1980)

Fig 1: Marriage & Divorce Rates, age 20-24
to 24.6. Prior to 1970, there had actually been a marked decline in these ages for men and for women.

Divorce statistics in the same source paint a gloomy picture: the percentage of all married men aged 20-24 who experienced divorce in 1988 was 5.59, while that for women was 4.63 -- both up from about 3.35% in 1970. The pattern in these statistics is not as clear as that in the marriage statistics, however: the percentages for women were relatively stable from 1985-88.

The only other ‘official’ source of marriage and divorce statistics is the Annual Summary of Births, Marriages, Divorces, and Deaths: United States, 1994, which indicates the following substantial decline in marriage rates between 1987 and 1994:

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<tr>
<td>marriages</td>
<td>9.9</td>
<td>9.8</td>
<td>9.7</td>
<td>9.8</td>
<td>9.4</td>
<td>9.3</td>
<td>9.0</td>
<td>9.1</td>
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<tr>
<td>divorces</td>
<td>4.8</td>
<td>4.8</td>
<td>4.7</td>
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¹ provisional data

The figures in this source, however are not age-specific: they simply represent a crude count per 1,000 total U.S. population. Since the median age at first marriage falls somewhere in the range 23-25, for both men and women, and since the ‘baby bust’ cohorts have been entering those prime marriage years while the baby boom cohorts have been entering their 40s, counts per 1,000 in the total population can present a very misleading picture of the trend in age-specific rates.

The aim of this paper is to test the two primary cohort size hypotheses regarding
Fig. 2a: Marriage Squeeze Variable (population 20-24 relative to population 18-22)

Fig. 2b: Male Relative Income Variable (earnings of young males/income of older families)
marriage and divorce rates, and in order to do so it was necessary first to determine what has been happening to age specific rates in the period since 1988, and then to analyze trends since the 1960s. This has been accomplished by identifying a time series of age-specific rates from another source: the annual March Current Population Survey (CPS) for the years 1964-95. Individuals report their marital status in this survey, along with other socio-economic data, and it is a carefully controlled random sample which is designed to give accurate estimates of characteristics of the total U.S. population. The individual records in the CPS data have been aggregated in this study to provide proportions married and divorced/separated by age, for ages 20-34. These data have also been analyzed by education and work status within age groups.

Models were then estimated, using the ‘marriage squeeze’ and ‘relative income’ measures presented in Figure 2. The derivation of these measures is explained later in this paper (section C.2): for now it is important only to note the marked changes which have occurred in both of these variables since about 1985 — in the one case after almost 15 years of monotonic decline, and in the other after a similar period of monotonic increase. These two variables have been almost perfect inverses of each other in the period since the mid 1970s.

Section B describes some of the patterns indicated by the data, while section C examines the patterns in an attempt to determine whether they are cohort or period related, and section D presents results from estimating a model of marriage rates for males aged 20-24. Section E discusses trends in enrollment and fulltime work, and the relationship between relative cohort size and male relative income, and Section F summarizes the paper’s main conclusions.

B. Proportions Married and Divorced, 1964-95
Fig. 3a: Proportions Married (3 yr. moving ave.), Fulltime Workers Aged 20-22

Fig. 3b: Proportions Divorced (3 yr. moving ave.), Ever-Married Men & Women Working Fulltime Aged 20-22

Fig. 3c: Proportions Married (3 yr. moving ave.), Fulltime Workers Aged 20-24

Fig. 3d: Proportions Divorced (3 yr. moving ave.), Ever-Married Men & Women Working Fulltime Aged 20-24
The newly derived CPS data presented here do not describe age-specific marriage and divorce rates, as in the published Vital Statistics. Rather, they describe proportions married and divorced each year: a stock measure, rather than a flow. This is to some extent a more informative measure, since it circumvents problems associated with minor changes in the timing of marriage and divorce within the 15-24 age group. The proportion married is simply that: the number in a group who are currently married, with spouse present, divided by the total population in the group. The proportion divorced, however, is adjusted since it is, of course, dependent on the proportion married: if the proportion married is declining rapidly over time, the unadjusted proportion divorced may decline as well -- simply because there are fewer people available to move into the divorced state. Since the purpose here is to give some idea of what has been happening to the divorce rate (calculated as the number of divorces per 1,000 married persons in a particular group), I have divided the proportion divorced/separated in each year by the proportion who were married in the group in the prior year.

Figure 3 illustrates the proportions married and divorced/separated in the period 1964-1995, among males and females aged 20-22 and 20-24 who are working fulltime year round. In Figures 3a and 3c, we can see that a long-term secular decline in the proportions married, which proceeded unabated for over fifteen years from the early 1970s, appears to have been

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1 It should be noted that these proportions, although cumulative measures, may still be analyzed in cohort terms since they simply describe the proportion of a particular age cohort who have married by a given year -- and remain married -- and of those who have married, the proportion who have divorced: these are comparable to cohort measures of cumulative fertility.

2 That is, the adjusted proportion of 23 year old males who are divorced in year t is derived by dividing the observed proportion of 23 year old males who are divorced in year t by the observed proportion of 22 year old males who were married in year t-1. For disaggregated statistics (within education and work status groups), this calculation is done within the respective group.

3 All proportions married and divorced are presented in three year moving averages.
Fig. 4a: All Men Aged 20-22, Prop. Married, & Prop. Divorced (of Ever-Married Men)

Fig. 4b: All Men Aged 20-24, Prop. Married, & Prop. Divorced (of Ever-Married Men)
arrested in the late 1980s, and that this phenomenon appears to be stronger among males than among females. These patterns are almost eerily similar to that of the two cohort-related variables in Figure 2. The proportions divorced (Figures 3b and 3d) show strong signs of reversal — especially among the males: in this case, the long increase in the proportions divorced (up to levels of about 20% of those previously married) since the early 1970s has begun to give way to marked declines.

Figure 4 illustrates these same proportions among all males aged 20-22 and 20-24 (regardless of work status). Here, the proportions divorced have plateaued at the same time that the proportions for fulltime workers in Figures 3b and 3d declined, but the ‘bottoming out’ of marriage rates seen in Figures 3a and 3c for fulltime workers since the late 1980s shows up here only as an attenuation of the decline in rates for males aged 20-22 (Figure 4a). Little if any effect can be detected in the marriage rate for males aged 20-24 (Figure 4b).

Figure 5 presents marriage rates for all males aged 20-22 and 20-24 (regardless of work status) by level of completed education. Here we can see that the ‘bottoming out’ of marriage rates shows up to some extent among individuals at nearly all levels of education. It appears, then, that the absence of a strong ‘bottoming out’ in the marriage rate for all men aged 20-24 is due to compositional effects: changes in the proportions working fulltime year round, and in the mix of educational levels, which can mask within-group changes at the aggregate level. Thus it is important to consider both the education/work status composition of young males and females -- and the factors which affect this composition -- as well as the factors directly affecting propensity to marry, in attempting to understand movements in age-specific marriage and divorce rates.
Fig. 5a: Proportions Married, All Males Aged 20-22, by education level

Fig. 5b: Proportions Married, All Males Aged 20-24, by education level
C. Cohort or Period Effect?

Andrew Cherlin (1992), in writing on marriage and divorce in the U.S., acknowledges the possibility of cohort effects in the trends we have observed in the post WWII period, but comes down fairly firmly on the side of period effects: "something in the air". His reasoning relies on the fact that many age specific rates appear to have moved in parallel when major transitions were occurring.

I. What Has Been Happening at the Cohort Level?

The changes in marriage rate trends identified in the previous section have so far shown up primarily in the youngest age group: those up to and including age 22 and in some cases through about age 24. This is evident in Figure 6, which illustrates proportions married in the 20-22, 20-24, 25-29 and 30-34 age groups, for fulltime year round workers. The bottoming out of rates occurred first among those 20-22, in about 1988, and then appeared in the 20-24 group in about 1990. And there is some hint that this bottoming out began to show up in the 25-29 group in 1994. The 30-34 age group, on the other hand, exhibits monotonically decreasing proportions married right up through 1995 -- with even an increase in the rate of decline in the 1990s.

In addition, as mentioned earlier, Figure 3 indicates that the changes in proportions married and divorced appear to have manifested themselves more strongly with respect to males in these age groups, than females. Both this differential male-female effect, and the fact that the changes began first with the youngest cohorts, suggest strongly that cohort effects may be
Fig.6: Proportions Married, All Men Working Fulltime by Age Group
operating here - and that these cohort effects are strongest on young males. Some sub-groups of young males who turned 20-22 in the first few years of the 1990s chose to marry at increasing rates, and selected partners in various (for the most part younger) female cohorts so that their rates increased, as well, but in a less clear pattern. (That is, a twenty year old male may select a partner his own age, or in any one of a number of other cohorts. To the extent that he chooses from a younger cohort in year $t$, the proportion of women aged 20-22 who are married will only show an increase in years after $t$, and this effect will be diluted to the extent that males choose partners in various age groups. Because the observed effects appear to have been strongest (or to have occurred first) for young males, the following discussion focuses primarily on them, and on factors affecting their propensity to marry.

2. Relative Cohort Size Theories

If it is a cohort effect, what has been happening to young males at the cohort level to cause such a change? The most obvious explanation relates to cohort size, since the relative size of cohorts reaching age 20-24 (that is, the number of persons aged 20-24 relative to the number in their ‘parental cohort’ aged, say, 50-54) began declining in 1984, after a steady rise since the 1960s. The two most widely cited theories which suggest a relationship between relative cohort size and marriage rates, are the ‘marriage squeeze’ theory of Glick, Beresford and Heer (1963) and Richard Easterlin’s (1980) relative cohort size hypothesis.

Glick et al. suggest that because women tend to look for marriage partners among males approximately two years older than themselves, periods of mismatch between any given cohort and the cohort two years older -- such as those during the baby boom -- will affect marriage
rates due to imbalances between the demand for and supply of mates. This suggests that cohorts born during the first half of a baby boom will exhibit declining marriage rates for females, while cohorts born during the latter half of such a boom will exhibit increasing marriage rates for females, because of this tendency for women to marry older men.

It is not entirely clear, however, how this variable is expected to affect male marriage rates: arguments of symmetry would imply that the marriage rates of young men should move inversely to those of young women in the same birth cohort, since a period of abundant mates for women will coincide with a period of scarcity for young men, and vice versa. Guttentag and Secord (1983) suggest in their 'too many women' hypothesis that this will not be the case, since young men in periods when potential mates are in abundance will not feel compelled to marry: thus male and female marriage rates will move together, declining(increasing) in periods of increasing(declining) relative cohort size. Figure 2a presented the pattern of a 'marriage squeeze' measure (population aged 20-24 relative to population aged 18-22) for the period since the 1960s, which will be used to test this theory in Section D.

Easterlin's theory postulates that as the relative size of cohorts reaching age 20-24 rose in the period up to 1985, their large numbers in the labor market relative to the number of older workers caused their earnings in the aggregate, relative to those of older males, to fall, since younger and older workers are not perfect substitutes. And this, in turn, should have caused their earnings in the aggregate to fall relative to their material expectations, assuming that their expectations are a function of the income of the families in which they were raised.  

It is important to note here that the effects of cohort size discussed by Easterlin do not require absolute changes in the earnings of young males — only changes relative to their expectations. In this sense, the generally poor market conditions in the late 1980s and early 1990s which have led to severely diminished expectations, can actually enhance the effects of cohort size — and rising relative income — predicted by Easterlin.
Easterlin's hypothesis with regard to the effects of relative cohort size on young males' relative earnings has been supported by the findings of numerous researchers, including Welch (1979) and Berger (1984, 1985 and 1989). The pattern of young males' earnings relative to the income of older families with children was presented in Figure 2b. It has been calculated using, in the numerator, the real average annual earnings of all currently (at time t) unenrolled males in their first five years of work experience (calculated as years of completed schooling minus age minus six), and in the denominator the fifth lag (that is, the observed value five years earlier at time t-5) of the real average total income of all families with children, with head (of either sex) aged 45-54.

The fifth lag is used in the denominator in order to approximate family income while the young men were still living in their parents' homes. Unenrolled young males in their first five years of work experience are used in the numerator, rather than young males aged 20-24, in order to incorporate the changing levels of educational attainment which are assumed to result from changing male relative income, as young males attempt to close the gap between earnings and expectations by increasing their level of completed education. As Figure 2b indicates, this relative income measure began to increase after 1985 -- shortly before the changes we have observed in the proportions married and divorced -- and then resumed declining in 1992. As pointed out earlier, this pattern seems to be mirrored, in a muted fashion, in the marriage rates of young men and women aged 20-22, in Figure 3a.

Easterlin hypothesized that changes in male relative income would bring about a series of adjustments as young males attempted to improve their standing relative to their material expectations. If a young male's income is low relative to his material expectations, he is more
likely to postpone marriage and family formation -- or, if he does marry, he will be more likely to experience marital difficulties due to the stresses induced by financial difficulties. The converse would be true when male relative income is high.

Similarly, Easterlin's relative income hypothesis suggests that young males who anticipate a decline in the earnings they can achieve as high school graduates relative to their expectations, are more likely to invest in higher education in order to improve their earnings relative to those expectations. This effect is documented for the period 1964-1991 using time series data in Macunovich (1993b), although it is offset by a negative effect of relative cohort size on the return to a college education, as documented for example by Welch (1979) and Berger (1989). Achieving a higher level of education than one's father is an obvious adjustment which can be made by young males when their potential earnings slip relative to those of their fathers. Thus one might expect to see the relative income effect on enrollments exhibited even more strongly in cross-section data, than in time series, where the tendency to enroll at higher rates due to relative income effects will be countered by a declining college wage premium with increasing enrollments.

3. College Enrollment as a Deterrent to Marriage

Cherlin (1991) refers to Easterlin's cohort size theory in his analysis of marriage patterns, but places little weight on it largely because of results reported in MacDonald and Rindfuss (1981), which "provide no support for Easterlin's hypothesis that marriage will occur earlier when young men judge their economic prospects favorably with respect to their parents' income". This MacDonald and Rindfuss result bears closer examination, however. Their
analysis made use of a detailed cross section describing Wisconsin 1957 male high school graduates. This dataset was unusual in the depth of its information: it contained Social Security earnings records for the young males, together with income tax reports for their parents and follow-up information through 1975. A year-by-year logit analysis was conducted, of the probability of marriage among these young males, in the years 1958-1965. The results show a strong positive effect of the young males' own income, on their propensity to marry, but little or no effect of the parents' income. MacDonald and Rindfuss interpret this as indicating that the relative income concept is not only erroneous, but "obfuscates important effects of young men's current earnings".

However, one of the strongest effects identified by MacDonald and Rindfuss in their analysis, is that of current enrollment in school: "young men's earnings and the time spent in schooling to increase them were found to be important influences on marriage timing. Additional schooling had little effect net of the time it absorbed". That is, they found that being in school had a significant deterrent effect on the probability of marriage -- but that additional schooling on its own, after controlling for the time taken to acquire it, had no significant impact on marriage probabilities.

This finding of MacDonald and Rindfuss is extremely significant in that it appears to have ignored the potential connection between relative income and enrollment rates. If enrollment rates as well as marriage rates are a function of relative income, then their model produced only a partial estimate of the effect of relative income on marriage, ignoring the effect which operated through enrollment rates.

That is, the relative income model suggests that young males, in making a decision
regarding higher education, will compare their potential earnings as high school graduates to those of their fathers, and will assess the potential returns to education relative to the size of the ‘gap’ between these two. In this context, the males in the MacDonald and Rindfuss sample who enrolled in college would be composed largely of those who had judged their potential high school earnings to be deficient relative to their aspirations. In this sense, then, their sample of wage earners was actually truncated, and truncated in a way which would have removed just those young males whose parental earnings were ‘high’ and thus caused postponement of marriage. The postponement which Easterlin would have identified as due to relative income effects, in the MacDonald and Rindfuss sample, showed up as simple college enrollment -- which did, indeed, have a strong negative effect on propensity to marry. Thus it seems possible that a re-estimation of the MacDonald and Rindfuss equation in a two-stage model with college enrollment as a function of relative income, would have shown support for Easterlin’s relative income hypothesis, and hence for cohort size effects on male marriage rates.

D. Estimating a Model of Male Marriage Rates

In this section, a model is estimated to test the hypothesis that marriage rates of males aged 20-24 are a function of relative cohort size. A ‘marriage squeeze’ variable -- the number of persons aged 20-24 at time t relative to the number aged 18-22 at time t -- is used in order to test the Glick et al. hypothesis, and a male relative income variable is used to test the Easterlin hypothesis. The relative income variable used is that presented in Figure 2b: it is expected to have a positive effect on young men’s marriage rates. The ‘marriage squeeze’
variable as formulated here and presented in Figure 2a would be expected to demonstrate a negative effect on the marriage rates of males aged 20-24, if the ease of finding a suitable partner -- which is inversely related to the size of one's own cohort relative to the size of the cohort two years younger -- increases marriage rates. If, however, the 'too many women' hypothesis holds, then this variable will exert a positive effect on young men's marriage rates.

Several other variables are needed as well, to encompass the full range of effects hypothesized to operate on male marriage rates. Becker (1981) and others have hypothesized that increases in women's real wages act to reduce a woman's economic need to marry in the first place, and provide a source of support which makes divorce more feasible once she has married. In addition, the temporary deterrent effect of enrollment rates which MacDonald and Rindfuss identified must be controlled for. And finally, Thomas (1927) and Kirk (1960) provided strong evidence that marriage is pro-cyclical: thus we should see a negative effect of the male unemployment rate. This would be consistent with the results of a recent study (Oppenheimer et al, 1993) which finds that young males' propensity to marry is strongly affected by their labor market experiences.5

Table 2 presents the results of regressions testing these two hypotheses, with the logistic transformation of the marriage rate for males aged 20-24, as reported in Vital Statistics, as the dependent variable. (Note that this variable is only available through 1988). Columns 1-4 of Table 2 present results testing the male relative income variable, in combination with the female real hourly wage, the proportions of males and females enrolled in college, and the male

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5 A time trend was tested along with all the other variables, but its coefficient was found to be insignificant and its presence had no significant effect on the estimated coefficients of other variables, apart from the 'marriage squeeze' variable as reported later in this section.
unemployment rate. The fit of this model is extremely good, as indicated in column 1, with an $R^2$ of .998 and a Durbin-Watson score of 2.5, and the second column indicates that these results are unaffected when the numerator and denominator of the relative income term are entered separately. The absolute values of the coefficients on the numerator and denominator are not significantly different from each other, even at the 1% level, thus supporting the use of the relative income ratio.

Because the trend being fitted in Table 2 is a relatively smooth curve, any time series fit is suspect. As a result, the model has also been fitted using first differences of these same variables: the results of this estimation are presented in columns (3) and (4) of Table 2. Even in differenced form, each of the models (with relative income entered as a ratio, and with numerator and denominator entered separately) has an $R^2$ of over .75, and all of the variables are highly significant and carry the same signs as in the undifferenced results. In both sets of results we can see a significant positive effect of male relative income operating on marriage rates of these young males -- but this positive effect is countered by strong negative effects of the female wage, enrollment levels, and male unemployment shocks. In results not presented here, the effects of each of these variables appears to be quite stable across sub-periods, with no statistical difference between the coefficients, for example, in the two sub-periods 1969-1983 and 1979-1988.

The 'marriage squeeze’ variable was tested first on its own, in combination with the real female wage, male and female enrollment rates, and the male unemployment rate. This result is presented in column 5 of Table 2, where it can be seen that the variable is estimated to have a significant negative coefficient. The coefficients on the other variables are somewhat
nonsensical, however: the coefficient on the female wage is positive (although not significant), and the coefficient on male enrollments is also positive -- and significant -- contrary to accepted theory. In addition, when the same model is estimated with a time trend, the coefficient on the marriage squeeze variable becomes insignificant (column 6 results). Column 7 presents results with a model containing both the male relative income and marriage squeeze variables, and here again the coefficient on the marriage squeeze variable is insignificant -- while the coefficients on all other variables remain virtually unchanged from those estimated in column 1. This suggests that the relative income term is the most appropriate one for capturing effects which have been operating on male marriage rates during this period.

When the dependent variable in the analysis is changed to the proportion of 20-24 year olds married, rather than their marriage rate, very similar results to those presented in Table 2 are obtained, as shown in the first two columns of Table 3. The only difference in the results here is the reversed sign and loss of significance on the male unemployment variable (which in this regression is a lagged three-year moving average to correspond to the cumulative nature of the dependent variable, whereas in the marriage rate regression it is the current unaveraged rate). This is probably due to the fact that the dependent variable here is a stock measure, and the unemployment rate is cyclical: low(high) unemployment in the three year period prior to time \( t \) would be a strong indicator that unemployment at time \( t \) would be high(low).\(^6\)

Also shown in Table 3 (columns 4 and 5) are the results of a regression with the logistic transformation of the divorced proportion of males aged 20-24 as the dependent variable. The

\(^6\)This hypothesis is supported by the fact that the smoothed unemployment rate loses significance, and then becomes significantly negative in its effect on the proportion married, if we move from a one year lag to a zero lag, and then a lead of this variable.
same explanatory factors are used here, as in the marriage regressions, and the overall fit is good, with an $R^2$ of 0.985. Here, as might be expected, the sign on each of the variables is the opposite of its counterpart in the marriage regressions -- although enrollments are not significant in these divorce equations. Male unemployment is significant (and negative) in this regression, and here again the cyclical nature of unemployment is felt to be the cause of this counterintuitive result.

When the marriage proportions equation was tested in differenced form, only the relative income variable was found to be significant, as shown in column 3 of Table 3, and it explained about 15% of the year-to-year variation in the proportion married among 20-24 year old males. The differenced results for the proportion divorced were much stronger, however: only the relative income and female wage variables were significant, as shown in column 6 of Table 3, but they explained about 44% of the year-to-year variation in the proportion divorced among 20-24 year old males.

E. Patterns of Enrollment, Work Status and Relative Income

The tendency of marriage rates to stabilize -- and for divorce rates to decrease -- as depicted in section B of this paper, was most marked among young men who were working fulltime, full year -- who were not currently enrolled in school. What has been happening to the proportion working fulltime, and enrollments, in recent years? And how well does the pattern of male relative income fit recent trends in relative cohort size?

The proportion of all currently unenrolled male high school graduates in their first five
Fig. 7a: Proportion Working Fulltime Full Year, All male HS grads in first five years of work experience.

Fig. 7b: Proportions Enrolled, men and women aged 20-24.
years of work experience who are working fulltime year round tends to be a highly cyclical variable, but in general it has exhibited a gradual upward trend in the period since the mid 1960s, as shown in Figure 7a in data taken from the March Current Population Survey. The proportion declined from about 59% in 1989 to about 50% in 1992, but appears to be on the rebound again in 1993 and 1994. A time trend fitted to the data shows a 10% increase in this proportion over the period, from about 49% in 1963, to 54% in 1994. Thus the tendency to marry has been increasing -- or at least stabilizing -- in a segment of the population which has also been increasing.

It might be thought that, similarly, because of the beneficial effect of declining cohort size on male relative income, and the negative relationship between male relative income and college enrollment, that the proportion not enrolled in college will also increase in coming years. However, cohort size has been demonstrated to affect enrollments not only through relative income, but also through the college wage premium (see, for example, Freeman 1975; Katz and Murphy 1991). This effect is generally attributed to declining substitutability between young and old workers, with rising levels of education: a new high school graduate can more easily replace his more experienced counterpart in blue collar labor, than can a new college graduate in white collar work. Given these differential levels of substitutability, a boom cohort will see the relative wages of young college graduates decline more rapidly than those of high school graduates, with the opposite occurring for baby bust cohorts.

In addition, as Figure 2b shows, after a long decline in the 1970s and early 1980s (while

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7 A similar pattern is exhibited in the proportion working fulltime among all males in their first five years of work experience from 1968-1994, but data for the 1963-67 period are unreliable in this group because of differences in methods used to classify currently enrolled students.
relative cohort size was increasing), male relative income improved briefly as relative cohort size began to decline in the late 1980s. But since 1990 male relative income has declined sharply, despite declining relative cohort size in most of these recent years. It appears that in this most recent period the documented beneficial effects of declining cohort size have been outweighed by the negative effects of other factors such as a strongly increasing level of imports, which some researchers (Murphy and Welch, 1991; Katz and Murphy, 1992; Wood, 1994) have shown to have a stronger negative effect on younger less skilled workers, than on more skilled workers.

Thus both a decline in male relative income and a rise in the college wage premium resulting from smaller cohort sizes, appear to have contributed to a steadily increasing level of male (and female) enrollments throughout the 1980s and into the 1990s, as shown in Figure 7b. There have been a number of studies which have examined the effects of cohort size on schooling choice (Ahlburg et al, 1981; Wachter and Wascher, 1984; Falaris and Peters, 1985; Nothaft, 1985; Connelly, 1986; Stapleton and Young, 1988; Macunovich, 1993b). Although each of these studies has taken a slightly different approach to the question of cohort size effects, (some examining boom vs. non-boom, and others examining pre- vs. post-boom), a general consensus emerges which points toward a tendency for post-boom cohorts to invest more heavily than boom cohorts. And certainly this, too, is consistent with the steep increases in enrollment rates observed in recent years.

F. Conclusions

There are a number of signs indicating that the marriage rates of young men and women
aged 20-24 have stabilized (and that divorce rates have begun to decline) since the middle of the 1980s, holding education constant. These changes in trend have appeared most strongly only for the youngest age groups thus far -- those aged 20-22 -- suggesting that cohort effects are operational. Even more, the effects appear to operate most specifically on young males, suggesting that recent changes in male relative income may be responsible. Regression results strongly support this hypothesis, even in differenced form, where over 75% of year-to-year variation in marriage rates of young males is explained with a relative income model.

The 'marriage squeeze' hypothesis -- which is directly related to cohort size -- did not stand up well to tests, however, and in addition it appears somewhat questionable whether movements in male relative income in recent years have corresponded to changes in relative cohort size. Thus, while male relative income appears to explain well the historical pattern of marriage and divorce rates for males aged 20-24, it is debatable whether relative cohort size per se will remain a good predictor of these patterns.
References


### Table 2: Regression Results for the Marriage Rate as Reported in Vital Statistics, for Males Aged 20-24 (Vital Stats marriage rates only available through 1988)

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<td>(1)</td>
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<tr>
<td>log male relative income (RY)</td>
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<td>denominator of male RY</td>
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<tr>
<td>time trend</td>
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<td>0.9983</td>
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* Results using a Cochrane Orcutt procedure produced insignificant estimates of rho.

Dependent variable in each case is the log-odds ratio -- logistic transformation of the actual marriage rate, or $\log(rate/(1-rate))$. Source: Vital Statistics.

Male relative income is average real annual earnings at time $t$ of unenrolled males in their first five years of work experience divided by the average total annual income of all families with children, with head of either sex aged 45-54 at time $t-5$ (in constant 1967 dollars). Source: March CPS.

Male (female) enrollments are the three year moving averages of the proportion of all males (females) aged 20-24 enrolled in college, relative to the total male (female) civilian noninstitutional population Source: Current Population Reports, Series P-20.

The female real wage is the average real female hourly wage of all unenrolled women working fulltime in their first five years of work experience, holding education constant at its 1968 level. Source: March CPS.

Male unemployment is the rate for all unenrolled males in their first five years of work experience. Source: March CPS.

Marriage squeeze variable is population aged 20-24 relative to population aged 18-22.

Years of work experience in all of the above are calculated as years of completed schooling minus age minus six. $t$-statistics in parentheses.
Table 3: Regression Results for Proportions Married and Divorced, Males Aged 20-24

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<td>log male relative income (RY)</td>
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<td>(young men's earnings)</td>
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<td>denominator of male RY</td>
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<td>(parents' income)</td>
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<td>(5.1)</td>
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<td>log female hourly wage</td>
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<td>(at time t-1, smoothed)</td>
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<td>adj. R-square</td>
<td>0.9640</td>
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+ Results using a Cochrane Orcutt procedure produced insignificant estimates of rho.

Dependent variable in each case is the log-odds ratio -- logistic transformation of the actual proportion, or log(proportion/(1-proportion)). Proportions married and divorced are three year moving averages of the annual proportions, to correct for noise in sample data. Source: March CPS.

Male relative income is average real annual earnings at time t of unenrolled males in their first five years of work experience divided by the average total annual income of all families with children, with head of either sex aged 45-54 at time t-5 (in constant 1967 dollars). Source: March CPS.

Male (female) enrollments are the three year moving averages of the proportion of all males (females) aged 20-21 enrolled in college, relative to the total male (female) civilian noninstitutional population Source: Current Population Reports, Series P-20.

The female real wage is the average real female hourly wage of all unenrolled women working fulltime in their first five years of work experience, holding education constant at its 1968 level. Source: March CPS.

Male unemployment is a three-year moving average of the rate for all unenrolled males in their first five years of work experience. Source: March CPS.

Years of work experience in all of the above are calculated as years of completed schooling minus age minus six. t-statistics in parentheses.
Appendix: Mean and Median Age at Marriage  
(source: Vital Statistics – Marriage and Divorce)

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<th>Year</th>
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