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**RELATIVE INCOME AND PRICE OF TIME: EXPLORING THEIR EFFECTS ON U.S. FERTILITY AND FEMALE LABOR FORCE PARTICIPATION, 1963-1993**

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November, 1994 (revised April, 1996)



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Research Paper No. 174
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**Abstract:**

This paper attempts to review and synthesize the literature on the economics of fertility in order to develop a model which combines the 'price of time' and the 'relative income' concepts underlying the two most well-known models. Emphasis is placed on the need to work with an exogenous measure of the female wage, and on the possibility that the income effect of the female wage may be stronger than has been assumed in most 'price of time' models. A model is formulated and tested using aggregated data from the March Current Population Surveys, 1964-1995, and is used to demonstrate the changing net effect of the female wage on fertility during this period. The model appears to explain well the observed pattern of fertility, even in differenced form, and produces a strongly positive effect of male relative income and a strong underlying negative price effect of the female wage. Interaction terms show that the price effect of the female wage has been declining over time, while its income effect has been increasing. Consistent with these findings for fertility, male relative income is found to exert a strong negative effect on female labor force participation and enrollment, while the female wage exerts a strong positive effect on these variables.

A recent paper in this journal (Pollak and Watkins, 1993) presents an excellent discussion of cultural and economic approaches to fertility analysis. The authors question the disciplinary boundaries between economics and sociology, and distinguish between two competing versions of the economic 'rational actor' model. One, the 'price of time' model (Mincer, 1963), part of the 'Chicago-Columbia model' initially articulated by Gary Becker (1960), assumes fixed and exogenous preferences and explains changes in behavior on the basis of shifting opportunities. The other, the 'relative income' or 'Pennsylvania school model' introduced in Easterlin (1966) and elaborated and formalized in Easterlin, Pollak and Wachter (1980), assumes that preferences are endogenous and thus can vary systematically, along with opportunities.[1] Pollak and Watkins point out that "[r]ecent work in the Chicago-Columbia tradition has been increasingly congenial to notions of preference formation and change, although these notions have not yet appeared in Chicago models of fertility." (p.489) They state that "many economists still regard preference formation and change as the business of other disciplines", and emphasize that they find "this intellectual division-of-labor argument unpersuasive." Pollak and Watkins go on to state that "because estimation presupposes a correctly specified model, empirical estimators cannot ignore preference change. If an investigator assumes fixed preferences when in fact they are changing, then all of the coefficient estimates, even those of the narrowly specified economic variables, are inconsistent. Thus, if preference change is taking place, economists cannot ignore it unless they are prepared to abandon empirical analysis and reconstitute economics as a purely deductive enterprise." (p.491)

Perhaps the most common problem cited in critiques of these two models, is that their predictions have failed badly in the 1980s. The 'price of time' model predicted continuing declines in fertility, while the 'relative income' model predicted an upturn during this same period. In contrast, as is fairly well known, fertility rates have remained fairly stagnant since the mid 1970s.

This paper is presented in the spirit of the Pollak and Watkins article, in an attempt to illustrate the type of results which can be obtained when preferences are assumed to vary systematically over time, and account is taken not only of male income, but also of the female wage.

Unfortunately, the term 'fertility' encompasses a wide range of concepts, all of which cannot be incorporated in one model (e.g., timing and spacing, completed family size, period versus cohort). However, as Morgan (this volume) emphasizes, probably the most crucial factor in completed family size, as well as in period measures of fertility, is the timing of the first birth. The earlier this occurs, the higher will be completed family size, on average, and if one cohort brings its first birth forward to overlap with delayed births in an earlier cohort, then period fertility rates will rise even in the absence of any change in completed cohort fertility. This is fortunate for empirical estimation since Easterlin's 'relative income' effect is expected to be strongest for young adults just entering the labor market and household formation stage. As a result, the work in this paper focuses on the behavior of young adults aged 20-24.[2]

## THE CONTEXT

While the bulk of socio-economic research to date has tended to assume that fertility and female labor force participation are somehow causally related, researchers have been hard-pressed to determine the *direction* of such causation. (See, for example, Cramer, 1980 and Lehrer and Nerlove, 1986). Easterlin's relative income model, on the other hand, assumes that these two factors are caused simultaneously by a third factor, male relative income, so that observed correlations between them are spurious. Easterlin postulates a systematic shift in preferences resulting from the fact that each successive generation, under economic development, experiences a successively higher parental standard of living. "In effect, a... 'subsistence level' constraint is added to the analysis of [fertility behavior] along with the budget line and production constraints." (Easterlin, 1978, and Ahlburg, 1984). Because of this 'subsistence level' constraint, economic or demographic fluctuations could cause periodic reversals in the secular downtrend in fertility, such as that observed in the developed countries in the postwar period. In the full Easterlin model, relative birth cohort size (the size of one's birth cohort relative to the size of one's parents' birth cohort) inversely affects cohort earning potential relative to material aspirations, which in turn affects cohort fertility and female labor force participation.

According to the Easterlin theory, a large birth cohort meets unfavorable labor market conditions which reduce the earning potential of young males relative to their aspirations. In

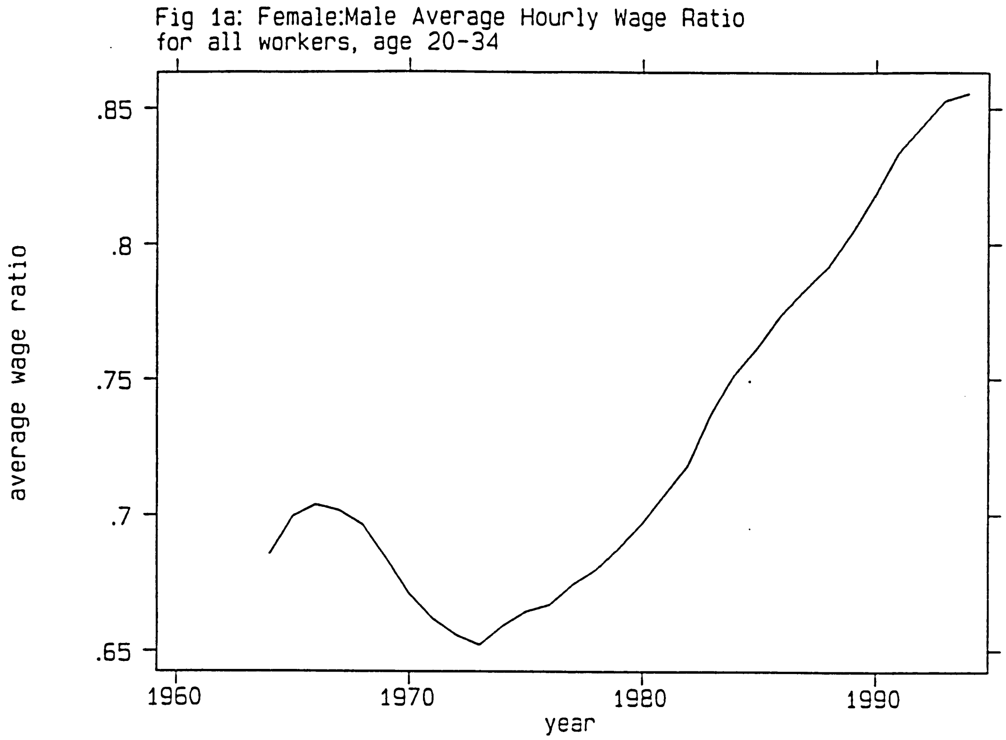


Fig 1a: Female:Male Average Hourly Wage Ratio
for all workers, age 20-34
average wage ratio
.85
.8
.75
.7
.65
1960
1970
1980
1990
year

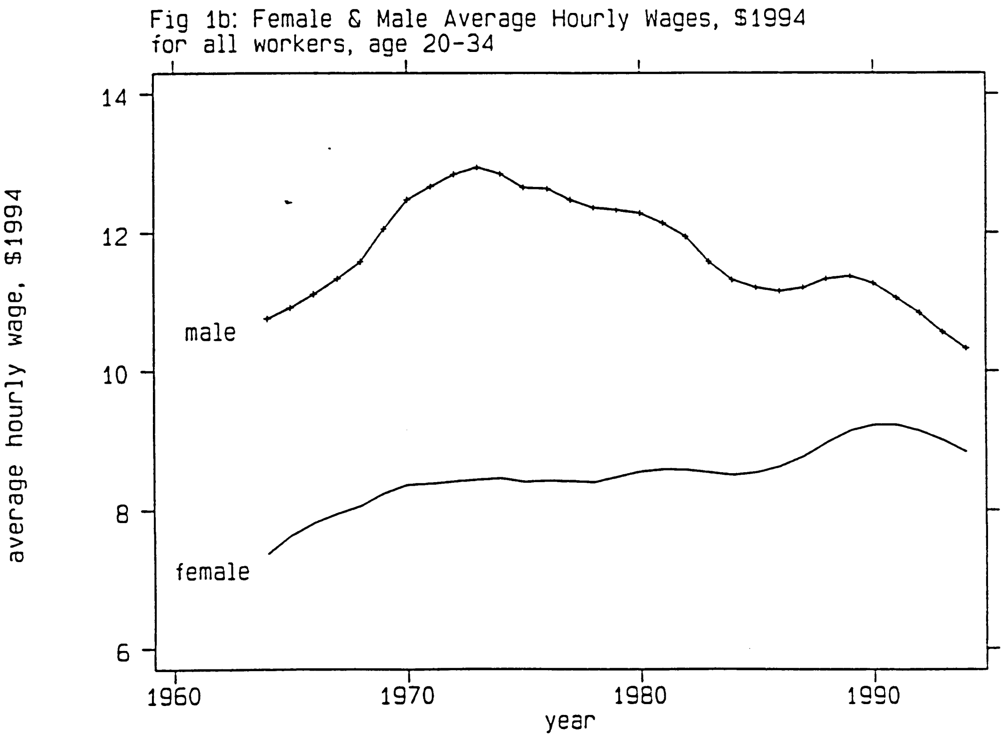


Fig 1b: Female & Male Average Hourly Wages, $1994
for all workers, age 20-34
average hourly wage, $1994
14
12
10
8
6
male
female
1960
1970
1980
1990
year

an attempt to close the gap between income and aspirations, members of such a cohort will tend to make a number of adjustments including increased female labor force participation and delayed/reduced marriage and childbearing. In addition, because young women in large cohorts will anticipate higher levels of labor force participation, they will tend to enroll in college at increased rates. In this formulation the driving force behind both increased female labor force participation and enrollments, and reduced fertility, is the desire of a large cohort to improve relative economic status, with parental income as the measure of that cohort's material aspirations.

What has been happening to some of the primary economic factors in these two models: male and female wages and male relative income? And what has been happening to some of the variables they're thought to affect -- marriage, divorce, fertility and female labor force participation rates -- for the age group focused on in this paper, young adults aged 20-24?[3]

Although the media and many academics would have us believe that female wages have been rising dramatically over the past twenty years, this is true only in a relative sense -- that is, relative to male wages. This observed relative female wage is presented in Figure 1a. An examination of its components, as presented in Figure 1b, shows a somewhat less rosy picture: there, it can be seen that the relative female wage rose largely because the male wage fell.

In addition, when we speak of the average female wage we must take care to correct for effects of changing average levels of education and experience, in a period of rapid increases in female labor force participation. (Without such corrections, we would not be observing an 'exogenous' female wage. Rather, we would be observing changes in the wage which women could have brought about at any time, simply by increasing their levels of education and labor force experience). Figure 2a shows the average real female wage for women in their first five years of work experience (thus to a great extent controlling for experience) and holding average education constant.[4] Here we see a disappointing picture since 1973: a real female wage which has not made any gain at all in a twenty year period. It exhibits only cyclic fluctuations -- and even appears to have declined in the last few years.

Figure 2b presents the average real weekly wage for female college graduates

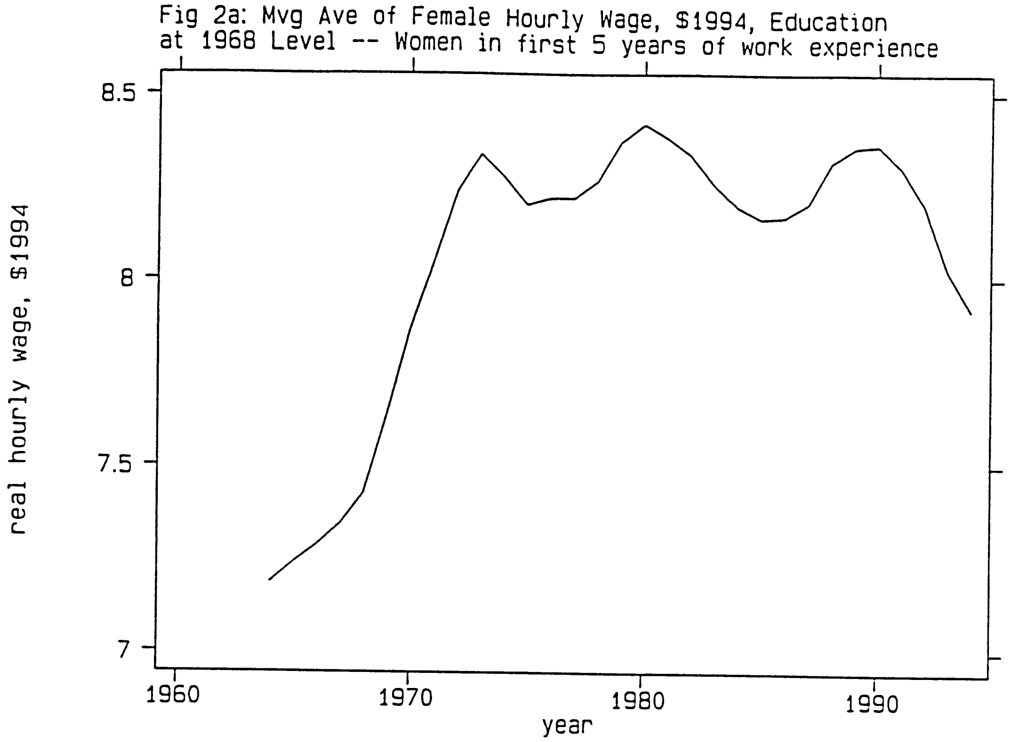


Fig 2a: Mvg Ave of Female Hourly Wage, $1994, Education at 1968 Level -- Women in first 5 years of work experience
8.5
8
7.5
7
real hourly wage, $1994
1960
1970
1980
1990
year

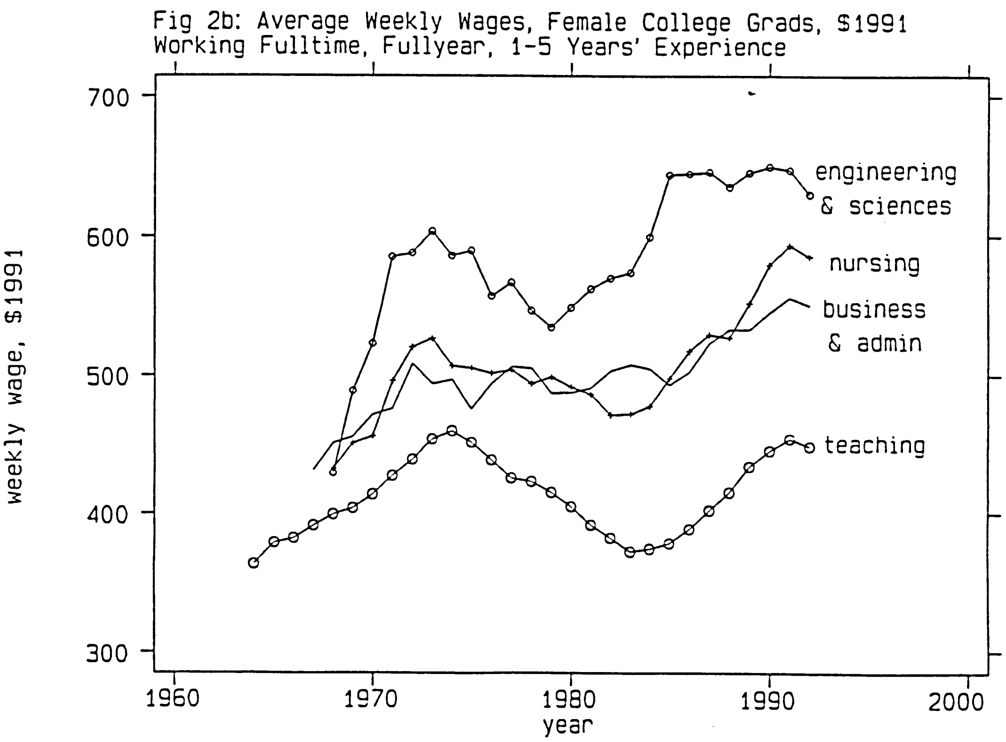


Fig 2b: Average Weekly Wages, Female College Grads, $1991 Working Fulltime, Fullyear, 1-5 Years' Experience
700
600
500
400
300
weekly wage, $1991
engineering & sciences
nursing
business & admin
teaching
1960
1970
1980
1990
2000
year

working fulltime, full year in traditional and not-so-traditional occupations during the same period. Here again, it is difficult to discern the steady upward trend that we are often led to imagine.

However, as shown in Figure 3a, male relative income (*RY*) as defined by Easterlin[5] improved temporarily in the late 1980s, after nearly twenty years of steady decline. While the data presented in Figure 3a are for all young men in their first five years of work experience, the same pattern emerges even in subgroups by education level and income quintile: this overall pattern of decline in *RY* from about 1970 through 1985, with a short but marked turnaround occurring after that date, appears to be a widespread phenomenon throughout the population, even during a period of rising earnings inequality. At the same time, measures of female labor force participation among women aged 20-24 appeared to top out and even begin to decline in the late 1980s and early 1990s, as shown in Figure 3b. Fertility rates for women aged 20-24 appeared to increase marginally in the late 1980s after the low levels achieved in the middle of that decade, as shown in Figure 3c, and marriage rates among fulltime workers aged 20-22 began to rise in the same period after approximately 20 years of decline (Figure 3d) while divorce rates fell (Figure 3e).

Fair and Macunovich (1993) present an aggregate quarterly time series model of enrollment and labor force participation rates for women aged 20-24 which appears to explain well the movements in these series since the 1950s using concepts of *RY* and the female wage derived from the aggregate wage. Similarly, Macunovich (1993b) presents a model of the marriage rate for males aged 20-24 based on the *RY* measure presented in Figure 3a and the female wage presented in Figure 2a, which explains both the levels and changes in marriage rates since the 1960s. It is in this context, then, that this model of fertility for women aged 20-24 is formulated, and shown to coincide with a model of female enrollment and labor force participation rates using the same explanatory variables.

**BACKGROUND**

Olsen (1994) presents a review of much of the literature on the 'relative income' and 'price of time' models, and concludes that "the recent stability of the fertility rate suggests these theories do not have good short term predictive power." Similarly, Pampel and Peters (forthcoming) review the literature on the relative income model, and find only a contingent

Fig 3a: Male Relative Income Variable in unlogged form
Males with 1-5 years exp relative to families with head 45-54

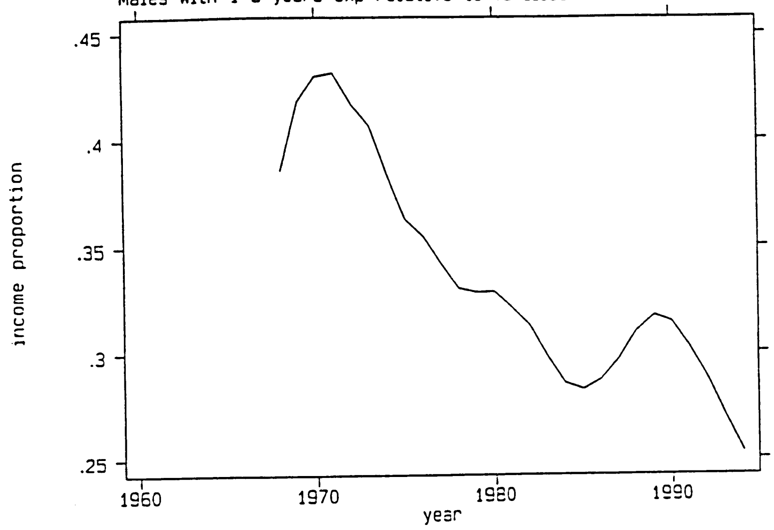



Fig 3b: Labor Force Participation Rates, Women Aged 20-24

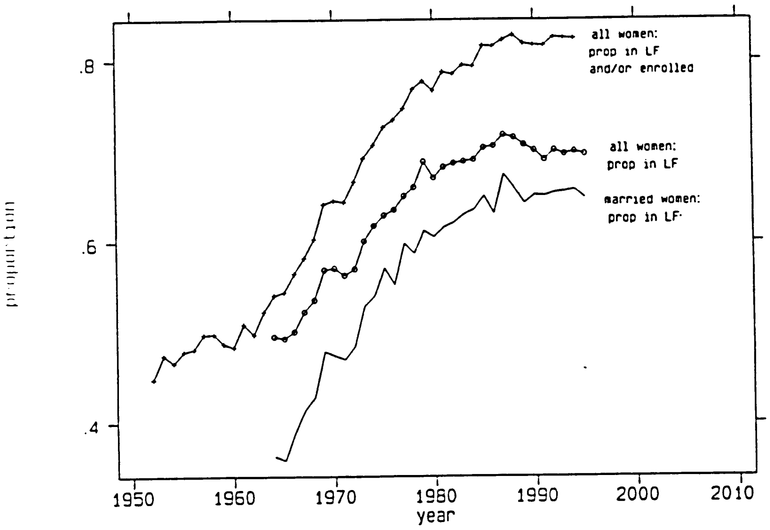



Fig 3c: Fertility Rate unlogged form
Women Aged 20-24, Births per 1,000

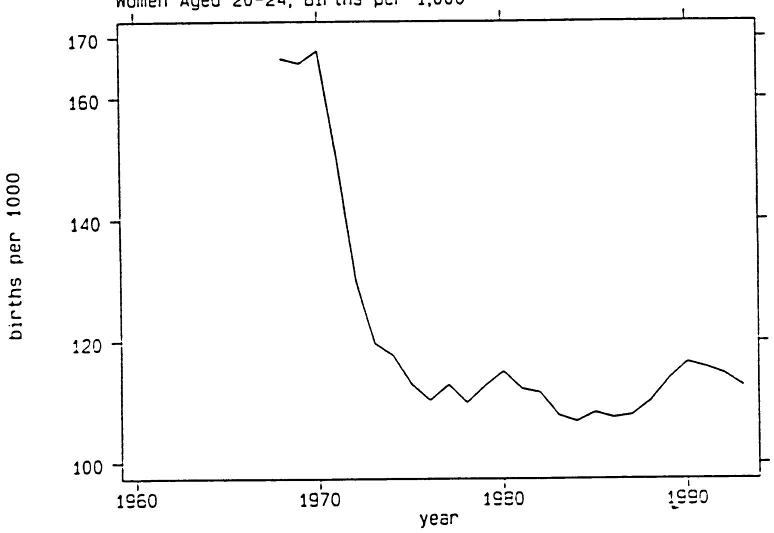



Fig 3d: Proportions Married (3 year moving average),
Fulltime Workers Aged 20-22

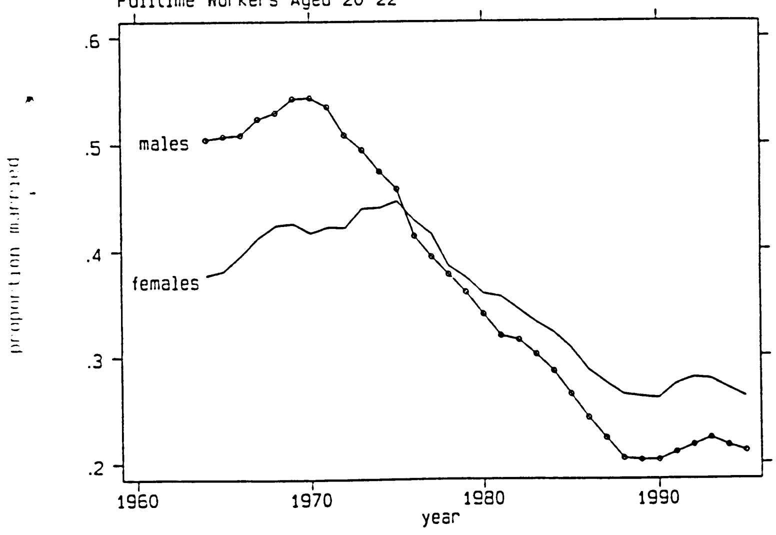
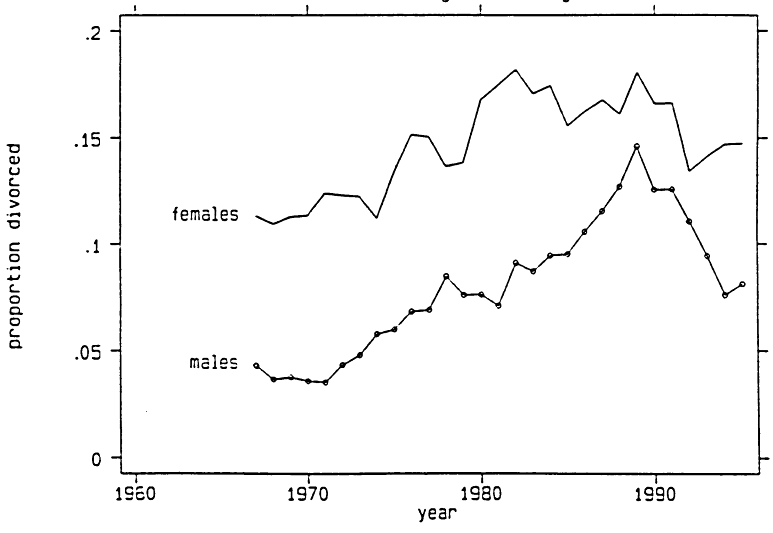



Fig 3e: Proportions Divorced
Ever-Married Men & Women Working Fulltime Aged 20-22



effect on fertility at best. They conclude "[t]he evidence for the Easterlin effect proves mixed at best and plain wrong at worst. . .[T]he predictions of the Easterlin effect best fit aggregate, period trends from 1945-1980 in the United States."

However, it should be noted that Olsen did not include several studies which have presented evidence in support of the relative income model for the U.S., such as Moffitt (1982), Devaney (1983), Schapiro (1988) and Behrman and Taubman (1990) -- the last using micro level data. Similarly, the Pampel and Peters review (seen only in draft) excluded favorable results presented in Devaney (1983), Behrman and Taubman (1990) and Danziger and Neuman (1989) -- the last two based on micro level data.

Olsen's and Pampel and Peters' conclusions accord well with those expressed in the review article by Montgomery and Trussell (1986), who point out that although "Easterlin's own work with aggregate census data. . .track U.S. fertility levels over the baby boom and bust rather well. . .in tests based on micro data, the evidence. . .is weak." (p. 252) With regard to the 'price of time' model, Montgomery and Trussell note that "[t]he economic logic of the argument is impeccable; unfortunately, econometric tests of the proposition have not been nearly so convincing." In particular, these authors note that 'price of time' models which show a negative effect of the female wage on fertility "rely heavily on female education as a driving variable" while those which do not rely so heavily on education "have not succeeded in establishing a negative wage effect." (p.264)

The attempt here -- rather than to reproduce the work presented in these recent reviews -- will be to discuss potential shortcomings in the formulations of the models previously tested, in order to lead toward a more useful combined version. These shortcomings will be addressed under the following three headings:

1) difficulties in formulating appropriate measures of relative income
2) exclusion of the female wage, or use of an endogenous measure
3) failure to control for the changing net effect of the female wage over time

The discussion presented below draws heavily on the more detailed review presented in Macunovich (1996 forthcoming).

**Measures of Relative Income**

Easterlin (1987) has characterized his concept of relative income as (earnings potential)/(material aspirations). He operationalizes this measure using ''the recent income experience of a young man relative to the past income of the young man's parents.''[6] The concept is hypothesized to operate at the micro level; that is, young adults are assumed to set their desired standard of living based on the standard of living experienced in their own parents' homes.

However, there has been little research on the formation of material aspirations,[7] so that the formulation of a relative income measure at the micro level is fraught with hazards. What is an appropriate relative income measure for an unmarried woman? And when looking at married couples, do we need a comparison of earnings of the husband or the wife, or both -- and relative to his parents, to her parents, or both? Not only have none of these questions been answered to date: in general we lack micro datasets which provide information on first and second generation income and fertility behavior for both partners in a union.

Thus any tests of the model at the micro level have either used gross proxies for the relative income measure (such as occupational data or responses to questions such as ''How well-off are you?''-- see MacDonald and Rindfuss, 1978), or used data only for the male in the couple, assuming that his preferences dominated in the household -- or they have done both. In addition, micro level formulations treat aspirations as being *only* a function of parental income, whereas it is likely that peer groups and other larger-scale social phenomena also operate on individual aspirations.

Such problems do not occur in aggregate time series analyses. At the aggregate level we don't have to associate individuals with their own parents, or with their own spouses, but rather a young generation with its parental generation. We can use average household income in the parental generation, with a lag, as a measure of the average desired standard of living for both males and females in the younger generation, and this measure will better approximate both parental and other influences on a cohort's aspirations.

In addition, working at the aggregate level of analysis in a relative income context helps us address the problem of delayed/postponed marriage as it affects extra-marital fertility -- something impossible in models of marital fertility such as the 'price of time' model. That is, as Korenman and Okun (1992) point out, there is increasing 'slippage' between marital and extramarital fertility. This makes it difficult to specify the effects of male relative income on females at the micro level. An aggregate relative income model of fertility accounts for such slippage, in that low male relative income prompts males to delay/postpone marriage, leaving more young women at risk of extramarital fertility. Incorporating the female wage in such an aggregate model, with allowance for its changing income effect, permits the female wage to exert the primary income effect on fertility in periods of low male relative income -- as it would for a population containing large numbers of single women.

**Endogeneity in the Female Wage**

We have seen in the previous section that the relative income model as traditionally formulated can be faulted not only for its omission of any measure of the female wage -- theoretically the 'price of time' in childcare -- but also in many cases for the use of inappropriate measures of male relative income. In this section we focus on 'price of time' models, and find that here the majority can be faulted not only for their exclusion of any measure of material aspirations -- the denominator in measures of male relative income -- but also for the use of erroneous and/or endogenous measures of the female wage.

As noted earlier, Montgomery and Trussell (1986) emphasize the possible endogeneity of female wages used in price of time models -- particularly in cases which find a significant negative effect of the female wage on fertility. That is, in periods of rapid increase in female labor force participation such as the last 30 years, the observed average female wage will reflect rising average levels of experience (and education) which must be purged in order to identify an exogenous female wage. Without such a purge, it is hardly surprising that empirical tests find a positive relationship between the wage and labor force participation: the real question should be about the direction of causation. Similarly, an observed negative effect between an endogenous wage and fertility is to be expected because of the observed inverse relationship between fertility and education/labor force participation. Lee and Gan (1989) show that this is a problem in Ogawa and Mason (1986), while Abeysinghe (1993) questions the results in Hyatt and Milne (1991) for the same reason. Similarly, Ermisch (1980), Smith and Ward (1984, 1985), Heckman and Walker (1989) and Whittington et al (1990) use observed female wages which have not been purged of education and experience effects. Schultz (1986) uses a predicted female wage, but uses education as the primary explanatory variable in his wage equation. Ermisch (1979, 1989) used expected men's real wages (age 21 and over) as a proxy for female wages.

Perhaps the most well-known and often-cited application of the price of time model is presented in the work of Butz and Ward (1979). However, as demonstrated in Macunovich (1995), their model suffers from a combination of the above problems. The female wage they used was an estimated one, and is demonstrated to have been erroneously estimated, in addition to being endogenous. Substitution in their model of an observed female wage series calculated from the March Current Population Survey produces insignificant coefficients on the female wage. In addition, even using the original Butz-Ward data, Macunovich found that their model produces insignificant and/or perverse results for the entire period after 1955 -- which is the period in which Butz and Ward felt that the female wage exerted its strongest effect on fertility.

**Changing Net Effect of the Female Wage**

In theory, the female wage is expected to exert both a positive (income) and a negative (price of time) effect on fertility. In practice, however, proponents of the Chicago-Columbia model have expected that any positive effect would be greatly outweighed by its negative effect. In addition, they have tended to ignore the fact that theory dictates an increasingly dominant income (positive) effect with rising wages and hours worked.[8] Thus, for example, Mincer (1962) estimates the wage elasticity of female labor supply in cross section and then applies that constant elasticity to explain changes in the level of female labor force participation between 1890 and 1960 -- a period when both the wage and hours worked are thought to have risen rapidly. Smith and Ward (1985) explain an apparently biased coefficient in an estimate produced using cohort data for the period 1950 to 1980 by referring to the "rapid rise in hours at young ages that we are not capturing with our controls". (p.S87)

In addition to the standard theoretical explanation underlying an increasingly positive effect of the female wage on fertility over the last century, it is possible to imagine at least two others. The first has to do with a declining negative -- price of time -- effect of the wage. This negative effect is posited on the assumption that a woman is the primary provider of childcare: every hour she spends in childcare will 'cost' her the foregone wage. However, to the extent that alternative (purchased) methods of childcare are both available and socially acceptable, this negative price effect will be diminished: a woman can work the extra hour and pay for a replacement for her time in the home. We have certainly observed the development of such conditions over the past thirty years. This is the type of effect described in Rindfuss and Brewster (this volume). Ermisch (1989) has been able to document this effect using British data for 1980. Similarly, findings in Cramer (1979), Lehrer (1989), Joshi (1990) and O'Connell and Bloom (1987) are all supportive of such an effect.

Secondly, the positive -- income -- effect of the female wage will tend to be higher, all other things equal, the higher the woman's material aspirations. Why? Because the perceived value of her nonearned income (income from sources other than her own wages) will appear diminished relative to those higher material aspirations, and her participation in the labor force will thus be higher, all other factors equal. In addition, to the extent that marriage is delayed/foregone in a period of low male relative income, fewer women will have husbands as a source of nonearned income. In that event, even the absolute value of women's nonearned income will decline, on average -- resulting, again, in higher female labor force participation rates, all other things equal. Women's greater number of hours in the labor force in both cases will tend to increase the income effect of a wage increase, since income equals the hourly wage times number of hours worked. A graphic example of this effect, using a Cobb-Douglas production function, is presented in Macunovich (1993c).

It is probably one of the most significant shortcomings of price of time models to date, that they have not allowed for any of these changing net effects of the female wage on fertility. In literature on female labor force participation, however, successive studies find increasingly smaller positive -- and even insignificant or negative -- net effects of the female wage on female labor force participation.[9]

## FORMULATING A COMBINED 'RELATIVE INCOME' AND 'PRICE OF TIME' MODEL OF FERTILITY

Combining these two models is not difficult in practice -- only in theory! It is assumed that individuals maximize utility over children and consumption goods, and that as in the 'price of time' model, the optimal number of children will be an increasing function of total income and a decreasing function of the price of children. Also drawing on that model, it is assumed that one of the most significant measures of the price of children will be the female wage, while total income will be a function of a woman's wage and her nonearned income (from wealth, a husband, or other sources). The departure from the 'price of time' model in this formulation is the conversion of total income into relative income by assuming that the optimal number of children will be a decreasing function of the individual's material aspirations, proxied by the income in the parental home five years earlier.

Then in addition, as discussed in an earlier section, we must allow the net effect of the female wage to vary, possibly over time (to capture effects of increasing availability and acceptability of purchased childcare), but certainly as a function of relative income. Two different formulations will be tested, in order to allow for these two types of variation. Variation as a function of male relative income (*RY*) can be approximated by including an interaction term between the female wage and the *RY* figure: as *RY* declines (rises), the income effect of the female wage, relative to its substitution effect rises (declines). Or, one might choose to interpret this interaction term as indicating that the importance of *RY* declines (rises) as the female wage rises (declines): the more a woman is able to support herself with her own wage, the less likely she will be to look first to male sources of income, for support.

Variation in the net effect of the female wage over time, resulting perhaps from the increased availability of purchased childcare, can be approximated by including a time trend and an interaction between the time trend and the female wage. This will be tested as an alternative formulation.

It is assumed that individuals operate within a utility maximizing framework in which expectations regarding earnings and wages and family income are formed using some type of moving average of those measures as observed in the years prior to and including the decision point. Shocks to this system of expectations formation are assumed to occur, which introduce uncertainty regarding the applicability of the model to future conditions. The unemployment rate is assumed to be a major indicator of this type of shock.

The data used in this study are, with the exception of the fertility rate, drawn exclusively from the March CPS for the years 1964-1995.[10] Individual data records were aggregated in order to identify labor force and population averages within various categories.

The fertility rate used is the age-specific rate for all women aged 20-24 in the U.S., as reported in *Vital Statistics* (but with data from 1980-1993 updated using the 1994 *Annual Summary of Births, Marriages, Divorces, and Deaths*). The rates for all women, rather than for married women, are used because it is postulated that *RY* and the female wage operate directly on fertility, whether marital or otherwise. The fertility equation has been modeled using a logistic transformation of the dependent variable, in order to provide a more meaningful interpretation of the intercept term.

Ideally, the real female wage used in this study should measure the market reward holding all characteristics constant over time: education, experience, tenure and training. The approximation to this ideal used in this study is the average real hourly wage of unenrolled females working fulltime in their first five years of potential work experience (measured as age minus completed schooling minus six), holding education constant over the period by taking real average wages by education level in each year and then combining them using the 1968 proportions of women at each educational level.[11] In addition, to make this an 'expected' wage, it is multiplied within each experience-year group by the average employment rate of the group (measured as one minus the unemployment rate in the CPS survey week). Finally, a five year moving average of this expected hourly wage is developed, in order to approximate an expectational model. This is the wage presented in Figure 2a.

The numerator of the male relative income measure (*RY*) is approximated using the real average annual earnings of all unenrolled males in their first five years of potential work experience. Males in their first five years of potential work experience are used, rather than males within a specific age category, in order to incorporate the effects of changing education levels which are assumed to result from young males' attempts to improve their earnings relative to those of their parents. All males are used, rather than males working fulltime only, in order to incorporate the effects of cohort size on workers' hours and weeks worked (Welch, 1979). Similarly, in order to make this a true 'expected' earnings figure, it is multiplied within each experience-year group by the activity rate of that group, where the activity rate is measured as the number employed divided by the total population in the group.[12] This helps to incorporate the observed effects of cohort size on labor force participation and unemployment rates. Finally, as with the female wage, a five year moving average of this earnings series is used. The average age of males in this group is about 22.1 in the early years, rising to 22.5 throughout the remainder of the period.

The denominator of *RY* is the real average annual income of all families with children under 18 and with a head (of either sex) aged 45-54. It is assumed that these families are representative of families of the young men and women who are the focus of this study. The use of family income allows for changing trends in terms of average number of labor market participants per family.[13] A five year moving average of this family income measure is used, and it is lagged five years in order to approximate economic conditions when the young men and women were still in their parents' households.[14] *RY* is derived by combining these measures of male wages and family income is presented in Figure 3a.

The unemployment rate used in this model to approximate shocks to the system is the unemployment rate of all women aged 20-24 as measured in the CPS survey week. The rate for women is used, rather than the rate for men, because earlier tests have demonstrated it to have the strongest effect on fertility (Macunovich and Easterlin 1988). This series is *not* averaged, since it is meant to represent the strength of recent shocks to the system.

All independent variables are lagged one period, to approximate the lag between conception and birth. Thus the fertility equation used in this analysis is the following:[15]

$$\Lambda(B_t) = \beta_0 + \beta_1 logRY_{t-1} + \beta_2 logW_{t-1} + \beta_3 logRY_{t-1}{*}logW_{t-1} + \beta_4 U_{1,t-1} \qquad (1)$$

where:

$$RY \equiv (Y_{1\text{-}5,t}{*}AR_{1\text{-}5,t})_{MA5} \,/\, (FY_{45\text{-}54,t\text{-}5})_{MA5} \qquad (2)$$

$$W \equiv (W_{1\text{-}5}{*}ER_{1\text{-}5})_{MA5} \qquad (3)$$

where equations (2) and (3) are identities and the subscript $_{MA5}$ denotes a five year moving average and:

$AR_{1\text{-}5,t}$ — is the activity rate (total employed divided by total population) of all unenrolled males in their first five years of potential work experience, at time t

$\Lambda(B_t)$ — is a logistic transformation of births per woman aged 20-24 at time t

$ER_{1-5}$ is the employment rate (total employed divided by total in the labor force) of all women working fulltime in their first five years of potential work experience, with educational levels held constant at 1968 levels

$FY_{45-54,t-5}$ is the real average annual family income of all families with children under 18 and with a head (of either sex) aged 45-54 ($1967), at time t-5

$U_1$ is the unemployment rate of all women aged 20-24

$W_{1-5}$ is the real average hourly wage of women working fulltime in their first five years of work experience, with educational distribution held constant at 1968 levels ($1967)

$Y_{1-5,t}$ is the real average annual earnings of all unenrolled males in their first 1-5 years of potential work experience ($1967), at time t

In addition, because fertility and female labor force participation are seen to be highly correlated, we present here as well a simple model of labor force and enrollment rates for women aged 20-24. It is assumed here, as in the standard neoclassical model, that an individual's labor force participation is a function of both her own wage and her nonearned income. Two departures are made here, from the standard neoclassical model -- as in the fertility model presented above. The first is to assume that a woman's labor force participation/enrollment decision is a function of her *relative*, rather than absolute, nonearned income, and the second is to allow for variation over time in the net effect of her own wage on her labor force participation. In this case, the variation over time in the net effect of her wage results from the increasing income effect of the wage as her hours worked increase.

Because it is assumed that college enrollment anticipates future labor force participation, the labor force participation rates used in this study are the traditional labor force data obtained from the Bureau of Labor Statistics, *supplemented* with data on women who are enrolled in college who are not in the traditional labor force. To obtain the latter data, the March Current Population Survey data were used for the years 1964-1995, and this information was supplemented with enrollment data from the *Current Population Reports, Series P-20* for the years prior to 1964. Using this combined series circumvents the problems normally associated with determining the effect of increasing enrollment rates on labor force participation rates. Thus the model tested for female labor force participation and enrollment rates is the following:

$$LFPE_t = \gamma_0 + \gamma_1 RY_{t-1} + \gamma_2 W_{t-1} + \gamma_3 LFPE_{t-4} {}^* W_{t-1} + \gamma_4 LFPE_{t-4} \qquad (4)$$

where *RY* and *W* are defined as in the fertility equation (1), and *LFPE* represents the proportion of all women aged 20-24 who are enrolled in school and/or participating in the labor force.

## RESULTS

The results obtained are extremely encouraging, in that both the fertility model presented in equation (1) and the female enrollment/labor force participation model in equation (4) appear to provide very close fits with observed trends during this period.

### The Fertility Model Results

The fertility model presented in equation (1) explains 99% of the variance in observed fertility for women aged 20-24 during this period, and the estimated coefficient on male relative income (*RY*) is positive and highly significant (indicating that children are *not* an inferior good), while the coefficient on the female wage term is negative and also highly significant. There is little if any serial correlation in the error terms, as indicated by the Durbin-Watson statistic, so that estimation with a Cochrane Orcutt correction has little effect on the coefficients or their significance. These results are presented in Table 1, and the fitted results are compared with the actual data in Figure 4a, where it can be seen that the model fits not only the marginal upturn in fertility which occurred in the late 1980s, but also its attenuation in the beginning of the 1990s.

A novel finding here is that, consistent with the theoretical model, the net effect of the female wage on fertility has indeed changed over time: when *RY* is high, the income effect of the female wage is reduced, so that the net effect of the female wage is negative, but when *RY* is low the income effect of the female wage strengthens. This is demonstrated through the use of an interaction term between the female wage and *RY*: the coefficient on this term is significantly negative. Another way of describing the effect of this interaction term would be to say that the strength of *RY* varies inversely as a function of the level of female wages: as female wages rise, women tend to focus less exclusively on male income in

Fig 4a: Fitted and Actual Fertility Rates
Women aged 20-24, using 1969-93 data

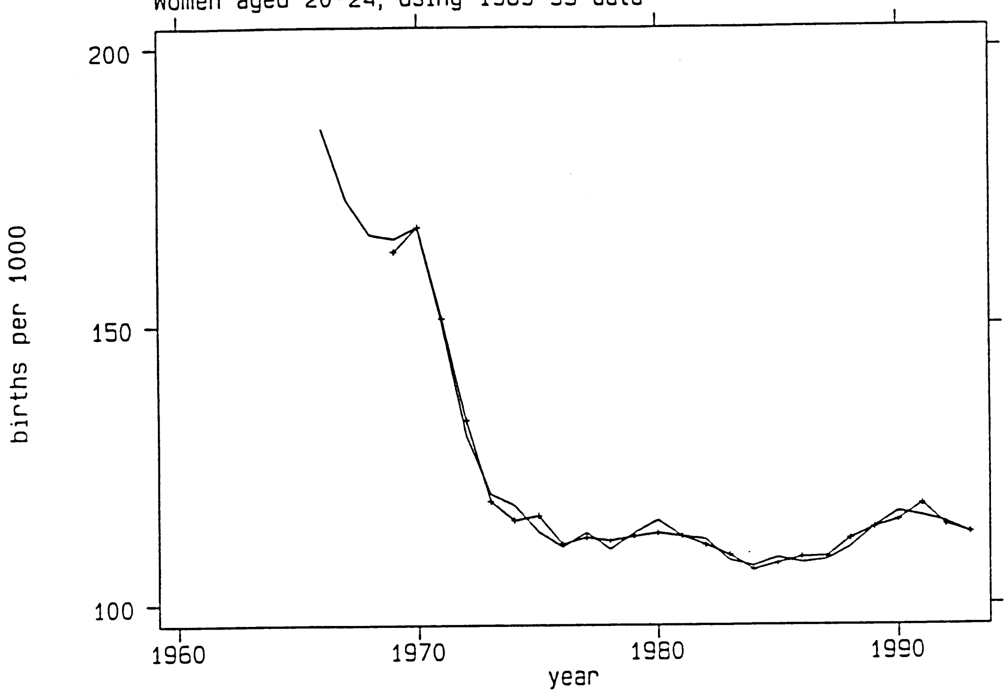
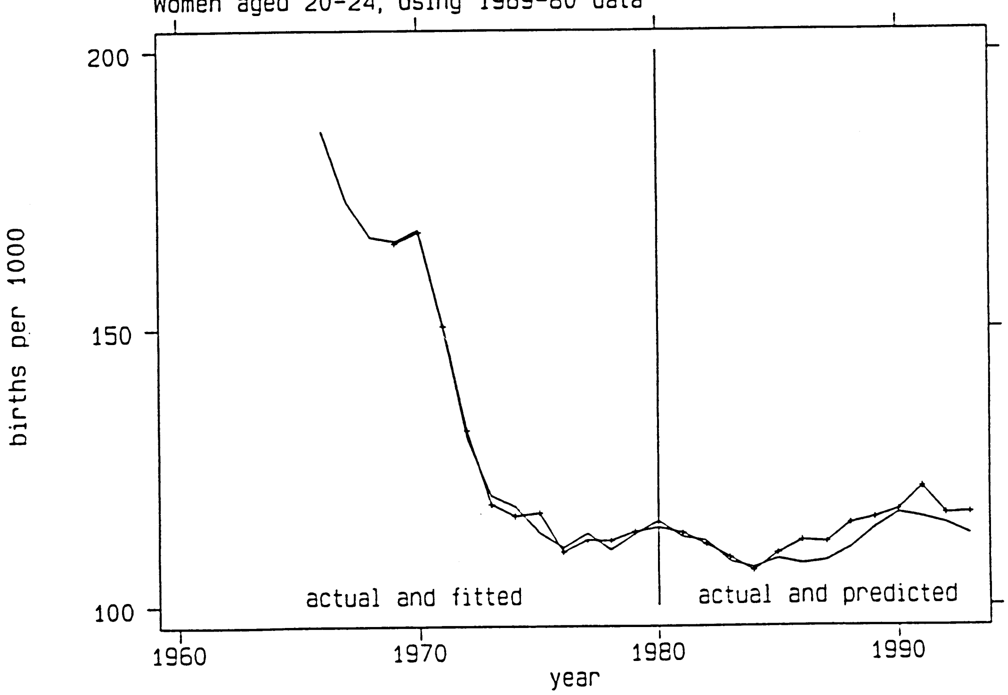



Fig 4b: Predicted and Actual Fertility Rates
Women aged 20-24, using 1969-80 data



childbearing decisions. Similarly, when *RY* is low, marriage may be postponed, ended, and/or foregone, and women will tend to focus more on their own earning potential in making childbearing decisions.[16]

It is significant, as well, that the effect of increased female unemployment is estimated to be *negative* on fertility, indicating that the reduction in the price of women's time during such periods is more than counterbalanced by the expectation of lost income.

These estimated coefficients appear to be highly stable within the sample: in both the full sample and in a 1969-80 subsample, the model explains over 99% of the variance in the time series, and coefficients remain relatively stable (columns 1 and 5 of Table 1). Even in the period 1975-93, following the marked decline in fertility and rise in the female wage, the model produces very similar results (column 4 of Table 1). As a result, the estimated coefficients from the 1969-80 model can be used to predict within sample through to 1993, and it can be seen (Figure 4b) that the fit through 1993 is quite good, with the recent slight upturn and attenuation mirrored in the predicted figures. In addition, column (3) of Table 1 indicates that the model performs extremely well in differenced form, as well as in levels: all coefficients remain highly significant with the same signs as in the levels equations, and the $R^2$ is a very high (for differenced data) 0.77.[17]

**Testing the Numerator and Denominator of Relative Income Separately**

Of theoretical interest is the fact that when the numerator and denominator of *RY* are entered separately, the signs and significance of the coefficients are in accord with the theory (positive on the male earnings and negative on the parental family income). In addition, the coefficients on the two are not statistically different from each other, thus supporting their use in ratio form. The separate interactions with the female wage, too, produce strongly significant coefficients with the expected signs (positive in the interaction with parental family income and negative in the interaction with male earnings). Here, too, the coefficients are not significantly different from each other. These results are presented in Table 2.

**The Changing Net Effect of the Female Wage on Fertility**

The changing effect of the female wage on fertility is illustrated in Table 3 -- in panel 1 using the full regression results and in panel 2 using regressions excluding the interaction term, estimated for subperiods at the beginning and end of the total sample period (1969-79 and 1974-93). In the early subperiod -- the 1970s, corresponding to part of the period covered by the Butz and Ward (1979) analysis -- the effect of the female wage on fertility was negative and significant; but in the later period -- the period since the Butz-Ward analysis -- this effect was reversed, becoming positive and significant, despite the fact that the effect of male relative income (*RY*) was positive and significant in both subperiods. The gross (uncompensated) effect of the female wage was significantly negative when *RY* was high (*RY* reached a peak in this dataset of 0.433 in 1972) and is significantly positive now with *RY* at an historic low (0.252 in 1995). This is consistent with the hypothesis that the marginal value of a woman's earned income is an inverse function of *RY* -- or, alternatively, that the marginal value of *RY* is an inverse function of the female wage.

**Testing the Alternative Assumption Re: Changing Net Effect of Female Wage**

Although the model presented in the previous section appears to fit the data well, and to be stable over the period examined, it is worthwhile to examine as well the alternative formulation described in the previous section. In this formulation, it is assumed that the net effect of the female wage on fertility has changed over time due largely to a *declining price effect*, as purchased childcare became widely accessible and socially acceptable. In this alternative formulation, the interaction term between the female wage and *RY* is replaced by a time trend and an interaction between the time trend and the female wage. The model in this formulation is:

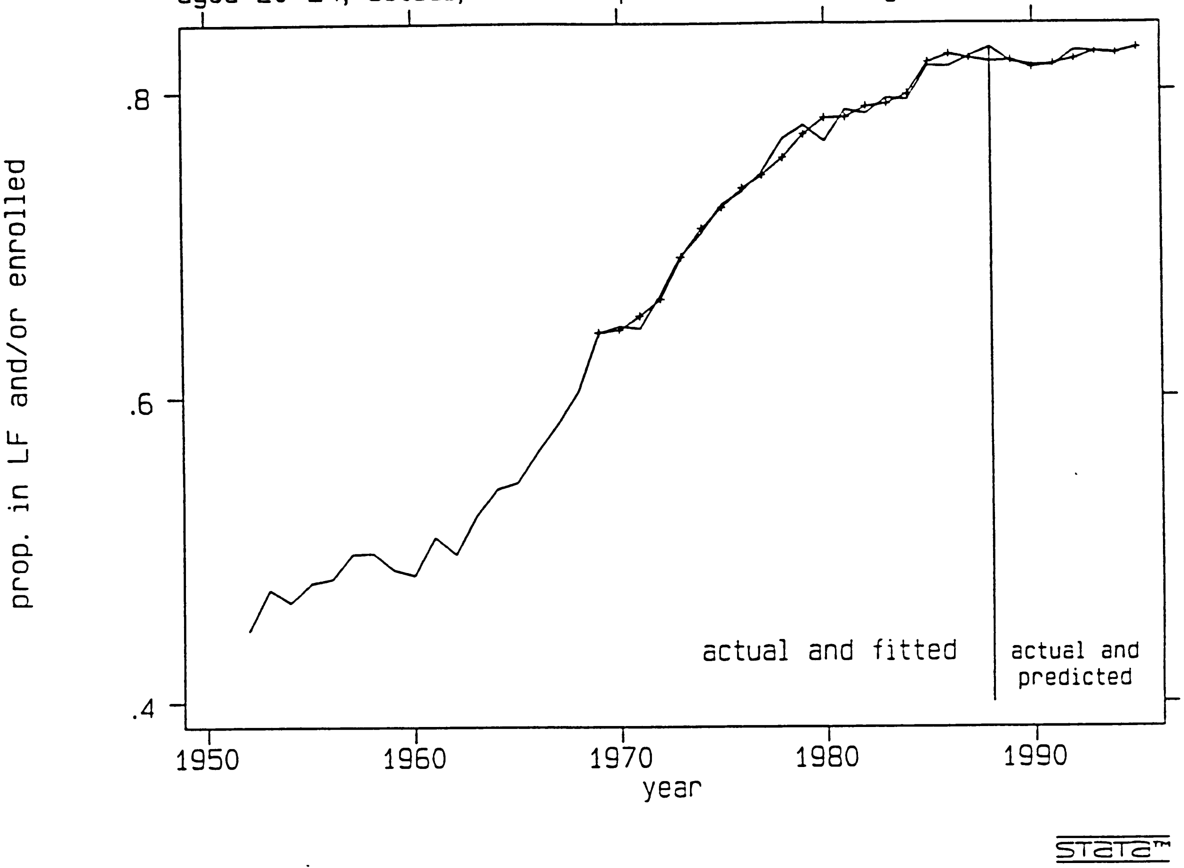
$$\Lambda(B_t) = \epsilon_0 + \epsilon_1 \log RY_{t-1} + \epsilon_2 \log W_{t-1} + \epsilon_3 U_{1,t-1} + \epsilon_4 time_{1963=0} + \epsilon_5 time_{1963=0} * \log W_{t-1} \quad (5)$$

The results of this alternative formulation are presented in column (1) of Table 4.

The results in Table 4 indicate that although the time trend and its interaction with the female wage are significant (the time trend with a negative effect, and the interaction term indicating an increasingly positive effect of the female wage over time), this specification does not explain as much variance as the formulation presented in Table 1 (96% as compared to 99%). However, the results presented in column (2) of Table 4 indicate that the time trend and its interaction with the female wage might be thought to retain some significance (*t*-statistics of 1.6) even when included in the full model presented in equation (1). The model in this formulation is:

$$\Lambda(B_t) = \eta_0 + \eta_1 \log RY_{t-1} + \eta_2 \log W_{t-1} + \eta_3 U_{1,t-1} + \eta_4 time_{1963=0} + \eta_5 time_{1963=0} * \log W_{t-1} + \eta_4 \log RY_{t-1} * \log W_{t-1} \quad (6)$$

Fig 5: Proportion in labor force and/or enrolled, Women aged 20-24, actual, fitted & predicted -- using 1969-88 data



In this formulation, the time trend and its interaction with the female wage do appear to add some explanatory power to the regression, suggesting that there might be validity to both of the hypothesized reasons for an increasingly positive net effect of the female wage on fertility.

**Female Labor Force Participation and Enrollment Rates**

The results of our simple model of female labor force participation are presented in Table 5 and in Figure 5, where it can be seen that once again *RY* and the female wage appear to explain very well observed trends over the past twenty-five years. The model accounts for about 99% of the variance in the time series during this period, with *RY* exerting a strong negative effect on female labor force participation, consistent with Easterlin's theory, and the female wage exerting a positive effect. However, there is in addition a positive effect of lagged female labor force participation rates, as indicated by the significant positive coefficient on that variable. At the same time, as indicated by the significant negative coefficient on the interaction between lagged *LFPE* and the female wage, the positive effect of the female wage on labor force participation appears to have been diminishing over this period, due to the strengthening income effect of the wage as hours worked increase.

This model, like the fertility model, is extremely robust, retaining all coefficients virtually unchanged when estimated for the subperiod 1969-1988 (columns 3 and 4 in Table 5), and when estimated using a Cochrane-Orcutt correction for either period (columns 2 and 4 in Table 5). The fitted data presented in Figure 5 were estimated using the model as estimated with only 1969-88 data: the years 1989 through 1995 are an in-sample prediction.

**Implications for Fertility in the Future**

It is, of course, dangerous to place too much reliance on models fitted on only 25 time series observations -- but in the spirit of this Conference it seems necessary to explore their ramifications. In order to do so, the fertility model in equation (1) has been estimated using two different levels of *RY* (depicted in Figure 6a) in combination with three different levels of the female wage in the period up to the year 2000. *RY* is allowed to rise to about 0.34, and to fall to about 0.22, from its 1993/4 level of 0.29, and the female wage is allowed to rise to $8.92 and fall to $6.92 from its 1993/4 level of about $8.10 (all in

Fig.6a: Actual, Fitted and Projected Male Relative Income
with variation in relative cohort size and net imports

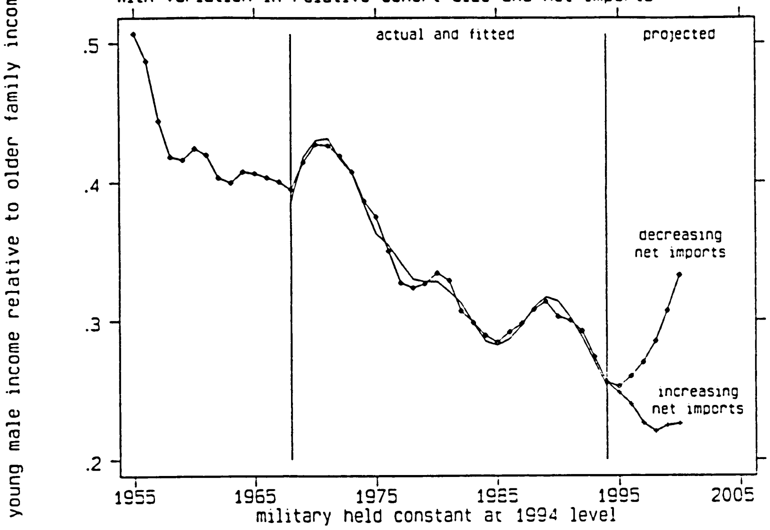



Fig 6b: Actual and Projected Fertility Using Eq.1, Age 20-24
assuming continued decline in male relative income

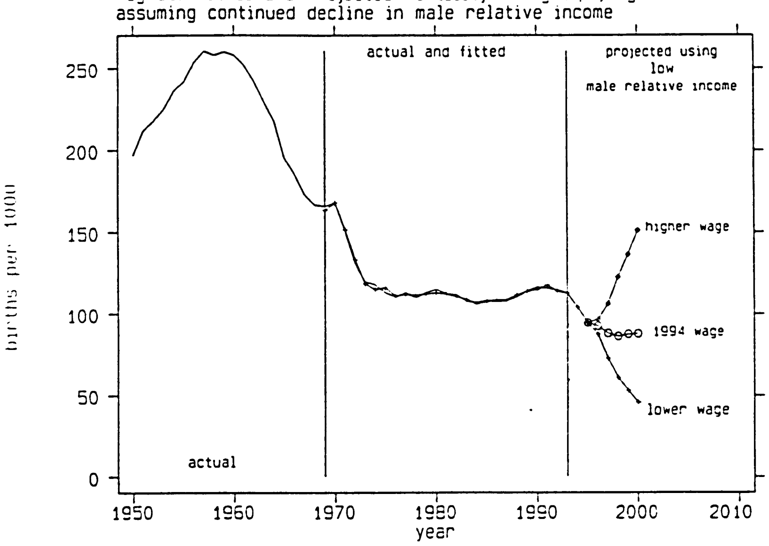



Fig 6c: Actual and Projected Fertility Using Eq.1, Age 20-24
assuming future increases in male relative income

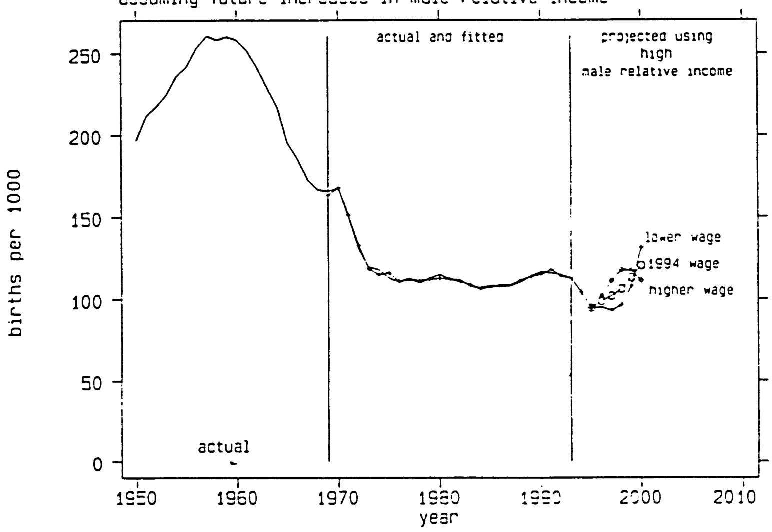
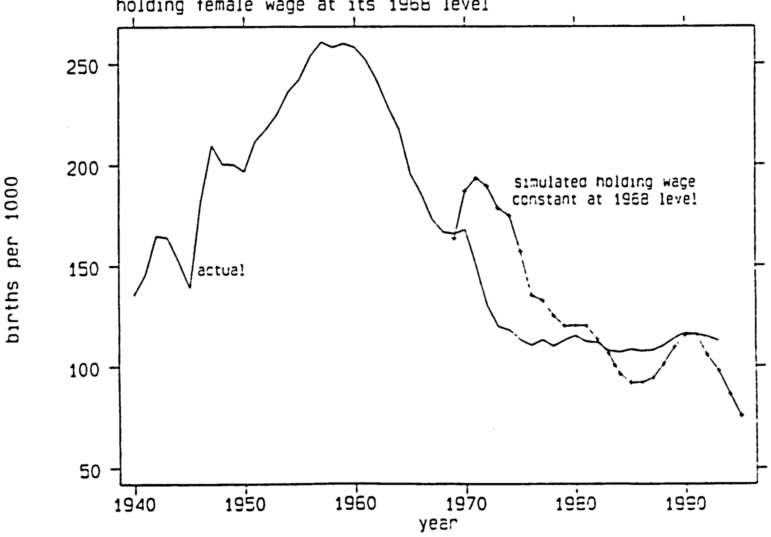



Fig. 6d: Actual and Simulated Fertility Using Eq.1, Age 20-24
holding female wage at its 1968 level



constant 1994 dollars),[18] as well as to remain at its 1993/4 level. The results of this estimation are presented in Figures 6b and 6c.

A comparison of these two figures indicates that variations in the effect of the female wage on fertility are greatest when *RY* is low: these are the times when the income effect of the wage is at its strongest. In this scenario, fertility (in terms of births per thousand women aged 20-24 in the year 2000) varies from about 45 (with the low wage) to about 150 (with the high wage). In contrast, when *RY* is high this large variation in the female wage produces little variation in fertility: the range here is only from 110 to 130 births per thousand -- *and in this case the higher wage produces the lower level of fertility*, because of the wage's reduced income effect when *RY* is higher.

In results not presented here, similar projections were prepared using equation (5) in which the interaction between the female wage and *RY* is replaced with a time trend and its interaction with the female wage -- on the assumption that the most important factor in the changing income effect of the female wage on fertility, has been the increased availability/acceptance of purchased childcare. These results are similar to the ones presented above -- although more muted -- except that in this case the highest fertility (about 140 births per 1000 in the year 2000) occurs with a combination of high female wages and high *RY*, while the lowest (about 65 per 1000) occurs when a low female wage is combined with low *RY*.

**Speculating on What Might Have Been**

Again, in the spirit of generating renewed interest in the study of U.S. fertility, it might be useful to examine the realism of these findings by looking at the effect of the strong increase which occurred in the real female wage in the period prior to 1974, since this is the only major change we observe in the wage during this period. Figure 6d simulates U.S. fertility for the 20-24 age group through the 1970s and 1980s as it would have been had the female wage remained constant in real terms, at its 1968. At this lower level of the female wage, the effects of changing *RY* would have produced *higher* than observed fertility in the late 1960s and early 1970s (because the female wage exerted primarily a negative price of time effect when male relative income was so high). But without the buffering effect of a higher female wage in the 1980s, the U.S. would have experienced fertility levels comparable to those experienced in European countries -- on the order of 92 births per 1000 women aged 20-24 in 1985, as opposed to our low of 108.3 -- before reversing again in the late 1980s. Even more dramatic is the buffering effect of the female wage currently in the 1990s, when we would be observing fertility levels of about 75 per 1000 if the wage were at its 1968 level, because of the current low levels of *RY*.

**CONCLUSIONS**

The analyses presented in this paper have attempted to develop a comprehensive framework for understanding the dramatic changes in fertility, female labor force participation and female enrollments which we have observed in the 20-24 age group in the U.S. over the past 25 years. This has been accomplished using a blending of the two primary economic models which have been developed for that purpose: the Easterlin 'relative income' and the 'price of time' models. This combination, together with the assumption of a changing strength of the income effect of the female wage, has produced models which have extraordinarily good explanatory power for the period since the mid 1960s. The models are extremely stable, producing similar results using virtually any subset of data from the 1969-93 period, and the results are relatively insensitive to variations in the formulation of *RY* (male relative income).

The results presented here are strongly supportive of the hypothesis that *RY* has been a dominant influence on many of the most significant socio-economic changes we have observed in the past three decades. As *RY* declined, it prompted young adults to make significant changes in the 'traditional' patterns: delaying family formation, increasing the number of two-earner households, and increasing women's market-oriented human capital accumulation in anticipation of increased labor market attachment.

These results support economists' longstanding belief that children are *not* an inferior good, and that the opportunity cost of time needed to raise children is a significant factor in the fertility decision. They also indicate, however, that it is necessary to measure male income in *relative* rather than absolute terms -- and that it is essential to consider not only the price effect of the female wage, but also its income effect. Considerations of child quantity/quality tradeoffs become extraneous in the model, when relative rather than absolute male income is used, presumably because the issue of child quality is subsumed in the overall desired standard of living proxied in the denominator of *RY*.

In addition, the results presented here do not support the hypothesis of 'countercyclical fertility': although the price of time effect is strong, periods of high unemployment appear to have a stronger effect in disrupting a woman's expectations regarding future income streams, than they do in providing 'windows of opportunity' for pregnancy. Countercyclical fertility would only be a possibility in the absence of any income effect of a woman's wage on fertility.

Ironically, however, while the blending of these two models suggests that each model was correct in its primary insights, it also suggests that the effects they hypothesized for fertility have largely canceled each other out over the past twenty years. The female wage reached a level in the U.S. at which it was able to counterbalance the large fluctuations which we saw in *RY* during that period. This effect was so strong that it buffered the U.S. from the extremely low levels of fertility experienced in other Western countries: if the female wage had remained at its 1968 levels, we would have experienced fertility in the 20-24 age group approximately 25% lower than we did in the 1980s.

The model suggests that the effect of the female wage on fertility is strongest when *RY* is low. At low levels of *RY*, young people learn to look to the income-producing power of female earnings: the income effect on fertility of a woman's earnings increases at such times. We have observed this happening in our society over the last decade when *RY* has been the lowest: women who are on their own because of delayed, avoided and/or broken marriages look to their own earnings, rather than those of a male partner, in making childbearing decisions. Conversely, when *RY* rises, this income effect of the female wage diminishes.

This indicates that the direction of fertility over the next decade, as a result, is highly dependent on the path of the female wage, and on the reasons for its changing net effect on fertility. Three different assumptions have been tested in this paper, regarding these reasons:

- equation (1) assumed that the income effect of the female wage has changed over time solely as a result of changing male relative income: when $RY$ is low(high), the income effect of the female wage is small(large).

- equation (5) assumed that the price effect of the female wage has diminished over time as a result of increased availability/acceptance of purchased childcare.
- equation (6) assumed that both of these effects have occurred.

If the female wage rises again as it did in its heyday of the 1960s, in the context of current low levels of male relative income, and if equation (1) is the correct model, we could see a 50% increase in the fertility rate of women aged 20-24 (to about 150 per 1000) over the next decade, while a falling wage with falling male relative income could take us to fertility levels of only about 45 per 1000 women aged 20-24. If, on the other hand, equation (5) is the correct model, we would experience the highest levels of fertility (about 140 births per 1000 in the year 2000) if a *rising* female wage is combined with *rising* male relative income. It must be emphasized here that the wage under discussion is the real female hourly wage controlling for education, tenure and experience.

If these models are properly specified (admittedly a tenuous assumption, given the relatively short time period under consideration), we can expect marked changes in fertility in the 20-24 age group over the next decade, unless the female wage remains at its current level indefinitely. It appears then to be important to conduct further work to determine why the net effect of the female wage on fertility might have changed over the past few decades, in the context of a relative income model. Such research is difficult using aggregate time series, although a disaggregation by race would be informative.

This work -- with its highly speculative projections -- is presented in the hope that it will generate interest in research on this phenomenon. Work at the micro level -- possibly with data from the Panel Study of Income Dynamics, or the recently expanded Wisconsin Longitudinal Study of Social and Psychological Factors in Aspiration and Attainment -- would make a valuable contribution, to the extent that measures of both sets of parental incomes can be included in the analysis, and to the extent that macro level indicators of male relative income are included.

## NOTES

1 The 'price of time' or 'Chicago-Columbia' model is also known as the 'new home economics model'.

2 A referee has suggested that "the time series pattern of births to women aged 15-19 is essentially the same as the pattern for births to women aged 20-24. But the time-series patterns of relative income and female wages would correspond to cohorts born 5 years later," suggesting that the similarity in these patterns contradicts the theory. However in the following analyses, relative income and the female wage are calculated *for males and females in their first five years of work experience.* Thus, included in the calculation are college graduates at ages 22-26 as well as high school graduates at age 18-22 -- and for that matter, high school dropouts at even younger ages -- so it is in fact encouraging that the fertility of both 15-19 year olds and 20-24 year olds appear to have moved in tandem with these measures.

3 Data presented in this section, except for fertility rates, were developed by the author from March Current Population Survey data tapes. Fertility rates are taken from *Vital Statistics (Natality)*, and from the 1994 *Final Report on Births, Marriages, Divorces and Deaths.*

4 See the third main section of this paper for a full description of this wage and the method used to control for education.

5 The measure of relative income shown here is the expected real annual earnings in year $t$ (earnings multiplied by the age-specific employment rate) of unenrolled males in their first five years of work experience, divided by the real annual household income in year $t-5$ in all families with children and a head (of either sex) aged 45-54.

6 Specifically, in his aggregate approximations of the measure, he uses the average income of families with head aged 14-24 in a given five-year period to that of families with head aged 45-54 in the previous three-year period.

7 But see MacDonald and Douthitt (1992), for an encouraging start in this direction.

8 This is due to the fact that *total income* rises as hourly wages and hours worked rise -- and at the same time, because one is working longer hours, the marginal utility of non-work hours/activities rises. An individual in these circumstances has more income to spend, and one of the most important things to spend it on will tend to be time for non-work activities.

9 See, for example, the excellent reviews by Killingsworth and Heckman (1986) and Mroz (1987), and estimates by Nakamura, Nakamura and Cullen (1979), Nakamura and Nakamura (1981, 1985), Robinson and Tomes (1985), Smith and Stelcner (1985, 1988) and

Stelcner and Smith (1985).

10 For the years 1964-67, the source is the uniform series of March CPS files created under the direction of Robert D. Mare (University of Wisconsin) and Christopher Winship (Northwestern University) with financial support from the National Science Foundation through grant SOC-7912648. For female labor force participation rates, I have used the rates provided by John Stinson of the Bureau of Labor Statistics, in order to conform as closely as possible with published data.

11 For a more detailed description of the methodology used to derive wage and earnings figures for this analysis, see Macunovich (1993a).

12 This multiplication is not double-counting: it is necessary here because average annual earnings here have been calculated over the population who worked in a year and who had positive earnings in that year.

13 This is consistent with the argument presented by Oppenheimer (1976).

14 Macunovich (1993a) presents an appendix devoted to the presentation of results using alternative formulations of the male relative income variable, to test the sensitivity of results to the choice of age group in the denominator, and to the use of family income rather than older males' income in the denominator. Results using the earnings of married males with 25-34 years of work experience in the denominator, instead of family income, produced virtually identical results, in terms of both signs and significance. However, this formulation produced a DW statistic of 2.7, indicating the possibility of mis-specification. Models using different choices of age groupings in the denominator -- earnings of males with 30-39 years of work experience, or income of families with heads aged 40-49, for example -- indicated little sensitivity to such changes, within reasonable bounds. Similarly, the sensitivity of results to the use of the 'expectational' form of earnings (i.e., average earnings multiplied by the activity rate) was minimal: minor changes in levels of significance were the only noticeable effect (with the expectational form producing marginally more significant results).

15 It should be noted that the Chicago-Columbia model often includes some measure of the desired *quality* of children, and that in the relative income model material aspirations are often thought to be a function of *per capita* income in the parental home, with some models (see Behrman and Taubman 1990) assuming in addition that there will be an *intergenerational correlation in number of children*. These effects can be incorporated in this combined model simply by testing the significance of a variable measuring the average number of children in the parental home. That is, if we assume that one's desired child quality will be a function of one's material aspirations, then no additional control for desired child quality will be required, other than the measure of material aspirations. Per capita income in the parental home is a decreasing function of the number of children in that home, so that fertility in the second generation would be expected to be an increasing function of number of siblings based on this effect alone. Any intergenerational correlation in

preferences for children will be additive with this per capita income effect of siblings. These hypotheses were tested by including a variable indicating the number of children in the parental family, but the coefficient on this variable was found to be insignificant, in results presented and discussed in Macunovich (1993a).

16 Concern was expressed by a referee about the fact that there is "a severe multi-collinearity problem" in the model, in that the interaction term is highly correlated with male relative income (0.976) and the female wage (-0.66). However this is to be expected, by construction, in an interaction term since it is derived by multiplying the two original variables. As stated in Pindyck and Rubinfeld (1981:89), ". . .while [the coefficients on the multi-collinear terms] will remain unbiased estimators, the reliance that we can place on the value of one or the other will be small. . ." However, as stated in Judge *et al.* (1985:897), "Despite the difficulty in isolating the effects of individual variables from such a sample, accurate forecasts may still be possible even outside the sample. This is only true, however, if the pattern of interrelationships among the explanatory variables is the same in the forecast period as in the sample period." By construction, of course, the latter requirement is fulfilled with an interaction term.

17 The intercept term of 12.85 exhibited in column (2) of Table 1 indicates that if male relative income were equal to one (e.g., young males could exactly reproduce their parents' income, and therefore young couples' desired standard of living), the female wage were equal to only $1/hour ($1967) and female unemployment were zero (e.g., no shocks to people's expectations), there would be virtually a zero probability (a probability of about .0002%) that women aged 20-24 would attempt to restrict conception: we would be in a 'natural fertility' regime.

18 It should be noted that according to March Current Population Survey data, *RY* fell from 0.29 to 0.25, and the real female wage fell from $8.20 to $7.92 (in 1994 dollars), between 1993 and 1995. These declines are built into the estimates presented in Figure 6. An earlier version of this paper postulated increases in male relative income for this period -- but those figures assumed a constant or even an improving balance of trade, while the opposite has occurred over the past few years.

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**Table 1: Fertility Regression Results**
**Using Female Wage*Relative Income Interaction[1]**

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| **using data for:** | 1969-93 | | | 1975-93 | 1969-80 | |
| | OLS Levels | Cochrane Orcutt Levels | OLS First Differences | OLS Levels | OLS Levels | Cochrane Orcutt Levels |
| relative income: $logRY_{t-1}$ | 12.36 (10.6) | 12.85 (12.9) | 11.71 ( 5.0) | 13.62 ( 2.7) | 10.49 ( 4.6) | 10.71 ( 3.0) |
| interaction: $logRY_{t-1}*logW_{t-1}$ | -16.87 (10.3) | -17.56 (12.5) | -16.16 ( 4.6) | -18.71 ( 2.7) | -14.35 ( 4.5) | -14.66 ( 2.9) |
| female wage: $logW_{t-1}$ | -19.38 (12.2) | -19.90 (14.8) | -18.24 ( 5.6) | -20.86 ( 2.6) | -16.92 ( 5.7) | -17.18 ( 3.8) |
| unemployment: $U_{t-1}$ | - 1.61 ( 8.1) | - 1.59 ( 9.5) | - 1.47 ( 4.2) | - 1.33 ( 6.7) | - 2.26 ( 5.0) | - 2.22 ( 4.9) |
| intercept | 12.28 (10.9) | 12.65 (13.3) | - 0.00 ( 0.6) | 13.23 ( 2.3) | 10.51 ( 4.9) | 10.70 ( 3.3) |
| rho | | - 0.20 ( 1.0) | | | | - 0.33 ( 1.0) |
| # of obs | 25 | 24 | 24 | 19 | 12 | 11 |
| Adjusted $R^2$ | 0.988 | 0.991 | 0.776 | 0.815 | 0.989 | 0.992 |
| Durbin-Watson | 2.35 | | | 2.34 | 2.61 | |

[1] Dependent variable is the logistic transformation of age 20-24 fertility rate. Absolute values of $t$-statistics in parentheses.

**Table 2: Alternative Specifications of Fertility Regression in Table (1), Entering Numerator and Denominator of Relative Income Term Separately**[2]

| | (1) OLS Levels | (2) Cochrane-Orcutt Levels | (3) OLS Levels |
|---|---|---|---|
| male earnings: $logY_{t-1}$ | 9.91 | 8.51 | 0.92 |
| | ( 5.4) | ( 2.7) | ( 2.6) |
| family income: $logFY_{t-6}$ | -13.27 | -13.07 | - 0.17 |
| | (10.7) | (11.9) | ( 1.4) |
| interaction: $logFY_{t-6}*logW_{t-1}$ | 18.06 | 17.86 | |
| | (10.6) | (11.8) | |
| interaction: $logY_{t-1}*logW_{t-1}$ | -13.60 | -11.63 | |
| | ( 5.1) | ( 2.6) | |
| female wage: $logW_{t-1}$ | -56.72 | -71.01 | - 4.06 |
| | ( 3.1) | ( 2.3) | ( 5.7) |
| unemployment: $U_{t-1}$ | - 1.57 | - 1.57 | - 2.02 |
| | ( 8.1) | ( 9.5) | ( 4.3) |
| intercept | 46.06 | 50.53 | - 4.93 |
| | ( 3.1) | ( 2.3) | ( 1.4) |
| rho | | - 0.26 | |
| | | ( 1.2) | |
| # of obs | 25 | 24 | 25 |
| Adjusted $R^2$ | 0.990 | 0.992 | 0.935 |
| Durbin-Watson | 2.63 | | 1.02 |

[2] Dependent variable is logistic transformation of age 20-24 fertility rate. Absolute values of $t$-statistics in parentheses. Columns (1)-(3) present results in which the coefficients on the numerator and denominator of the relative income variable (and its interaction terms) are allowed to vary. Column (3) presents results in which the interaction term is omitted.

**Table 3: Illustrating the Changing Net Effect of the Female Wage on Fertility**[3]

*Panel 1:*

| year | coefficient on female wage, given male relative income |
|---|---|
| 1969 | -3.31 |
| 1970 | -4.67 |
| 1971 | -5.14 |
| 1972 | -5.21 |
| 1973 | -4.62 |
| 1974 | -4.21 |
| 1975 | -3.23 |
| 1976 | -2.29 |
| 1977 | -1.91 |
| 1978 | -1.29 |
| 1979 | -0.70 |
| 1980 | -0.60 |
| 1981 | -0.61 |
| 1982 | -0.22 |
| 1983 | 0.22 |
| 1984 | 1.03 |
| 1985 | 1.76 |
| 1986 | 1.94 |
| 1987 | 1.68 |
| 1988 | 1.13 |
| 1989 | 0.39 |
| 1990 | -0.02 |
| 1991 | 0.15 |
| 1992 | 0.79 |
| 1993 | 1.62 |
| 1994 | 2.69 |
| 1995 | 3.81 |

*Panel 2:*

| regression results for subperiods: | 1969-79 | 1974-93 |
|---|---|---|
| relative income: $logRY_{t-1}$ | 0.35 | 0.17 |
| | ( 2.6) | ( 4.7) |
| female wage: $logW_{t-1}$ | - 3.63 | 0.79 |
| | ( 9.8) | ( 2.0) |
| unemployment: $U_{t-1}$ | - 2.17 | - 1.16 |
| | ( 2.7) | ( 5.6) |
| intercept | 1.13 | - 2.34 |
| | ( 4.5) | ( 5.6) |
| # of obs | 11 | 20 |
| Adj $R^2$ | 0.972 | 0.785 |

[3] In Panel 1, the figures represent the total estimated effect of the female wage, taken to be the sum of the coefficient on the female wage plus the sum of the coefficient on the interaction term multiplied by the male relative income. That is, if

$$\Lambda(B_t) \quad = a*logRY_{t-1} + b*logW_{t-1} + c*logRY_{t-1}*logW_{t-1} + d*U_{1,t-1}$$

then we could also express this as

$$\Lambda(B_t) \quad = a*logRY_{t-1} + \propto*logW_{t-1} + d*U_{1,t-1}$$

where

$\propto = b + c*logRY_{t-1}$, and the values of $\alpha$ are presented in Panel 1.

In Panel 2, results of a simple regression excluding the interaction term, for subperiods at the beginning and end of the study period are presented. Dependent variable is the logistic transformation of age 20-24 fertility rate. Absolute values of $t$-statistics are in parentheses.

**Table 4: Alternative Specifications for Fertility Regression Using Female Wage*Time Interaction**[4]

| | using data for 1969-93 | |
|---|---|---|
| | **Eqn. (5)** | **Eqn. (6)** |
| time trend: $time_{1963=0}$ | - 0.16 | - 0.04 |
| | ( 4.3) | ( 1.6) |
| interaction: $logW_{t-1}*time_{1963=0}$ | 0.22 | 0.06 |
| | ( 4.3) | ( 1.5) |
| relative income: $logRY_{t-1}$ | 0.34 | 10.94 |
| | ( 2.8) | ( 7.4) |
| interaction: $logRY_{t-1}*logW_{t-1}$ | | -14.96 |
| | | ( 7.1) |
| female wage: $logW_{t-1}$ | - 5.20 | -17.95 |
| | (8.8) | ( 9.9) |
| unemployment: $U_{t-1}$ | - 1.57 | - 1.70 |
| | ( 3.1) | ( 6.3) |
| intercept | 2.20 | 11.23 |
| | ( 5.3) | ( 8.7) |
| # of obs | 25 | 25 |
| Adjusted $R^2$ | 0.961 | 0.989 |
| Durbin-Watson | 1.57 | 2.67 |

[4] Dependent variable is the logistic transformation of age 20-24 fertility rate. Absolute values of $t$-statistics in parentheses.

**Table 5: Female Labor Force Participation and Enrollment Regression Results (Eqn. 4)[5]**

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | using data for 1969-94 | | using data for 1969-88 | |
| | OLS | Cochrane-Orcutt | OLS | Cochrane-Orcutt |
| lagged LF and enrollment: $LFPE_{t-4}$ | 0.24<br>( 2.2) | 0.28<br>( 2.2) | 0.24<br>( 1.8) | 0.28<br>( 1.3) |
| interaction: $W_{t-1}*LFPE_{t-4}$ | - 0.01<br>( 3.8) | -0.01<br>( 3.6) | -0.01<br>( 2.9) | -0.01<br>( 2.8) |
| relative income: $RY_{t-1}$ | - 0.65<br>( 6.2) | -0.61<br>( 4.8) | -0.65<br>( 4.7) | -0.62<br>( 2.7) |
| female wage: $W_{t-1}$ | 0.29<br>( 5.4) | 0.27<br>( 5.8) | 0.28<br>( 4.3) | 0.27<br>( 4.0 |
| intercept | 0.23<br>( 3.8) | 0.21<br>( 2.5) | 0.24<br>( 2.8) | 0.27<br>( 1.4) |
| # of obs | 26 | 25 | 20 | 19 |
| Adjusted $R^2$ | 0.990 | 0.992 | 0.987 | 0.989 |
| Durbin-Watson | 2.42 | | 2.29 | |
| rho | | -0.14<br>( 0.7) | | -0.15<br>( 0.6) |

---

[5] Dependent variable is the proportion of women aged 20-24 who are in the labor force and/or enrolled in school. The interaction term is applied only in the years up to 1985, when female labor force participation rates were increasing and thus the income effect of the female wage would have been changing. Absolute values of $t$-statistics in parentheses.

## Appendix: Data Used in the Analysis

| year | $(B_t)$ | $logY_{t-1}$ | $logFY_{t-1}$ | $logRY_{t-1}$ | $logW_{t-1}$ | $U_{t-1}$ | $logN_{t-1}$ | $logP_{t-1}$ | $LFPE_t$ |
|---|---|---|---|---|---|---|---|---|---|
| 60 | . | . | . | . | . | . | . | . | 0.4837 |
| 61 | . | . | . | . | . | . | . | . | 0.5085 |
| 62 | . | . | . | . | . | . | . | . | 0.4974 |
| 63 | . | . | . | . | . | . | . | . | 0.5229 |
| 64 | . | . | . | . | . | . | . | . | 0.5404 |
| 65 | . | . | . | . | . | . | . | . | 0.5445 |
| 66 | . | . | . | . | . | . | . | . | 0.5647 |
| 67 | . | . | . | . | . | . | . | . | 0.5823 |
| 68 | . | . | . | . | . | . | . | . | 0.6031 |
| 69 | -1.616 | 8.191 | 9.143 | -0.952 | 0.615 | 0.063 | 0.716 | 1.483 | 0.6409 |
| 70 | -1.601 | 8.262 | 9.133 | -0.872 | 0.643 | 0.062 | 0.732 | 1.495 | 0.6455 |
| 71 | -1.734 | 8.306 | 9.150 | -0.844 | 0.673 | 0.070 | 0.737 | 1.500 | 0.6437 |
| 72 | -1.899 | 8.334 | 9.174 | -0.840 | 0.696 | 0.091 | 0.738 | 1.504 | 0.6660 |
| 73 | -1.995 | 8.320 | 9.195 | -0.875 | 0.719 | 0.093 | 0.738 | 1.508 | 0.6917 |
| 74 | -2.014 | 8.331 | 9.231 | -0.899 | 0.731 | 0.079 | 0.743 | 1.516 | 0.7060 |
| 75 | -2.060 | 8.325 | 9.282 | -0.957 | 0.723 | 0.087 | 0.738 | 1.519 | 0.7256 |
| 76 | -2.088 | 8.308 | 9.321 | -1.013 | 0.714 | 0.127 | 0.732 | 1.522 | 0.7336 |
| 77 | -2.061 | 8.312 | 9.347 | -1.036 | 0.717 | 0.112 | 0.726 | 1.524 | 0.7463 |
| 78 | -2.092 | 8.309 | 9.381 | -1.072 | 0.717 | 0.109 | 0.714 | 1.522 | 0.7680 |
| 79 | -2.062 | 8.300 | 9.407 | -1.107 | 0.722 | 0.096 | 0.703 | 1.519 | 0.7772 |
| 80 | -2.040 | 8.306 | 9.419 | -1.113 | 0.735 | 0.087 | 0.693 | 1.518 | 0.7663 |
| 81 | -2.068 | 8.317 | 9.430 | -1.113 | 0.741 | 0.089 | 0.680 | 1.517 | 0.7871 |
| 82 | -2.075 | 8.310 | 9.446 | -1.136 | 0.736 | 0.101 | 0.667 | 1.516 | 0.7846 |
| 83 | -2.113 | 8.290 | 9.452 | -1.162 | 0.731 | 0.113 | 0.654 | 1.517 | 0.7952 |
| 84 | -2.124 | 8.251 | 9.461 | -1.210 | 0.721 | 0.122 | 0.636 | 1.513 | 0.7938 |
| 85 | -2.108 | 8.219 | 9.472 | -1.253 | 0.714 | 0.102 | 0.615 | 1.510 | 0.8159 |
| 86 | -2.118 | 8.214 | 9.478 | -1.264 | 0.710 | 0.088 | 0.588 | 1.502 | 0.8154 |
| 87 | -2.112 | 8.226 | 9.474 | -1.248 | 0.710 | 0.091 | 0.560 | 1.492 | 0.8224 |
| 88 | -2.089 | 8.251 | 9.467 | -1.216 | 0.715 | 0.082 | 0.540 | 1.485 | 0.8280 |
| 89 | -2.053 | 8.289 | 9.461 | -1.172 | 0.728 | 0.081 | 0.519 | 1.476 | 0.8189 |
| 90 | -2.026 | 8.316 | 9.464 | -1.148 | 0.733 | 0.074 | 0.500 | 1.466 | 0.8169 |
| 91 | -2.034 | 8.318 | 9.476 | -1.157 | 0.734 | 0.057 | 0.487 | 1.454 | 0.8164 |
| 92 | -2.045 | 8.307 | 9.502 | -1.196 | 0.727 | 0.077 | 0.471 | 1.443 | 0.8260 |
| 93 | -2.064 | 8.289 | 9.534 | -1.245 | 0.715 | 0.070 | 0.455 | 1.429 | 0.8252 |
| 94 | . | 8.249 | 9.557 | -1.308 | 0.693 | 0.078 | 0.440 | 1.419 | 0.8240 |
| 95 | . | 8.199 | 9.574 | -1.375 | 0.679 | 0.084 | 0.433 | 1.411 | |

where:

$(B_t)$ = *logistic transformation of fertility rate for women aged 20-24*
$logY_{t-1}$ = *log of male earnings*
$logFY_{t-1}$ = *log of parental family income*
$logRY_{t-1}$ = *log of male relative income*
$logW_{t-1}$ = *log of female wage ($1967)*
$U_{t-1}$ = *unemployment rate for all women aged 20-24*
$logN_{t-1}$ = *log of number of children in parental family*
$logP_{t-1}$ = *log of number of persons in parental family*
$LFPE_t$ = *proportion of women aged 20-24 who are enrolled and/or in the labor force*

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