Explaining the Labor Force Participation of Women 20-24

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Abstract

Between the mid 1960s and the late 1970s there was a remarkable rise in the labor force participation of women and then a leveling off that has persisted through the mid 1990s. This paper attempts to explain the labor force participation of women 20-24 over this period. A variable is constructed measuring the potential wage rate of women 20-24 that can be taken to be exogenous to the labor supply decision, and a potential relative income variable is constructed, based on Easterlin's (1980) relative income hypothesis, that can also be taken to be exogenous. Both variables are estimated using Easterlin's "cohort wage" hypothesis, and both are found to be important in explaining labor force participation. The basic equation estimated does well in various tests that were performed on it, and it appears to explain well the rapid rise and then leveling off of the labor force participation of young women.

1 Introduction

Between the mid 1960s and the late 1970s there was a remarkable rise in the labor force participation of women and then a leveling off that has persisted through the mid 1990s. For example, Figure 1 presents a plot of the labor force participation of women aged 20-24 for the 1952.1–1995.3 period, where this pattern is quite apparent.[1] Can this pattern be explained using economic variables? Studies that have focused only

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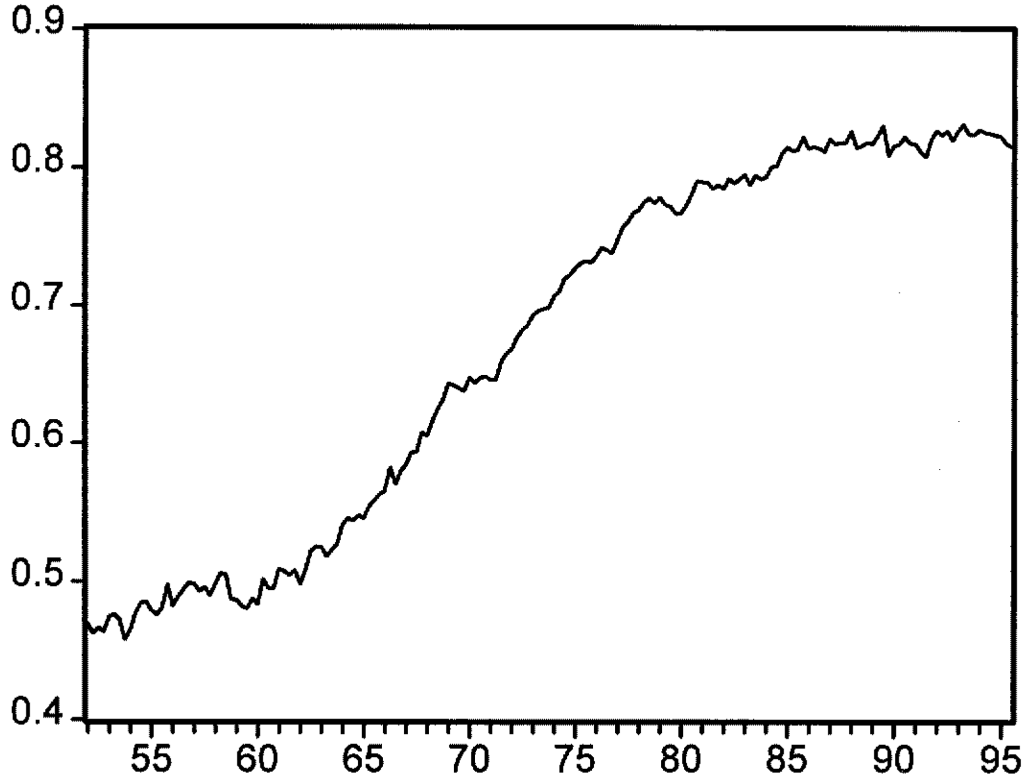
[1] The variable plotted in Figure 1 is the total labor force of women 20-24, including those in the armed forces and those enrolled in college, divided by the total population of women 20-24.

FIGURE 1
Labor Force Participation of Women 20-24



Period: 1952:1-1995:3

on traditional wage-rate and income effects on labor force participation have been unable to account fully for the rise in the 1960s and 1970s. For example, Smith and Ward (1985) were able to account for only 58 percent of the total increase in the labor force participation of women 20-64 between 1950 and 1980 using the female wage rate and male income. Also, some have questioned the general view that the sharp rise in female labor force participation was driven primarily by the large rise in the female wage rate. Killingsworth and Heckman (1986), for example, point to evidence that "the compensated and uncompensated wage elasticities of women workers are little different from those of men; indeed, in this work, the female uncompensated elasticity is often estimated to be negative." Others—for example, Mroz (1987)—have questioned the exogeneity of wage rates used to obtain the various estimates of wage-rate elasticities. The wage rates used have frequently been uncorrected—or improperly corrected—for changing levels of education and work experience.

This paper examines whether the rise and the subsequent leveling off can be accounted for by 1) using a *relative potential* income measure, in the spirit of Easterlin's (1980) relative income hypothesis, in place of more traditional absolute income measures and 2) using a *potential* wage-rate variable in place of more traditional wage-rate variables. As discussed below, both our relative potential income variable and our potential wage-rate variable can be taken to be exogenous to the labor supply decision. The construction of these variables is based on another Easterlin hypothesis, which will be called the "cohort wage" hypothesis.

Easterlin (1980, p. 42) defines the "relative income" of a couple to be the ratio of the earnings potential of the couple to the material aspirations of the couple. In general terms the relative income hypothesis states that a change in relative income leads young adults to make various adjustments in their lifestyles. A reduction in the relative income of young adults, for example, may lead to postponement of marriage and family formation and when marriage occurs to an increased tendency toward the formation of two-earner households. A fall in relative income of young adults may thus lead, among other things, to an increase in their labor force participation because it will tend both to increase the proportion of young women who are single (and thus have a higher participation rate) and to increase the participation of married women.[2]

Easterlin's cohort wage hypothesis is that relative cohort size affects relative wage-rate potential. For example, a large relative cohort size, such as exists for baby boomers, has, according to this hypothesis, a negative effect on the cohort's wage-rate potential relative to that of other cohorts. Easterlin suggested crowding—in the home, the school, and the labor market—as the basis of this effect, and researchers such as Welch (1979) have found supporting evidence of imperfect substitutability between older and younger workers. We use this hypothesis in a key way below in constructing our relative potential income and potential wage-rate variables.

We focus our attention in this paper on women aged 20-24. A labor force par-

[2] A few other studies have attempted to examine relative income effects. Fair and Dominguez (1991) set out to test Easterlin's hypothesis as part of a larger analysis of age distribution effects in macroeconomic models, but their equations do not contain a relative income term, only an absolute wage-rate term that they allow to vary with cohort size. Wachter (1972) attempted to approximate a relative income term in labor supply equations for secondary workers by using the ratio of the current aggregate wage rate to a ten year moving average of the same aggregate wage rate. This approximation missed one aspect of Easterlin's theory, which we discuss next, namely the effects of cohort size on age-specific wage rates. Devaney (1983) found a significant negative effect of relative income on the labor supply of women 20-44 between 1957–1977. Her definition of relative income, which differs from Easterlin's, is the deviation from trend of young males' age-specific income. Shapiro (1988) found a negative effect of relative income on the labor force participation of women 25-34 between 1950–1985, although he constrained the effect of the female wage rate to be zero in his equations.

ticipation equation is estimated and tested for this group. The sample consists of quarterly time series data for the 1952.1–1995.3 period, and the estimation period is 1956.4–1995.3. We include in the labor force those in the armed forces and those enrolled in college. Those in the armed forces are clearly participating in the labor force, and so they should be counted. Most of those enrolled in college are likely to enter the labor force, so they were also counted.

Regarding the definition of relative income, Easterlin (1980, p. 42) proposes to approximate the ratio of the earnings potential of a couple to the material aspirations of the couple by the ratio of the "recent income experience of [a] young man" to the "past income of [the] young man's parents." The use of past income of parents is based on the idea that material aspirations depend on the standard of living of parents and are formed when people are still living at home.

2 The Model

As noted above, we are dealing with quarterly time series data. Let f denote female, m male, i age group i, and t quarter t. Let L_{fit} be the labor force participation rate of women in age group i, and let W_{fit} and W_{mit} be some measures of the average real potential wage rate of women and men, respectively, in age group i. Finally, let Q_t be some measure of aggregate labor market tightness. As noted above, we focus on the age group 20-24, which will be denoted age group 1. Age group 2 will be taken to be the age group of the parents of people in age group 1.

The following labor force participation equation is postulated for women 20-24:

$$\begin{aligned} \log L_{f1t} &= \alpha_0 + \alpha_1 \log L_{f1t-1} + \alpha_2 Q_t + \beta_1 \log W_{f1t} \\ &\quad + \beta_2 \log(W_{m1t}/W_{m2t-r}) + \epsilon_{1t} \end{aligned} \tag{1}$$

The lagged dependent variable and Q_t are entered to pick up dynamic and aggregate cyclical effects, a common procedure in the specification of labor force participation equations in the macro literature. The two other explanatory variables are the own potential wage rate, W_{fit}, and the ratio of the potential wage rate of men aged 20-24 at time t to the potential wage rate of men in age group 2 at time $t-r$, W_{m1t}/W_{m2t-r}.

We take W_{m1t}/W_{m2t-r} as an approximation to Easterlin's concept of relative income, using potential wage rates in place of potential income.[3] For the rest of this paper we will call W_{m1t}/W_{m2t-r} "potential relative income." For the main results below we have assumed that the average age of parents at the birth of their children is 30 and that the material aspirations of children are formed at age 18. If we take the average age of people in our sample (ages 20-24) to be 22, then this group's material aspirations are assumed to have been formed on average four years (16 quarters) ago. The value of r in equation (1) is thus 16. Age group 2 is 46-50, since this is the age range of the parents four years before time t, when their children were on average aged 18. Although we have used the potential wage rate of men in both the numerator and denominator of the potential relative income variable, it will be seen below that our results do not really discriminate between the use of men versus the use of men plus women.

Equation (1) cannot be directly estimated because quarterly data on potential wage rates by age groups are not available. Data are available on the actual aggregate wage rate, which we will denote W_t, and on the percentage of people of age group

[3] Regarding the use of potential wage rates in place of potential income, the observed trends in these two variables are remarkably similar during this period, for older men and for families with older heads. Work in Macunovich (1996) using annual March Current Population Survey data indicates that the use of older males' earnings in place of family income in the denominator of the relative income term has only a small effect on estimated coefficients in equations explaining fertility, college enrollment, and labor force participation of women 20-24 for the period 1963–1993.

i in the total population, which we will denote p_{it}. How does one go from data on W_t and p_{it} to data on W_{fit} and W_{mit}? We do this by using Easterlin's cohort wage hypothesis. In particular, we postulate that

$$\log(W_{fit}/W_t) = \gamma_{0i} + \gamma_1 logp_{it}, \qquad \gamma_1 < 0 \tag{2}$$

$$\log(W_{mit}/W_t) = \gamma_{0i}^{'} + \gamma_1^{'} logp_{it}, \qquad \gamma_1^{'} < 0 \tag{3}$$

Equations (2) and (3) state that a cohort's potential wage rate relative to the aggregate wage rate is a negative function of the relative size of the cohort, other things equal. This is consistent with the imperfect substitutability between cohorts by age/experience as identified by Welch (1979). People in relatively large cohorts have relatively small potential wage rates. It is important to realize that equations (2) and (3) pertain to potential, not actual, wage rates. A relatively large cohort has a relatively small potential average wage rate, but not necessarily a relatively small actual average wage rate. We are interested in the potential wage rate since it is independent of any adjustments individuals might make in response to it—adjustments that will change the actual (observed) wage rate. We are assuming that this potential wage rate, estimated as a function of cohort size, provides us with an exogenous variable that can be used as an explanatory variable in the labor force participation equation.

Using equations (2) and (3), equation (1) becomes:

$$\begin{gathered}\log L_{f1t} = [\alpha_0 + \beta_1\gamma_{01} + \beta_2(\gamma_{01}^{'} - \gamma_{02}^{'})] + \alpha_1 \log L_{f1t-1} + \alpha_2 Q_t + \beta_1 \log W_t \\ + \beta_1\gamma_1 \log p_{1t} + \beta_2 \log W_t + \beta_2\gamma_1^{'} \log p_{1t} - \beta_2 \log W_{t-r} \\ - \beta_2\gamma_1^{'} \log p_{2t-r} + \epsilon_{1t}\end{gathered} \tag{4}$$

Sufficient data are available to estimate this equation. In other words, the use of equations (2) and (3) allows us to estimate the β coefficients in equation (1) using only data on the aggregate wage rate and on the proportions of the age groups in the total population.

Our formulation does not suffer from the usual wage-rate endogeneity problems that haunt the labor supply literature.[4] First, the aggregate wage rate, W_t, can be taken to be exogenous to the labor supply decision. This is because, since women 20-24 make up small fraction of the total population, their decisions regarding education and labor force participation have a trivial effect on the aggregate wage rate. Second, p_{1t} can be taken to be exogenous, since it is not affected by education and labor force decisions.

We have gotten around the wage-rate endogeneity problem by using the potential wage rate rather than the actual wage rate in equation (1). W_{f1t} is meant to measure how women 20-24 perceive their labor market opportunity. If, for example, they are in a large cohort, they perceive a lower opportunity than do those in a smaller cohort (conditional on the aggregate wage). This perception then influences their decisions regarding education, labor force participation, family formation, and the like. These

[4]Endogeneity is a problem in this literature because an increase in labor force participation rates may induce higher levels of human capital accumulation and bring about higher average levels of experience and tenure, thus leading to higher average wages. In explaining the labor supply of women 20-44 for the 1957–1977 period, Devaney (1983) attempted to get around this problem by using a predicted female wage. However, the base series used for her wage regression were those prepared by Butz and Ward (1979), which Macunovich (1995) has found to be flawed due to the need to estimate a female wage rate using total annual income of all women (both in and out of the labor force) and average hours worked of all workers (both male and female) in the retail trade. Trends in the two data series do not follow the actual trends for female workers. Perhaps as a result, Devaney found an insignificant effect of the female wage. Blau and Grossberg (1991) used the median annual income of all women working year round full time as their proxy for the wage of married women, and they attempted to control for the endogeneity of this measure by using a 2SLS estimation procedure in which the percentage of the female labor force with 4+ years of college—a variable that might be considered endogenous—was used as an exogenous regressor.

decisions in turn influence the actual wage rates that they receive (so the actual wage rates are endogenous), but not W_{f1t}. Similar arguments apply to our use of potential relative income. Our potential relative income variable is meant to measure how women and men 20-24 perceive their relative income opportunity, which affects their decisions, which affect their actual relative income. These decisions do not, however, affect our potential relative income variable. We have thus constructed wage-rate and potential relative income variables that may affect the labor force decisions of women 20-24, but that are not themselves affected by these decisions.

Note that in (2) and (3) we have allowed for the possibility that the cohort effect may differ for women and men, i.e., γ_1 may differ from $\gamma_1^{'}$. The hypothesis that the two are equal is tested below.[5] Note that $\gamma_1^{'}$ appears in equation (4) because we have used the male potential wage rate in the relative income variable in equation (1). If instead we had used the female potential wage rate, γ_1 would replace $\gamma_1^{'}$ in equation (4) and we would not need equation (3). If γ_1 is equal to $\gamma_1^{'}$, then we cannot distinguish between the use of the male versus female potential wage rates in the relative income variable because equation (4) is the same in both cases.

3 The Data

The population data by age and sex, which are needed to create p_{it}, are from the Bureau of the Census. Prior to 1980 the data are annual and are from the *Current*

[5]Easterlin is ambiguous regarding the relative size of cohort effects for men and women. In his original 1980 text, which is reproduced on page 27 of his 1987 edition, he refers to a larger effect of cohort size on the earnings of full time, full year workers for females aged 20-24 than for males aged 20-24 during the years 1955–1977. But on page 171 in his 1987 "Epilogue" he indicates a smaller effect on similar earnings for women aged 25-34 than for men aged 25-34 during the years 1968–1982. Fortunately, we do not have to take a stand on this issue because we can estimate both γ_1 and $\gamma_1^{'}$.

Population Reports, Series P-25, Numbers 311, 519, and 917. Quarterly data were created from these data by interpolation. The first quarterly observation was for 1952:3. Since 1980 the data are either quarterly or monthly and are available on diskette and on the Web (Series PPL-21). The population series used from PPL-21 was "resident population plus Armed Forces overseas."

Regarding the labor force data, a few years ago the Bureau of Labor Statistics (BLS) stopped publishing data on the total labor force and total noninstitutional population. It now publishes only the civilian counterparts to these. We need data on the labor force of women 20-24 including those in the armed forces, and we constructed these data as follows. Our starting point was the most recent data on the civilian labor force and civilian noninstitutional population of women 20-24 from the BLS. These data are monthly, and we used them for January 1952 through September 1995.[6] For years prior to 1980 we added to these numbers the old BLS estimates of the number of women 20-24 in the armed forces. This created a total labor force series and a total noninstitutional population series for the period prior to 1980. For 1980 on, we added armed-forces estimates that we were able to calculate from the Census data mentioned in the previous paragraph.[7]

Finally, we need to add to the total labor force the number of women 20-24 enrolled in school who are not counted in the traditional labor force. We used the March Current Population Survey public use microdata to get the proportion of civilian noninstitutionalized women 20-24 who were enrolled in college but who were not in the traditional labor force. This information was available annually

[6] From 1971:4 back, the civilian noninstitutional population data were multiplied by 1.0157 to splice these data to the data beginning in 1972:1.

[7] The armed forces figure for each age was computed as the difference between the "resident population plus Armed Forces overseas" and the "civilian population."

for the 1964–1993 period. We supplemented this information with enrollment data from the *Current Population Reports*, Series P-20, for the years prior to 1964 by assuming a constant ratio between total enrollment (which includes some people in the traditional labor force) and enrollment of those not in the traditional labor force. We took the values for 1994 and 1995 to be the same as the value for 1993. The annual proportions were interpolated to obtain quarterly values, which were then multiplied by the quarterly civilian noninstitutional population values to obtain quarterly enrollment values. L_{f1t} is the ratio of the sum of the total labor force and enrollment to the total noninstitutional population. From now on we will call L_{f1t} the "labor force participation rate," where participation includes people enrolled in school.

The aggregate wage rate used (W_t) is variable WA/PH in Fair (1994), updated through the third quarter of 1995. It is a real, after tax wage rate. WA is constructed as total after-tax compensation of workers in the economy divided by total paid hours adjusted for overtime. PH is a price index for household expenditures. The labor market tightness variable (Q_t) is variable Z in Fair (1994). Z is constructed as $min(0, 1 - JJP/JJ)$, where JJ is the ratio of the total number of paid hours in the economy to the total population 16 and over, and JJP is a series constructed from peak-to-peak interpolations of JJ. Z is a labor constraint variable in the sense that it is zero or close to zero when the aggregate worker hours-population ratio is at or near its peak and gets progressively larger in absolute value as the ratio moves below its peak.

4 The Results

The Use of 2SLS

In the estimation of equation (4) we have treated the aggregate wage variable, $\log W_t$, and the labor market tightness variable, Q_t, as endogenous. It may be that aggregate shocks contemporaneously affect these variables and the error term, ϵ_{1t}, in the equation. The equation was thus estimated by two stage least squares (2SLS), where the first stage regressors that were used are the main predetermined variables in the US model in Fair (1994).[8]

It is important to note that the present use of the 2SLS estimator is not an attempt to get around the standard wage-rate endogeneity problem in the labor supply literature. We have done this by the use of the potential wage-rate variables as discussed above. Rather, the use of the 2SLS estimator is just to account for the possibility that, say, some aggregate shock affects both W_t and ϵ_{1t}. This might happen even though the labor supply decisions of women 20-24 have a trivial affect on W_t and thus W_t can be treated as exogenous to these decisions.

The Basic Equation

The results of estimating equation (4) by 2SLS are presented in the top half of Table 1. The estimation period begins in 1956:4, which is the first quarter available for estimation given the need for lagged values. The equation is nonlinear in coefficients and was estimated using nonlinear 2SLS.

The estimate of the lagged dependent variable coefficient (α_1) in Table 1 is .825

[8] The first stage regressors include: $\log L_{f1t-1}$, $\log p_{1t}$, $\log W_{t-r}$, $\log p_{2t-r}$, $\log W_{t-1}$, a constant, a time trend, and a number of lagged endogenous variables in the US model in Fair (1994). The right-hand side endogenous variables in equation (4) are Q_t and $\log W_t$.

Table 1
2SLS Results for Equation (4)

$$\log L_{f1t} = cnst + \alpha_1 \log L_{f1t-1} + \alpha_2 Q_t + \beta_1 \log W_t$$
$$+\beta_1\gamma_1 \log p_{1t} + \beta_2 \log W_t + \beta_2\gamma_1^{'} \log p_{1t}$$
$$-\beta_2 \log W_{t-r} - \beta_2\gamma_1^{'} \log p_{2t-r} + \epsilon_{1t}$$

Coef.	Coef. Est.	t-stat.	
cnst	-.111	-2.29	
α_1	.825	21.74	
α_2	.044	0.96	
β_1	.207	4.51	
β_2	-.105	-2.93	
γ_1	-.204	-3.34	
$\gamma_1^{'}$	-.388	-1.92	
SE	.00914		
R^2	.9976		
DW	2.10		
χ^2 Tests:	χ^2	**df**	**p-value**
Lags	10.45	6	.107
$RHO = 4$	7.90	4	.095
T	3.16	1	.075
Leads +1	1.75	1	.186
Leads +2	4.49	2	.106
$\log PH_t$, $\log PH_{t-16}$	3.94	2	.140
$\log W_t^*$	2.99	1	.084
Stability Test:			
AP	T_1	T_2	λ
6.21	1972:1	1980.4	2.492

Estimation period is 1956.4–1995.3
PH = price level
W^* = 40 quarter moving average of the real wage rate

and is highly significant. The estimate of the coefficient of the labor market tightness variable (α_2) is positive, as expected, but not significant, which suggests a small or non existent cyclical effect on the labor force participation of women 20-24.

The estimates of γ_1 and $\gamma_1^{'}$ are negative, as expected, and significant, with the coefficient for men ($\gamma_1^{'}$) being larger in absolute value. The cohort effect on the potential wage rate is thus estimated to be larger for men than for women. The estimate of β_1, the own potential wage-rate coefficient, is positive (.207). The long-run potential wage-rate elasticity is 1.18 $(.207/(1-.825))$.[9] The estimate of β_2, the coefficient of the potential relative income variable, is negative ($-.105$) and significant. The long run elasticity is $-.60$. Women 20-24 are thus estimated to participate more when their own potential wage rate rises and when potential relative income declines.

χ^2 Tests of the Equation

It is important to see how well equation (4) does in various tests. Various single-equation χ^2 tests are presented in the second half of Table 1. These tests, which are discussed in Fair (1994, Chapter 4), consist of adding various variables to the equation and testing whether the addition is significant. In the following discussion a χ^2 value will be said to be insignificant if its p-value is greater than .05. An insignificant χ^2 value means that the equation has passed the test.

The first test is to add the lagged values of all the explanatory variables to the

[9] How does the elasticity of 1.18 compare to those from previous studies? Few studies have estimated wage-rate elasticities using time series data; the usual approach is to estimate them using cross section data. Also, of those that have done so, such as those mentioned in the last footnote, absolute income rather than relative income has been used as the income variable and so the estimated elasticities may be biased. Other studies have also not handled the endogeneity problem in the way we have. For what it is worth, however, the 1.18 estimate accords fairly well with other results. Blau and Grossberg (1991) estimated wage-rate elasticities ranging from .96 to 1.35 for married women in the period 1956–1986. Smith and Ward (1985) estimated a wage-rate elasticity of .82 for women 20-64 between 1950 and 1980. Goldin (1991, p.152) estimated a wage-rate elasticity "greater than one in absolute value" (and positive) for married women from 1890 to 1980. (These latter two studies used pooled cross-section time series data.)

equation and test their joint significance. This addition encompasses many different types of dynamic specifications, and so it is a fairly general test of the dynamic specification of the equation. The variables added are log L_{f1t-2}, Q_{t-1}, log W_{t-1}, log W_{t-17}, log p_{1t-1}, and log p_{2t-17}. As can be seen in the table, the χ^2 value is not significant, and so the test is passed.

The second test is to estimate the equation under the assumption of a fourth order autoregressive process of the error term, another test of the dynamic specification. This test was also passed.

The third test is quite important in the present context; it is to add a time trend to the equation. This is a test to see if there is a trend in the labor force participation of women 20-24 that has not been accounted for by the variables in the equation. Again, this test was passed, which means that the time trend was not significant. This result suggests that the trend in the labor force participation of young women is accounted for by the female potential wage rate and the potential relative income variable.

For the next two tests values of the aggregate wage rate one or more periods *ahead* were added to the equation. These tests can be looked upon as tests of the expectation mechanism. If the forward values are significant, this is evidence in favor of the rational expectations hypothesis.[10] The forward values are not significant, and so the two tests are passed.

The aggregate wage-rate variable (W_t) used in this paper is a real wage-rate variable, variable WA/PH in Fair (1994), where WA is the nominal wage rate and PH is the price level. For the sixth test, the logs of PH unlagged and lagged 16

[10] See Fair (1994, Chapter 4) for a discussion of this test. This test requires that the equation be estimated using Hansen's (1982) method of moments estimator, which was done here.

quarters were added to see if the restrictions imposed by the use of the real wage rate rather than the nominal wage rate and the price level separately was supported by the data. The test was passed, and so the real wage-rate restriction is supported.

The seventh test is a test of Wachter's (1972) relative income term, which is the ratio of the current aggregate wage rate to a ten year moving average of the same aggregate wage rate. From Wachter's perspective, the ten year (40 quarter) moving average of W_t belongs in our equation. Call this 40 quarter moving average W_t^*. As can be seen, when the log of W_t^* was added to the equation, it was not significant, and so this test is also passed. (The estimation period in this case began in 1963:1 to handle the lagged values.)

The last part of Table 1 presents results of a stability test. This test is due to Andrews and Ploberger (1994) and is also discussed in Fair (1994, Chapter 4). This test does not require that a break point be chosen *a priori*, just a range in which the structural break occurred if there was one. The range used for this test was 1972:1–1980:4. The AP value was 6.21, which is not significant at the 5 percent level, and so the stability test of no break is passed.

These test results are thus quite favorable to the equation. The equation seems to have adequately captured dynamic and trend effects, and it seems stable over time.

Other Tests

Equation (4) was also estimated under the assumption that material aspirations are formed at age 17 rather than 18. This means that r is 20 rather than 16 and that age group 2 is 45-49 rather than 46-50. The estimation period for this work began in 1957:4 instead of 1956:4 to account for the longer lags. The basic equation was also reestimated for the shorter estimation period for comparison purposes. The results for the two versions were very similar, with the basic version having a slightly smaller standard error. The overall conclusions are not sensitive to the use of $r = 16$ versus $r = 20$.

We also tested the hypothesis that γ_1 equals $\gamma_1^{'}$, which is the hypothesis that the cohort effect on the potential wage rate of men and women is the same. When equation (4) was estimated under this restriction the estimate of γ_1 (and $\gamma_1^{'}$) was $-.152$ with a t-statistic of -2.48. This estimate compares to the separate estimates of $-.204$ and $-.388$, respectively. Imposing the restriction had only a small effect on the other coefficient estimates. The estimate of β_1 was .194 (t-statistic = 4.27), which compares to .207 in Table 1, and the estimate of β_2 was $-.122$ (t-statistic = -3.56), which compares to $-.105$ in Table 1. Testing the hypothesis of equality resulted in a χ^2 value of 2.32, which has a p-value of .136. The hypothesis is thus not rejected at the 5 percent level. This means, as discussed at the end of Section 2, that the data do not distinguish between the use of the male versus female potential wage rate in the relative income variable in equation (1). In spite of the fact that the hypothesis of equality was not rejected, we have chosen to focus on the equation without the restriction imposed in the next section. The separate estimates seem sensible, the other coefficient estimates are little affected, and the hypothesis is close to being rejected.

Men 20-24

In the initial work for this paper we tried postulating an equation like (1) for men 20-24, where the own wage was W_{m1t}. (We constructed data on the labor force participation of men 20-24 using the same methodology employed for women 20-24.) We do not report these estimates here because they were not satisfactory. For

Figure 2
Labor Force Participation Women 20-24: L_{f1t}

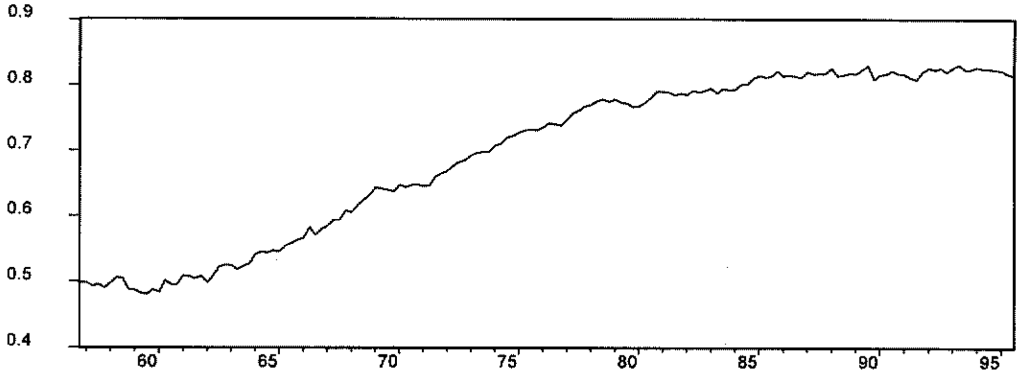


Figure 3
Estimated Opportunity Wage Rate of Women 20-24: W_{f1t}

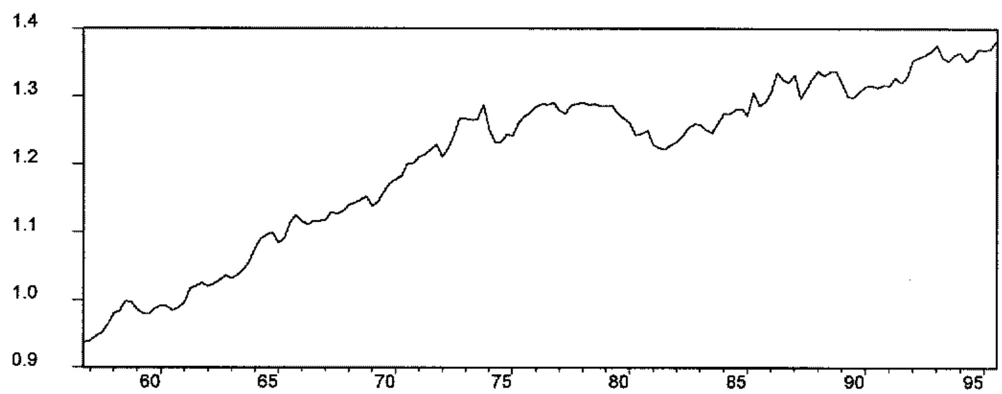
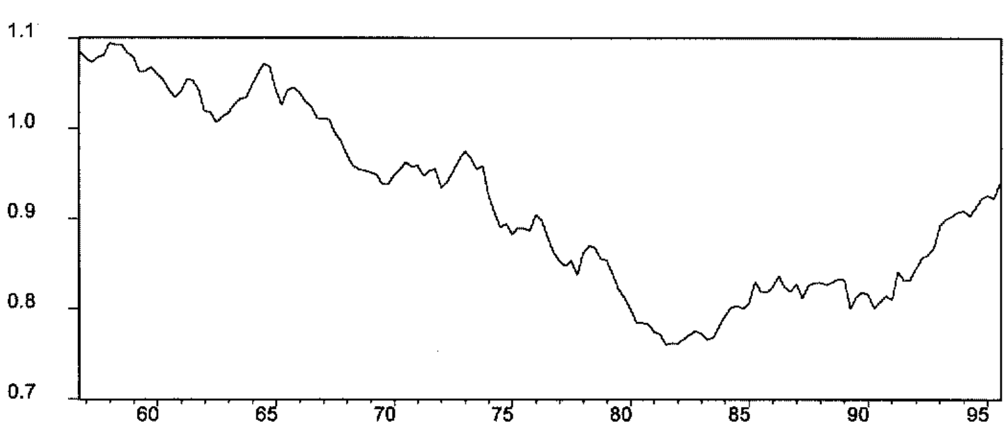


Figure 4
Estimated Opportunity Relative Income: W_{m1t}/W_{m2t-16}



Period: 1956:4-1995:3

example, the estimates of β_1 were always very small and insignificant, as were the estimates of $\gamma_1^{'}$. The results were also not good when the equations for women and men were jointly estimated by three stage least squares. The labor force participation of men 20-24 was high throughout the entire period, and it does not follow the same pattern as that for women. The negative results here suggest that whatever fluctuations there are in young men's participation, they cannot be explained using the model in this paper.

5 Implications of the Estimates

Given data on W_t and p_{1t} and given an estimate of γ_1, one can use equation (2) to compute the potential wage rate of women 20-24 (W_{f1t}) up to a proportionality factor. Similarly, given data on W_t and p_{1t} and given an estimate of $\gamma_1^{'}$, one can use equation (3) to compute the potential wage rate of men 20-24 (W_{m1t}) up to a proportionality factor. Finally, given data on W_t and p_{2t} and given an estimate of $\gamma_1^{'}$, one can use equation (3) to compute the potential wage rate of men 46-50 (W_{m2t}) up to a proportionality factor. From W_{m1t} and W_{m2t}, the potential relative income variable in equation (4) can be computed up to a proportionality factor.[11]

Figure 2 shows a plot of L_{f1t} for the estimation period 1956:4–1995:3, and Figure 3 shows a plot of W_{f1t} for the same period, where W_{f1t} is computed using the estimate of γ_1 in Table 1 (and taking γ_{0i} to be zero). Figure 4 shows a plot of W_{m1t}/W_{m2t-16}, the potential relative income variable, for the same period, where W_{m1t} and W_{m2t-16} are computed using the estimate of $\gamma_1^{'}$ in Table 1 (and taking $\gamma_{0i}^{'}$ to be zero).

[11]These calculations are only up to a proportionality factor because γ_{0i} and $\gamma_{0i}^{'}$ in equations (2) and (3) are not separately estimated in equation (4) because they are not identified.

Figure 3 shows a fairly sharp rise in W_{f1t} until the early 1970's, essentially no further rise until the mid 1980's, and then a modest rise from the mid 1980's on. Figure 4 shows a negative trend in the potential relative income variable until the early 1980's, a very small positive trend until the early 1990's, and then a much larger positive trend from the early 1990's on.

It is interesting to see how the pattern of L_{f1t} in Figure 2 is explained by the patterns of W_{f1t} and W_{m1t}/W_{m2t-16}. The period of most rapid growth in Figure 2 is between about 1964 and 1978; between 1963:4 and 1978:4, L_{f1t} grew by 47.0 percent. In this same period W_{f1t} grew by 21.9 percent and W_{m1t}/W_{m2t-16} fell by 17.3 percent. Using the long run potential wage-rate and potential relative income elasticities of 1.18 and $-.60$, respectively, the predicted change in L_{f1t} from the potential wage rate increase is 25.8 percent (1.18 times 21.9) and from the potential relative income decrease is 10.4 percent ($-.60$ times $-.173$), for a total of 36.2 percent. The potential wage-rate and potential relative income movements thus explain a fairly large fraction of the total increase in labor force participation over this period.

There was much smaller growth in L_{f1t} between 1978 and 1984; between 1978:4 and 1984:4, L_{f1t} grew by 4.7 percent. In this same period W_{f1t} fell by 0.3 percent and W_{m1t}/W_{m2t-16} fell by 6.5 percent. Again, using the long run elasticities the predicted change in L_{f1t} from the potential wage-rate decrease is -0.4 percent and from the potential relative income decrease is 3.9 percent, for a total of 3.5 percent. The smaller growth rate in L_{f1t} is thus attributed to no further growth in the own potential wage rate and less of a decline in potential relative income.

Finally, L_{f1t} essentially did not grow at all between 1984 and 1995. Between

1984:4 and 1995:3, L_{f1t} grew by 0.6 percent, the potential wage rate grew by 7.8 percent, and potential relative income grew by 17.4 percent. Using the long run elasticities, the predicted change in L_{f1t} from the potential wage-rate increase is 9.2 percent and from the potential relative income increase is -10.5 percent, for a total of -1.3 percent. The small growth in participation since 1984 is thus attributed to offsetting effects: a positive effect from the growth of the own potential wage rate and a negative effect from the growth of potential relative income.

Note that although we can estimate W_{f1t} up to a proportionality factor, as in Figure 3, it would not be sensible to compare this estimate with data on actual wage rates. W_{f1t} is a measure of the average potential wage rate facing women 20-24, not the average actual wage rate.

6 Conclusion

The results in Table 1 support the hypothesis that relative cohort size affects potential wage rates: the estimates of both γ_1 and γ_1' are significant. The results also support the hypothesis that potential relative income affects labor force participation of young women in that the estimate of β_2 is significant. Young women's participation is estimated to respond negatively to changes in potential relative income. The overall test performance of the equations is quite good. In particular, the time trend test suggests that the trend in labor force participation of women 20-24 has been explained well by the potential wage-rate and potential relative income variables. The analysis in the last section shows that the rapid rise in participation in the 1964–1978 period is attributed to a combination of a rise in the own potential wage rate and a fall in potential relative income. The much smaller rise in the 1978–1984 period is attributed to the absence of a further rise in the own potential wage rate and a continuing fall in potential relative income. Finally, the flattening out in the 1985-1995 period is attributed to the opposing effects of an increase in the own potential wage rate and an increase in potential relative income.

If the results in Table 1 are to be trusted, they say that fairly subtle concepts have been estimated using quarterly time series data. Picking up cohort effects on potential wage rates and behavioral responses to potential relative income changes is not necessarily something that one would expect of this kind of data. Because of this, the results should be interpreted with some caution even given the good test results.

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