LABOR FORCE PARTICIPATION, WAGE RATES, AND MONEY ILLUSION

Ray C. Fair *

I Introduction

IN view of the numerous studies of labor force participation in the past few years, it is surprising that so little attention has been given to the question of the effect of wage rates on participation rates over time. The results of the cross-section studies of Mincer [4] and Cain [1] indicate that substitution effects dominate income effects for married women, which means that wage rates have a net positive effect on the labor force participation of this group. One would expect that this effect might appear in time-series data as well, but most time-series studies have not considered this possibility.[1]

There are two main problems that must be considered when analysing the effect of wage rates on participation rates over time: the choice of the appropriate wage rate variable and the specification of the appropriate lag distribution of wage rates on participation rates. It will be seen in section II that these two problems are closely related; both concern the question of whether labor force participants respond more to the money wage or the real wage in the short run and thus whether there is any element of money illusion in the short run. It will also be seen from the work in section II that it is possible to *estimate* the degree of money illusion on the part of labor force participants. In section III these estimates are made for sixteen age-sex groups using quarterly United States time-series data for the 1956I–1970II period. The results in section III indicate both the degree to which participation rates respond to wage rates over time and how much of this response is due to money illusion.

II The Effect of Wage Rates on Participation Rates

The Case of No Money Illusion

Only one aggregate money wage variable has been considered in this study. Mincer and others, using cross-section data, distinguish between the wage rate and potential income of the husband and the wage rate and potential income of the wife or other members of the family. This then enables them to separate income and substitution effects. Unfortunately, this cannot be done using aggregate time-series data, and for the work here it has to be assumed that the average wage of each age-sex group moves closely with the aggregate wage. Under this assumption, an increase in the aggregate wage increases the wage rate and potential income of both the husband and wife, as well as any other working members of the family, and it is thus not possible to isolate the pure substitution effect from the income and cross substitution effects. Only the *net* effect of the wage rate on the various participation rates can be estimated.

Since people are likely to differ in the timing of their response to changing wage rates, it seems likely that group participation rates will be a function of a distributed lag of past wage rates. In addition, the behavioral response of a single individual may be such that he can be considered to be responding to a distributed lag of past wage rates. The lag distribution considered in this study is the quadratic distribution.[2] The distribution was constrained to be zero at the beginning and end of the lag, which meant that the resulting distribution was a function of only one parameter. The parameter was estimated using the Almon technique.

Previous time-series studies have indicated that participation rates are a function of the

* I would like to thank Orley C. Ashenfelter, Daniel S. Hamermesh, and Dwight M. Jaffee for their helpful comments on a rough draft of this paper.

[1] See, for example, Tella [7] and Dernburg and Strand [2]. Officer and Andersen [5], using Canadian time-series data, do report that a wage rate variable was tried in their equations, but that per capita income gave better results. Per capita income is interpreted by them as measuring "standard of living" effects (p. 283).

[2] A truncated Pascal distribution was also considered in this study, and the results using this distribution were similar to the results using the quadratic distribution. These results are available in a mimeographed version of this paper. The mimeographed version also contains an appendix explaining how the truncated Pascal distributions were estimated.

Reprinted from
THE REVIEW OF ECONOMICS AND STATISTICS
Published by Harvard University

Vol. LIII, No. 2, May 1971

tightness of the labor market, as measured by the unemployment or employment rate, and this effect was taken into account in this study as well. Let LF_i denote the labor force of group i, POP_i the population of group i, M/POP the aggregate employment-population ratio, and WR the real wage rate. Then in the absence of money illusion, the equation explaining the labor force participation of group i is taken to be:

$$\log \frac{LF_{it}}{POP_{it}} = a_0 + a_1 t + a_2 \log \frac{M_t}{POP_t} + a_3 L(\log WR_{t-j}) + \mu_t, \quad (1)$$

where the t subscripts refer to period t and where $L(\log WR_{t-j})$ denotes the distributed lag of log WR. A time trend is added to the equation to pick up any trend effects not captured by the other explanatory variables. The reason the log form of the equation is used will be clear below. The error term in equation (1), μ_t, is assumed to be first order serially correlated, and all of the equations below were estimated to take this correlation into account.[3]

The Case of Money Illusion

In the short run it may be that labor force participants do not respond directly to changes in the real wage, but instead respond in a more complicated way to changes in the money wage and the price level. Depending on the speed with which information on wage and price changes becomes available, people may respond more quickly to one change than to another. Only in the long run may it be the case that the response is primarily to the real wage. If, for example, people respond more quickly to money wage changes than to price changes, they can be considered to be suffering from short-run money illusion. If, on the other hand, people respond more quickly to price changes, they can be considered to be suffering from "negative" short-run money illusion (to be referred to as short-run price illusion). Fortunately, the degree of short-run money or price illusion can be estimated.

Let W denote the money wage rate and P the price level. Then the following equation can be estimated:

$$\log \frac{LF_{it}}{POP_{it}} = a_0 + a_1 t + a_2 \log \frac{M_t}{POP_t} + a_3 L_1(\log W_{t-j}) + a_4 L_2(\log P_{t-j}) + \mu_t, \quad (2)$$

where $L_1(\log W_{t-j})$ denotes the distributed lag of log W and $L_2(\log P_{t-j})$ denotes the distributed lag of log P. By definition WR equals W/P, or $\log WR = \log W - \log P$, and if there is neither short-run nor long-run money and price illusion, then the distributed lag price term in equation (2) should be the exact negative of the distributed lag money wage term. If there is short-run but not long-run money (price) illusion, then the sum of the coefficients of the two distributions should be equal in absolute value, but the distributed lag money wage term should have larger (smaller) coefficients at the beginning of the distribution and smaller (larger) coefficients at the end in absolute value than the distributed lag price term does. If there is also long-run money (price) illusion, then the sum of the coefficients of the money wage distribution should be larger (smaller) than the sum of the coefficients of the price distribution. The degree of money or price illusion can thus be estimated by estimating the two separate lag distributions and observing their properties.

III The Results

The Data and Period of Estimation

All of the equations were estimated for the 1956I–1970II period, excluding observations for the third and fourth quarters of 1959 and the first quarter of 1960 because of the steel strike and for the fourth quarter of 1964 and the first two quarters of 1965 because of the automobile strike. The period of estimation thus consisted of 52 observations.

Sixteen age-sex groups were analyzed: males and females ages 16–17, 18–19, 20–24, 25–34, 35–44, 45–54, 55–64, and 65+.[4] The data on LF_i and POP_i are household survey data and were collected directly from the Bureau of Labor Statistics (BLS). Both LF_i and POP_i

[3] The equations were estimated by the Cochrane-Orcutt iterative technique within the context of the Almon method.

[4] Because of the cross-section work of Mincer and Cain on married women, it would have been of interest to treat married and single women separately. Time-series data are not available to do this, however, but since over 80 per cent of women between the ages of 25 and 54 are married (see, for example, U.N. Demographic Yearbook [8]), the results achieved in this study for women of these ages should be similar to the results that would have been achieved for married women of these ages had the data been available.

include people in the armed forces, and LF_t is seasonally adjusted. The data on M, POP, and W were also collected from the BLS. The data on M are primarily establishment based data and refer to the total level of employment in the private nonfarm sector. M and W are seasonally adjusted. The data on P refer to the quarterly average of the consumer price index. Establishment based data were used for the employment variable because of possible measurement errors in the household survey data. Household survey measurement errors are likely to show up in both the employment and labor force data, which will cause measurement error bias in equation (2) if household survey data are used for the employment variable. This bias can be corrected by using a two-stage least squares or instrumental variable technique, as was done for one of the labor force participation equations in Fair [3], but for present purposes it is somewhat easier to use the establishment based data for the employment variable.

The Basic Results

When using the Almon technique, one must specify the length of each lag distribution. In this study five lengths were tried for each distribution: 4, 6, 8, 10, and 12 quarters. Since there are two distributions under consideration (one for log W and one for log P), this meant that $5^2 = 25$ regressions had to be run for each of the 16 age-sex groups. The summary results of these regressions are presented in table 1. The first equation presented for each group is the equation that gave the best fit. For 6 of the 16 groups quite different results were obtained for similar fitting equations, and for these 6 groups a second equation is presented in table 1. This second equation is the best fitting member of the set of equations that gave quite different results from the best overall fitting equation. The sum of the lag coefficients for each distribution is also presented in table 1, as is the estimate of the first order serial correlation, ρ, for each equation.

Looking at the results for females first, there definitely appears to be a positive wage rate effect for females 20–24 and 25–34: the money wage rate has a positive and the price level a negative effect on the participation rates for these two groups. This result is consistent with the cross-section results of Mincer and Cain. There is also evidence of a positive wage rate effect for females 45–54 and 65+, although the size of the estimates for females 65+ is somewhat suspect. For females 65+ the estimate of the coefficient of the time trend is quite large in absolute value, and for all of the regressions run for this group there were indications of strong collinearity among the estimates of the time trend and of the coefficients of the price and wage distributions. For females 35–44 and 55–64, the results are ambiguous. For females 35–44 the best fitting equation gives a negative effect for the money wage rate and a positive effect for the price level, with lags of length 12 and 4 quarters respectively, but a similar fitting equation gives the opposite effect, with lags of length 6 and 12 quarters. In the first equation the time trend has a significantly positive effect, and in the second equation it has a significantly negative effect. There appears to be enough collinearity among the estimates of the time trend and of the coefficients of the wage and price distributions to prevent any definitive conclusions from being made as to whether the real wage rate has a positive or negative effect on the participation rate of this group. Likewise, for females 55–64 it is difficult to draw any conclusions. Finally, for females 16–17 and 18–19 the results indicate a negative wage rate effect, although none of the coefficient estimates of the wage and price terms are significant.

For males between the ages of 20 and 64, wage rate effects are small and on the whole negative. While the results for males 20–24 and 25–34 are somewhat ambiguous, the overall results nevertheless indicate a small negative effect for males aged 20–64. The effects for the other males appear to be larger, with a definite positive wage rate effect for males 65+ and an uncertain effect for males 16–17 and 18–19.

The coefficient estimate for the employment variable is positive and quite significant for females 16–17, 18–19, 25–34, 35–44, and males 16–17, 65+. For females 16–17 and males 16–17, 65+, the coefficient estimate is greater than one, which means that a given percentage change in the aggregate employment-population ratio results in a larger percentage change in the

TABLE 1. — ESTIMATES OF EQUATION (2)

Group	Const.	t	Coefficient Estimates for $\log(M_t/POP_t)$	$L_1(\log W_{t-j})$	$L_2(\log P_{t-j})$	ρ	R^2	S.E.	Length of lag W	Length of lag P	Sum of Lag Coefficients W	Sum of Lag Coefficients P
Females												
16–17	−.20	.013	2.69	−.38	1.05	.22	.822	.0387	10	4	−2.80	3.34
	(−0.04)	(0.62)	(7.69)	(−0.96)	(1.50)	(1.66)						
18–19	.35	.010	.64	−.29	.56	.21	.527	.0217	8	4	−1.71	1.78
	(0.15)	(0.92)	(3.60)	(−1.08)	(1.37)	(1.57)						
20–24	−3.90	−.012	.27	.46	−.28	.54	.967	.0145	6	8	2.11	−1.63
	(−3.30)	(−2.47)	(1.15)	(2.83)	(−2.10)	(4.59)						
25–34	−3.54	−.008	.77	.23	−.10	.39	.977	.0127	8	12	1.35	−.83
	(−3.71)	(−3.03)	(5.55)	(3.37)	(−1.64)	(3.06)						
35–44	.12	.011	.43	−.16	.43	.29	.974	.0092	12[a]	4[a]	−1.40	1.38
	(0.10)	(2.43)	(4.57)	(−2.37)	(3.38)	(2.16)						
	−2.80	−.004	.44	.15	−.03	.38	.972	.0097	6	12	.67	−.25
	(−3.94)	(−2.40)	(3.96)	(2.68)	(−0.74)	(2.92)						
45–54	−5.46	−.012	−.31	.21	−.32	.55	.971	.0087	12[a]	4[a]	1.82	−1.02
	(−3.51)	(−2.01)	(−2.54)	(2.37)	(−1.95)	(4.79)						
55–64	.80	.013	.24	−.33	.09	.27	.985	.0102	4[a]	12[a]	−1.05	.75
	(1.21)	(9.24)	(2.09)	(−4.96)	(2.56)	(2.03)						
	−2.52	−.007	.11	.22	−.55	.27	.984	.0103	12[a]	4[a]	1.88	−1.77
	(−2.02)	(−1.52)	(1.05)	(2.93)	(−3.99)	(2.01)						
65+	−17.88	−.071	.23	1.06	−1.42	.37	.590	.0038	12[a]	6[a]	9.15	−6.50
	(−3.00)	(−2.66)	(0.61)	(2.53)	(−2.37)	(2.89)						
Males												
16–17	4.92	.019	1.85	−.36	.78	.42	.912	.0180	12[a]	4[a]	−3.11	2.50
	(1.83)	(1.21)	(8.57)	(−2.30)	(2.69)	(3.38)						
	−.31	−.013	1.92	.25	−.16	.46	.907	.0185	8	12	1.47	−1.37
	(−0.20)	(−3.12)	(8.50)	(2.24)	(−1.65)	(3.78)						
18–19	3.49	.019	.27	−.37	.60	.58	.892	.0145	12	6	−3.19	2.74
	(1.09)	(1.40)	(1.34)	(1.71)	(1.97)	(5.15)						
	−1.76	−.009	.04	.26	−.06	.64	.889	.0146	4	12	.82	−.55
	(−1.09)	(−2.46)	(0.15)	(1.51)	(−0.75)	(6.01)						
20–24	−.20	.001	.08	−.08	.05	.12	.844	.0069	6	12	−.35	.43
	(−0.52)	(1.20)	(1.35)	(−2.68)	(2.46)	(0.91)						
	−1.26	−.007	.07	.09	−.18	.12	.837	.0071	12	4	.78	−.58
	(−1.73)	(−2.29)	(1.24)	(2.07)	(−2.20)	(0.90)						
25–34	−.09	−.001	−.01	.02	−.07	.29	.495	.0024	12	4	.20	−.21
	(−0.30)	(−1.05)	(−0.24)	(1.32)	(−1.98)	(2.16)						
	.33	.001	−.01	−.03	.01	.32	.486	.0025	8	12	−.15	.08
	(1.94)	(2.74)	(−0.57)	(−2.10)	(0.91)	(2.45)						
35–44	.26	.002	−.01	−.03	.04	.15	.716	.0021	12	8	−.27	.23
	(1.12)	(1.24)	(−0.43)	(−1.39)	(1.40)	(1.07)						
45–54	1.00	.004	−.02	−.07	.10	.17	.910	.0021	10[a]	4[a]	−.51	.32
	(4.18)	(3.34)	(−1.25)	(−3.41)	(2.86)	(1.23)						
55–64	−.46	.000	−.11	−.03	.04	.53	.964	.0038	12	12	−.22	.32
	(−1.25)	(0.01)	(−2.19)	(−0.75)	(0.99)	(4.45)						
65+	−6.98	−.023	1.15	.27	.03	.53	.986	.0165	6	12	1.22	.30
	(−4.59)	(−6.60)	(4.75)	(2.20)	(0.41)	(4.56)						

[a] Significantly different at the 95 per cent confidence level from any of the equal length distributions estimated.
(t-statistics are in parentheses.)

participation rates of these groups. The coefficient estimate is negative and significant for females 45–54. For the other groups the estimate is generally small and not significant.

The time trend is generally either significantly negative or insignificant in table 1. Only for males 45–54 does it appear to be significant and unambiguously positive (although even here the effect is small). For females 35–44, 55–64, and males 25–34 the time trend is significantly positive for one of the equations, but not for the other. For these three groups, collinearity problems prevent any definitive conclusions from being drawn about the influence of the time trend. The overall results with respect to the time trend are thus somewhat striking and indicate that the unexplained trend in labor force participation rates is either zero or negative. In particular, it is interesting to note that the rapid increase in the labor force participation of groups like females 20–24 and 25–34 in the last half of the 1960's appears capable of being explained by rising real wage rates and the generally rising employment-population ratio.

Tests for Short-Run and Long-Run Money or Price Illusion

It is quite easy to test for the existence of long-run money or price illusion. Given the estimate of the variance-covariance matrix of the coefficient estimates and given the particular lag distributions used, one can test in a fairly straightforward manner the hypothesis that the sum of the coefficients of the wage distribution is equal in absolute value to the sum of the coefficients of the price distribution. If the sums are significantly different from one another in absolute value, then this is evidence that there is long-run money or price illusion.

For none of the equations in table 1 could the hypothesis that the sums are equal in absolute value be rejected at the 95 per cent confidence level. There is thus little evidence of long-run money or price illusion. The sums for the money wage rate tend to be somewhat more dominant, but not enough to call into question the basic conclusion of no long-run money or price illusion.

Testing for the existence of short-run money or price illusion is somewhat more complicated, but the following test can be made. Compare the final equation chosen (i.e., the equation that gave the best fit) with equations estimated under the assumption that the length of the wage distribution is equal to the length of the price distribution — in the present case either 4, 6, 8, 10, or 12 quarters. If an F test reveals that relaxing the restriction that the lengths of the wage and price distributions be equal does not significantly increase the fit of the equation, then accept the hypothesis that the lengths are equal; otherwise accept the hypothesis that the lengths are not equal. If the lengths are not equal, then this is evidence that people respond more quickly to one variable than to another and thus that there is short-run money or price illusion.

For 10 of the 16 groups — females 16–17, 18–19, 20–24, 25–34, males 18–19, 20–24, 25–34, 35–44, 55–64, 65+ — the fit of the best fitting equation was not significantly different at the 95 per cent confidence level from the fit of one or more of the equations estimated under the assumption of equal lag-distribution lengths. For the other 6 groups — females 35–44, 45–54, 55–64, 65+, males 16–17, 45–54 — the fit of the best fitting equation was significantly better than the fit of any of the equal length equations. For all but 1 of these 6 groups — females 55–64 — the length of the money wage distribution is longer than the length of the price distribution. The results for females 55–64 are again ambiguous, since the fit of both equations in table 1 was better than the fit of any of the equal length equations and yet the first equation in table 1 has lengths of 4 and 12 quarters respectively for the wage and price distributions while the second has lengths of 12 and 4 quarters respectively. In summary, therefore, for 5 of the groups there appears to be a significantly faster response to the price level than to the money wage rate and thus some element of short-run price illusion; for 10 of the groups there is little evidence of either short-run price or money illusion; and for 1 of the groups the results are ambiguous. Overall, the results indicate that the existence of short-run money or price illusion is not very pronounced.

REFERENCES

[1] Cain, G. G., *Married Women in the Labor Force* (Chicago: The University of Chicago Press, 1966).

[2] Dernburg, T., and K. Strand, "Hidden Unemployment, 1953–62: A Quantitative Analysis by Age and Sex," *The American Economic Review*, LVI (March 1966), 71–95.

[3] Fair, R. C., *A Short-Run Forecasting Model of the United States Economy* (Lexington: D.C. Heath and Company, 1971).

[4] Mincer, J., "Labor Force Participation of Married Women," in *Aspects of Labor Economics, A Conference of the Universities*, National Bureau of Economic Research (Princeton: Princeton University Press, 1962).

[5] Officer, L. H., and P. R. Andersen, "Labour-Force Participation in Canada," *The Canadian Journal of Economics* II (May 1969), 278–287.

[6] Solow, R. M., "On a Family of Lag Distributions," *Econometrica*, XXVIII (April 1960), 393–406.

[7] Tella, A., "Labor Force Sensitivity to Employment by Age, Sex," *Industrial Relations*, IV (February 1965), 69–83.

[8] United Nations, *United Nations Demographic Yearbook 1968*.