The Importance of Family Members in Determining the Labor Supply of Puerto Rican, Black, and White Single Mothers*

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The authors examine whether access to informal child care resources through residential family members is important in determining labor supply. It is necessary to control for the characteristics of co-resident adults as the average costs and revenues associated with family size and composition are unique to each sample. Co-resident adults can have either positive or negative effects, depending upon what aspect of the labor supply decision is being considered.

Studies show that the composition of a family's household has a significant impact on determining the employment of married and single mothers. Mothers who reside with relatives or other adults are more likely to enter employment (Heckman, 1974; Tienda and Glass, 1985; Rexroat, 1990). One explanation is that family members provide low-cost child care and produce household goods and services that the mother would otherwise provide.

However, the literature reveals inconsistent results on the employment effect of kin networks among racially and ethnically diverse groups, with some studies indicating that kin networks are more important in facilitating the employment of black or Hispanic mothers, while others find few differences between black and white mothers (Tienda and Glass, 1985; Rexroat, 1990). Parish, Hao, and Hogan (1991) found that kin network use among black and white young women does not necessarily increase employment, suggesting that sharing a residence with family members does not lower child care costs.

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This paper contributes to the debate by examining whether access to informal child care resources through residential family members is important in determining the labor-supply decisions of Puerto Rican, non-Hispanic black, and non-Hispanic white single mothers. Specifically, this study tests the validity of the hypothesis that greater numbers of co-resident relatives and/or other adults (controlling for their employment and gender status) increase the probability that a single mother will be employed *ceteris paribus*. A second analysis is made to determine whether living with a large number of relatives and/or other adults significantly increases the hours worked per week or the weeks worked per year of *already* employed single mothers. The results from this second test should provide an alternative check of the cost-of-child-care argument by analyzing the impact of family members not only on labor force participation, but on the hours and weeks decisions made by the already employed.

Section one reviews the literature on the labor-market experiences of single mothers, with particular attention given to the predicament of Puerto Rican single mothers. Puerto Ricans living in the United States have the worst socioeconomic status of all minority groups and the situation of female-headed families is particularly critical. The second section describes the methodology and data used in the labor supply estimations. The final section summarizes the empirical results, which show that within each group of single mothers, those who live with other adults are more likely to be employed. Indeed, for Puerto Rican and black single mothers already in the work force, the presence of other adults increases their work effort.

The Labor Market Behavior of Single Mothers

Finding affordable child care is a major obstacle to the successful employment of many single mothers (Cattan, 1991). Child-care costs have been estimated at from 21 to 25 percent of income for low-income households that pay for care (U.S. Bureau of the Census, 1990). Puerto Rican and black single mothers are more likely to head low-income households, where child care costs may present a formidable barrier to employment. Puerto Rican and black single mothers are also more likely to be out of the labor force when compared to white single mothers, although the degree to which this is due to problems in arranging for child care is not known. A low earnings capacity (due to education and experience deficiencies) and inadequate child support payments are important factors which discourage the greater labor market involvement of single mothers (Ellwood, 1988).

The literature shows that the category of female heads of households captures several diverse individuals, some of whom lack on-the-job training, education, assistance from the absent parent, and general support from both inside and outside the home (Kamerman and Kahn, 1988:22). Female heads of household may be single mothers because of widowhood, divorce, or separation. During the last two decades the young and never-married cohort has increased significantly as a proportion of all single mothers. Smith (1988) noted in his research on poverty and the family, "Women heading families today are unlikely to be random draws from the population, but are more likely to come from impoverished backgrounds" (p. 161).

There is evidence that white single mothers as a group enjoy higher rates of economic well-being when compared to black and Hispanic single mothers. Overall, white single mothers exhibit higher rates of labor force participation and lower poverty rates than do black or Hispanic single mothers. Zandvakili (1991) demonstrated that existing income inequality among female heads of households was largely due to racial differences in conjunction with educational achievement. Dissimilarities between white and nonwhite single mothers in the number of children at home and the duration of nonparticipation in the labor force were factors which contributed to an increase in the long-run (1976–86) income inequality between these two groups.

Today, women without adequate skills and experience can work full-time jobs and still find themselves near poverty at existing wage rates (U.S. General Accounting Office, 1991). Thus, many low-wage single mothers may rationally calculate that the implicit value of their eligible benefits, which includes Medicaid, food stamps, and possibly housing subsidies, is greater than the net earnings they could expect to receive from a full-time job. Having the sole responsibility for organizing and carrying out household production tasks increases both the time and money costs associated with market work and effectively lowers the real wage these mothers would receive if they were to enter employment.

The evidence shows that if single mothers *are* employed, they are significantly more likely to be working full-time, full-year when compared to married mothers (U.S. General Accounting Office, 1991; Ellwood, 1988). Single mothers who work less than a full year or who don't work at all may receive means-tested transfers such as Aid to Families with Dependent Children (AFDC). These women may have a lower earnings capability relative to those single mothers who do not receive aid, as the AFDC program imposes high marginal tax rates on earnings, and thus reduces work incentives for those who could expect to earn only minimal wages. Slack labor market conditions, in which workers are laid off or work hours are reduced, also diminish earnings and increase poverty for those affected. The percentage of female householders who receive AFDC payments varies by racial and ethnic classification, but Puerto Rican female-headed families are more likely than all other families to be in poverty and receiving this form of income assistance during the 1980s (Bean and Tienda, 1987).

Focus on Puerto Rican Single Mothers. In 1990 a high 64 percent of all Puerto Rican female-headed families were poor compared to 56 percent of black female-headed families and 45 percent of white female-headed families (U.S. Bureau of the Census, 1991a). Puerto Rican women had lower labor force participation rates during the 1970s and 1980s than other groups of women, and in 1991 only 42 percent of Puerto Rican women 16 years old and over participated in the labor force, compared to 57.4 percent of all non-Hispanic women (U.S. Bureau of the Census, 1991b). Although the labor force participation rate for all Puerto Rican women increased between the years 1970 (29.2 percent) and 1991 (42 percent), their consistently low rate of labor force activity is considered anomalous because their trends of increasing education, lowered fertility rates, increases in the percentage born in the United States, and decreases in the number of women of primary child-bearing age should have more significantly increased their rate of participation over time (Cooney and Colon-Warren, 1979; Tienda, Donato, and Guzman, 1990).

The increased number of Puerto Rican female heads of household (in 1991 the rate of female headship among Puerto Ricans families was 43 percent) and the tendency for these women to be out of the labor force has reinforced the low labor force participation rate of Puerto Rican women as a group. In a paper which examined intercity variations in the 1970 labor force participation rates for Puerto Rican women, Cooney (1979) cited the rate of female headship as an important predictor of variations in city-specific employment rates. She found that the increase in females heading families depressed the rate of female work force participation, especially among Puerto Ricans in New York City, and that job scarcity and low wages may have encouraged female householders to take the "welfare option" as their job prospects worsened.

Puerto Rican women have historically been located in jobs associated with part-year work, and seasonal unemployment has exacerbated their employment and earnings problems over time (Tienda, Donato, and Guzman, 1990). The restructuring of regional economies in the Northeast and the corresponding decline in the number of light-manufacturing jobs during the last two decades has disproportionately affected the employment of Puerto Rican women because of their geographical concentration in these areas and their overrepresentation in operative occupations (Ortiz, 1991).

Research has also underlined the important of English-language skills, education, and job-specific experience in providing Puerto Rican women with better employment opportunities and wages (Carliner, 1976). The educational gap between Puerto Ricans and non-Hispanics remains wide. Four-fifth (80.5 percent) of non-Hispanics aged 25 years and over have completed high school, whereas little more than half (58 percent) of all Puerto Ricans 25 years old and over have done so (U.S. Bureau of the Census, 1991b). Thus, despite gains in education, many Puerto Rican women lack the necessary credentials to get a job in the post-industrial service economy. Additionally, Tienda, Donato, and Guzman (1990: 36) found that the influence of education on hiring is conditioned by race and Hispanic origin, with employers ranking black and Puerto Rican women at the bottom of the labor queue. Ethnic discrimination in association with displacement by other workers appears to have increased the employment barriers faced by Puerto Rican women.

Cattan (1991) found that Hispanic mothers were more likely to be out of the labor force because of child care problems. It is poor, single mothers with children under the age of six who are more likely to give "family reasons" as the cause of their limited work year or their total lack of work outside the home (Littman, 1989). Although the significance of child care problems in determining the employment choices of Puerto Rican single mothers is not known, it is likely that cost considerations and Spanish language concerns associated with formal child care arrangements, as well as cultural attitudes that may encourage mothers to rely heavily on other family members for child care needs, have deterred Puerto Rican single mothers from seeking employment (New York State Department of Social Services, 1990: 80). The model described in the next section will incorporate these factors of labor market conditions, human capital endowments, and child care arrangements in determining the labor supply of Puerto Rican single mothers.

A Model for Estimating Women's Labor Supply

Working mothers must earn enough to pay for day care, cooking, cleaning, and many other services that they would traditionally perform if they stayed at home. The single mother at home cannot be easily replaced with goods and services acquired in the marketplace, especially if she is the sole parent and nurturer. During school vacation or closings, working single mothers may be pressed to leave because of a lack of alternative child care facilities, or because they hold jobs that do not offer paid vacations or sick leave. Some mothers may prefer to leave work and be with their children during these times (Blank, 1988). Therefore, entering the labor force to work only a few hours a week or a few weeks during the year may not make economic sense. As a result of weekly fixed costs associated with transportation, child care costs, and the like, the labor force participation decision will no longer be continuously linked with the weekly labor supply decision. Therefore, a reservation wage/reservation hours model which allows for this discontinuity in the labor supply function will be used.

Description of Data and Labor Supply Variables. The data for the analysis were taken from the 1980 U.S. Census 5 percent public use microdata sample (PUMS) for New York City. In 1980 over half of all Puerto Ricans living on the U.S. mainland resided in New York City. Our decision to use this data set was based on previous research on the unique labor market experiences of Puerto Rican female New Yorkers (Ortiz, 1991; Cooney, 1979) as well as the need to use a data set that would allow enough cases with detailed information on the household composition of Puerto Rican families to be obtained.[1] Three samples were created from the data for all Puerto Rican, non-Hispanic black, and non-Hispanic white women living in New York City by distinguishing all female family heads aged 18–64 who had children under the age of 18 living at home.

In a single-period decision model such as is used here, the decision to work and the number of hours or weeks that an individual works are the result of both supply and demand factors.[2] On the demand side in this specification, market wages (equation (1)) are assumed to be given independently of hours or weeks and are determined by a semi-logarithmic earnings function which includes years of completed schooling, English proficiency, a proxy for potential labor market experience and its square, and a control for work disabilities.[3] These variables are defined in Table 1.

The first equation on the demand side shows

$$\ln(W_i) = X_{1i}\, b + u_i, \tag{1}$$

where $\ln(W_i)$ is the natural logarithm of the wage offer available to individual i, and X_{1i} is a row vector of observed individual characteristics with the associated parameter vector b. The mean-zero random disturbance term u_i represents the effects of unobserved factors (e.g., motivation) on market wages and is assumed to be a normal variate with classical properties for all i.

On the supply side we have

$$\ln(W^*_i) = X_{1i}c^* + X_{2i}d^* + u^*_i \tag{2}$$

where $\ln(W^*_i)$ is the ith individual's reservation wage. Working women maximize their utility by combining household production time with market goods and services. The amount of labor supplied will depend on the value of the reservation wage, and this is a function of individual characteristics contained in X_{1i}, where c^* is the associated coefficient vector. Variables thought to influence the reservation wage include exogenous income, the number, employment, and gender characteristics of other household members, the number and ages of children living in the household, and taste

[1] Pooling recent Current Population Surveys (CPSs) was not a feasible alternative as the sample size of Puerto Rican single mothers living in New York City is inadequate for detailed analysis of household composition and income. The 1990 Census PUMS was not yet available.

[2] Legal constraints and employer stipulations concerning the number of hours per week and weeks per year that must be worked are relaxed in the model.

[3] Most empirical labor market analysis assumes that wages are unaffected by hours of work. In a test for differences in the wages of part-time and full-time women workers, Blank (1990) concluded that "there is no simple way to characterize the effects of part-time work on women's wages" (p. 144). Without more research, it seems reasonable to keep the identifying assumption used in the paper. Wage regression results are available upon request.

TABLE 1

Definitions of Variables

Kin16–64	= number of relatives aged 16–64 living in household
Kin≥65	= number of co-resident relatives age 65 or over
FemKin	= proportion of relatives who are women
EmpKin	= proportion of relatives who are employed
Nonkin	= number of co-resident nonrelatives aged 16–64
EmpNonkin	= proportion of co-resident nonrelatives who are employed
FemNonkin	= proportion of co-resident nonrelatives who are women
Kids≤6	= number of children six years or younger living in household
Kids7–11	= number of children 7–11 years old living in household
Kids12–17	= number of children 12–17 living in household
Education	= number of years of education completed by respondent
NoEnglish	= 1 if respondent reported poor English proficiency; 0 otherwise
UnempRt	= the civilian county unemployment rate for each racial/ethnic group
NevMarr	= 1 if never married; 0 otherwise
Separate	= 1 if separated from spouse; 0 otherwise
Widow	= 1 if widowed; 0 otherwise
Married	= 1 if married with spouse absent; 0 otherwise
Divorced	= 1 if single mother was divorced; 0 otherwise
NotAble	= 1 if respondent reported work or transportation disability; 0 otherwise
ForBir	= 1 if born outside of U.S. mainland; 0 otherwise
LnExogInc	= natural logarithm of household income excluding labor earnings of respondent and any public assistance payments
PotExp	= potential experience calculated as age of respondent minus years of completed education minus 6
PotExpSq	= potential experience squared
Employed	= 1 if respondent worked at least one week during 1979 and reported positive wage and salary income; 0 otherwise
AnnHrs	= 1979 hours worked per week × weeks worked in 1979
WksPerYr	= weeks worked in 1979
HrsPerWk	= number of hours worked per week in 1979
WageRt	= 1979 wage and salary income/annual hours worked in 1979, truncated from below at $2.60

factors relating to time spent at home. These variables are contained in the row vector X_{2i}, and d^* is the associated coefficient vector. The random disturbance term u^*_i refers to unobservable factors and is assumed to be a normal variate with classical properties for all i. The availability of low-cost child care as proxied by the number of relatives and nonrelatives in the household should lower the reservation wage and make it more likely that a woman with access to these networks will be employed. The control for the proportion of relatives and nonrelatives who are female should capture whether these women are more likely to be substitutes for the mother.

Having a greater number of employed adults living in the household should increase the ability of the mother to purchase day-care services, or expand her access to job information networks and the world of work. By controlling for the employment status of relatives and nonrelatives, we can also determine whether the increased cost of an additional household member is counterbalanced by their value as either a child care provider or as an income earner. The control for the number of relatives over age 64 in the household should positively affect employment because of the need to generate income to support elderly dependents, or because these family members engage in household and child care activities.

The final supply equation in this model shows that the number of hours supplied is a discontinuous function of the market wage,[4] where

$$\begin{aligned} H_i &= a[\ln(W_i)] + X_{1i}c + X_{2i}d + e_i \qquad \text{for } \ln(W_i) > \ln(W^*_i); \\ &= 0 \qquad \text{for } \ln(W_i) \leq \ln(W^*_i). \end{aligned} \tag{3}$$

The annual-weeks-of-work equation is modeled similarly.[5]

Sample Characteristics. Key descriptive data in Table 2 show that non-Hispanic white single mothers living in New York City had the highest proportion employed (59.1 percent) when compared to black (50.5 percent) or Puerto Rican (20.6 percent) single mothers. A cross tabulation of weekly hours and annual weeks shows that among employed Puerto Rican single mothers 54.7 percent worked full-year/full-time (more than 46 weeks a year and 40 hours or more a week), and among working black householders and white householders fully 59.5 percent of both samples worked full-time/full-year. Consistent with the theory of discontinuities at the point of zero labor supply, less than 2 percent of the women from any of the three samples worked 10 hours or less in combination with 12 weeks of work or less.

Labor Force Participation Findings

Table 3 presents the coefficients from the probit model of labor force participation. The partial derivative for each independent variable evaluated at the sample means can be obtained by multiplying each coefficient by the constant of proportionalities, given in Table 3 in the first row for each

[4]The relationship between equations (2) and (3) arises from the fact that since reservation wage W^*_i equals the greatest wage offer consistent with zero hours of labor supply, then $c^* = -c/a$, $d^* = -d/a$, and $u^*_i = -e/a$. See Prescott, Swidinksy, and Wilton (1986: 136).

[5]The first-stage probit estimates used the entire sample of working and nonworking single mothers to establish the probability of being in the employed sample. The coefficient estimates from the probit were used to form a measure of the "selection bias" variable lambda for each observation. The second-stage estimates of the wage function for workers contained the set of regressors included in the vector X_1, but no right-hand-side endogenous variables. In order to correct for the simultaneous equation bias of the OLS estimator, imputed wages obtained from the selection-bias-corrected wage regression were used as an instrument for actual wages, in the estimation of the annual hours and weeks equations. The parameters of the annual hours and weeks equations were then estimated, again correcting for the possibility of selectivity bias by including the lambda variable in the equations. Despite the unidirectional dependence between the endogenous variables in equation (3), the system is not recursive because of the assumed correlation between the disturbances in the market wage and annual hours or weeks equations.

TABLE 2

Means (and Standard Deviations) of Variables for Working and Total Populations

	Puerto Rican		Non-Hispanic Black		Non-Hispanic White	
Variable	Working	Total	Working	Total	Working	Total
Kin16–64	0.342	0.303	0.441	0.415	0.323	0.309
	(0.675)	(0.670)	(0.845)	(0.819)	(0.667)	(0.662)
Kin≥65	0.013	0.009	0.024	0.019	0.036	0.027
	(0.114)	(0.098)	(0.169)	(0.147)	(0.201)	(0.174)
FemKin	.032	.032	.043	.041	.020	.017
	(0.204)	(0.241)	(0.287)	(0.275)	(0.187)	(0.175)
EmpKin	.179	.123	.223	.185	.219	.207
	(0.462)	(0.394)	(0.572)	(0.517)	(0.550)	(0.547)
Nonkin	0.090	0.056	0.089	0.082	0.070	0.068
	(0.296)	(0.243)	(0.313)	(0.305)	(0.271)	(0.272)
EmpNonkin	.036	.072	.059	.070	.056	.063
	(0.194)	(0.265)	(0.255)	(0.272)	(0.249)	(0.258)
FemNonkin	.004	.006	.012	.012	.004	.006
	(0.070)	(0.081)	(0.126)	(0.130)	(0.066)	(0.077)
Kids≤6	0.485	0.764	0.423	0.629	0.227	0.360
	(0.676)	(0.871)	(0.632)	(0.823)	(0.502)	(0.635)
Kids7–11	0.501	0.633	0.481	0.544	0.449	0.479
	(0.678)	(0.801)	(0.659)	(0.743)	(0.629)	(0.666)
Kids12–17	0.734	0.782	0.813	0.799	0.839	0.795
	(0.909)	(1.02)	(0.902)	(0.960)	(0.774)	(0.812)
Education	10.5	8.99	12.0	11.42	13.0	12.3
	(3.17)	(3.36)	(2.36)	(2.36)	(2.79)	(2.93)
NoEnglish	.144	.321	.079	.007	.013	.021
	(0.352)	(0.466)	(0.088)	(0.083)	(0.115)	(0.143)
UnempRt	12.0	12.3	11.3	11.4	5.3	5.4
	(1.69)	(1.42)	(1.31)	(1.31)	(0.251)	(0.253)
NevMarr	.256	.337	.344	.405	.059	.093
	(0.437)	(0.472)	(0.475)	(0.491)	(0.236)	(0.291)
Separate	.335	.371	.332	.327	.244	.255
	(0.472)	(0.483)	(0.471)	(0.469)	(0.429)	(0.436)
Widow	.056	.048	.067	.076	.151	.164
	(0.231)	(0.215)	(0.250)	(0.265)	(0.358)	(0.370)
Married	.039	.041	.039	.035	.019	.032
	(0.195)	(0.200)	(0.195)	(0.184)	(0.138)	(0.177)
Divorced	.311	.200	.216	.155	.526	.454
	(0.400)	(0.412)	(0.362)	(0.500)	(0.498)	(0.498)
NotAble	.063	.180	.038	.117	.040	.098
	(0.244)	(0.384)	(0.193)	(0.321)	(0.196)	(0.297)
ForBir	.755	.198	.233	.160	.147	.144
	(0.430)	(0.398)	(0.423)	(0.366)	(0.354)	(0.352)
LnExogInc	2.75	2.25	2.91	2.81	5.25	4.85
	(3.75)	(3.62)	(3.83)	(3.84)	(3.95)	(4.15)
PotExp	17.3	18.1	17.5	17.2	19.7	19.7
	(9.24)	(10.2)	(8.78)	(9.49)	(8.16)	(8.86)
PotExpSq	387	432	385	388	456	468
	(389)	(465)	(365)	(401)	(361)	(398)
Employed	—	.206	—	.505	—	.591
		(0.404)		(0.500)		(0.491)
AnnHrs	1,359	299	1,444	764	1,536	934
	(728)	(667)	(700)	(896)	(658)	(919)
WksPerYr	41.0	8.57	43.2	22.0	44.6	27.2
	(15.7)	(18.4)	(14.3)	(23.8)	(13.1)	(24.0)
HrsPerWk	33.1	7.4	33.4	17.8	34.4	21.3
	(10.6)	(15.2)	(10.8)	(19.0)	(9.94)	(18.7)
WageRt	5.06	1.31	7.44	3.76	7.63	4.51
	(1.76)	(4.61)	(7.78)	(6.66)	(6.93)	(6.52)
n	755	3,660	3,133	6,202	1,486	2,512

TABLE 3

Probit Model Coefficients for Labor Force Participation of Single Mothers in New York City

Variable	Puerto Rican	Black	White
$f(\overline{X}\beta/\sigma)$[a]	*.2394*	*.3988*	*.3862*
Constant	−0.173	−0.042	0.608
	(0.285)	(0.224)	(0.407)
Kin16–64	−0.106+	−0.113**	−0.037
	(0.063)	(0.034)	(0.085)
Kin≥65	−0.242	−0.010	0.469**
	(0.247)	(0.125)	(0.183)
FemKin	−0.048	0.036	0.124
	(0.128)	(0.067)	(0.160)
EmpKin	0.283**	0.241**	0.068
	(0.092)	(0.052)	(0.099)
Nonkin	−0.002	−0.187+	−0.444+
	(0.184)	(0.113)	(0.262)
EmpNonkin	0.513*	0.421**	0.396
	(0.226)	(0.136)	(0.286)
FemNonkin	−0.168	−0.098	0.193
	(0.355)	(0.153)	(0.485)
Kids≤6	−0.494**	−0.502**	−0.591**
	(0.045)	(0.028)	(0.060)
Kids7–11	−0.283**	−0.261**	−0.240**
	(0.039)	(0.025)	(0.045)
Kids12–17	−0.186**	−0.135**	−0.094*
	(0.034)	(0.022)	(0.042)
Education	0.086**	0.125**	0.095**
	(0.010)	(0.008)	(0.011)
NoEnglish	−0.424**	−0.380+	−0.076
	(0.067)	(0.203)	(0.213)
UnempRt	−0.060**	−0.067**	−0.189**
	(0.016)	(0.013)	(0.056)
NevMarr	−0.318**	−0.437**	−0.432**
	(0.075)	(0.057)	(0.103)
Separate	−0.213**	−0.291**	−0.145*
	(0.069)	(0.056)	(0.068)
Widow	−0.148	−0.493**	−0.262**
	(0.130)	(0.081)	(0.087)
Married	−0.182	−0.199*	−0.661**
	(0.141)	(0.106)	(0.158)
NotAble	−0.822**	−1.12**	−0.979**
	(0.087)	(0.062)	(0.097)
ForBir	−0.089	0.649**	0.186*
	(0.069)	(0.052)	(0.083)
LnExogInc	0.014+	−0.007	−0.010
	(0.008)	(0.005)	(0.007)
PotExp	0.033**	0.038**	0.015
	(0.011)	(0.007)	(0.014)

TABLE 3—continued

Variable	Puerto Rican	Black	White
PotExpSq	−.05E−02** (.02E−02)	−.08E−02** (.01E−02)	−.05E−02+ (.03E−02)
Log likelihood	−1,519	−3,415	−1,403
n	3,660	6,202	2,512

NOTE: The maximum likelihood estimation method was used for the analysis, with Employed as the dependent variable. Standard errors are given in parentheses.

[a]Multiply coefficients by this factor to obtain slopes at the variable means.

+$p < .10$ level.

*$p < .05$ level.

**$p < .01$ level.

sample. While at least one of the family-member variables did exert a positive impact on labor force participation, the findings indicate that variations in the membership of these households across the samples operate to condition which adults are more important in increasing employment probabilities.

Among Puerto Rican single mothers a one-unit increase in the number of kin aged 16–64 diminished participation probabilities by 2.5 percent. Yet, an interaction term capturing the influence of an additional relative aged 16–64 and a child 6 years and younger significantly improved the participation of Puerto Rican mothers indicating that the cost-of-child-care argument is most relevant for women with young children. Participation increased significantly, by 6.7 percent, if there was a one-unit increase in the number of employed relatives in the residence and by 12.2 percent if an additional nonrelative was employed. Employed relatives and nonrelatives (which may include partners of consensual unions) must provide other forms of assistance (job information, income for purchasing formal child care services or household goods and services) that encourage the participation of these single mothers.

Among black women, a one-unit increase in the number of resident relatives negatively influenced work probabilities by 4.5 percent *ceteris paribus*. Yet, having an additional *employed* relative in the household significantly increased this group's participation by 9.6 percent. A one-person increase in the number of nonrelatives decreased employment chances by 7.4 percent, while a one-person increase in the number of employed nonrelatives increased work force probabilities by 16.7 percent. As in the Puerto Rican case, the impact of other adults on participation probabilities is mediated by their employment status.

Among white single mothers the presence of an additional adult relative

65 years of age or older increased their participation by 18.1 percent. White single mothers who lived with older relatives may be more likely to work in order to support these elderly family members. A one-person increase in the number of nonkin residents diminished participation probabilities by 17.1 percent, although the statistical significance of this factor is low. Tests for interaction between the number of young children at home and the kin and nonkin variables indicated that the cost-of-child-care hypothesis could not be supported.

Other findings in Table 3 affirm previous research on the labor force determinants for single mothers. The impact of an additional child 6 years or younger on the participation of Puerto Rican mothers is to decrease their chance of participation by 12 percent, while the probabilities for white and black single mothers declined by 23 percent and 20 percent, respectively. Interestingly, in contrast to conventional assumptions made about Hispanic mothers, Puerto Rican single mothers are comparatively less likely to be deterred from entering employment due to the presence of young children at home.

An additional year of completed education raised the probability of Puerto Rican work force participation the least (by 2 percent), while black mothers saw an increase of 5 percent and white mothers an increase of 4 percent. Poor English-language proficiency curtailed job participation by 15 percent for non-Hispanic black single mothers, significantly decreased the participation of Puerto Rican mothers by 10 percent, and did not significantly affect white mothers' participation. Foreign birth had a positive impact on work force probabilities in both the black and white samples and was not significant for Puerto Ricans.

A one-unit increase in the group-specific civilian unemployment rate by county resulted in a significant 7 percent decline in the participation of white mothers, a 3 percent decline for black mothers and a small 1 percent for Puerto Rican mothers. In another study we also found that Puerto Rican women exhibited a smaller "discouraged worker effect" when compared to black and white women (Melendez and Barry Figueroa, 1992). Lastly, in comparison to divorced single mothers, those who claimed alternative marital statuses participated in the labor force at significantly lower rates across the three samples, although marital status was far less important in distinguishing labor force probabilities among Puerto Rican women.

Annual-Hours- and Weeks-Worked Findings

The distinction between annual hours of work and annual weeks worked per year provides two unique labor supply variables. The third implicit variable captured in this model is the number of hours worked per week, a nonlinear combination of the other two labor supply variables. Table 4 shows that the independent estimates of weeks and hours, conditional on

TABLE 4

Estimates of Annual Hours and Weeks, Conditional on the Labor Force Participation of Working Single Mothers

	Puerto Rican		Non-Hispanic Black		Non-Hispanic White	
Variable	AnnHrs	WksPerYr	AnnHrs	WksPerYr	AnnHrs	WksPerYr
Constant	625	25.2+	387	12.8	1,997	45.9**
	(653)	(13.7)	(509)	(10.3)	(401)	(8.15)
Kin16–64	194**	3.97**	82.9**	1.24**	35.4	1.08
	(67.6)	(1.90)	(25.9)	(0.530)	(51.3)	(1.04)
Kin≥65	396+	9.01*	140*	1.90	19.6	0.165
	(227)	(4.77)	(72.9)	(1.48)	(95.6)	(1.94)
FemKin	112	2.50	−49.7	−1.02	30.9	1.67
	(129)	(2.71)	(43.8)	(0.894)	(90.8)	(1.84)
EmpKin	−231**	−3.08	−37.4	0.358	20.7	−0.128
	(99.0)	(2.07)	(38.0)	(0.777)	(59.9)	(1.21)
Nonkin	−133	−3.93	−15.4	−1.63	246	4.72
	(181)	(3.82)	(84.9)	(1.73)	(211)	(4.30)
FemNonkin	558+	8.04	−4.54	−1.09	28.7	−0.321
	(329)	(6.90)	(100)	(2.05)	(219)	(4.46)
EmpNonkin	−1.79	0.904	42.3	2.55	−110	−2.44
	(214)	(4.50)	(99.7)	(2.04)	(219)	(4.46)
Kids≤6	−7.61	−3.97	−100*	−4.11**	−125	−3.55*
	(90.5)	(1.90)	(45.1)	(0.923)	(85.2)	(1.73)
Kids7–11	62.4	0.519	−39.0	−179**	−97.6**	−1.99**
	(59.9)	(1.26)	(27.1)	(0.554)	(39.7)	(.807)
Kids12–17	17.9	0.130	12.6	−0.184	50.1+	0.771
	(45.4)	(0.955)	(20.1)	(0.411)	(29.3)	(0.597)
Wage	509	11.0	738**	1.0**	182	3.43
	(381)	(8.03)	(281)	(5.74)	(204)	(4.16)
UnempRt	18.2	0.525	−5.26	−0.405**	−109**	−0.852
	(16.5)	(0.347)	(9.57)	(0.195)	(40.8)	(0.830)
Widow	3.20	0.761	29.6	0.774	−211**	−3.16**
	(119)	(2.51)	(66.1)	(1.35)	(65.2)	(1.32)
NotAble	70.8	0.646	−64.5	−3.81	−315*	−7.37**
	(149)	(3.14)	(117)	(2.40)	(155)	(3.16)
ForBir	126*	2.06	47.9	2.09*	−13.2	−0.482
	(64.0)	(1.35)	(50.5)	(1.03)	(50.6)	(1.03)
LnExogInc	−12.0+	−0.504**	−17.5**	−0.488**	−18.4**	−0.355**
	(8.68)	(0.182)	(3.94)	(0.080)	(4.85)	(0.098)
Lambda	−348*	−6.09	−185	2.06	−72.9	1.30
	(184)	(3.88)	(149)	(3.05)	(242)	(4.92)
R^2	.10	.15	.08	.08	.08	.06
n	755	755	3,133	3,133	1,486	1,486

SOURCE: 1980 United States Census, five percent PUMS for New York City.

NOTE: Standard errors are given in parentheses. Marital status variables (never married, married, separated) were included in the analysis, but were not significant at <.05 in any of the equations.

+ p < .10 level.

* p < .05 level.

** p < .01 level.

labor force participation, provide same-signed, significant coefficients in both equations, although the magnitudes are not proportional. The findings indicate that the two variables measure similar but separate decision-making processes.

The number of relatives aged 16–64 and over age 64 positively increased both the weeks and hours worked for already employed Puerto Rican single mothers. Since average weekly hours for this group was 33.1 (Table 2), the effect of annual weeks as given by the coefficient associated with the kin (aged 16–64) variable implies an increase of 131 annual hours (33.1 × 3.97). Since the coefficient associated with this kin variable in the annual hours equation is larger at 194, the impact of having an additional relative aged 16–64 at home is concentrated on increasing the supply of weekly hours. Using the same procedure for the kin over age 64 variable, we find that the effect is once again to increase the supply of weekly hours. Alternatively, the greater the number of employed relatives at home, the smaller was the number of annual hours worked, indicating an income effect that allowed working mothers to devote more of their time to household activities. The effect of an increase in the number of female nonkin residing with the single mother was to increase annual hours by 558. However, this variable had a low level of significance, presumably because of the small percentage of single mothers who had female nonrelatives residing with them. The kin variables increment each other, and the sum of these coefficients shows an overall positive impact on the number of annual hours worked. The nonkin variables also increment each other, and their impact on annual hours is positive. The findings indicate that family members play an important role in assisting the single mother with child care and household tasks so that full-time/full-year work is more of an option.[6]

Among black women, increased numbers of relatives aged 16–64 significantly increased the annual weeks and hours of work with the effect concentrated on the weekly supply of hours. An increased number of elderly kin also encouraged the mother to work a greater number of hours over the year. An interaction variable combining the effect of increased numbers of kin 16–64 and the number of children 6 years and younger proved significant in increasing the number of weekly hours and annual weeks, although this variable was not important in determining participation. This confirms that among working black mothers with young children, increased access to child care resources lowers the fixed costs of working.[7] Increases in the wage also significantly increased the weekly hours worked by these women.

[6] In the Puerto Rican sample the coefficient on the selection bias variable lambda was negative and significant in the annual hours equation. This means that unmeasured factors (such as a desire for a social life outside of the home, implicit valuations of available social programs) which tend to increase the probability of participation for Puerto Rican householders act to decrease the number of hours worked over the year.

[7] Results from the interaction models estimated for each sample are available from the first author upon request.

For white mothers, none of the variables reflecting family membership were significant in determining the intensity of their work effort. Household composition does not appear to affect child care costs or to otherwise condition the weeks and hours worked decisions. Other factors such as work disability status, the county unemployment rate, the number of children aged 7–11, and the level of exogenous income available to the household were more important determinants of the work schedule for this sample.

Summary and Conclusions

It is necessary to distinguish the characteristics of family members to better understand their impact on defining the labor supply decisions of single mothers. For single mothers with children, the composition of the household will affect both the fixed time and money costs associated with work. Average costs and revenues associated with family size and composition are unique to each sample. The impact of a relative or nonrelative on labor supply can differ depending on whether one is considering the decision to enter the labor market or the decision among existent participants to work more hours. For example, increased numbers of kin aged 16–64 had a negative impact on the labor force participation decision across the three samples, but a positive effect on the number of annual hours and weeks worked.

The results for the Puerto Rican and black samples indirectly confirm the cost-of-child-care argument. Interaction tests revealed that the combination of having relatives aged 16–64 and young children at home significantly improved the probability of labor force entry for Puerto Rican single mothers and increased the number of hours and weeks worked among black single mothers. In both samples, living with other employed adults was a positive and significant factor in determining labor force participation. We do not know whether employed family members contribute income to help purchase child care services, or if they help generate employment contacts. The latter form of assistance would be particularly useful for these women, given the high unemployment rates prevailing in their communities.

Despite the importance of extended families in helping to construct the labor supply of both Puerto Rican and black single mothers, the number of women who live in these types of households has been declining over the years (Bean and Tienda, 1987). This suggests that a major factor for improving the income and time poverty facing these single mothers may be disappearing rapidly. Thus, formal or institutional mechanisms through which these women can access child care services have become all that more important. Our findings suggest that many Puerto Rican single mothers who want to work or who are already employed will continue to depend on coresident family members for help in caring for the children until affordable quality child care is more readily available. **SSQ**

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