Gender and the Short-Run Economic Consequences of Marital Disruption*

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Abstract

Past studies on gender differences in the economic consequences of divorce have been limited to presenting descriptive statistics. This article examines sources of gender differences in the economic ramifications of marital disruption for young non-Hispanic white, black, and Hispanic adults separating or divorcing in the 1980s. Using data from the 1979-88 waves of the National Longitudinal Survey of Youth, the results show that even among a less-advantaged subgroup, marital disruption has more serious consequences for women than men. Although young men, particularly minority men, are not faring well economically in absolute terms, women's postdisruption economic welfare is significantly lower than men's for all race-ethnic groups. Multivariate analyses reveal that this disparity stems, either directly or indirectly, from women's roles as primary child caretakers.

The growth of mother-only families over the last few decades, rising from 9% of all families with children in 1960 to over 20% in 1987, is of central concern to policymakers and to researchers of social stratification and inequality (e.g., Bane 1986; Ellwood 1988; Garfinkel & McLanahan 1986; Kammerman & Kahn 1988; Ross & Sawhill 1975). This concern is primarily due to the indisputably high poverty rates among mother-only families; roughly 50% of women and children in these families live in poverty (U.S. Bureau of the Census 1988). Additionally, low income is an important cause of disadvantages faced by children of mother-only families. Compared to children of two-parent families, those from mother-only families are more likely to drop out of high school, form single-parent families themselves, and be poor as adults (e.g., McLanahan 1985; McLanahan & Bumpass 1988).

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Marital disruption continues to be a major source of the growth of mother-only families (e.g., Bianchi & Spain 1986), and numerous studies have directly linked the event of marital disruption to women's postdisruption economic vulnerability (Corcoran 1979; Duncan & Hoffman 1985; Hoffman 1977; Hoffman & Duncan 1988; Morgan 1989, 1991; Mott & Moore 1978; Nestel, Mercier, & Shaw 1983; Peterson 1989; Smock 1993; Stirling 1989; Weiss 1984; Weitzman 1985; see Holden & Smock 1991 for a review of this literature). Perhaps the most well-known finding from this body of research is that the economic costs of marital disruption are quite unequally distributed by gender. While women typically undergo marked declines in income and measures of economic status that take family size into account, men undergo minimal income loss and even experience improvements in family size-adjusted measures of economic status (Burkhauser et al. 1991; Duncan & Hoffman 1985; Hoffman 1977; Smock 1993; Sørensen 1992; Weitzman 1985).

However, past studies on gender differences in the economic consequences of divorce have been limited to presenting descriptive statistics on men's and women's economic status before and after separation or divorce (e.g., Burkhauser et al. 1991; Duncan & Hoffman 1985; Hoffman 1977; Smock 1993; Sørensen 1992; Weitzman 1985). To date, we have no direct evidence on the sources of the gender disparity. In this study I extend previous research by investigating the bases of gender inequities in the economic impact of marital disruption. I relate gender differences in levels of postdisruption economic welfare to key work- and family-related characteristics. Specifically, I examine the extent to which women's postdivorce economic disadvantage is due to children's living arrangements and/or to differences in men's and women's employment-related characteristics (i.e., human capital). Whether such factors — and which ones — can account for gender differences in postdivorce economic outcomes has important implications for the broader issues of women's economic vulnerability outside of marriage and the economic hardship of mother-only families.

By focusing on a less-advantaged population — young white, black, and Hispanic adults separating or divorcing in the 1980s — I also explicitly question the common assumption that men fare well economically upon marital disruption, and so much better than women. Whether this is the case for young adults is unclear for at least two reasons. First, those who marry at young ages tend to obtain less education than those who delay marriage, and, more generally, have lower economic prospects (Bianchi & Spain 1986). Indeed, among divorced women, those who are young and less educated are the most economically vulnerable (e.g., Meyer 1993). Thus, given the current emphasis on child support as the remedy to the economic difficulties of single mothers (e.g., Garfinkel & McLanahan 1986), understanding the economic impact of divorce on young men is crucial. Men are likely to be economically vulnerable themselves, particularly in view of deterioration in young men's employment prospects and earnings over recent decades (Levy & Michel 1991; Phillips & Garfinkel 1993). Second, one might expect gender inequities in the economic ramifications of marital disruption to be moderate, or even small, for young adults; women are increasingly likely to be employed during marriage and their childbearing years, and younger men and women are experiencing greater equality in wages, labor force attachment, and orientations towards market work (e.g., Bianchi & Spain 1986; Moen 1992).

Analysis Plan

The next section describes the data and measures. The following section reports results, presented in two parts. First, I present descriptive statistics on two measures of men's and women's economic status before and shortly after marital disruption: personal income and per capita income. The choice of these two measures reflects the importance of both labor market position and familial position in understanding gender economic inequality (Curtis 1986). The first furnishes information about men's and women's own economic resources independent of living arrangements. Prior studies have tended to report family income, rather than personal income, an approach which obscures that much of women's predisruption family income is, in fact, their husbands', and that many young people move in with their parents or other relatives after marital disruption (Bumpass & Sweet 1991; Smock 1993). The other measure I use, per capita income, does take into account pre- and postdisruption living arrangements, along with the economic resources of any coresident family members.

Second, I estimate ordinary least squares (OLS) regressions of postdisruption personal income and per capita income for men and women together, my intent being to explain overall differences between men and women in levels of economic well-being after separation or divorce. I examine the extent to which two central sets of independent variables account for gender inequalities in economic outcomes: (1) predisruption "human capital" characteristics (i.e., educational attainment and work experience), and (2) the presence of children.

Theoretically, human capital differences could explain a good deal of the gender disparity in postdisruption economic well-being, particularly in personal income. Human capital theory posits that economic rewards from market work depend upon skills proxied by such factors as educational attainment and work experience (Becker 1975). Women tend to accumulate less work experience than men within marriage, in large part because they are primarily responsible for housework and child care (Berk 1985; Coverman & Sheley 1986; Robinson 1988). When marriage dissolves women may have neither jobs paying living wages, nor the employment experience to acquire such jobs. For per capita income, women's customary role as primary child caretakers (both during and after marriage) is likely to account directly for their postdivorce economic disadvantage beyond the indirect influence of children on women's human capital. Children are dependents who increase the need for income, but do not generally contribute to this income, and women are far more likely to assume physical custody of their children after marital disruption (e.g., Maccoby, Depner & Mnookin 1988; Meyer & Garasky 1993).

Data and Measures

DATA

I use high quality panel data from the National Longitudinal Survey of Youth (NLSY), 1979-88. This is a nationally representative sample of 12,686 men and women aged 14–21 in 1979, including oversamples of blacks and Hispanics.[1] NLSY retention rates are quite high; roughly 90% of respondents remained in the sample through 1988.

These data are quite well suited for this research. They provide a rich set of background information and annual measures of an extensive array of variables: marital status, labor force participation, earnings and nonearnings income, and characteristics of spouses and other household members. The yearly measures permit me to capture change in economic well-being and other characteristics coincident with marital status change.

I focus on the relatively short-term economic consequences of marital disruption, largely to minimize case loss from remarriage or cohabitation (see below). While not ideal, past research shows that short-term economic consequences approximate longer-term consequences for women unless remarriage occurs (Duncan & Hoffman 1985; Morgan 1991; Stirling 1989; Weiss 1984). There is some evidence of modest income improvement over the years following marital disruption for men, particularly for less-advantaged men whose initial postdisruption incomes were low (Duncan & Hoffman 1985; Phillips & Garfinkel 1993); the results presented here may therefore slightly underestimate men's eventual economic status, and, correspondingly, gender differences in postdisruption economic status.

I draw on subsamples of men and women observed married in at least one survey and subsequently reporting themselves as either separated or divorced. These are separate samples of men and women separating or divorcing, not primarily cases where the men and women were married to each other.[2] When a respondent reports being married in one or more survey years between 1979 and 1986 and separated or divorced in a subsequent survey before 1988, I define this as a marital disruption. In this study, $T-1$ represents the predisruption observation (the last year of marriage) and $T+1$ the postdisruption observation, where marital disruption is first recorded in survey year T. I rely on information ascertained at the $T + 1$ survey, rather than T, to measure postdisruption economic status because income questions in the NLSY reference the prior calendar year. Thus, income reported in $T+1$ generally references the first full postdisruption year.

Not surprisingly, a substantial minority of men and women are remarried or cohabiting by $T + 1$.[3] Remarriage is common at young ages and can occur quite quickly; there has also been a sharp rise in nonmarital cohabitation over recent years and previously married individuals are particularly likely to cohabit (Bumpass & Sweet 1989; Sweet & Bumpass 1987). I exclude these cases from analyses because my primary purpose is to examine gender differences in economic vulnerability outside of marriage or marriagelike relationships. Analyses not reported here indicate no evidence of sample selection bias from this exclusion.[4] After eliminating cases of rapid remarriage/cohabitation and those with missing income data, final sample sizes include 452 separating and divorcing young women and 254 men.[5]

INCOME MEASURES

The NLSY measures several components of income: earnings, farm-business income, aid from relatives, unemployment compensation, income from Aid to Families with Dependent Children (AFDC), educational benefits, disability income, supplemental security income, alimony/child support (both received and paid), each component measured separately for the respondent and spouse (if present). If the respondent lives with adult family members other than a spouse or children, the survey also ascertains the incomes of these family members.

I measure personal income as the sum of all earnings and nonearnings income (e.g., public and private transfers) of the respondent. I also adjust income for any child support or alimony received or paid.[6] Per capita income uses family income as the numerator (the sum of personal income and the income of any coresident family members) and the number of coresident family members as the denominator. In multivariate analyses I use the logarithmic transformation of these measures because of the skewness of the income distribution. All income amounts are adjusted by the consumer price index and are presented in 1987 dollars.

INDEPENDENT MEASURES

Multivariate analyses are ordinary least squares regressions of the natural logarithm of postdisruption personal income and per capita income. Selection of independent variables was guided by past research (e.g., Peterson 1989; Smock 1993), and by my motivation to explain the overall effect of gender with variables from the key domains of children and employment.[7] Because analyses pool the samples of women and men, I tested for significant interactions with gender for all independent variables. Although none was statistically significant in a model with all independent measures, I retain one for substantive reasons (see below).[8]

Predisruption Human Capital

Human capital characteristics such as educational attainment and work experience are expected to have positive effects on postdisruption economic welfare. Men and women with higher levels of schooling and more work experience will be far more likely to be employed after marital disruption, and, if so, to receive higher wages. Although levels of educational attainment are not likely to differ substantially between young maritally disrupted men and women, as they do not for the population of young people as a whole (e.g., Bianchi & Spain 1986), differences in amounts of work experience during marriage could potentially account for a major portion of gender differences in postdisruption economic welfare.

I measure educational attainment at T-1 as years of schooling completed at the T-1 survey. I use two measures of work experience. First, I include total years of market employment, either full-time or part-time, in all years before T-1. This measure need not represent continuous years of work experience; all weeks worked are summed and divided by 50. Second, I include a variable capturing work experience immediately before marital disruption: number of weeks worked in T-1. The two variables together represent total years of work experience before T. Weeks worked in T-1 is likely to have a positive effect on postdisruption well-being, even net of previous work experience. Those who have higher work involvement in the year before marital disruption should be better off economically when marriage ends, both because they are less likely to need to seek new jobs and are gaining additional and more immediate work experience.

Children

A central reason for gender disparity in postdisruption economic well-being is likely to stem, either directly or indirectly, from the presence of children. Women are generally the primary child caretakers both during and after marriage; outside child care is costly, resulting in an inherent conflict between parenting and women's labor market work. This implies that much of the effect of children on women's postdisruption personal income will be captured by predisruption work experience and weeks worked. Childless women, for example, are likely to accumulate more work experience before separation and divorce, and will be more able to enter or continue employment than mothers. Mothers who do maintain strong market work attachment while still married are likely to be in a better position to do so after marriage than those who have not, presumably because working mothers' wages were sufficient to make employment economically feasible while raising children.

Children are likely to have a negative effect on postdisruption per capita income, even controlling for predisruption human capital. Although this is probably true for both men and women who continue to live with at least one of their children, prevailing child custody patterns suggest that this variable could account for much of the gender difference in postdisruption per capita income.

Capturing the effect of children is not straightforward. After preliminary analyses using various measures, I chose to simply include the respondent's number of children in the household at $T+1$ in the equation. Among women, this variable is, not surprisingly, strongly and positively correlated with number of children before separation or divorce ($r = .89$); one measure is sufficient to more broadly represent having had children at all for women. Still, this choice could have been potentially problematic for two reasons. First, children commonly reside with their mothers after marital disruption, perhaps leading to an extremely high correlation between gender and children. Fortunately, about 12% of men in this sample have at least one child living with them following separation or divorce, and there is a substantial minority of women with no children (43%). Thus, the correlation between being female and number of children following marital disruption is only moderately strong ($r = .45$). Second, using number of children at $T+1$ might not adequately represent men's experiences. That is, while only a minority of men have children living with them after separation or divorce, many have children living with their former wives, and men's postdisruption incomes are to an extent reduced if they pay child support. Analyses not shown, however, indicate no statistically significant effect of children prior to marital disruption on either measure of men's postdisruption economic welfare.

Finally, it may be that the negative effect of children is limited to women. It is not clear, for example, that children impede employment, either before or after marital disruption, among men who obtain custody. Indeed, analyses not shown suggest that those men who do have at least one of their children living with them after divorce are a relatively select subgroup with, in fact, higher earnings and personal income than other men. Therefore, I also allow for an interaction between gender and children, by including in some models two separate variables in lieu of the measure for overall number of children: (1) number of children if the respondent is male, and (2) number of children if the respondent is female. This specification is equivalent to a conventional interaction term, but the two coefficients represent the effect of children for men and for women, respectively, relative to the omitted category, rather than one coefficient representing the "extra" effect of children for one sex.

Control Variables

I include other variables in the equation as controls. First, because sample sizes are not large enough to support separate analyses by race-ethnicity, I include dummy variables for black and Hispanic respondents, with non-Hispanic whites being the omitted category. Interactions between race-ethnicity and independent variables were tested but none proved statistically significant. I also include a dummy variable indicating if the respondent is part of the white, disadvantaged oversample. Marital duration is measured in years; I do not include age at separation or divorce in the equations because it is strongly and positively correlated with years of work experience ($r = .58$ for women and $r = .70$ for men). Past research suggests that predisruption economic status has positive effects on postdisruption economic status (Duncan & Hoffman 1985; Morgan 1991); predisruption income is correlated with earnings, and, for those not employed, earnings potential. For example, men with relatively high incomes tend to marry women with high earnings or earnings potential (e.g., Treas 1987). I measure predisruption economic status as total family income, in thousands of dollars, at $T - 1$. Finally, I include a dummy variable indicating whether the respondent is living with any adult relatives at $T + 1$ in the per capita income analysis. Although coresidence is probably not exogenous to economic well-being, in that the decision to coreside among young adults is often a result of financial difficulties (e.g., Bumpass & Sweet 1991; Glick & Lin 1986; Hogan, Hao & Parish 1990), it is a necessary control in per capita income analyses. The expected relationship to per capita income is unclear because it depends on whether the income contributed by other family members exceeds the increase in the number of people who share this income.

Results

DESCRIPTIVE STATISTICS

Table 1 displays mean personal and per capita income surrounding the time of marital disruption by gender and race-ethnicity, along with predisruption and postdisruption "gender ratios" in the last two columns (women's mean pre- or postdisruption economic status divided by men's).[9] The top panel shows results for personal income. Prior to marital disruption, personal income is sharply higher among men than women because women's labor supply and earnings in marriage are usually lower than men's. Personal income while still married is roughly $17,000 for white men compared to just $7,500 for white women. Hispanic men's personal income is $13,000 compared to $7,000 for their female counterparts, and black men's income is about $10,000 in contrast to just $4,300 for women. The fourth column shows that the ratio of women's to men's predisruption personal income ranges from 43% to 55%.

After marital disruption, women's personal incomes increase quite sharply. White women's incomes increase, on average, from $7,500 to $12,800, black women's from $4,300 to $8,200, and Hispanic women's from $7,000 to $10,700. These aggregate figures mirror changes on the individual level as well; typically, women experience increases in personal income of more than 50% between $T-1$ and $T+1$ (not in table).[10] These increases are primarily a result of increased earnings. Overall, earnings account for almost 70% of women's postdisruption personal income. A minority of women do begin to receive AFDC after marital disruption (25% of white, 27% of Hispanic, and 45% of black women), and, for this subgroup, AFDC payments account for the majority of postdisruption personal income.

In contrast to women, men experience little change in income upon marital disruption; men's mean pre- and postdisruption incomes are quite similar. At the individual level, too, the typical change experienced by men ranges only from -3% to +13% (not in table).[11] As the last column of the table indicates, while women certainly do not attain parity with men, their postdisruption personal incomes are only moderately lower than men's. The gender ratio is .75 for whites, .85 for blacks, and .88 for Hispanics. That the ratio is substantially higher for blacks and Hispanics underscores the relatively low personal incomes, both before and after marital disruption, of young minority men. Hispanic men's personal incomes are not substantially higher than those of maritally disrupted white women, and black men's incomes are in fact lower.

The bottom panel of Table 1 indicates that gender differences in postdisruption per capita income are more dramatic than they are for personal income. Before marital disruption, per capita incomes are similar for men and women, although in most cases slightly higher for women. Because age at marriage is generally a few years older for men than women, the spouses of the women sampled here were somewhat older, and had slightly higher earnings, than the sample of separating and divorcing men.

After marital disruption, however, per capita income is dramatically higher than predisruption levels for all subgroups of men, and lower than predisruption levels for women, consistent with prior research (see, e.g., Holden & Smock

TABLE 1: Mean Personal Income and Per Capita Income of Men and Women Experiencing Separation or Divorce at Time *T*, by Race-Ethnicity[a]

	Men		Women		T - 1 Gender Ratio	T + 1 Gender Ratio
	T - 1	T + 1	T - 1	T + 1		
Personal income						
All	15,975	16,051	7,035	12,047	.44	.75
White	16,815	16,962	7,455	12,818	.44	.75
Black	10,065	9,676	4,339	8,222	.43	.85
Hispanic	12,716	12,167	6,971	10,698	.55	.88
Per capita income						
All	9,210	14,103	9,725	8,217	1.05	.58
White	9,659	15,445	10,624	9,125	1.10	.61
Black	5,204	7,634	5,534	4,203	1.06	.55
Hispanic	8,281	10,330	8,232	6,779	.99	.66

[a] All income amounts are in 1987 dollars and use weighted data. *T* represents the survey year of divorce or separation; *T* - 1 represents the survey year before, *T* + 1, the survey year after. Sample restricted to maritally disrupted men and women who are not remarried or cohabiting at *T* + 1. Sample sizes include 258 white, 110 black, and 84 Hispanic women, and 167 white, 44 black, and 43 Hispanic men. The gender ratio is defined as women's mean pre- or postdisruption economic status divided by men's mean pre- or postdisruption economic status.

Source: National Longitudinal Survey of Youth 1979-88

1991; Sørensen 1992). For example, per capita income rises from $9,700 to $15,000 for white men, from $5,200 to $7,600 for blacks, and from $8,300 to $10,300 for Hispanics. Among women, per capita income declines from $10,600 to $9,100 for whites, from $5,500 to $4,200 for blacks and from $8,200 to $6,800 for Hispanics. Analysis of changes at the individual level yields similar patterns. On average, men experience substantial increases in per capita income upon marital disruption, ranging from 18% for Hispanic men to 61% for white men. Women uniformly experience substantial decreases; white and Hispanic women experience declines of about 20% and black women of about 35% (not in table). Correspondingly, the gender ratio in the last column shows large gender disparities. The ratio is .61 for whites, .55 for blacks, and .66 for Hispanics.

In sum, descriptive statistics yield two main conclusions. First, even among young men and women, separation or divorce is associated with considerable gender disparity in economic outcomes. The inequality is relatively small for personal income, but striking for per capita income. Second, although sample sizes for minority men are small, it is clear that young minority men are quite economically vulnerable themselves. Their postdisruption personal and per capita incomes are as low, and for black men lower, than those of white separated and divorced women.

MULTIVARIATE RESULTS

Table 2 displays the means and standard deviations of the dependent and independent variables for men and women separately. As expected, several independent variables differ significantly by gender. Average predisruption work experience is 4.4 years for men, in contrast to slightly over 3 years for women. Men worked almost 40 weeks in the year before separation or divorce, compared to 32 weeks for women. Educational attainment differs only slightly for men and women (11.75 vs. 11.98), however, with the modal category for both men and women being a high school diploma. Almost 50% of men and women have 12 years of schooling; 21% of men and a slightly higher 25% of women have more than 12 years of education (not in table). As expected, women are much more likely than men to have children in their household after marital disruption. The average number of children is .89 for women in contrast to only .19 for men. This difference is not primarily a result of differing proportions having had children; prior to marital disruption over 40% of men and 57% of women have children, but only 12% of men report at least one child living with them after separation or divorce (not in table).

Statistics for the control variables indicate that the marriages of these men and women were brief ones, lasting just three to four years. Although not included in the equation, it is also important to note that the ages of these men and women are similar; women, on average, are slightly over 23 years old at $T - 1$ and men about 24 years old. Men have somewhat lower predisruption family incomes than women ($23,000 vs. $26,000). Finally, a substantial minority of both young men and women are living with relatives, primarily parents, after marital disruption: 41% and 35% among men and women respectively.

Personal Income

Table 3 shows the effects of independent variables on the natural logarithm of postdisruption personal income. Column 1 displays the equation with only gender and the control variables entered as independent variables. The positive and significant coefficient for gender, with women being the omitted category, indicates that being male leads to higher postdisruption personal incomes.

Columns 2 and 3 add the child measures to the equation. As column 2 shows, once children are taken into account the effect of gender diminishes substantially and becomes statistically insignificant. Broadly speaking, this finding implies that differences between men and women in their parenting roles — proxied by the presence of children following marital disruption — fully accounts for gender disparities in income.

However, column 3 reveals an important distinction in the effect of children for men and women. Among women, children are negatively and significantly associated with postdisruption income, while this is not apparently the case for men. The coefficient for number of children present among men is not statistically significant, and even positive; those young men who have custody of at least one child experience no shortfall in income. The insignificant

TABLE 2: Means and Standard Deviations of Variables by Gender[a]

	Men	Women
Dependent variable		
Ln (Personal income)	9.35 (.99)	9.12 (.84)
Ln (Per capita income)	9.18 (1.00)	8.58 (1.01)
Control variables		
Race (0 = white)		
Black	.09 (.28)	.14 (.35)
Hispanic	.06 (.24)	.08 (.27)
Oversample	.08 (.27)	.07 (.25)
Marital duration (years)	3.04 (1.81)	3.56 (2.26)
Family income at T - 1 (in $1,000s)	23.21 (14.19)	26.32 (17.71)
Lives with relative(s) at T + 1 (1 if yes)	.41 (.49)	.35 (.48)
Predisruption human capital		
Work experience by T - 1 (years)	4.40 (2.43)	3.06 (2.21)
Weeks worked in year T - 1	39.74 (17.65)	31.94 (19.86)
Educational attainment (years)	11.75 (2.11)	11.98 (1.88)
Children		
Number of children living with respondent at T + 1	.19 (.58)	.89 (.97)
Unweighted N	254	452

[a] Sample restricted to maritally disrupted men and women not remarried or cohabiting at $T + 1$, the survey year after separation or divorce. Income variables are coded in constant dollars with 1987 as the base year. Statistics are weighted.

Source: National Longitudinal Survey of Youth 1979-88

TABLE 3: Effects of Characteristics on the Natural Logarithm of Post-disruption Personal Income[a]

	Control Variables	Children		Human Capital	Children Plus Human Capital	
	(1)	(2)	(3)	(4)	(5)	(6)
Independent variable						
Constant	8.581**	8.686**	8.721**	7.209**	7.219**	7.259**
	(.093)	(.097)	(.098)	(.190)	(.197)	(.198)
Gender (0 = female)	.221**	.103	.033	.052	.047	.009
	(.068)	(.075)	(.080)	(.064)	(.070)	(.074)
Control variables						
Black	-.254**	-.186*	-.165	-.147	-.144	-.132
	(.086)	(.087)	(.087)	(.077)	(.078)	(.078)
Hispanic	-.138	-.121	-.106	-.047	-.046	-.039
	(.090)	(.089)	(.089)	(.080)	(.080)	(.080)
Oversample	-.280**	-.264**	-.248**	-.167*	-.167*	-.159*
	(.088)	(.087)	(.087)	(.078)	(.078)	(.078)
Marital duration	.027	.042**	.041**	.012	.013	.012
	(.015)	(.015)	(.015)	(.014)	(.014)	(.014)
Predisruption family income	.019**	.017**	.017**	.008**	.008**	.008**
	(.002)	(.002)	(.002)	(.002)	(.002)	(.002)
Predisruption human capital						
Educational attainment				.096**	.095**	.094**
				(.016)	(.015)	(.015)
Work experience				.059**	.058**	.058**
				(.016)	(.016)	(.016)
Weeks worked in T - 1				.011**	.011**	.011**
				(.001)	(.001)	(.001)
Children						
Number of children living with respondent at T + 1		-.130**			-.006	
		(.038)			(.035)	
Number of children if female			-.160**			-.024
			(.039)			(.037)
Number of children if male			.136			.137
			(.115)			(.103)
R^2	.157	.171	.178	.345	.345	.347

[a] Sample restricted to those not remarried or cohabiting at $T + 1$, the survey year after separation or divorce.
Source: National Longitudinal Survey of Youth 1979-88

* p < .05 ** p < .01

coefficient for gender in column 3 also implies that those women without children fare about as well as men; women with children are at an economic disadvantage.

The model in column 4 shows that predisruption human capital variables alone can also account for overall gender differences in postdisruption personal income. The gender coefficient becomes statistically insignificant once predisruption employment-related variables are included in the equation because men have more work experience and worked more weeks prior to marital disruption than women (see Table 2). Net of such factors, men have no postdisruption personal income advantage. This also suggests a key source of the finding from column 3 (i.e., that the negative effect of children is limited to women); on average, children impede women's labor force attachment but not men's.

Indeed, column 5, showing the effects of the full set of independent variables, indicates that children have no impact on personal income net of predisruption human capital. Similarly, the equation in column 6, allowing the effects of children to vary by gender, suggests no net effects of children for either men or women. Controlling for work experience prior to disruption, women — whether or not they have children — fare about as well as men. In reality, of course, women's work experience is certainly influenced by children. Analyses not shown indicate that women with children have roughly 2.5 years of work experience, compared to 3.5 years for those without children. And whereas childless women worked 39 weeks in the year before separation or divorce, virtually identical to the average for men, women with children were employed just 23 weeks. But assuming that men and women have equivalent levels of work experience, women suffer no personal income disadvantage relative to their male counterparts.

The effects of other variables are as expected. Schooling, work experience, and predisruption weeks worked are each associated with higher levels of postdisruption income once marital disruption occurs. This is true for both men and women; as noted previously, interactions with gender were tested but none proved significant. At least among these young, relatively less advantaged men and women, men do not appear to be receiving greater "income returns" to their human capital. Predisruption family income has a positive effect on postdisruption personal income; prior family income tends to be positively correlated with earnings, or earnings potential, and is therefore associated with higher levels of well-being when marriage dissolves. Finally, although Table 1 revealed relatively sharp race-ethnic disparities, black and Hispanic men and women do no worse than whites following marital disruption net of other characteristics.

Per Capita Income

Table 4 displays the results for postdisruption per capita income. Column 1 shows a positive and statistically significant effect for gender indicating that men have a fairly strong per capita income advantage when marital disruption occurs. Adding the living arrangements of children to the equation in column 2, the coefficient for gender diminishes dramatically (from .537 to .146) and becomes statistically insignificant. The presence of children is associated with lower levels of per capita income, as would be expected because there are more individuals who must share family income, and fully explains the gender effect.

Like the results for personal income, however, the results in column 3 indicate that the negative effect of children on per capita income is limited to women. As noted previously, this appears to at least partially stem from the fact that those men living with at least one of their children after marital disruption are a select subgroup better off economically than other men.

Column 4 includes human capital characteristics in the equation. Unlike the case for personal income, differences between men and women in their human capital simply cannot account for men's postdisruption per capita income advantage. Comparing columns 1 and 4, the gender coefficient does diminish somewhat — from .537 to .385 — but remains statistically significant. This finding is not surprising. Unlike personal income, per capita income is strongly and directly influenced by the number of dependents, and women are far more likely than men to remain with their children after marital disruption. Thus, even assuming equivalent levels of labor force attachment among men and women, women's customary role as the primary child caretakers after marital disruption places them at a distinct disadvantage in measures of economic well-being adjusted for household size.

The inclusion of the full set of independent variables in column 5 indicates that young maritally disrupted men have no net per capita income advantage. Adding the presence of children to the equation in column 4 reduces the gender coefficient from .385 to .109, and it is no longer statistically significant. This underscores that a good deal of women's per capita income disadvantage compared to men is directly associated with children's living arrangements; children continue to exert a negative effect on postdisruption per capita income even controlling for human capital characteristics. Yet, as column 6 shows, this is apparently the case only for women. Men who have custody of at least one child have per capita incomes as high as those who do not, even when their human capital characteristics are taken into account.[12]

Predisruption economic status, educational attainment, and weeks worked are each associated with higher levels of per capita income after marital disruption for both men and women, effects that probably operate largely through the numerator of per capita income (i.e., by increasing personal income). Living with relatives after marital disruption is associated with lower levels of per capita income. It is not possible here to make a causal argument about this effect, because an individual's choice of living arrangements is likely affected by expectations of economic well-being. But this result suggests that the families of these young men and women have limited resources themselves (i.e., the additional earnings from these family members do not offset the increase in the number of people sharing income).[13] Finally, unlike the case for personal income, black men and women do suffer a net per capita income disadvantage. This appears to be due to the fact that, although the equation controls for living with relatives, the income reported from other family members is somewhat lower for black men and women than for whites and Hispanics (data not shown).

TABLE 4: Effects of Characteristics on the Natural Logarithm of Post-disruption per Capita Income[a]

	Control Variables	Children		Human Capital	Children Plus Human Capital	
	(1)	(2)	(3)	(4)	(5)	(6)
Independent variable						
Constant	8.306**	8.682**	8.719**	7.086**	7.612**	7.661**
	(.109)	(.107)	(.108)	(.229)	(.226)	(.228)
Gender (0 = female)	.537**	.146	.080	.385**	.109	.066
	(.077)	(.079)	(.085)	(.077)	(.079)	(.085)
Control variables						
Black	-.594**	-.366**	-.345**	-.494**	-.339**	-.326**
	(.097)	(.093)	(.093)	(.092)	(.089)	(.089)
Hispanic	-.200	-.131	-.116	-.129	-.088	-.079
	(.103)	(.096)	(.096)	(.096)	(.092)	(.092)
Oversample	-.311**	-.256**	-.241**	-.207*	-.188*	-.179*
	(.099)	(.092)	(.092)	(.093)	(.088)	(.089)
Marital duration	-.035*	.015	.013	-.045*	-.003	-.004
	(.017)	(.016)	(.016)	(.017)	(.016)	(.016)
Predisruption family income	.024**	.019**	.018**	.014**	.012**	.012**
	(.002)	(.002)	(.002)	(.002)	(.002)	(.002)
Lives with relative at T + 1 (1 if yes)	-.320**	-.381**	-.387**	-.281**	-.332**	-.337**
	(.076)	(.070)	(.070)	(.072)	(.068)	(.068)
Predisruption human capital						
Educational attainment				.082**	.068**	.066**
				(.019)	(.018)	(.018)
Work experience				.035	.034	.034
				(.019)	(.019)	(.019)
Weeks worked in T - 1				.012**	.008**	.008**
				(.002)	(.002)	(.002)
Children						
Number of children living with respondent at T + 1		-.438**			-.347**	
		(.040)			(.040)	
Number of children if female			-.467**			-.367**
			(.042)			(.042)
Number of children if male			-.183			-.181
			(.122)			(.117)
R^2	.247	.355	.360	.352	.415	.417

[a] Sample restricted to those not remarried or cohabiting at $T + 1$, the survey year after separation or divorce.
Source: National Longitudinal Survey of Youth 1979-88

* $p < .05$ ** $p < .01$

Conclusions

Separation and divorce have far more serious economic consequences for women than men, even among young, less-advantaged adults. Unlike men, whose personal incomes undergo little if any change upon marital disruption, women dramatically increase their personal incomes. But these efforts are not enough to put them on par with men; the postdisruption personal incomes of women as a percentage of men's range from 75% among whites to 85% to 88% among blacks and Hispanics. Women are markedly worse off than men on economic status measures that take into account children's postdisruption living arrangements. Although men and women have similar per capita incomes before separation or divorce, women's postdisruption per capita incomes are just 55% to 66% of those of their male counterparts.

Indeed, multivariate analyses show that women's economic disadvantage after marital disruption stems, either directly or indirectly, from their responsibilities towards their children. At least for this sample of young adults, childless women fare about as well as men; mothers are at the distinct economic disadvantage. For example, the gender gap in personal income following separation or divorce can be explained in two ways. First, differences in men's and women's human capital fully account for gender differences in postdisruption income. Men and women with similar levels of work attachment before separation or divorce tend to have similar postdisruption personal incomes, largely because prior work experience predicts employment and earnings after marital disruption. But the reality, of course, is that few women with young children can achieve this level of labor force attachment. Childless women have significantly more years of work experience compared to women with children, and childless women worked as many weeks as men in the year before marital disruption.

Taking into account women's disproportionate responsibility for children also explains overall gender differences in postdisruption personal income. I measured this responsibility in the analysis using the number of coresident children after separation or divorce, because it is extraordinarily difficult to disentangle the influence of having had children before marital disruption from that of living with the children afterwards. This distinction is somewhat meaningless, in fact, because the two continue to covary quite strongly in the U.S. Not surprisingly, the gender disparity in per capita income cannot be explained solely by gender differences in human capital. Women's disadvantage in this measure, a crucial measure because it reflects the amount of income available to each family member, chiefly stems from children being far more likely to live with their mothers than with their fathers after separation or divorce.

Overall, these results are quite consistent with prior research showing that gender differences in labor supply are the major source of married women's economic dependency (Sørensen & McLanahan 1987). The current study shows how this dependency, stemming from the division of labor in marriage and, particularly, women's roles as primarily child caretakers, is also the major factor in the economic toll of marital disruption on women and children. Those men who do live with at least one child after separation or divorce appear to experience little or no economic disadvantage, either in personal or per capita income; admittedly, it is likely that the relatively strong economic prospects of these men were at least partially a cause of their custody of children. But the finding implies that the conflict between parenting and labor force attachment (either before or after marital disruption) is one borne almost invariably by women.

The focus of this study on a less-advantaged population has, in itself, important implications. On the one hand, these results confirm prior research using broader samples that marital disruption engenders precarious economic futures for women and children, and often better economic futures for men (e.g., Duncan & Hoffman 1985; Hoffman 1977; Sørensen 1992; Weitzman 1985). Even economically vulnerable young men tend to experience dramatic improvements in their per capita incomes when they separate or divorce (see Table 1).

On the other hand, this analysis challenges the notion of a two-tier society, the top consisting of men and the bottom of divorced women and their children (e.g., Weitzman 1985). This may be an accurate description of the ramifications of divorce for the middle class, but for less-privileged men and women, race-ethnic economic inequalities clearly crosscut, and are as pronounced, as those associated with gender. With postdisruption personal incomes of only roughly $12,000 for Hispanic men and $10,000 for black men (in 1987 dollars), Hispanic men are no more economically secure than white women who are separated or divorced, and black men are substantially less secure than white women. A similar pattern emerges even for postdisruption per capita income (see Table 1). To be sure, the economic difficulties of these young men tend to precede marital disruption, probably in fact contributing to marital instability (e.g., Hoffman & Holmes 1976), rather than being a "consequence" of marital disruption.

From a policy perspective, the focus here on economically vulnerable young adults also questions the emphasis in prior research on the dramatic improvement in men's standards of living upon separation or divorce. Public policy efforts to ameliorate the economic difficulties of single mothers aim to increase the incidence and amount of child support payments from nonresident fathers (e.g., Garfinkel & McLanahan 1986); thus, changes in men's economic status are less crucial than absolute levels of income. The small sample sizes here for minority men dictate some caution in extrapolating from my findings, but they suggest that many young men simply do not have the economic resources to provide ample financial support to their children. While Phillips and Garfinkel (1993) show that disadvantaged divorced fathers — initially poor or near-poor before divorce — experience rather considerable increases in income over a seven-year postdivorce period, this interim is a lengthy one, and especially so from a child's perspective. Child support policy proposals include the idea of providing a guaranteed minimum amount of support when the nonresident parent cannot pay (Garfinkel 1992; Garfinkel & McLanahan 1986). The results from this study underscore that such initiatives are imperative.

Notes

1. The survey also oversamples non-Hispanic white youth from disadvantaged geographic areas which I include in the analysis to increase sample size. Descriptive statistics are weighted; multivariate analyses include a control for this oversample. In preliminary analyses, I weighted the regressions. Apart from a loss of statistical power, results are similar to those presented here.

2. The NLSY did sample a small number of individuals married to each other in 1979 (n = 300 or 150 pairs). There are no more than a handful of these cases in the samples drawn here.

3. Approximately 37% of men and 34% of women in this sample are remarried or cohabiting at $T + 1$.

4. I estimated two-equation sample selection models to explore the possibility of sample selection bias. The first equation, a probit, predicted the likelihood of not being remarried or cohabiting at $T + 1$. The second equation predicted postdisruption personal income for those still single. The model allowed for correlations between the error terms of the equations. I estimated equations separately for men and for women. For both, the correlation was small and statistically insignificant, suggesting little evidence of sample selection bias.

5. I coded income as "missing" is any one of the following three sources was missing: (1) respondent's wage-salary income; (2) spouse's wage-salary income (if present); and (3) the income of any coresident family members (if present). Reported zeros for any of these categories were considered valid responses.

6. In this sample, slightly over half of the men with at least one child before separation or divorce report paying some child support. Among those that do, the median annual payment is $2,100 (in 1987 dollars), quite close to national statistics on child support amounts received by women aged 18–29 (U.S. Bureau of the Census 1990).

7. Estimating effects of characteristics on postdisruption economic outcomes by OLS is adequate only if women and men experience separation or divorce randomly, conditional on their observed characteristics (Heckman 1979). Studies show that while women whose marriages will disrupt are more likely to be employed, husband's low earnings and employment instability are positively associated the likelihood of separation or divorce (e.g., Hoffman & Holmes 1976). This suggests that the sample of maritally disrupted young men may be a relatively more "select" group than the women not only in measured ways, but perhaps in unmeasured ways as well. In other analyses, I estimated sample selection models of postdisruption economic welfare (similar to those I describe in Endnote 4) drawing also on subsamples of continuously married men and women in the NLSY. Again, I found no evidence of sample selection bias.

8. I examined whether the coefficient for an interaction between gender and each independent variable was at least twice its standard error in a full, pooled model.

9. Median levels of economic well-being are moderately lower for both young men and women because of the skewness of the income distribution. But general conclusions using medians are similar to those using means.

10. I compute changes at the individual level by dividing an individual's postdisruption personal (or per capita) income by predisruption personal (or per capita) income. I use the median of this distribution to illustrate the typical change in economic status experienced upon marital disruption.

11. Without deductions for alimony and child support, men's mean postdisruption incomes are slightly higher for all subgroups than those presented in Table 1. The modest deterioration in income between $T - 1$ and $T + 1$ among black and Hispanic men also disappears.

12. This statement may be somewhat misleading. As noted earlier, I found no statistically significant interactions between any variable (including number of children) and gender in a full, pooled model; column 6 is the full model. This implies that there is no statistically significant difference between -.367 and -.181 (the coefficients in column 6 of Table 4 for number of children for women and for men). However, the interaction coefficients between gender and number of children were statistically significant in the other models in Tables 3 and 4, supporting my basic interpretation of an important gender difference in the effect of living with children.

13. However, per capita income does not assume any economies of scale associated with larger household size. In other analyses I substituted the income-to-needs ratio for per capita income as the dependent variable, and the effect of living with relatives was not statistically significant.

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