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CHILD SUPPORT IN THE U.S.: CAN FATHERS AFFORD TO PAY MORE?

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This paper presents up-to-date estimates of the ability of non-resident fathers to pay child support. While no nationally representative data exist on the incomes of fathers, this issue has become more pertinent in recent years, as child support has become an important national issue. We find that fathers on average are able to pay nearly five times more in child support than they currently pay, and also that low income fathers can afford to pay substantially more than they actually pay. We also find that changes in nonmarital childbearing and the returns to education have had only minimal effects on trends in fathers' incomes.

INTRODUCTION

The past decade was an important period of reform for the child support system in the United States. During this decade a series of increasingly stringent laws was passed in an effort to secure more child support awards and to better enforce payment by non-resident parents. The issue of child support also gained a place in the national spotlight. Indeed, many policymakers who have recently taken aim at welfare have touted child support as the private alternative to support poor children.

While it is unclear whether the child support system can serve as a credible alternative to welfare, there is some reason to believe that better enforcement will lift many children out of poverty. The economic status of children has declined substantially over the past two decades, and much of this decline can be attributed to the increasing percentage of children living in female-headed households (Bane and Ellwood, 1989; Fuchs, 1986; Eggebeen and Lichter, 1991). It has been well-established that women's economic status declines following marital separation (Duncan and Hoffman, 1985). The poverty rate in 1990 for children living in female-headed families was 54 percent.

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Currently, however, a majority of these families are not receiving support from the non-resident parent. In 1989 nearly 11 million women and their families were eligible to receive child support, but only 60 percent of these women had child support awards. Moreover, among those women with awards, only half received full payment from the non-resident father. Many of the recent reforms are aimed at increasing both award rates and payment rates. The Family Support Act of 1988, for example, increases federal funding for the establishment of paternity and requires states to improve their paternity establishment rates, establish presumptive, or mandatory, child support guidelines and institute immediate wage withholding for all new child support cases.

Thus these new policies have the potential to significantly improve the well-being of children. Over half of all poor children live in female-headed families and thus are potentially eligible to receive child support. Garfinkel and Oellerich (1989) estimate that in 1983 a perfectly enforced child support program would have increased child support payments by over $20 billion.

An important assumption underlying child support reform, however, is that most absent fathers, particularly low income fathers, who pay little or no child support are able to pay more than they do. This assumption is by no means a given. In fact, two trends over the past decade would suggest that fathers today are on average less able to pay child support than were fathers 10 years ago. First, as a result of trends in non-marital fertility, the percentage of custodial mothers who are never married increased from 19 percent in 1979 to 30 percent in 1990. Previous research indicates that never married men earn less than their ever married counterparts (Korenman and Neumark, 1991). This trend will be expected to lower the mean and alter the distribution of non-resident fathers' incomes. The second factor to affect fathers' incomes is the continued increase in male earnings inequality that began in the early 1970s (Blackburn *et al.*, 1990; Levy, 1993). Levy, for example, finds that male earnings inequality increased fairly rapidly from 1979 to 1989 and that this increase was driven in part by the declining economic status of less educated men. The increase in earnings inequality highlights the importance of calculating more recent estimates of low income fathers' ability to pay child support.

Previous estimates of fathers' incomes and potential child support payments under the Wisconsin guidelines (Garfinkel and Oellerich, 1989) are based on data from the mid-1970s and may, for the reasons listed above, be biased. However, these estimates may also be criticized on methodological grounds. Since there are no nationally representative data on non-resident fathers, it is necessary to indirectly estimate fathers' incomes. We use the methodology developed by previous researchers, but improve upon it in several ways. First, our estimates will be derived from what we will argue is a more appropriate sample, and are, therefore, subject to less potential bias. Second, our estimates will be based on a sample of custodial families obtained in 1990, rather than in 1979. As noted above, the composition of custodial families has changed substantially since then. Finally, we reduce the number of steps and independent samples used in the estimation, as compared with the previous method, to reduce the chance of error.

This paper attempts to provide additional information on the current status and potential of the child support system by presenting up-to-date estimates of the income and ability to pay child support of non-resident fathers. We examine income differences by marital status and race, as well as the distribution of ability to pay child support. We focus particular attention on whether low income fathers can afford to pay more child support than they currently pay. We also examine the effect on fathers' incomes over the past decade of trends in non-marital fertility and changes in the returns to education. Additionally, given that this paper represents a methodological improvement over previous estimates, we also discuss, in some detail, the methodology and assumptions behind our estimates.

The paper proceeds as follows. Section One presents the findings and limitations of previous estimates. Section Two discusses the data and methodology used in our analysis and presents evidence on assortative mating among divorced couples. Section Three presents estimates of fathers' incomes and ability to pay child support under the Wisconsin guidelines. Section Four presents a decomposition of changes over time in fathers' incomes. Section Five concludes.

Previous Research

The ability of an absent father to pay child support is determined by his income and the number of children owed support. Currently, however, no nationally representative data source exists containing both types of information.[1] Using a sample of divorced men for this purpose, for example, will include men who are not non-resident fathers and exclude both married and never married men who are fathers.

Garfinkel and Oellerich (1989) employ an indirect method in which they predict the non-custodial father's income using the characteristics of the custodial family. They first use the Current Population Survey to estimate the relationship between a husband's income and his wife's race, age, years of schooling, and geographic location. Information on this relationship is then used to predict an initial value of the non-custodial father's income for each custodial family from the 1979 CPS-Child Support Supplement. The initial estimate is then adjusted for earnings differences among married men with children, divorced and separated men with children, and never married men. The differentials are estimated, net of other demographic characteristics, from a sample of men in the 1976 Survey of Income and Education (SIE). The SIE is also used to estimate the proportion of divorced men that remarry. Yet another data set and regression is used to further reduce the estimates for ever married men whose children are receiving welfare and to increase (somewhat) the estimates for ever married men whose children are not receiving welfare.

Using the adjusted estimate of the nonresident father's income and the Wisconsin child support guidelines—17 percent of gross personal income for one child; 25 percent, 29 percent, 31 percent, and 34 percent for two, three, four, and five or more children, respectively—Garfinkel and Oellerich estimate that non-resident fathers in 1979 were able to pay a total of $28.5 billion in child support

[1] A pilot survey of absent parents was initiated in the mid-1980s, but the full survey was never conducted (see Sonenstein and Calhoun, 1988).

(in 1983 dollars).[2] However, both the numbers and demographic composition of families potentially eligible for child support have changed substantially since 1979. More recent calculations for 1988, which are also estimated indirectly using data on custodial families for the Survey of Income and Program Participation, yield an estimate of $49 billion (Kim, 1993).

There are a number of weaknesses to the Garfinkel–Oellerich estimate. First, much of the data used are nearly two decades old. Trends in the timing and incidence of marriage, for example, suggest that income differentials between married and never married men may not be the same today as they were in 1975. In addition, current remarriage rates may differ from those estimated from the 1976 SIE. Second, even though each step in the estimation process seems sensible, the sheer number of steps, each relying upon a different data set, may introduce some error. Third, the method implicitly assumes that assortative mating is similar for married and divorcing couples and that, controlling for demographics, the incomes of never married men with children are similar to the incomes of never married men without children. We will test these assumptions in this paper. Finally, the estimate Garfinkel and Oellerich provide—potential aggregate child support payments—gives no indication of the potential payments to different types of custodial families or of the distribution of fathers' ability to pay child support.

More recently, Sorenson (1993) attempts to identify a sample of non-resident fathers from the 1990 Survey of Income and Program Participation (SIPP) by including (1) men who have been divorced or separated and report making child support or financial payments to children under 21 who live elsewhere, and (2) men who report having had more children than are currently living with them. One problem with these criteria is that the first group includes fathers supporting their college children and the second includes fathers with grown children living apart from them. To address these issues, Sorenson excludes men above certain ages and men who have been married, divorced or separated for at least 16 years. While these restrictions improve the accuracy of the sample, an unknown number of men remain inappropriately included. In addition, Sorenson's sample of non-resident fathers is only 85 percent as large as the number of custodial mothers in the SIPP. In order to adjust for the difference, she (1) assumes that the number of non-resident fathers equals the number of resident mothers, (2) assumes assortative mating and uses the race and age distributions of custodial mothers to obtain an estimate of the race and age of the missing fathers, and (3) assumes that the incomes of the missing fathers are identical to the incomes of the observed fathers of similar race and age. Sorenson also finds, and corrects for, under-reporting of fertility among black males, especially those in their teens and twenties.

To the extent that assortative mating holds, these adjustments are likely to compensate for the inappropriate inclusion of men discussed above. Indeed, Sorenson's estimate of aggregate ability to pay child support using the Wisconsin

[2]Garfinkel and Oellerich test the robustness of their methodology by utilizing it to predict the incomes of selected sub-samples of non-resident fathers from other data sets that have data on the incomes of these fathers. In only one case do they overpredict income and then by only two percent. In only two cases is the underprediction large—20 percent and 25 percent. For a review of other studies of the incomes of non-resident fathers, see Phillips and Garfinkel (1993).

child support guidelines is $53 billion, only 8 percent higher than the updated estimate using the Garfinkel-Oellerich methodology. As with Garfinkel and Oellerich, however, Sorenson provides no test of her assumption of assortative mating.

This paper uses the methodology found in Garfinkel and Oellerich to estimate the incomes of non-resident fathers, except that we estimate the relationship between a husband's income and his wife's characteristics from a sample of divorced couples.[3] Thus we eliminate any potential bias caused by differences between married and divorced couples and the need to adjust for such bias. In addition, we estimate the incomes of never married men directly and test for differences in income between those with and without children.

Data and Estimation

The data source most frequently used to examine trends in child support is the series of April Child Support Supplements to the Current Population Survey (CPS–CSS). In 1979 the Census Bureau surveyed a sample of women from the March supplement who had children under 21 whose fathers were absent from the household. These women were asked several questions pertaining to child support, such as whether they were owed child support, the amount they were owed, and the amount they actually received during the previous year. The April supplement was then merged with the March Annual Demographic File and thus contains a wealth of demographic information on custodial families eligible to receive child support. Child Support Supplements were repeated in 1982, 1984, 1986, 1988 and 1990. With the exception of the 1979 survey, none of the surveys collected information about the absent father.[4]

We use the indirect method mentioned above, coupled with the 1990 CPS–CSS, to calculate up-to-date estimates of fathers' incomes. We first estimate the relationship between a mother's characteristics and the income of her children's father. The estimated coefficients are then applied to the sample of CPS–CSS custodial families to predict the income of the non-resident father.

In order to predict the incomes of ever married non-resident fathers we estimate a regression in which a husband's post divorce income is a function of his wife's characteristics at the time of their divorce.

$$Y_m = X_w \beta + \varepsilon. \tag{1}$$

Y_m is the husband's income in the period following the divorce, and X_w is a vector of standard demographic characteristics of the wife in the period preceding the divorce.

Equation (1) is estimated using data from the Survey of Income and Program Participation (SIPP). The SIPP consists of a series of 28 to 32 month panels, the first of which was initiated in 1984, in which a nationally representative sample

[3]Michalopoulos and Garfinkel (1989) apply this method to a sample of divorced couples from the 1984 SIPP in order to estimate potential child support payments for single mothers receiving AFDC.

[4]In the 1979 survey the women were asked about the income of the non-custodial father. However, only 20 percent of the women provided a response.

of households is interviewed every four months for the duration of the panel.[5] The survey obtains detailed information on monthly income, program participation, and household composition. For each of the panels, from 1984 to 1990, we create a sample of couples who divorced or separated during the survey period and had children at the time of their separation. We then continue to follow the husband for the remainder of the survey and record his income for each month subsequent to the divorce. Thus, some men will be followed for up to 31 months, while others will be followed for only one month. Men for whom we cannot observe at least one month of post-divorce income are excluded from the sample. We then pool the samples from each panel for a total of 1,055 observations.[6] Sample means are presented in Appendix A.

The results of estimating two variants of equation (1) for the sample of divorced couples are shown in columns 1 and 2 of Table 1. The dependent variable is the husband's annual income—12 times average post-divorce monthly income—in 1990 dollars.[7] The characteristics of the wife in the period preceding the divorce are included as regressors, as are dummy variables indicating whether the observation had missing values for age or education.[8] In the first column of Table 1, we include no interactions and specify age as a linear function. The coefficients are all in the expected direction and generally statistically significant. In the second column, age is specified as a quadratic function and education is interacted with both age and race-ethnicity. Many individual coefficients are now insignificant. For the prediction in the second stage, however, we use the coefficients from this specification because it is more consistent with previous research and the signs of the interaction terms conform with previous findings.

In order to further explore the husband's post-divorce income, we take advantage of the longitudinal nature of the SIPP. Instead of calculating an average of monthly post-divorce income for each father, we stack the monthly SIPP observations—in effect treating each person-month as if it were an independent observation. We then re-estimate a variant of equation (1) with monthly income as the dependent variable and include a variable for the number of months since the divorce. The coefficient on months since the divorce (not reported) is positive and statistically significant—but not very large ($6 per month). This result provides evidence that mens' post-divorce income increases over time, suggesting that, in this respect, our estimates of fathers' incomes can be considered conservative.

To the extent that never married fathers are poorer earners than ever married fathers (Garfinkel and Oellerich, 1989; Korenman and Neumark, 1991), using

[5]The 1984 and 1990 panels each contain about 20,000 households, while panels initiated in each of the intervening years consist of approximately 12,000 households. The 1989 panel is currently unavailable from the Census Bureau.

[6]Approximately 15 percent of the men (and 7 percent of the women) dropped out of the survey immediately following divorce and were, therefore, not included in our analysis. A calculation of sample means revealed that these men were somewhat more likely than the included men to be black or Hispanic. They were also slightly younger and less educated.

[7]A log-linear model was estimated but did not fit the data as well as the linear model. We also excluded dummy variables indicating the year of the panel because they were not statistically significant and their inclusion did not change the other coefficients.

[8]Although we do not indicate whether the man is remarried or in a first marriage, earlier estimates by Garfinkel and Oellerich (1989) indicate no difference in earnings between married and remarried men.

TABLE 1

REGRESSION ESTIMATES FOR HUSBAND'S ANNUAL INCOME (1,000's)

	All	All	Black	White	Hispanic
Intercept	−21.19	13.90	−29.61	−20.69	−18.62
	(3.71)	(14.51)	(7.38)	(2.94)	(5.98)
Age	0.501	−0.160	1.457	1.732	1.603
	(0.075)	(0.647)	(0.335)	(0.139)	(0.327)
Age^2	—	−0.010	−0.022	−0.035	−0.029
		(0.009)	(0.005)	(0.002)	(0.005)
Education	2.301	−1.634	0.540	−1.643	−1.105
	(0.255)	(1.104)	(0.604)	(0.237)	(0.412)
Education × Age	—	0.115	0.040	0.107	0.071
		(0.031)	(0.019)	(0.008)	(0.015)
Education × (Black or Hispanic)	—	−0.426	—	—	—
		(0.296)			
Black	−6.919	−1.680	—	—	—
	(1.867)	(3.733)			
Hispanic	−7.200	−2.576	−1.518	—	—
	(2.102)	(3.599)	(2.356)		
Urban	2.780	2.855	1.561	3.131	2.064
	(1.088)	(1.086)	(0.976)	(0.341)	(1.024)
R^2	0.17	0.18	0.28	0.36	0.31

Note: (1) Columns 1 and 2 are estimated using the SIPP divorced sample ($N = 1{,}055$). The wife's characteristics are used as regressors. Also included are dummy variables for whether the respondent's age and education are missing. (2) Columns 3–5 are estimated using a sample of never married men ages 17 to 55 from the 1989 March CPS ($N = 981$, 8,387, 1,008, respectively). (3) Region dummies are included in all models. (4) Standard errors are in parentheses.

the coefficients obtained from equation (1) will overestimate the incomes of never married men. In order to obtain first-stage estimates for never married non-custodial fathers, we estimate a traditional earnings equation for a sample of never married men in which the man's income is regressed on his characteristics.

$$Y_m = X_m \gamma + \varepsilon. \tag{2}$$

This equation is estimated using three samples of black, Hispanic, and white-non-hispanic never married men ages 17 to 55 from the 1989 Current Population Survey. The results are presented in columns 3 through 5 of Table 1. The dependent variable is the respondent's annual income in 1990 dollars. The regressors include age, education, race and place of residence. All of the coefficients are of the expected sign, and nearly all are statistically significant.

In the second stage of the estimation, we use the coefficients obtained from equations (1) and (2) along with the characteristics of the custodial mothers from the 1990 Child Support Supplement to predict the absent father's income. The 1990 CPS–CSS contains information on 4,342 women who were eligible to receive child support in 1989. We also make the following simplifying assumptions. First, we assume that there is only one absent father associated with each custodial mother, i.e. no custodial mother has children from two fathers who both owe child support, and no man has fathered children by two different mothers and owes them both child support. Second, we assume that no absent fathers have been incarcerated or have died without the knowledge of the custodial mother. While it is unclear how a violation of the first assumption will bias the results, a violation of the second assumption will clearly lead to an overestimate of fathers' incomes. This bias is likely to be most pronounced when we predict absent fathers' incomes for young black women, given the high rate of incarceration among young black men. We also assume that the marital status of the custodial mother is the same as that of the non-resident father, i.e. if the mother is ever married (never married), then the father is ever married (never married). (Since some of the fathers of children of never married mothers will be married, this assumption will lead to an underestimate of non-resident father's income.)

Thus, for each custodial family in which the mother is ever married, we use the coefficients obtained from equation (1) combined with the characteristics of the woman to obtain a predicted level of income for the ever married absent father. For never married mothers, we assume that the father is also never married and use the coefficients from equation (2) along with the custodial mother's characteristics to predict the income of the non-custodial father. We also assume that positive assortative mating occurs: The father is the same race and ethnicity as the mother, he is 2.66 years older than the mother, and he has 0.126 more years of schooling than the mother (see discussion below). We adjust each never married mother's age and education accordingly before predicting the father's income. Thus our estimates for never married fathers depend upon two critical assumptions: (1) that positive assortative mating, as described above, occurs and (2) that the incomes of never married fathers are the same as the incomes of never married men without children, net of other demographic characteristics. We examine these two assumptions below.

Assortative mating is a common assumption in this area of research, and one on which previous estimates are also based (see section 2). In the following section we present evidence on assortative mating patterns among divorced and married couples. This evidence will serve two purposes. First, it will support our argument that we are using a more accurate sample than previously used to obtain first-stage estimates. In particular, we argued above that the use of a sample of married couples to derive the relationship between the husband's income and the wife's characteristics will yield biased estimates if mating patterns differ between married and divorced couples. Second, these estimates will provide us with assortative mating patterns to use in our analysis of never married couples.

Assortative Mating Among Divorced Couples

According to the economic theory of marriage (Becker, 1981), positive assortative mating is the optimal outcome in the marriage market, given that the gains from marriage for a given couple are higher if the two individuals are relatively similar in terms of personal characteristics.

Empirical evidence on marriage patterns is consistent with the theory. For a majority (60 percent) of married couples, the husband is the same age as the wife or only one to three years older. Additionally, over 50 percent of couples are in the same educational category (Sweet and Bumpass, 1989). No research to our knowledge, however, has examined these differences among couples who separate.

The existence of marital separation rests on the argument that the marriage market is characterized by imperfect information (Becker *et al.*, 1977). The primary cause of separation is newly acquired information subsequent to the marriage. Thus there is little reason to expect couples who divorce to be relatively less similar in terms of age and education, given that information on these characteristics of the spouse is available at the time of the marriage. However, Becker *et al.*, cite evidence that "persons who marry out of their race, religion, education, and age group have relatively high probabilities of divorce." [See Levinger (1965) for an early review.] Since assortative mating is optimal, they argue, the gains from marriage are lower for couples who are sufficiently different from each other. People who marry dissimilar mates are assumed to have relatively high marital search costs (for some reason) and are, therefore, more likely to accept a less-than-perfect match. An alternative explanation is that the match is not less-than-perfect, but rather that couples who are relatively dissimilar may be more likely than other couples to become different as they age. This newly acquired information increases the likelihood of separation.

Table 2 presents data on the characteristics of couples from the 1984–90 panels of the SIPP. The first column reports age and education differences between spouses for the sample of divorced couples, while the second column presents similar data for a combined sample of couples from each panel who were married and with children in the first wave. The husbands in the married sample are on average not as many years older than their wives as the divorced men, although the dispersion of this difference, as measured by the coefficient of variation, is greater for the divorced sample. On the other hand, married men tend to be relatively more educated than their wives as compared with divorced men, and the dispersion of these differences is smaller for the married men. The mean differences for education level are statistically significant at the 5 percent level.

An alternative measure of assortative mating is the correlation between the spouses' age and education shown in the bottom two rows of Table 2. This correlation can be thought of as indirectly related to the coefficient we would obtain from a regression of the husband's income on, for example, the wife's age. The ages of the husband and wife are highly correlated among married couples (0.878) and less so for couples who divorce (0.788). Similarly, the education levels of the spouses are less correlated among the divorced couples. Differences between the correlation coefficients are statistically significant at the five percent level.

Thus, although the differences in sorting between married and divorced couples are not of a substantial magnitude, the estimates in Table 2 do suggest that divorced fathers' incomes can be more accurately predicted from the estimated relationship between mothers' characteristics and fathers' incomes found from a sample of divorcing couples. In order to test this hypothesis more directly, we add to our sample of SIPP divorced couples with children a sample of married couples with children and re-estimate equation (1), including a dummy variable for divorce and interactions terms between divorce and several of the demographic characteristics of the wife. Table 3 presents the results. Two results are noteworthy. First, the coefficient on divorce is negative and statistically significant, indicating that net of other factors, divorced men earn less than married men. Equally interesting, the coefficient on the interaction between years of schooling and

TABLE 2
AGE AND EDUCATION DIFFERENCES BETWEEN SPOUSES: MARRIED AND DIVORCED COUPLES WITH CHILDREN

	Couples who Separate During Survey	Couples Married in First Wave
Age of husband– age of wife	2.66 [1.91]	2.52 [1.64]
Education of husband– education of wife	0.126 [18.37]	0.278 [8.66]
Correlation between husband'sand wife's:		
Age	0.788	0.878
Education	0.545	0.638
N	1,006	7,056

Source: SIPP 1984–90.
Note: Sample weights used. Coefficient of variation in brackets.

divorce status is significant, providing some evidence of differences in returns for divorced and separated fathers. Additionally, the coefficient on the interaction between race and divorce is positive, though marginally significant. Taken together these findings suggest that estimating the incomes of formerly-married fathers from a sample of formerly-married fathers is superior to estimating their incomes from a sample of married fathers.[9]

For never married fathers, given that we have no comparable separating sample, we use coefficients from income equations estimated for samples of never married men. Thus, we use the mean age and education differences from the SIPP sample reported in Table 2 to adjust the never married mothers' characteristics. While there is little information, to our knowledge, on sorting patterns among never married couples, unpublished data from the National Center for Health Statistics indicate that assortative mating with respect to age is similar among never married couples to that reported in Table 2. In California (one of three states in which 80 percent of births to unmarried women are accounted for in the data), for example, the unmarried women who had children in 1990 were on average 2.4 years younger than the fathers of these children.

Never Married Fathers

In addition to assortative mating, we have also assumed that, net of demographic characteristics, there are no earnings differences between never married fathers and never married men who have not fathered children. Robertson (1995), however, uses data from the National Longitudinal Survey of Youth and finds that never married men who have fathered children earn less than their childless

[9]At the very least, the significant divorce coefficient indicates that if one starts with a sample of married couples, then it is necessary to adjust the income estimate downward, as did Garfinkel and Oellerich. The coefficients on the interaction terms provide evidence that even an ad hoc adjustment of this kind may err.

TABLE 3

REGRESSION ESTIMATES
DEPENDENT VARIABLE: LOG (ANNUAL INCOME)

	Married and Divorced Couples with Children
Intercept	4.745 (0.141)
Divorced	−0.366 (0.156)
Divorced × Education	0.021 (0.011)
Divorced × Age	−0.002 (0.004)
Divorced × Black	0.127 (0.095)
Age	0.092 (0.008)
Age^2	−0.001 (0.0001)
Education	0.073 (0.004)
Black	−0.396 (0.037)
Hispanic	−0.273 (0.035)
Urban	0.209 (0.020)
Northeast	0.018 (0.028)
West	0.030 (0.026)
South	−0.092 (0.023)
Education missing	0.475 (0.177)
Age missing	2.43 (0.249)
Income set to $1	−7.38 (0.090)
R^2	0.51
N	8,137

Note: (1) The model is estimated for a sample of married and divorced couples with children from the SIPP. Year of panel dummies are also included. (2) Standard errors are in parentheses.

counterparts.[10] He estimates an earnings equation for never married men, controlling for a variety of demographic characteristics, and reports coefficients on non-custodial father status of −245, −7,815, and −5,657 for black, white, and Hispanic fathers, respectively. We use these coefficients to reduce estimated income for each never married father, by $245 for black fathers, $7,815 for white fathers, and $5,658 for Hispanic fathers. Some caution is in order, however, given the findings from previous research that never married fathers underreport their fertility (Cherlin *et al.*, 1983). If the likelihood of underreporting is correlated with some unobservable aspect of earnings capacity, then these coefficients will be biased. Unfortunately, we cannot test this hypothesis.

ESTIMATED INCOMES OF NON-CUSTODIAL FATHERS

We predict the income of the absent father for each custodial family from the 1990 CPS–CSS by using the coefficients reported in Table 1. Table 4 presents estimates of the average incomes of non-resident fathers. We also "build back" the variance of the observed distribution into the predicted distribution. The

[10]For details on the sample and variable construction see Robertson (1995). The regressions that Robertson reports in his dissertation include a number of additional independent variables. The results reported here are not reported in the dissertation, but were conducted by Robertson especially for this paper.

variance of predicted income does not account for the error term in the regression and, therefore, underestimates the variance of actual income. In order to build back in the variance, for each observation we multiply the standard error of the first-stage regression by a generated random number from a standard normal distribution. This number is then added to the value for predicted income, and the standard deviation of this series is calculated.[11]

Estimated average income for all non-custodial fathers is $22,998. Average income in 1990 for all men ages 15 to 64 is $27,032. Thus, non-resident fathers' earn about 15 percent less than other similar aged men. Our estimate of average income is very similar to those provided by other studies, only two percent higher than the updated Garfinkel–Oellerich estimate and no different from Sorenson's estimate of $23,300.

TABLE 4

ESTIMATED MEAN INCOME OF NON-CUSTODIAL FATHERS

	All	Black	Non-black
Remarried	28,226	21,798	28,621
	(17,227)	(16,089)	(17,163)
Divorced	27,603	23,112	28,562
	(17,752)	(16,754)	(17,701)
Separated	23,079	19,945	24,357
	(17,128)	(14,946)	(17,787)
Never-married	13,621	15,465	11,301
	(11,905)	(11,553)	(11,848)
All	22,998	17,835	24,916
	(17,284)	(14,312)	(17,946)

Notes: Calculations from the 1990 CPS–CSS. $N=4{,}144$. Standard deviations in parentheses. All figures in 1990 dollars.

The data indicate that while there is substantial within-group variance, as shown by the standard deviations, the between-group differences are also large. Black non-resident fathers' income is 28 percent less than white fathers' income. The mean income of never married black fathers, however, is 29 percent lower than that of their remarried counterparts. This difference for white fathers is 60 percent.[12]

[11]We can compare our final estimated income for ever married fathers ($26,882) to the actual annualized incomes of the fathers from the SIPP ($25,200). Although the latter is somewhat lower than our final estimate, the SIPP fathers are about three years younger on average than the sample of all ever married fathers, given that they come from a sample of very recent divorces. As an additional comparison, using the first few months following the separation we obtain an average monthly child support payment per child by the SIPP fathers of $85 (or $1,020 annually). This figure is consistent with actual payments per child reported for all ever married fathers ($993).

[12]The income difference between never married and ever married men that we obtain is lower than that obtained by Garfinkel and Oellerich. Our estimate for this difference using the 1979 CPS–CSS is also considerably lower than their estimate. As mentioned above, they calculate income for never married fathers by applying an adjustment (determined by a coefficient on marital status in an earnings equation for all men) to the equation used to impute income for all fathers. The coefficient they use to adjust income was obtained from 1976 data. Separate regressions using the 1976 and 1989 Current Population Surveys (not reported) indicate that the percentage income difference between married and never married men has fallen by 40 percent since 1976. Additionally, they apply an independently estimated coefficient to adjust for AFDC receipt by the custodial family. Thus they may overadjust downward the incomes of never married men.

Fathers' Ability to Pay Child Support

In this section we apply the Wisconsin child support guidelines to the estimated income of each non-custodial father. These guidelines are the easiest to implement given that the amount of support due is based only on the father's gross personal income and the number of children owed support.[13] This information allows us to estimate potential child support payments, assuming that fathers pay 17 percent of their income for one child, 25 percent, 29 percent, 31 percent and 34 percent for two, three, four, and five or more children, respectively. Using this standard, we estimate that aggregate ability to pay child support is $45.4 billion and that the average annual potential payment per family is $4,833 (in 1990 dollars). Despite the fact that our estimate of fathers' average income is similar to Sorenson's estimate, our aggregate figure is approximately 14 percent lower than her estimate of $53 billion. The discrepancy partly reflects an undercount by the CPS of eligible mothers. The CPS–CSS surveys do not define women as child support eligible if they have had children out-of-wedlock but have subsequently married. Thus our calculations are based on a sample of custodial mothers that is approximately seven percent smaller than Sorenson's sample.

Table 5 presents the average potential payment per child for different characteristics of the custodial family. Potential payments, using the Wisconsin guidelines, are compared with actual child support payments as reported by the custodial mother. We also compare each of these estimates with the actual amount of child support owed, among women owed support, in order to examine the extent to which the discrepancy between the amount paid the potential amount paid is due to the low award level.

Panel 1 of Table 5 presents estimates for all women. The figures illustrate the dramatic difference between what absent parents on average are reported to actually pay in child support ($764 per child) compared with what we estimate they would be able to pay under the Wisconsin guidelines ($3,369 per child). On average, non-resident fathers pay only 23 percent of what they could potentially pay in child support, while never married non-custodial fathers pay only 10 percent.[14] Panel 1 also indicates, however, that a substantial part of the child support "gap," as proxied by the difference between payments under a perfectly enforced system using the Wisconsin guidelines and current payments, can be attributed to the low average amount of child support due, or the award level.[15] While fathers actually pay only 23 percent of what they could potentially pay under the Wisconsin guidelines, they are paying 67 percent of what they actually owe in child support. The amount they actually owe, in turn, is only 34 percent of what they would owe under the Wisconsin guidelines.

[13]All states are free to develop and use their own guidelines, and the Wisconsin percentage-of-income guideline is but one of many available. Thus, while we have chosen this guideline for its simplicity, our results would be somewhat different if we used, for example, an income-shares guideline. Garfinkel and Oellerich (1989) found that aggregate ability to pay using a Colorado income-shares guideline was $30.1 billion (in 1983 dollars), compared with $28.5 billion using the Wisconsin guideline.

[14]See Weiss and Willis (1985) for an analysis of why absent fathers may have little incentive to pay child support.

[15]This finding is consistent with previous research on award levels (Oellerich, Garfinkel, and Robins, 1991).

The estimates make clear, however, that the child support system could potentially collect much more than it is currently collecting from absent fathers. Moreover, children in poverty and the children of never married parents stand to benefit from a perfectly enforced child support system. Although the non-resident fathers of these children are able to pay only about 64 percent of what all fathers can pay on average, this amount is roughly 7 to 10 times what they are reported to actually pay.

TABLE 5

Potential and Reported Child Support Payments

	(1) All Fathers			(2) Fathers Paying Some Child Support			(3) Fathers Paying No Child Support
	Potential	Owed	Received	Potential	Owed	Received	Potential
All	3,369	1,138	764	3,909	2,344	1,976	3,028
Ever married	3,966	1,471	993	4,150	2,443	2,059	3,796
Never married	1,948	307	197	2,154	1,558	1,312	1,912
Black	2,471	505	295	2,765	1,687	1,398	2,393
Non-black	3,702	1,372	938	4,108	2,458	2,076	3,367
In Poverty	2,283	593	334	2,835	1,671	1,300	2,093
Above Poverty	4,054	1,481	1,036	4,283	2,575	2,208	3,854

Source: 1990 CPS–CSS. All figures in 1990 dollars.

The estimates from Panel 1 of Table 5 are for all women eligible to receive child support and thus are somewhat misleading. Many of these women do not have child support awards and, consequently, are not due any child support. The very low average payment received by custodial families is driven in part by the fact that most families are receiving no payments. In addition, the low average award level as a fraction of potential payments is caused by the fact that many women are owed no child support. In Panels 2 and 3 of Table 5 we divide the sample into those women receiving payments and those women not receiving payments. The results indicate that fathers who do pay child support tend to pay the majority of what they owe.

Among those fathers who are paying child support (Panel 2), they pay on average 51 percent of what they could potentially pay and 84 percent of what they actually owe. These payment rates are also very similar across characteristics of the custodial family. Again, the estimates indicate that the majority of the child support gap is due to the low award level relative to what the award would be under the Wisconsin guidelines. Panel 3 of Table 5 reports potential payments for custodial families who are not receiving child support. Actual payments and the amount owed and are not relevant for this sample, but the estimates of potential payments do indicate that fathers who pay child support have higher average incomes than fathers who do not.

To further examine the causes of the aggregate child support gap, we present the following exercise. The total child support gap of \$35.2 billion (\$45.4 in potential payments less \$10.2 in actual payments) can be attributed to low compliance, low award levels, and low award rates. Using the aggregate amount potentially owed, actually owed, and actually paid among families with child support awards, we estimate that 31 percent of the payment gap is attributable to low award levels, or the difference between potential and actual awards. Another 14 percent of the gap can be attributed to low compliance, or the difference between actual award levels and actual payments. While the remaining gap might be fully attributed to low award rates, we know that families without awards would not, if they were to receive awards, receive potential award amounts or complete payments. Thus, using the average payment rate and the average actual award level as a fraction of the potential award level among families with awards, we estimate that obtaining a child support award for every family would reduce the payment gap by 33 percent. The remaining 22 percent of the gap can be allocated between low compliance and low award levels. Thus, these aggregate calculations accord with the conclusions drawn from Table 5, in that the bulk of the payment gap can be attributed to low award levels and low award rates.

We look further at the distribution of ability to pay in Table 6. This table presents a comparison of potential payments and actual payments by quintile of estimated potential payments. We also present estimates of potential payments that arise after the variance has been added back to the income distribution.[16] Adding back the variance helps to approximate the true distribution of income. Thus, the results for the variance-added estimates can be considered lower (upper) bounds for the bottom (top) quintiles. The true distribution of ability to pay most likely lies somewhere in between the randomized and non-randomized estimates.

The results indicate that fathers in the bottom quintile can afford on average to pay from \$321 to \$1,209 per child in child support. These two estimates suggest that these fathers, who are actually paying an average of \$167, are paying from 14 percent to 52 percent of what they can pay. This finding that these fathers are paying up to one-half of what they potentially should pay is consistent with Meyer's finding from the National Survey of Families and Households that, as an upper bound, fathers with incomes less than \$10,000 are paying 43 percent of what they would be paying under the Wisconsin guidelines (Meyer, 1995). Fathers in the top quintile, in contrast, are paying from one-fifth to one-quarter of what we estimate they could pay. The estimates also indicate that fathers in the middle three quintiles can pay up to four times more than what they currently pay.

The results suggest that the answer to one question raised in the introduction is that child support reform will raise collections from the entire distribution of fathers, although less so from poor fathers. Our estimates of income and child support payments are in agreement with previous studies in that they suggest that most absent fathers have the resources to pay substantially more child support than they are currently paying; on average they are able to pay nearly five times more in child support than they actually pay. However, the average payment each eligible child would potentially receive varies considerably with her race, poverty status, and the marital status of her mother. Additionally, a substantial part of

[16]Adding a random error component to each predicted income allows us to approximate the true variance of the distribution and the mean of potential payments within each quintile. However, this procedure also eliminates much of the correlation between potential and reported payments. Actual payments by quintile reported in Table 6 are therefore estimated based on a ranking of predicted incomes for which the variance has not been added back. Thus, although we can compare potential (variance-added back) payments to actual payments within a given quintile, they are not calculated over exactly the same samples.

TABLE 6

AVERAGE POTENTIAL CHILD SUPPORT PAYMENTS BY QUINTILE

Quintile	(1) Actual Payment per Child	(2) Potential Payment[1] per Child (cols 1/2)	(3) Potential Payment[2] per Child (cols 1/3)
1	167	1,209 (0.14)	321 (0.52)
2	339	2,401 (0.14)	1,776 (0.19)
3	756	3,352 (0.23)	3,126 (0.24)
4	991	4,202 (0.24)	4,667 (0.21)
5	1,569	5,850 (0.27)	7,687 (0.20)

Source: 1990 CPS–CSS. All figures in 1990 dollars.

[1]Column 2 payments are based on an estimate of the income distribution for which the variance has not been added back.

[2]Column 3 payments are based on an estimate of the income distribution for which the variance has been added back.

the average gap in child support payments is due to the large proportion of mothers without awards and to the low levels of current award as compared with potential awards under the Wisconsin guidelines.

As a final look at the potential effects of a perfectly enforced child support system, we examine child poverty rates under a perfect system. According to tabulations from the 1990 CPS–CSS, approximately 43 percent of all children eligible to receive child support lived in poverty in 1990. The total poverty gap, or the total amount by which incomes would have to be raised to meet the poverty line, was $25.2 billion. A perfectly enforced child support system using the Wisconsin guidelines, on the other hand, would result in only 31 percent of children living in poverty, with a corresponding poverty gap of $15.0 billion. Thus while the child support system has the potential to significantly improve the economic status of children, cutting poverty rates by up to 30 percent, a substantial fraction of children would remain poor.

CHANGES OVER TIME IN FATHERS' INCOMES

This section presents an analysis of the change in fathers' average income between 1978 and 1989. As mentioned in the introduction, trends in non-marital fertility and male income inequality over the past decade may have had an important impact on the mean and distribution of fathers' incomes. While the previous section established that never married fathers are able to pay in child support only half of what ever married men can pay, the percentage of women eligible for child support who are never married has increased from 20 percent to 30 percent since 1980. Additionally, the population of non-resident fathers has most certainly been affected by the increase in male earnings inequality and the declining economic status of less-educated males.

We examine the effect of both of these trends on fathers' incomes by presenting a decomposition of the change in income between 1978 and 1989. We use data from the 1979 CPS–Child Support Supplement to estimate fathers' incomes in 1978. Given that a decomposition involves allowing both the coefficients and the sample means to change over time, we modify our method of estimating fathers incomes from that presented in the body of the paper. We use samples of prime-aged males from both the 1979 and 1990 March CPS to estimate earnings equations by race for ever married and never married men. These equations are then used, along with the 1979 and 1990 Child Support Supplements and the appropriate adjustments to each mother's age and education, to predict fathers' incomes.[17] (The coefficients from the equations are available from the authors upon request.)

Table 7 reports the change in fathers' average income. Using the average characteristics of custodial families in each year and the corresponding coefficients from that year we estimate that the mean income of absent fathers has fallen by 18 percent, from $28,029 in 1978 to $22,998 in 1989.[18] In order to assess how trends in non-marital fertility have affected fathers' incomes, we recalculate average predicted income using the percentage of mothers who are never married from 1979, mean values for all other characteristics from 1990, and the coefficients from the 1990 equations. Had the proportion of mothers who are never married remained at its 1978 level, fathers' average income would have declined by only 11 percent by 1989, or to $25,089. Thus this trend had the effect of reducing fathers' incomes by 7.4 percent. A demographic change working to offset this

TABLE 7

A DECOMPOSITION OF CHANGES IN FATHERS' AVERAGE INCOME

	Predicted Income	% Change from 1979	Effect on Income Trend
Using 1979 means and coefficients	28,029	—	—
Using 1990 means and coefficients	22,998	−0.179	—
Using 1990 means and coefficients and			
% never married in 1979	25,089	−0.105	−0.074
Mean age and education in 1979	21,895	−0.219	0.040
Using 1990 means and			
1979 coefficients	28,651	0.022	−0.201
1979 education coefficients	22,430	−0.199	0.020

Source: CPS–CSS 1979 and 1990. All figures in 1990 dollars.

effect, however, was the increase in the mean age and education of all custodial mothers (see Appendix B for sample means). Table 7 indicates that the rise in mothers' age and education increased estimated fathers' income by 4 percent. In other words, if mothers' average age and education had remained at its 1979 level, mean income in 1989 would have been 22 percent lower than in 1978.

The third panel of Table 7 reports the effect on fathers' incomes of changes in the returns to individual characteristics. We estimate income by using sample

[17]Our samples of March CPS ever married men include never divorced men, who are known to earn more than ever divorced men. In order to account for this, we adjust our estimates downward by a factor determined by the ratio of the predictions from the SIPP sample to that from the March CPS sample (or 7.8 percent). While we include this adjustment, it is of no importance for the decomposition analysis.

[18]Our prediction of fathers' average income for 1990 was about one percent lower than that presented in the body of the paper and derived from the SIPP. This discrepancy is most likely caused by differences in samples and methodology. For consistency, we adjust upward our estimates for both 1979 and 1990 by one percent.

means in 1990 and coefficients from the 1979 equations, and thus allow only the coefficients to change over this decade. The table shows that changes in the coefficients had a substantial effect on fathers' incomes; average income would have been 2.2 percent higher, rather than 18 percent lower, had the 1979 coefficients prevailed in 1990. The next row indicates, however, that changes in the returns to education played virtually no part in this effect. Holding the returns to education fixed at their 1979 levels, estimated mean income in 1990 would have been very similar to that estimated using the 1990 coefficients. Further analyses (not reported) revealed that the effect of the coefficients on the change in fathers' income derived from two factors. First, a fall in the returns to age over the decade had an important negative effect on the trend in white fathers' incomes. Second, a decline in the economic status of black men in general, as measured by changes in the intercept term, had a large negative effect on black fathers' incomes.

The results indicate that while there has been a fairly substantial increase in the percentage of custodial mothers who are unmarried, this increase did not have a major impact on fathers' ability to pay child support. The fact that trends in non-marital fertility have had a relatively modest impact on incomes can be explained by examining the average characteristics of never married mothers in each year (see Appendix B). The percentage of never married mothers increased over time, but the never married women in 1990 are on average nearly three years older and somewhat more educated than their counterparts in 1979. Also, a greater fraction are non-black. These changes are consistent with a decline in the propensity to marry among the general population and suggest that future trends in non-marital fertility may have only a limited impact on absent fathers' incomes. Changes in overall economic conditions that have affected all men, on the other hand, have had a more important impact non-resident fathers' incomes.

Conclusion

In this paper we present new estimates of the incomes and ability to pay child support of non-resident fathers. Such estimates are important for predictions about the ability of a fully enforced child support system to improve the well-being of children. Previous estimates may be outdated, however, given recent trends in non-marital childbearing and income inequality. Our estimates are obtained using new data, which are not only more recent, but also superior, for the purpose at hand, to data used heretofore. The data also permit the use of less cumbersome and cobbled methods and allow for the examination of previously unverified assumptions.

Using an indirect method to estimate the average resources of non-resident fathers, we find that they are able to pay nearly five times more in child support than they currently pay. Furthermore, our estimates suggest that even the lowest income fathers can afford somewhat more than they are currently paying. Thus recent efforts at child support reform seem promising, as our estimates suggest that a perfectly enforced child support system can expect to collect much more in child support from the entire distribution of non-resident fathers. In addition, we find that much of the discrepancy between what fathers currently pay and what they would potentially owe is due to the absence of awards and the relatively low level of current child support awards. This finding highlights the importance of paternity establishment and child support guidelines as instruments of child support reform.

Finally, we examine changes over time in non-resident fathers' incomes. The evidence indicates that the rise in the proportion of women who are never married has had a negative, albeit modest, impact on fathers' average incomes. Changes in the overall economy, as they affect the earnings of all men, had more important effects on fathers' incomes.

Our results suggest that any effort to strengthen child support enforcement would undoubtedly improve the economic well-being of eligible children and their mothers. However, we also find that not all children would benefit equally from a perfect system. While fathers on average are able to pay more child support, the ability to pay differs substantially by the race, marital status, and poverty status of the custodial mother.

A final comment relates to the applicability of these methods to data from other countries, given that the United States is not alone in its limited data on non-resident fathers. While panel data on separating couples comparable to the SIPP are unlikely to be widely available, these methods could easily be implemented using data on married couples. In particular, the Luxembourg Income Study (LIS) contains household survey data for over 20 countries. The LIS includes information on the age, education level, earnings, ethnicity/nationality, and geographic location of all adults in the household. One could estimate income equations for married men using the wife's characteristics as regressors and then apply these coefficients to a sample of single mothers with children, using information on assortative mating to adjust these mothers' characteristics.[19] As our analysis indicated, however, the final estimates would need to be adjusted for any earnings differences that may exist between married and ever divorced men, which could also be determined with the LIS data. Thus as child support becomes more important in the array of policies for children, researchers may employ these methods to determine the potential impacts of child support enforcement.

Appendix A

Variable Means for the Divorced Sample

SIPP 1984–90

Age of wife	31.6
Education of wife	12.3
Urban	0.65
Northeast	0.12
West	0.19
South	0.36
Black	0.09
Hispanic	0.07
Husband's average monthly income (1990 dollars)	2,025
Number of months income observed	11.4

Note: Sample weights used, $N = 1{,}055$.

[19]Most surveys, unless they contain marital and fertility history information, will probably not allow for the identification of all eligible families.

APPENDIX B

SELECTED MEAN CHARACTERISTICS OF CUSTODIAL MOTHERS

1979 and 1990

	1979		1990	
	Ever Married	Never Married	Ever Married	Never Married
Age	35.1	25.4	36.0	28.0
Education	11.8	11.1	12.4	11.6
Black	0.17	0.63	0.15	0.56
Hispanic	0.07	0.09	0.09	0.14
Northeast	0.19	0.22	0.18	0.21
West	0.21	0.14	0.22	0.18
South	0.34	0.41	0.37	0.37
Never married	0.17		0.29	
N (unweighted)	2,710	553	2,999	1,133

Source: CPS–CSS 1979 and 1990.
Note: Sample weights used.

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