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An Economic Analysis of Co-Parenting Choices:
Single Parent, Visiting Father, Cohabitation, Marriage*

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Abstract

This paper sheds light on the determinants of choice between four co-parenting arrangements: father absence, father's non-residential visitations, cohabitation, and marriage. In our theoretical framework, we use an adaptation of Becker's Demand & Supply (D&S) model of marriage and a hierarchy of co-parenting arrangements--ranked in terms of degree of fathers' involvement in the lives of mother or child--as an observable price measure for women's work as mothers. We predict effects on co-parenting choice of factors such as welfare benefits, sex ratios, income, black versus white, or education, and black/white differences in these effects. We test our predictions with data from the Fragile Families and Child-Wellbeing Survey. Our findings include (1) the larger the grant amount in the state where the mother resides, the more it is likely that fathers will have some contact with their children, and the more it is likely that fathers will cohabit with the mothers; (2) fathers who have more children with other women are less likely to have contact with a given woman's children, but this discouraging effect of men's other children is smaller for blacks than for whites; and (3) employment in the last year reduces the likelihood that a white mother is married to her child's father, while increasing that likelihood among black mothers.

1. Introduction

"The most significant family, or social problem facing America is the physical absence of the father from the home." (Canfield 1996).

Almost 80% of Americans agreed with that statement in 1996. Fewer fathers at home often means fewer marriages, but no marriage does not necessarily imply father's absence. Parents may cohabit: cohabiting couples now account for about 40 percent of all unwed births and, by one estimate, about 26 percent of children born during the 1980's and 1990's will live with a cohabiting mother by age 14 (Bumpass and Lu 2000, Manning and Lichter 1996, and Graefe and Lichter 1999**)**. Fathers' involvement may also take the form of non-residential visitations by non-custodial fathers. Non-residential visitations by fathers who divorced or are separated from their children's mothers have been common since the divorce rate exploded. Non-residential visitations by unwed fathers have grown in importance more recently, especially among the poor and among African-Americans (see Mincy and Oliver 2000, Mincy et al. 2003, Neponshany 2003).

If co-parenting is defined as cooperation between parents raising a child, we thus observe that in the U.S. co-parenting in marriage has lost popularity relative to a number of arrangements that offer a continuum of father's involvement in a child's life, ranging from lone motherhood with no father's involvement to visiting fathers and co-parenting in cohabitation. Father's absence is a situation few people aspire to. Even among unmarried parents surveyed at birth by the Fragile Families and Child-Wellbeing Survey, an overwhelming majority of mothers expected fathers to be involved in the lives of their children (Reichman et al. 2001), and almost all unmarried fathers expected to have such involvement (Mincy and Dupree, 2001). However, by the time these children were 3 years old only 50 percent of unmarried, non-resident fathers had seen their children in the past month. This paper's principal goal is to shed more light on the determinants of men's involvement in co-parenting.

Since WWII, blacks have moved away from traditional co-parenting arrangements faster than whites. In 2000 the proportion of children raised by lone mothers was considerably higher among blacks (43.1 percent) than among whites (12.0 percent) (Dupree and Primus 2001). However, this racial gap in lone mothering does not necessarily indicate as big a gap in father's presence: visits by unwed, non-residential fathers as a co-parenting outcome are more common among blacks than among whites (Mott 1990, Waller 2002). In the U.S., among blacks the term "baby father" has become a popular way to call an unwed father who does not live with his child's mother.[1] This paper's additional goal is to help explain black/white differences in co-parenting arrangements.

One of the factors that apparently influences co-parenting choices is the level of welfare benefits a woman possibly qualifies for if she has children on her own. There is some evidence that higher welfare benefits discourage women from having children in marriage (see Moffitt 1992, 1998). More recent empirical studies have shown that welfare may affect co-parenting arrangements more generally, not necessarily via marriage. For example, Mincy and Dupree (2001) show that welfare benefits increase the odds that unmarried mothers plan to cohabit with the fathers of their children, but not the odds of marital plans, and Carlson, Garfinkel et al. (2004) show that welfare benefits increase the odds that unmarried mothers continue to cohabit with the fathers of their children. We investigate how welfare benefits influence choice between four possible arrangements: marriage, cohabitation, non-residential visitation, and total absence of the father.

The existing literature does not offer a theoretical foundation for analyzing effects of welfare benefits or other variables on such expanded choice set of co-parenting arrangements. We provide such a theoretical framework and derive testable predictions from it, including predictions regarding racial gaps in the effects of selected variables on co-parenting choices. We then test our predictions with data from the Fragile Families and Child-Wellbeing Survey. Our major findings include: (1) the larger the welfare benefits granted in the state where the mother resides, the more it is likely that fathers will have some contact with their children, and the more it is likely that fathers will cohabit with the mothers; (2) the higher the sex ratio in the city of residence, the more it is likely that women are married; (3) fathers with more children from other women are

[1] The term was originally introduced by social scientists researching black families in the Carribean, where blacks have experienced double-digit non-marital birth rates for many decades (Clarke 1957, Senior 1991,

less likely to have contact with the mother of their last child, but that this discouraging effect of men's other children is smaller among blacks than among whites; and (4) employment in the last year reduces the likelihood that a white mother is married to her child's father, while increasing that likelihood among black mothers.

2. Comparisons with the Existing Theoretical Literature

Traditional economic analyses of fertility and co-parenting—such as Becker (1960), Mincer (1963) and Willis (1973)—assume that a couple/household (typically married) has first been established, with co-parenting being one of its major stated goals. Then the household chooses to have children or not. Since the choice of having a child outside of marriage is one of the options we include in our analysis, these traditional economic analyses of co-parenting based on a unitary household model are of little use to us. In contrast, in our model, men and women express preferences for each possible co-parenting arrangement before they enter the dating and marriage markets to search for co-parenting partners.

Bargaining models of marriage such as Manser and Brown (1980), McElroy and Horney (1981), Lundberg and Pollak (1993), and collective models such as Chiappori's (1992), also diverge from the unitary household decision-making model. However, these models assume that a couple has first been formed and that spouses then individually express their preferences, including preferences for co-parenting. In contrast, in our model individuals consider forming a co-residential couple as one of their multiple routes towards becoming a parent.

Any model of co-parenting choices needs to analyze separately the choices of men and women who are considering each other as potential co-parents, while

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Brown et al. 1993).

recognizing that these choices are interdependent. This can be accomplished via a coordination mechanism established exogeneously, such as the one found in Becker's (1973) Demand & Supply (D&S) model of marriage: the price mechanism.[2] At sharing levels or prices set in a market equilibrium, individuals interested in working as co-parents (the supply side) will find satisfactory matches with others interested in obtaining the services of a co-parent and willing to compensate that co-parent for his or her work (the demand side). Prices help translate potential matches into actual matches.[3] We recognize American women's predominant responsibilities for raising children and place women on the supply side. Given that in our data set there are very few same sex couples, we place men on the demand side.[4]

We diverge from Becker's D&S model of marriage in that (1) we distinguish co-parenting from other productive functions often performed in marriage;[5] (2) our model recognizes that male/female cooperation in household production tasks such as parenting does not necessarily occur in marriage; and (3) we posit that a major element of the price

[2] Eugene Choo and Aloysius Siow (2004) call this model Becker's "transferable utility model of the marriage market". For a more detailed discussion of Becker's D&S model and how it evolved between 1973 and 1981, see Grossbard (2004).

[3] Willis' (1999) model of choice between co-parenting in marriage and lone motherhood does not rely much on prices as a coordination mechanism. Willis assumes that men search for the child-bearing capacity of women, while minimizing the transfer required to obtain a child of a given quality. The women in Willis (1999) choose between having a child alone and co-parenting a child in marriage. Willis incorporates a market analysis but prices or compensations don't play much of a role in guiding his individual decision-makers.

[4] Grossbard-Shechtman (1984) offers a more general analysis of markets for spouses' work in household production (not necessarily co-parenting workers) in the context of a general equilibrium model that includes (commercial) labor markets for men and women.

[5] This distinction was also made in Grossbard (1976). Considering married fertility only, Grossbard (1976) saw 'occupation: married mother' as including geneticial services, i.e. work related to reproduction and other parenting functions. Grossbard-Shechtman (1986) attempted to empirically separate geneticial services from other services that women often perform. Edlund (2002) analyzes markets for mothers as capital markets. We follow Becker in analyzing markets for mothers' work as labor markets.

of a spouse consists of non-monetary benefits such as commitment and protection against divorce.[6]

Our analysis shares some features with previous analyses of choice between two co-parenting arrangements, such as co-parenting in couple versus lone motherhood (see for instance, Heer and Grossbard-Shechtman (1981), Guttentag and Secord (1983), Akerlof, Yellen and Katz (1996), Willis (1999), Grossbard-Shechtman, Ekert-Jaffe and Lemennicier (2002), and Grossbard forthcoming), or co-parenting in marriage versus co-parenting in cohabitation (see Grossbard-Shechtman (1982), and Ekert-Jaffe and (1996)). Our model follows Grossbard (1976) and Grossbard (forthcoming) in that we use an occupational choice model along the lines of those found in labor economics, but in this case the occupation consists of mothers working for the benefit of fathers. Workers in this occupation are compensated in the form of father's time, monetary transfers, or commitment in the form of marriage. Men choose between remaining childless and paying a mother to have their child, under different possible arrangements varying in their costs to men.

Prices for spouses, fathers and mothers, are not readily available. This is one reason why previous D&S analyses of marriage have been limited in their applicability. Our model's central feature consists of its use of Mincy and Huang's (2003) index of father's involvement as a proxy for the unobservable 'mothers' wages' or compensations that women can expect in return for supplying co-parenting work.

Our theoretical framework consists of a conventional market analysis, in which we change one variable at a time, *ceteris paribus*. Some of the variables shift the supply

[6] In that respect we follow Grossbard-Shechtman (1982, 1993) and Grossbard-Shechtman and Neuman 1988).

of women willing to work as co-parents, some of the variables lead to shifts in men's demand for such work. Some variables lead to shifts in both these demands and supplies.

3. The theoretical framework

We view a co-parenting arrangement—be it marriage, cohabitation, or non-residential visitation--as a non-profit firm that has the care of shared children as its major objective. In contrast to commercial firms, the principal objectives of co-parentships are non-profit. We view the work that goes into co-parenting as an occupation. Usually one parent spends more time in this occupation than the other. We view the other parent who spends less time in parenting, and often pays more bills, as having a demand for the work of a co-parent. The incentives to work in co-parenting include monetary benefits and psychic benefits such as job satisfaction. Co-parenting is an occupation similar to teaching and almost everything that can be said about teachers can also be said about mothers (or fathers) acting as co-parents for the benefit of the other parent, except for the absence of monetary wages.

The Demand and Supply (D&S) analysis that guides us is similar to standard D&S analysis used in labor economics. In traditional labor market analyses workers have portable and firm-specific non-portable skills. We also assume that some co-parenting skills are portable and therefore affect an individual's value in the market for co-parenting. Other skills are couple-specific.

An index of father's involvement. Our major assumption is that observable co-parenting arrangements—unwed non-residential visitations; cohabitation; and marriage—correspond to points on a four-point non-pecuniary compensation scale that men 'pay' to the women who work as their children's co-parents. More precisely, we rank the four co-parenting arrangements in terms of the level of the following benefits to women working as co-parents: attention, money, and commitment directed at children, and attention, money, and commitment directed at the mothers. We expect women to find that (1) *non-residential visitation* offers more benefits than lone motherhood for it offers some degree of father's involvement with their children; (2) *cohabitation* offers higher benefits than a situation of no co-residence (lone motherhood or non-residential visitation), for cohabitation is likely to be positively associated with the time and money that men devote to both mother and child; and (3) *marriage* is preferable to cohabitation to the extent that it typically implies more commitment on the part of a man to the children and/or their mothers. However, the drawback of marriage is that it makes it more difficult for women to conceal the fathers' co-residence and income, and therefore marriage may reduce eligibility for welfare benefits (Moffitt et al. 1995, Primus and Beeson 2000, and Mincy and Dupree 2001).

We thus define an observable co-parenting arrangement variable, Y , that is allowed to take four discrete, ranked, observable values:

$Y = 0$: absent fatherhood

$Y =1$: non-residential visitation

$Y = 2$: non-marital cohabitation, and

$Y = 3$: marriage

When Y = 0, the father contributes zero non-monetary benefits. When Y>0, there is some father's involvement. The higher Y, the more the father is involved in the life of child and/or mother. We use this benefits scale acts as a price variable, which makes it easier to apply standard Demand and Supply (D&S) analysis.[7] This observable and discrete Y variable will be used as a proxy for an unobservable *yf** defined as the non-monetary benefits obtained by a woman *f* supplying co-parenting services. We posit that

(1) $Y = Y(yf^*)$, with the first derivative being positive.

This variable *yf** is similar to the share of the gain from marriage obtained by married women in Becker's (1973, 1981) D&S analysis of marriage, except that (1) Becker analyzes marriage markets and we analyze markets for co-parenting, and (2) this *yf** is an hourly non-monetary benefit that a woman may obtain for her co-parenting services.[8] For simplicity, *yf** is called the 'mothers' wage' in the remainder of this discussion.

D&S Analysis

The 'mothers' wage' is defined as:

(2) $yf^*_i = X_i\beta^*$, where X_i are variables that explain co-parenting arrangements. If markets for mothers' work are competitive, factors that increase men's demand for women's co-parenting services will be associated with an increase in the 'mothers' wage', and factors that increase women's supply of co-parenting services will be associated with a decrease in 'mothers' wage'. Next, we proceed with a comparative statics traditionally performed in markets for labor or goods.

Demand-shifting factors. If a factor's effect on men's aggregate demand is expected to be positive, that factor will be associated with a higher level of 'mothers' wage' and therefore we predict a more desirable co-parenting arrangement Y from the

[7] For a formal derivation of demands and supplies of mothering in the case of four co-parenting arrangements, see Grossbard-Shechtman and Mincy (2003).
[8] Note that Becker (1973), when presenting his D&S theory of marriage, uses terms from labor economics not found in the *Treatise on the Family* (Becker 1981), see Grossbard (2004).

perspective of a woman working at co-parenting. Men's demand is expected to be higher if women are more productive at co-parenting or due to income effects, substitution effects, or the effects of taste. The predictions following from this demand-side analysis are summarized in the first column of Table 1. Factors associated with productivity include maturity, parenting experience, characteristics of the family of origin, and education.

If women are extremely *young*, one does not expect them to be very productive at co-parenting and the demand for their services will be lower. As they age, one expects the demand to first increase and then decrease. Consequently, we predict a non-linear relationship between Y and mother's age: the first derivative of age is expected to be positive, and the second derivative negative.

Men are expected to have a lower demand for *women with children from previous unions*. This leads us to expect a lower equilibrium 'mothers' wage' for women who have children from previous unions, and a lower Y. However, some men may value women's past experience, especially if they like the way that women are raising their children from previous unions.

Women who come from *a two-parent family* are likely to be more productive at parenting and therefore men may have more of a demand for their work as mothers, which will raise their expected mothers' wage, yf^*, as well as Y.

If producing high quality children is a primary motivation for seeking women a co-parent, and women with more *education* are more productive at parenting, then men will have a higher demand for the work of educated co-parents. This will raise educated women's expected mothers' wage, yf^*, and lead to a positive association between women's education and Y. However, some men may not appreciate educated women, which would reduce the 'wage' of educated mothers. The net effect of education is unclear *a priori*. It is also possible that the effect of education is non-linear (see Grossbard-Shechtman 1993).

We don't have information on women's earnings, but we do know whether a woman was *employed* in the year prior to giving birth. Two reasons lead us to predict that employed women have a higher 'mothers' wage': they may have more desirable unobservable qualities that are appreciated by both potential fathers and employers (such as a higher I.Q.), and men may appreciate women's higher income. This leads us to expect a higher Y. However, employment could also indicate a lower productivity at being a mother, and that would be associated with a lower Y. There is no clear prediction here.

The higher the ratio of men to women (the s*ex ratio*) in a particular place at a particular time, the larger men's demand for women's work in co-parenting, the higher the equilibrium 'mothers' wage' and the higher we predict Y to be. Demographers typically define the sex ratio as the number of men divided by the number of women. Sex ratios vary by women's characteristics and will be computed separately for blacks, Hispanics, and whites.

Even though we don't have statistics on 'mothers' wage', there are at least two reasons to assume that the equilibrium 'mothers' wage' is lower for *black* women than for white women. First, sex ratios are much lower among blacks than among whites (see Wilson 1987), and the extent of low local sex ratios may not be captured in our city-wide measures.[9] Second, there may be racial discrimination against black women's mothers' work. Blacks also appear to be discriminated against by skin color in the labor market (see Goldsmith, Hamilton, and Darity 2004), so why would that not apply to markets for mothers' work? For these two reasons one expects black women to command lower equilibrium 'mothers' wages' and we predict a negative effect of 'black' on the observable Y, the co-parenting arrangement. To the extent that men discriminate against dark skin in markets for mothers' work, we don't expect sex ratios to capture the entire extent of black women's disadvantage in markets for mothers' work.

If a woman had *previous children with her child's father*, this indicates that more specific human capital has been invested. His demand for her services will be larger than other men's demand. Note that in contrast to the other factors we have discussed, this expected shift is not a market-level shift in demand, as the human capital in this case is match-specific and not general (portable). When specific skills are only appreciated by two people and don't have a market value, there is a dual monopoly bargaining situation reminiscent of that between firms and workers who acquired firm-specific skills. Nevertheless, we expect her 'mothers' wage' and Y to be higher.

The higher *men's income*, the higher men's demand for women's work as mothers is likely to be. Income and substitution effects are expected to go in the same direction. A rightward-shift in demand will lead to a higher equilibrium 'mothers' wage', yf^*, and therefore higher values of observable Y. We don't have information on income, but we have two variables that can be used as proxies for income: education and weeks worked.

[9] See Heer and Grossbard-Shechtman (1981) and Guttentag and Secord (1983) for earlier discussions of black/white differences in sex ratio and their effect on marriage and fertility.

Men with children from previous unions are expected to have a lower demand for mothers' work: they are expected to be less interested in taking responsibility for another child.

Men's demand for women's co-parenting services may be higher where *welfare benefits* are higher, as the mother may bring more cash with her. This leads to a higher 'mothers' wage', yf*. However, this effect is limited to the lower categories of index Y and does not apply to the odds of marriage versus lower values of Y, given that marriage is likely to disqualify women from welfare.

Supply-shifting factors.

Any factor that shifts the supply of women's co-parenting to the right will lead to a lower 'mothers' wage' and a lower Y. The predictions following from a supply-side analysis are summarized in the second column of Table 1. The third column predicts the total effect on the co-parenting arrangement index Y.

Women's willingness to have a child is expected to rise the more their biological clocks are ticking, thereby causing a shift to the right in the supply of mothering. This effect of *age* is expected to be non-linear: we don't expect substantial supply effects at younger ages. Adding demand and supply considerations leads to an unambiguous prediction that the effect of age will be positive at lower ages and negative at higher ages, i.e. the sign of age, squared, is expected to be negative.

Wo*men with children from previous unions* may be less motivated to supply their mothering services and their supply may lye to the left of that of women who have not yet experienced motherhood. However, women with children who need a father for their children from previous unions are expected to be more willing to supply their mothers' work. They may be especially eager to enter a co-parenting relationship with men indicating an interest in acting as a father. This lowers mothers' asking wage, when they are considered for a serious relationship, implying that women are willing to pay compensating differentials to get their child a father. The expected sign is ambiguous from the supply side, and the sign of the total predicted effect of this variable is also ambiguous.

Women who grew up in a two-parent family may be more willing to work as mothers, which will lower the yf* they are asking. They may also be more productive at work other than mothers' work, which will shift their supply to the left and imply a higher yf*. The total predicted effect of family of origin on yf* and Y from the supply side is ambiguous. Overall, given that from the demand side this factor is expected to push predicted mothers' wage up, we expect the net effect on yf* and Y to be ambiguous.

More *educated* women may be earning higher permanent wages in the labor force and therefore have a supply of mother's work that lies to the left of that by less educated women. *Ceteris paribus*, this raises the equilibrium 'mothers' wage'. However, it is also possible that higher education is associated with higher income and that women want to spend part of their income on working as mothers; in that case the supply shifts to the right and 'mothers' wage', yf*, falls. The net effect on the supply side is thus ambiguous, and therefore so is the net effect that combines demand and supply. We know that historically in the U.S. women's education was negatively associated with marriage and fertility (see Goldin 1990), but in the recent years there seems to be a positive effect of education on U.S. women's likelihood of being married (see Rose 2005).

Employed women tend to earn a higher wage than women not in the labor force. From the supply side an income effect and a substitution effect may discourage women from supplying their mothers' work, leading to a shift in supply to the left. This implies a higher yf* and a higher Y. The net effect of women's employment, taking account of supply and demand, is ambiguous.

If *black* women don't discriminate against or in favor of white men, we don't expect 'black' to matter on the supply side. The demand factors may dominate here, leading to the prediction of a negative effect of black on Y.

If a woman has *previous children with the father of the child under study*, more specific human capital has been invested in this co-parenting relationship. Her supply may include the non-monetary benefits of co-parenting the new child with this father, and therefore her supply lies to the right of that of a woman who does not have previous children with a particular child's father. This is not a market-level shift in supply, as the human capital in this case is specific to this match and not general (portable). Given the opposite effect on the demand side, the net effect of this variable on Y is ambiguous.

Men's education is expected to increase women's supply to the extent that women prefer to co-parent with more educated men. Any male characteristic that is valued by women who supply mothers' work is likely to be associated with compensating differentials: men who are considered more attractive by women can obtain mothering services at a lower 'mothers' wage', yf*. (This is the equivalent of compensating differentials at the workplace) The net impact of men's education (taking account of both D and S) on equilibrium 'mothers' wage', yf*, is ambiguous.

Women's supply of mothers' work to men *with children from previous unions* will lie to the left of that to men without children, as this reduces the chances of father's future involvement. This implies that women will ask for a higher Y from men with children from previous unions. The effects of this factor via demand and supply may cancel each other out: the net predicted effect is ambiguous.

The higher the *welfare benefits* available to a woman who has a child alone, the more the supply of women's work lies to the left: a woman's choice of having a child alone, without any father's involvement, is now more attractive. This implies a higher equilibrium value of y*f, 'mothers' wage', and a higher Y. This reinforces the expected effect of welfare benefits on the demand side. Therefore, welfare benefits are likely to be associated with a higher likelihood of co-parenting in comparison with lone mothering, and if there is co-parenting, it is more likely to occur with co-residence.[10] However, we don't expect higher welfare benefits to increase the odds of marriage, given that marriage generally makes it more difficult for mothers to qualify for welfare. [11]

Black/white differentials in the effect of explanatory variables X on co-parenting arrangement Y. Our analysis of black/white differences in expected co-parenting arrangement Y focuses on the demand side, for there is no compelling reason to assume that women's supply differs by ethnicity. We expect most effects on expected co-parenting arrangement Y to be less likely to hold for black women than for white women. The effect of any factor X on Y operates via its effect on unobservable y*f, 'mothers' wage'. We expect the effect of X variables on the unobservable mothers' wage of black

[10] This had also been predicted by Mincy and Dupree (2001).

[11] Although, unwed couples may cohabit while concealing the incomes of fathers from welfare authorities, this is much more difficult if the couple marries (Primus and Beeson, 2001, Moffitt, Revelle, and Winkler 1995, Mincy and Dupree, 2001).

women willing to work as mothers to be smaller than these variables' effect on the mothers' wage of white women, given that black women's mothers' wages are estimated to be more concentrated at low levels (see Grossbard-Shechtman 1995). Therefore, all demand side effects are expected to be weaker in the case of black women, and we predict that all interaction terms between a variable X_j and 'black' will have the opposite sign of the sign of the direct effect of that variable on Y. The only exception is the interaction term 'Black and total fertility with father'. How many children this particular couple had in the past does not entail an effect on market demand but rather on the negotiated value of Y as the result of intra-household bargaining. It is only when characteristics are portable that interaction terms with 'black' are expected to differ from zero and to take the opposite sign of the direct effect. These predictions are summarized in the lower part of Table 1.

The following interaction terms with 'black' are predicted to take the opposite value of the direct effect of that variable on Y: *Black X welfare benefits* [The interaction term is expected to be negative, as long as we are not comparing marriage and cohabitation]; *Black X woman has children from previous unions; Black X 'two-parent family at 15'*; '*Black X men's years of education*'; '*Black X weeks worked'* ; 'Black X men's children from previous unions"; '*black X education'* , '*black X employment'*. In the case of women's employment, a negative association between yf* and women's employment is more likely to be observed for white women than for black women. For whites, women's employment is less likely to be associated with a beneficial co-parenting arrangement, a high Y. White women are more likely to be paid to be wives and mothers of a man's child. The women not in the labor force are likely to be women who were offered attractive co-parenting work packages.

In contrast, for blacks, y*f is rarely high and therefore if employment is observed, this is more likely to indicate that the woman earns a high wage in the labor force than that she earns a low yf* in co-parenting. It is therefore more likely that among Black women we will observe a positive association between employment and 'mothers' wage', yf*.[12] If the direct effect of employment is negative, we predict that the interaction term '*Black X female employment* will be positive. It is possible that if education reflects wages in the labor force, we will also see more of a positive association between women's *education* and Y for blacks than for whites.

The interaction '*black X sex ratio'* is also expected to have the opposite sign to the sign of sex ratio. By introducing this variable into our regression, we also get some sense of the degree to which observed black/white differential effects are the result of black/white differences in sex ratio. Racial differentials after controlling for sex ratio and interactions between 'black' and sex ratio may indicate that in markets for mothers' work, men discriminate against black women willing to work in mothers' work.

We now investigate whether evidence from the Fragile Families and Child Well-being Survey supports the predictions of our model.

4. Data and Methodology

Data. The Fragile Families and Child Well-being Survey (FFCWS) is a national study designed to provide longitudinal data on the conditions and capabilities of new unmarried parents and the consequences for child well-being. The survey includes

[12] Similar arguments were made in Grossbard-Shechtman (1995) regarding the effects of income, welfare benefits, etc. on labor supply and welfare dependency.

information about fathers, the nature of the relationships between unmarried mothers and fathers, and the extent to which fathers are involved with children. The study follows a birth cohort of about 3,700 children born to unmarried parents in 20 U.S. cities, selected based on variations in their labor market conditions, generosity of welfare benefits and strictness of child support enforcement. The full sample is representative of all non-marital births to parents residing in cities with populations over 200,000. To permit comparisons across critical domains, a total of 1,100 married parents were interviewed in all 20 cities, in the full baseline survey. New mothers were interviewed in hospitals or birthing clinics within 48 hours after giving birth, and fathers were interviewed either in the hospital, birthing clinic, or elsewhere as soon as possible following the birth of their child.

Response rates for both mothers and fathers in the baseline FFCWS are encouraging: fully 85 percent of eligible mothers and 76 percent of eligible fathers participated in the study. However, response rates were higher for fathers who maintained some positive relationship with the mothers. Additionally, the interviewer asked the mother to provide some basic demographic information for use in situations in which the father was not interviewed. This allows larger samples to be used in the analysis, with control variables to account for missing data on some fathers. Our empirical work includes only mothers who reported their race as white, black, or Hispanic. Hispanics include both blacks and whites, but are mostly white.

Methodology. We cannot observe $yf^*_{i,}$. Instead, we can observe four discrete, ranked co-parenting outcomes Y using mothers' answers to questions about non-residential visitation and union status at 12 months. We posit that the Y function varies with yf* according to three cutting points θ as follows:

$Y_i = 0$ if $yf^*_i < \theta_1$

$Y_i = 1$ if $\theta_1 < yf^*_i < \theta_2$

$Y_i = 2$ if $\theta_2 < yf^*_i < \theta_3$

$Y_i = 3$ if $yf^*_i > \theta_3$,

Each cutting point separates two contiguous categories:

- cutting point 1 refers to the odds that the mother is in *any co-parentship* as opposed to absent fatherhood,
- cutting point 2 refers to the odds that the mother is in *a residential co-parentship* with her child's father (in a cohabiting or marital union) as opposed to the other two alternatives, and
- cutting point 3 refers to the odds that the mother is in a *marital co-parentship* with the child's father rather than in any of the three alternative co-parenting arrangements

Our model estimates the impact of various factors X on the probability that a mother chooses a particular co-parenting outcome Y_i versus all outcomes of lower rank. We will be estimating the effects of independent variables X on a co-parenting outcome falling above or below a given cutting point θ_i (i=1,3). Given that Equation 2 above defined 'mothers' wage' as $yf^*_i = X_i\beta^*$, where $X_i = [X_{1i}, X_{2i}, X_{3i,}]$ is a vector of right-hand side variables, it follows that the odds that yf^*_i falls between certain cutting points is also a function of these variables X, and therefore the odds of finding particular values of Y are a function of the same explanatory variables that play a role according to the D&S analysis presented in the previous section.

For a multinomial dependent variable in which the categories are ordered, the best statistical procedure is ordered logit regression (Amemiya 1981; Agresti 1984; and Peterson and Harrell 1990). This procedure estimates the independent variables' effects on a mother's outcome falling above or below a given cutting point. The more commonly used version of ordered logit assumes that the impact of each variable is the same for all cutting points, something known as the proportional odds assumption. While this method is useful in many situations, it appears unlikely that such an assumption would hold true in this situation. For example, a factor that may encourage a cohabiting mother to marry the father of her child would not necessarily have the same impact as a factor that encourages a mother with an absent father to allow non-residential visitation. Moreover, statistical tests did not support the proportional odds assumption. To estimate the hypothesized effects, we use a less restrictive method, known as generalized ordered logit, which produces three sets of coefficients that correspond to each cutting point.

Our sample. Table 2 shows descriptive statistics by race. Marriage and non-residential visitation vary substantially by race. While 31 percent of the mothers in the pooled sample were married to and living with the fathers of their children, the proportion of white mothers with this outcome (44 percent) was almost 3 times the proportion of black mothers (17 percent).[13] Similarly, although in the pooled sample 25 percent of mothers were unmarried and in non-residential visitation co-parentships with the fathers of their children, this masks substantial racial diversity. The proportion of white mothers in this arrangement (14 percent) was less than half the proportion of black mothers (36 percent). Very similar proportions of white mothers (31 percent) and black mothers (28 percent) were living with their children and their children's fathers. The proportion of white mothers in absent-father co-parentships (11 percent) was lower than the proportion of black mothers (18 percent) with this outcome. The right-hand side variables used in this analysis fall into one of three categories, [X_{1i}, X_{2i}, X_{3i}]: demographic characteristics of women, demographic characteristics of men, and policy-related characteristics specific to the state in which the mother resides. These are mostly the characteristics covered in the D&S analysis above.

The variables are listed in Table 2. Note that *White* (as reported by mother) is also non-Hispanic; a*ge* is reported age or is calculated age based on date-of-birth); y*ears of education* were transformed into a continuous variable [14], *mother worked prior to childbirth* is a dichotomous variable indicating whether or not the mother reported working during the year prior to the birth[15]; *total fertility with father* is the reported number of children in common with the father of the focal child; *multiple partner fertility* is a dichotomous variable indicating whether or not the mother (or father) had at least one child with a partner other than the co-parent of the focal child); father's *age* is reported age or calculated age based on date-of-birth [16]; *father's employment* (If the father was not interviewed, then the variable is based on the mother's report of his employment

[13] Note that we defined black Hispanics as Hispanic and not as black.

[14] The following transformations were made: gen myedu=.;replace myedu=0 if m_educ==1 (no formal school); replace myedu=6 if m_educ==2 (8th grade or less); replace myedu=10 if m_educ==3 (some high school); replace myedu=12 if m_educ==4 (high school diploma); replace myedu=12 if m_educ==5 (GED); replace myedu=14 if m_educ==6 (some college); replace myedu=14 if m_educ==7 (technical or trade school); replace myedu=16 if m_educ==8 (BA degree); replace myedu=18 if m_educ==9 (graduate school), where 'gen' means 'generate', m stands for mother, and yeduc stands for years of education. Same calculation was done for fathers.

[15] We use mothers' employment status during the year before the birth to avoid simultaneous equation bias involving employment and mothering arrangement in the same year.

[16] If the father was not interviewed, then the variable is based on the mother's report of the father's age.

status);[17] the *Grant Amount* is the level of welfare benefits in hundreds of dollars based on the state's TANF (Temporary Assistance for Needy Families) grant amount for a representative family of three as of 1997; and *Sex Ratio*, calculated from the five-percent Public Use Microdata Sample from the 2000 Census of Population PUMS for each of the twenty cities in our sample. Sex ratios were defined as the ratio of men to women in a given age and minority status group. We grouped ages into six categories, each including 5 years. We assume that men in the youngest male-age group, who are between 17 and 21 years, are paired with women in the youngest female-age group, who are between 15 and 19 years old. The other male-age groups are: 22-26 years old, 27-31 years old, 32-36 years old, 37-41 years old and 42-46 years old. The corresponding female age groups are: 20-24 years old, 25-29 years old, 30-34 years old, 35-39 years old; and 40-44 years old.[18] We computed separate sex ratios for non-Hispanic whites, non-Hispanic blacks, and Hispanics. Therefore, the sex ratio assigned to a given mother depends upon her age, minority status, and the metropolitan area in which she resides.[19]

The demographic characteristics and capabilities of FFCWS respondents, including the policy environments in which they live, have been discussed elsewhere (Garfinkel et al. 1999). However, few studies have disaggregated these data by race. Although whites are slightly older than blacks, the two sub-samples are quite similar with respect to educational attainment. More than two fifths of the black sub-sample (of mothers and fathers) has had a child with someone other than the parent of their newborn,

[17] We use fathers' employment status during the week before the birth to avoid simultaneous equation bias involving employment and mothering arrangement in the same year.
[18] We used a two-year age difference in view of the fact that at first marriage men are on average two years older than women (in that respect, we follow Grossbard-Shechtman 1993).
[19] With the exception of Indianapolis and Nashville, we calculate separate sex ratios for using *place* as the geographic type. For Indianapolis and Nashville we use the metropolitan statistical area.

while one quarter of the white sub-sample has done so. Notably, over 50 percent of mothers in the white sub-sample were in two-parent families at age 15, while only 30 percent of the mothers in the black sub-sample were in such families.

Differences in the proportion of white and black mothers who worked in the year before the birth are not statistically significant, but the proportion of white fathers who worked in the week before the birth is higher than the proportion of black fathers who did so. Nevertheless, 74 percent of the black FFCWS fathers were employed. Over half of the mothers in the white sub-sample describe their religious affiliation as Catholic, while only 5 percent of the black sub-sample does so, most likely the result of our inclusion of Hispanics, who account for 54 percent of the white sub-sample. Religious affiliation is expected to affect fertility preferences and preferences for co-parenting arrangements, so it is important we use it as a control variable.

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5. Results

Table 3 presents the results of our estimated model. The first panel shows results for white women. In the bottom panel we interact each coefficient with 'black'. We estimate effects of included variables on the odds of any contact with the father versus no contact with father ($Y \geq 1$ vs. $Y = 0$), the odds of co-residence with the father versus non-residential visitation or no contact ($Y \geq 2$ vs. $Y < 2$), and the odds of marriage to the father versus the three lower co-parenting arrangements ($Y = 3$ vs. $Y < 3$). We first examine the results in order to test whether the predictions we derived from our theoretical framework, summarized in Table 1, receive support from the data. Most of our predictions regarding direct effects were ambiguous, but we had a few clear predictions.

1/ Do our predictions hold?

We predicted a positive effect of age on Y, a negative effect of age, squared, a positive effect of ‘two-parent family of origin’, a negative effect of ‘black’, and a positive effect of welfare benefits on Y. It can be seen from Table 3 that a one-year increase in a non-black *mother’s age* increases the odds in favor of marital co-parentship (Y = 3) vs. the other alternatives (Y < 3) by 16 percent (1.1566 – 1.0). Black women’s odds of marriage are even more sensitive to age.[20] The effect of age for black women is obtained by multiplying the odds ratio for non-Hispanic white women with the interaction term of age and black, which gives 1.1566 x 1.5967, from which we deduct 1. This implies that each year increases a black woman’s odds of marriage by 85 percent. Unlike the results for white and Hispanic women, the odds ratio for the quadratic age term is statistically significant, indicating that the positive effect of age on marriage prospects declines as black women get older. The effects of age on other forms of co-parenting are not statistically significant.

As predicted, we find that compared to women who were in single-parent families at age 15, Y is higher for women who grew up in intact families. This effect is apparent at both the level of co-residence and the level of marriage: the odds of a residential co-parentship vs. non-residential visitation or absent fatherhood ($Y \geq 2$) are 45 percent higher for women who were in two-parent families at age 15. The odds of marital co-parenting vs. the other alternatives (Y = 3) are 24 percent higher, although this effect is only marginally significant.

We had predicted that for non-marital co-parentships (Y<3) higher welfare benefits in a state will be associated with more desirable forms of co-parenting from

[20] We estimated regression slopes for non-Hispanic whites and all Hispanics (white or black). We include interaction terms with ‘black’ but not with ‘Hispanic’. Preliminary tests indicated that the regression slopes

women's point of view, i.e. a higher Y. Indeed, we find that each additional $100 of welfare benefits increases the odds of any co-parentship vs. absent fatherhood ($Y \geq 1$) by 21 percent and the odds of residential co-parentships vs. non-residential visitation or absent fatherhood ($Y \geq 2$) by 10 percent. This is consistent with recent findings about the perverse effects of welfare benefits on family structure, when non-marital families are disaggregated (Mincy and Dupree 2001 and Carlson, Garfinkel et al. 2004). Note that the interaction of black with welfare is statistically insignificant, indicating that the effects of welfare on co-parenting arrangements for blacks do not differ from those for Hispanic and white women.

As predicted, we find a positive association between sex ratio and Y. More specifically, the sex ratio has a statistically significant positive effect on the odds of marriage versus other co-parentships ($Y = 3$). We find that an increase of 10 percent, say from 1 to 1.1 --an extra 10 men per hundred women--, increases these odds by 14.1 percent. However, the effects of sex ratio on the odds of lower co-parentships are not statistically significant.

Based on the coefficients of a dummy for 'black', we find that black women have odds of marital co-parentship versus other co-parentships ($Y = 3$) that are nearly 100 percent lower than those of (non-Hispanic) white women, but being black does not have a significant effect on the odds of any co-parentship vs. absent fatherhood ($Y \geq 1$) and the odds of residential co-parentships vs. non-residential visitation or absent fatherhood ($Y \geq 2$). Being Hispanic does not make a large difference: Hispanic women have odds of marital versus other co-parentships that are 27 percent lower than that of non-Hispanic white women, and there is no effect of Hispanic on the other forms of co-parentship.

are quite similar for Hispanics and non-Hispanic whites.

In summary, all our unambiguously predicted effects were confirmed for some of the observed cutting points in Y.

2/ Tests of predictions regarding black/non-black differentials

Our model investigates black/white differences in co-parentship in more detail by including interactions between black and every other variable in the model, i.e. we estimate a full interaction model.[21] This allows us to test for another one of our predictions: demand-side predictions for the sample as a whole will be accompanied by an interaction term between that variable and 'black' that takes the opposite sign, except for the variable 'total fertility with father.'

We find that, as predicted, 'black' interacted with some of the variables leads to an effect on Y that has a sign opposite of the effect of that variable on white women's Y. More specifically, we find that:

- Relative to a black woman who was in a single parent family at age 15, the odds of such a residential co-parentship ($Y \geq 2$) are only 5 percent higher for the black woman from an intact family (this variable has a much larger effect for white women). In contrast, being from an intact family raised a white (or Hispanic) woman's odds of residential co-parentship by 45 percent.
- For white and Hispanic women, the effects of education on $Y \geq 2$ and $Y = 3$ are negative, and the effects of education-squared on these two outcomes are positive, but that is not the case for any contact, $Y = 1$. The interactions of 'black' with women's education have positive effects on all three outcomes and the interactions with education-squared have negative effects.

[21] For simplicity, these slope differences are defined as 'black/white differentials'. In fact, they are differentials between black and non-Hispanic whites plus Hispanics (black or white).

- Pre-birth employment reduces the odds of marital versus other co-parentships ($Y = 3$) by 35 percent for white and Hispanic women. In sharp contrast, pre-birth employment of black women *increases* the odds of marital versus other co-parentships by 24 percent! We also find that pre-birth employment is not significantly associated with the odds of residential co-parenting versus non-residential visitation or absent fatherhood ($Y \geq 2$) for white and Hispanic women, whereas pre-birth employment increases the corresponding odds for black women by 37 percent.
- While we find direct effects of 'father's children from previous unions' to be negative for all three co-parentships, we find that the interaction terms with 'black' are positive for all three co-parentships. The net effects of men's fertility with previous mates on all three odds of Y are less negative for blacks than for whites. The racial differential is especially large in the case of odds of marital versus other co-parentships ($Y = 3$) effects of 39 percent for blacks and 63 percent for whites and Hispanics.
- Age of father has a positive effect on odds of all three levels of co-parentship for the sample as a whole, and father's age, squared, has a negative effect that is significant for $Y = 2$ and $Y= 3$. The interaction terms of these variables with black takes on the opposite sign, although it is only significant for the odds of any contact.

Most of the other interaction terms with 'black' go in the predicted direction but are not statistically significant. The interaction of black with 'total fertility with father' is not significant, as predicted. Overall, we consider these results supportive of our prediction that effects on odds of co-parentship will be weaker for black women than for white women.

3/ Other findings

Next, we discuss findings regarding the effects of variables for which we did not derive clear predictions in our theoretical framework.

We find that white *women's education* reduces the odds of residential co-parentships versus non-marital visitation and absent fatherhood ($Y \geq 2$) and the odds of marriage versus other co-parentships ($Y = 3$). We also find that the coefficients of *education-squared* are positive, implying non-linear effects of education. This indicates that relative to women with medium levels of education, women with the lowest and the highest education are most likely to be in couple.

We find that having *children from other men* reduces women's odds of residential co-parentship ($Y \geq 2$) by 20 percent and the odds of marital co-parentship ($Y = 3$) by 51 percent. Multiple partner fertility also reduces the odds of any co-parentship vs. absent fatherhood ($Y \geq 1$) by 20 percent, but this effect is only marginally significant. These results were the same for women from all three ethnicities. This finding helps explain why early unmarried childbirths often lead to subsequent unwed childbearing with multiple (sequential) partners. Thus, the demise of the shot-gun wedding may have led to growing complexity in American families (Furstenberg and King 1999, Mincy 2001). Moreover, by promoting *healthy* marriage among young unwed parents, policy makers may be attempting to restore and improve upon a lost tradition in the U.S. (Dion et al. 2003).

In contrast, *having previous children with the same man* has strong positive effects on the quality of co-parentship that a woman can expect. For each previous child that a woman has with the father of the focal child, the odds of any co-parentship versus absent fatherhood ($Y \geq 1$) rise by 25 percent; of a residential co-parentship vs. non-marital visitation and absent fatherhood ($Y \geq 2$) by 27 percent; and of marital versus other co-parentships ($Y = 3$) by 41 percent.

Men's multiple partner fertility is one of the major factors reducing Y, especially for white women. We find that if their child's fathers have at least one previous child from another woman, white women have 67 percent lower odds of any co-parentship vs. absent fatherhood ($Y \geq 1$), 56 percent lower odds of residential co-parentship vs. non-residential visitation or absent fatherhood ($Y \geq 2$), and 63 percent lower odds of marital versus other co-parentships ($Y = 3$). These large effects indicate that this variable's effect on the demand side dominates its effect on the supply side: women don't seem to be able to get more commitment from men who have an unattractive trait: ties to previous mothers and children. This finding underscores the warnings about premature fatherhood that some have tried to signal in the literature (Lerman 1993 and Knock 1998).

We also find that fathers' additional *years of schooling* reduce the odds of a married co-parentship until a certain level of education. After that, the effect of father's education on odds of a married co-parentship is positive (the square term has a positive coefficient). *Fathers' employment* has a positive effect on the odds of co-parentship: relative to men who did not work during the week prior to the birth, men who worked have 76 percent higher odds of any co-parentship vs. absent fatherhood ($Y \geq 1$); 87 percent higher odds of residential co-parentships vs. non-residential visitation or absent fatherhood ($Y \geq 2$); and 79 percent higher odds of marital co-parentship vs. the other alternatives ($Y = 3$).

We controlled for additional variables not discussed in our theoretical framework. This includes *religious affiliation*. We found that women affiliated with a Protestant, Catholic, or other faith have odds in favor of marital versus other co-parentships that are almost 88, 59 (marginal significance), and 235 percent higher, respectively, relative to women with no religious affiliation. The coefficients for the residential co-parentships vs. non-marital visitation and absent fatherhood are marginally significant and positive for Protestant, Catholic, and for other religions. However, religious affiliation does not have a significant impact on any contact between mother and father ($Y \geq 1$).

Racial homogamy has large effects on all three co-parenting outcomes. When a man is of the same race, women have 49 percent higher odds of any co-parentship ($Y \geq 1$); 39 percent higher odds of residential co-parentship ($Y \geq 2$); and 79 percent higher odds of marital co-parentship ($Y = 3$).

For each additional year of a *man's age* the odds of any co-parentship vs. absent fatherhood rise by 19 percent; the odds of a residential co-parentship vs. a non-residential visitation or absent fatherhood rise by 20 percent; and the odds of a marital co-parentship vs. the other alternatives rise by 17 percent. The effect of the father's age is quadratic, but like the effects of age for women, the effect of men's age remains positive within the range of our data. If men are one year older, the odds of any co-parentship vs. absent fatherhood rise by only 2 percent for black women.

6. Conclusions

This paper provides a simple theoretical framework that helps explain how parents chose among four co-parenting arrangements: marriage, cohabitation, visitation, or total absence of the father. This framework, based on Becker's Demand and Supply analysis of marriage, uses an observable price variable: an index based on the four co-parenting arrangements observed in our data.

Our theoretical framework yields some unambiguous predictions about the effects of factors such as age, race, welfare benefits, and sex ratios on the value of that co-parenting index. Tests of these predictions, using the Fragile Families and Child Well-being data, generally support our theoretical model. Consistent with our predictions we find that women from intact families, women who are neither too young nor too old, and white women are likely to co-parent in more desirable arrangements. We also find that more favorable co-parenting arrangements (from the mothers' vantage point) are likely to be found where welfare benefit levels are higher (these increase the odds of more favorable co-parenting arrangements up to cohabitation) and where sex ratios are higher.

Our model also leads us to predict that most factors (except the number of previous children a mother has with a child's father) will have larger effects on the co-parenting arrangements of white women than they will have on the arrangements of black women. We find larger effects (in absolute value) for white and Hispanic women in the case of the following factors: intact family of origin, female education, and number of father's children from previous unions, and father's age. Our model also helps explain why pre-birth employment reduces the odds of marriage for non-black women, while it increases the probability of marriage, relative to other arrangements, for black women.

In addition, the empirical model yields noteworthy results regarding the effects of education. We find that relative to women with medium levels of education, women with the lowest and the highest education are most likely to be in couple; that women with children from other men have lower odds of cohabitation and marriage; that total fertility with father has positive effects on all forms of co-parentship; that men with previous children have substantially lower odds of entering any form of co-parentship; and that father's employment has a positive effect on the odds of all forms of co-parentship.

Policy Relevance. Since family structure has become a direct target of policy efforts intended to reduce poverty and improve child well-being, our framework and empirical results carry important implications for public policy. We find some differences in the way blacks respond to personal or environmental change, relative to whites and Hispanics. This underscores the need for government policies that target each ethnic group when it comes to the encouragement of stable co-parenting relationships such as the healthy marriage initiative.

Furthermore, traditionally, the primary tool for policy intervention in this area has been welfare benefits. Previous studies have concluded that more generous welfare benefits have small, negative, but statistically significant effects on marriage. These have mostly been documented for white women only. We don't find these effects for the odds of marriage, but we find that more generous welfare benefits encourage non-residential visitation and cohabitation. More generous welfare benefits thus encourage father's involvement.

Recent debates have raised the possibility that policy efforts aimed at increasing employment among men may divert funds needed to increase self-sufficiency among single mothers. Our results suggest that for white and Hispanic women, mother's employment reduces the odds of co-parentship in the form of marriage, while having no impact on the odds of other forms of co-parentship. However, for black women we find that their own employment increases the odds of marriage and cohabitation. Helping women find employment may thus be an effective way of encouraging father's presence among blacks, but it may not be as effective for whites and Hispanics. As for father's employment, we find that pre-birth employment tends to increase all forms of father involvement: non-residential visitation, non-marital cohabitation, and marriage, and this holds for all three ethnic groups.

Our research also holds implications for faith-based initiatives intended at promoting marriage.[22] We find that religious women are significantly more likely to raise their children in co-residential co-parentships.

Practical implications. The idea that men may use non-residential visitations, cohabitation, or commitment in marriage to compensate women for having their babies also carries practical implications for individuals. Reframing in these terms may help men and women negotiate ways to better coordinate parenting. Consider the following case of a father who disappears from his child's life. For whatever reason, the mother has a low market mothers' wage that does not induce men to prefer marriage as a co-parenting arrangement. If the mother insists on marriage this amounts to requiring too high a 'price' for her mothering. One often overlooked potential strategy for getting the father more involved in her child's life is for the woman to lower her expectations and try to get the child's father to visit the child without demonstrating more commitment to her.

[22] Separate results for subsamples of black and non-black women indicate that religious affiliation had a positive effect on odds of co-parentship for non-blacks but not for blacks (see Mincy and Huang 2003).

This amounts to a lowering of her ‘asking mothers’ wage,’ in line with the concept of ‘asking wage’ in labor economics. Or consider men wanting to be in contact with their children but the mothers don’t cooperate. It may be helpful to apply our way of thinking and help men understand that if they want to play a role in their children’s lives they need to pay their children’s mothers a higher ‘mothers’ wage’ in the form of more attention or monetary transfers.

This paper may also serve as a reminder of the usefulness of Becker’s Demand and Supply (D&S) theory of marriage. When it comes to the choice between four forms of co-parenting that we address in this paper--no co-parent, father’s visitation, cohabitation, or marriage--D&S theories such as the one presented here are more applicable than bargaining theories assuming the pre-existence of a couple. We consider various co-parentship arrangements as a price variable that is perceived by all agents interesting in having children, before and after a relationship is established. It would be difficult for bargaining theories to accommodate this use of the price mechanism as a means of coordination decision-making by unrelated individual men and women. Bargaining between two agents does not require much coordination, and therefore bargaining theories typically don’t expect prices to accomplish much coordination or signaling.

Suggestions for further research. Future empirical research should attempt to specify the factors that affect co-parenting arrangements more carefully, allowing for competing interpretations of the role of critical factors. For example, men’s pre-birth employment, which we find to exert strong effects on all co-parenting arrangements, might be a proxy for other variables, such as the quality of couple relationships, which are correlated with employment. In addition, child support enforcement is a policy variable that was excluded from our analysis, yet it is likely to have important effect on co-parenting arrangements. We also hope that future research will shed more light on the effects of sex ratio on the odds of various types of co-parentship among blacks and other ethnic groups. Our model followed traditional gender roles. In the future, we hope to extend our analyses to markets for fathers where women are on the demand side and men on the supply side.

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Table 1: Predicted Effects on Co-Parenting Arrangement

	Effect via Demand	Effect via Supply	Net Predicted Effect
Mother's Characteristics			
Age	positive		positive
Age, squared	negative	negative	negative
Children from Previous Unions	negative	?	?
Two-Parent Family at Age 15	positive	positive	positive
Years of Education	?	?	?
Worked in the Year before Birth	?	positive	?
Black	negative		negative
Total Fertility with Father	positive	negative	?
Father's Characteristics			
Years of Education	positive	negative	?
Children from Previous Unions	negative	positive	?
Worked in the Week before Birth	positive	?	?
Environment			
Sex Ratio	positive		positive
Grant Amount from Welfare	pos. for Y<3	pos. for Y<3	pos. for Y<3
Interaction Terms			
Black (B) x Mother's Characteristics			
B x Age	negative		
B x Age, Squared	positive		
B x Children from Previous Unions	positive		
B x Two-Parent Family at Age 15	negative		
B x Education	opposite of direct effect		
B x Worked in the Year before Birth	opposite of direct effect		
B x Total Fertility with Father	?		
Black (B) x Father's Characteristics			
B x Children from Previous Unions	positive		
B x Education	negative		
B x Worked in the Year before Birth	negative		
Black (B) x State Environment			
B x Sex Ratio	negative		
B x Grant Amount from Welfare	neg. for Y<3		

Table 2: Descriptive Statistics for Main Variables

	All Sample	White	Black	F or X^2 test
Mother's Race				
Non-Hispanic White	23%	46%	---	
Non-Hispanic Black	49%	---	100%	
Hispanic	28%	54%	---	
Type of Co-Parentship				
No Contact	15%	11%	18%	434.9 ***
Non-residential Contact (spent 1+ hour last month)	25%	14%	36%	
Non-marital Cohabitation	29%	31%	28%	
Marriage	31%	44%	17%	
Mother's Characteristics				
Age	25.2 (6.0)	25.8 (6.2)	24.5 (5.8)	44.3 ***
Years of Education	12.1 (2.5)	12.2 (2.9)	12.1 (1.9)	1.5
Total Fertility with Father	1.6 (0.9)	1.6 (0.9)	1.5 (0.9)	1.3
Children from Previous Fathers	35%	25%	46%	176.7 ***
Two-Parent Family at Age 15	43%	56%	29%	264.1 ***
Religious Affiation				
No Religion	11%	9%	12%	9.9 **
Protestant	50%	30%	71%	635.4 ***
Catholic	28%	51%	5%	993.3 ***
Other Religion [1]	10%	9%	11%	3.4
Father and Mother Same Race	88%	81%	94%	140.2 ***
Worked in the year before the birth	69%	71%	68%	1.9
Father's Characterstics				
Age	27.8 (7.3)	28.3 (6.9)	27.3 (7.6)	15.3 ***
Years of Education	12.0 (2.6)	12.1 (3.1)	12.0 (1.9)	1.8
Children from Previous Mothers	35%	25%	45%	164.7 ***
Worked in the week before the birth	81%	88%	74%	123.3 ***
State Environment				
Grant Amount [$100]	3.32 (1.30)	3.29 (1.54)	3.34 (1.00)	1.0
Sex Ratio	.975(.205)	1.075(.140)	0.834(.091)	
N	3822	1938	1884	

Note:+ p<.10; * p <.05; ** p <.01; *** p <.001

[1] Include Jewish, Muslim, and Jehovah's witness, and other religion.

Table 3: The Generalized Logit Model of Co-Parentship Outcomes

	Odds Ratio		
	Y $\geq$ 1 vs. Y = 0	Y $\geq$ 2 vs. Y <2	Y = 3 vs. Y<3
	Any	Residential	Married
Mother's Characteristics			
Age	1.061789	1.11401	1.156681 *
Age-Squared	0.9993153	0.998194	0.998466
Years of Education	0.8974941	0.732136 **	0.735142 ***
Years of Education - Squared	1.007153	1.016479 ***	1.019832 ***
Fertility with Previous Fathers	0.8029916 +	0.803207 *	0.486089 ***
Multiple Partner Fertility	1.247892 *	1.272743 ***	1.406465 ***
Two-Parent Family at Age 15	1.277419	1.450204 ***	1.242397 +
Religious Affiliation			
Protestant	1.08591	1.309342 +	1.877391 **
Catholic	0.9892554	1.381114 +	1.586608 +
Other Religion [1]	0.8580355	1.542827 +	3.354265 ***
Black	2.02145	0.747985	0.000272 ***
Hispanic	0.7880893	0.843562	0.732154 ***
Father and Mother are of the same race	1.48968 **	1.392238 **	1.787564 ***
Worked in the year before the birth	0.8883899	0.945537	0.653212 ***
Father's Characteristics			
Age	1.193016 **	1.201373 ***	1.173713 +
Age-Squared	0.9975852 *	0.997507 **	0.997881
Years of Education	1.034705	0.968543	0.855655 +
Years of Education - Squared	1.000893	1.003717	1.012308 **
Multiple Partner Fertility	0.3274723 ***	0.441363 ***	0.366842 ***
Worked in the week before the birth	1.759671 **	1.866302 ***	1.790712 ***
State Environment			
Grant Amount [$100]	1.212386 ***	1.101772 ***	0.967876
Sex Ratio	1.253842	1.283128	2.410559 **

Interactions with Black

	Any	**Residential**	**Married**
Mother's Characteristics			
Age	1.159113	0.988569	1.596744 ***
Age-Squared	0.9981938	1.000191	0.991964 ***
Years of Education	1.623608 *	1.559347 **	1.73662 +

Years of Education - Squared	0.9806186	*	0.982566	*	0.977346	+
Total Fertility with Father	0.9116304		1.011249		0.936026	
Multiple Partner Fertility	1.10035		1.191043		1.16364	
Two-Parent Family at Age 15	1.097958		0.725327	*	0.947302	
Religious Affiliation						
Protestant	0.5970237		0.827083		0.649942	
Catholic	1.510736		1.317293		0.886952	
Other Religion [1]	1.041999		0.717238		0.584655	
Father and Mother are of the same race	0.9769012		0.914501		0.685423	
Worked in the year before the birth	1.209375		1.453264	*	1.892103	***
Father's Characteristics						
Age	0.8518009	*	0.963723		1.001351	
Age-Squared	1.002217	*	1.000392		0.999873	
Years of Education	0.6416322		0.70838		0.790448	
Years of Education - Squared	1.018218		1.014065		1.006268	
Multiple Partner Fertility	1.72258	***	1.462237	*	1.652895	***
Worked in the week before the birth	0.9675624		0.674995		1.004459	
State Environment						
Grant Amount [$100]	0.8560664		0.93066		1.01023	
Sex Ratio	1.064478		0.890063		0.646331	
N			3822			
Log Likelihood			-4268.0			
Pseudo R Square			0.17			

[1] Includes Jewish, Muslim, and Jehovah's witness, and other religion.
+ p < .10, * p < .05, ** p <.01, *** p < .001
Robust standard errors are used to take into account the fact that women live in same city.