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Family Values and the Value of Families:
Theory and Evidence of Marriage as an Institution

Working Paper Number E98-03

July 15, 1999

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Abstract

Children take considerable time and effort to "produce" and their production is overseen by their families. As a consequence, family *type* may have a significant effect on child outcomes. One would expect that the relative disadvantages of having unmarried parents would have diminished over the past few decades. The expansion of social welfare programs and greater social acceptance of alternate lifestyles should have reduced the burdens faced by non-married couple families; improved control over fertility should have reduced the likelihood of children being born into disadvantaged environments.

I present evidence that the opposite is true: the average difference between children whose parents are married and those whose parents are not has *increased*. This increase is consistent with an asymmetric information model of marriage (Murphy 1999). The apparent increase in the penalty faced by children growing up outside of married couple families reflects a composition effect: the pool of surviving marriages has changed. Expanded AFDC, lower divorce costs, and smaller penalties associated with out-of-wedlock births allow women to more easily avoid or escape bad marriages; as a consequence, surviving marriages are "better" on average.

*Comments and suggestions are welcome. I thank Sandy Black, Glenn Loury, Dilip Mookherjee, Alwyn Young, and seminar participants at Boston University, The College of William and Mary, The University of Massachusetts at Amherst, and Virginia Polytechnic Institute and State University for helpful comments. Discussion with and advice from Eli Berman and Kevin Lang has been particularly valuable in helping me think about empirical issues.

I Introduction

The debate over "family values" continues to be sharp and contentious. Many with "tolerant" views of abortion, divorce, and out-of-wedlock child bearing believe that tolerance and compassion are the least we can offer to those women and children caught in wrenching circumstances. Conversely, many with "strict" views claim that tolerance, while well intentioned, is ultimately counter-productive: it encourages socially costly behavior, exacerbating the plight of those it was intended to help. Many bystanders would take refuge in the belief that if the welfare of children is our primary concern, the question is largely an empirical one: to what extent do children benefit or suffer from differing family circumstances?

In this paper, I present evidence concerning a simple model of non-contractible paternal investment in children which captures both of these concerns: the difficulties faced by women and children "trapped" in difficult family circumstances, and the general equilibrium effects of greater tolerance on the average child. I show that this model has a counter-intuitive, but empirically verifiable implication for the estimated relationship between parents' marital status and child outcomes. When parents face a more tolerant social climate (i.e. readily available contraception, greater acceptability of divorce, and little stigma attached to out-of-wedlock child bearing), the average difference (e.g. in dropout likelihood) between children of married and unmarried parents should be *higher* than when parents face a strict climate. That is, the "divorce penalty" faced by children of unmarried parents should be larger today than, say, 40 years ago. While initially surprising, this effect has a simple explanation: as bad marriages become easier to leave or avoid, the pool of intact marriages becomes better.

The model focuses on the incentive effects of marital status: incentives that arise from having a particular status and consequently, the incentives to choose one status over another. It accounts well for several stylized facts about marriage and children and provides aggregate predictions consistent with U.S. experience of the past several decades. It also provides a convenient framework for highlighting the often conflicting effects of changes in marital status contingent payoffs such as AFDC, cost of divorce, and social stigma for out-of-wedlock births.

A Why model marriage*?*

Three stylized facts about children and marital status are striking:

- good outcomes for children are associated with intact marriages, yet even intact marriages do not guarantee good outcomes
- the "use" of intact marriages has steadily declined recently
- individuals have traditionally discriminated between friends and neighbors on the basis of their marital status

The differences between children growing up in different family types seem to persist even after accounting for parental resources (McLanahan and Sandefur 1994). Children of married parents are more likely to graduate from high school and college, more likely to successfully enter the labor force, and more likely to avoid teen child bearing, even after controlling for various background characteristics. However, the prevalence of traditional family arrangements has declined over the past few decades, a decline attributed by many to increasing economic independence for women.

One explanation for the recent decline of marriage is the improved economic position of women. But this labor market story is not an entirely satisfying explanation of family behavior, in part because income is not a good answer to the more basic question of: "Why marry?" Income sharing, household public goods, and gains to specialization concern optimal *household* formation rather than *marriage.* The gains implied by these motives are, arguably, available to unmarried households as well as married ones, yet traditionally couples have chosen to marry. To rescue the income hypothesis, one might claim, as Becker (1991) does, that couples seek gains from household formation and that marriage offers a *contract* for household and child-rearing services. For the vast majority of couples (excepting some temporally and geographically distant cultures cited by Becker), marriage is *not* an explicit, enforceable contract. It involves no detailed specification of performance nor of penalties that might be faced for breach. Some marriages are good and some are bad, yet couples know a good one only after the marriage has begun. And, whether the marriage is "good" or not can seldom be verified by a third party. Finally, marriage is seldom a private agreement between individual agents. Instead, it involves a public declaration; an important element is *social*.

The paper is organized as follows: Section II provides an overview of the marriage model (see (Murphy 1999) for a fuller discussion); Section III discusses changes in U.S. marital behavior over the past few decades in the context of model decision rules. Section IV discusses the estimated decision rule shifts relative to changes in likelihood of pregnancy and in reductions in the costs (social or legal) of divorce and out-of-wedlock child bearing (see (Akerlof, Yellen, and Katz 1996), (Anderson 1990), (Stern 1993)). Section V considers the effects of these changes and shows that the model has an unexpected implication: as fewer couples marry (and remain married), the children of unmarried parents should exhibit a larger *apparent* "penalty" relative to the children of married parents. Using samples from the US Census (Ruggles and Sobek 1995), I show that the apparent dropout risk associated with having unmarried parents almost doubled between 1960 and 1980. Section VI shows that instrumental variables estimates which take into account changes in marital status contingent payoffs are stable across time, unlike the naive estimates.

II Model

The model focuses on the social distinction between formal matches (marriages) and informal ones and is based on the difficulty of ensuring non-contractible investment by fathers. The model is a simple two-period game. Men and women are matched in couples and have two concerns: the quality of their current match and the man's decision to invest costly effort in his children. A match, if accepted, offers each individual a match specific payoff m (m is drawn from a continuous pdf $f(m)$ with associated cdf $F(m)$). If a match is accepted, the woman may become pregnant and have to care for their child (at cost to her of $\theta_W > 0$). Women prefer fathers to invest effort in their child. Some men find it costly to care for children and although they might promise to invest in a child, ex-post they may renege. The only recourse a woman has concerning a man who does not invest is to end the match. Individuals whose matches end during period one return to the pool of available mates for period two. Figure 1 shows the sequence of moves and payoffs *for one period* of this game.

If the couple stays together after observing the man's effort choice (after the woman has a child), the mother receives an additional payoff based on her partner's effort choice (in addition to $m-\theta_W$). She receives $w_H > 0$ from her child's increased welfare if her partner chose high effort and $w_L < 0$ if he chose low effort. She receives no additional payoff from her child's welfare if she and her partner do not stay together.

Consider two possible information regimes. In one, an available man's period one match status is unobservable in period two (i.e. a prior breakup is unobservable to prospective partners). In the second, his period one match status (either *single* or *divorced*) is observable. *Marriage* is modeled by allowing a woman to choose, at the beginning of a match between the two regimes. That is, women choose between informal (breakup unobservable) and formal (breakup observable matches). Period two generates the continuation payoffs that men receive, but I am mostly interested in period one strategies: both to characterize behavior and to compare decisions across information regimes. I concentrate on decision rules conditional on the woman becoming pregnant and focus, not on completely general rules, but on ones for plausible and interesting parameter values.

The equilibrium strategies of interest for women (when pregnant in period one) are characterized by an upper (divorce) bound, $\overline{m}$, and a lower (marriage) bound, $\underline{m}$. When match quality (m) is above the divorce threshold $\overline{m}$, a woman never leaves her partner, even if he chooses low effort, because her expected value

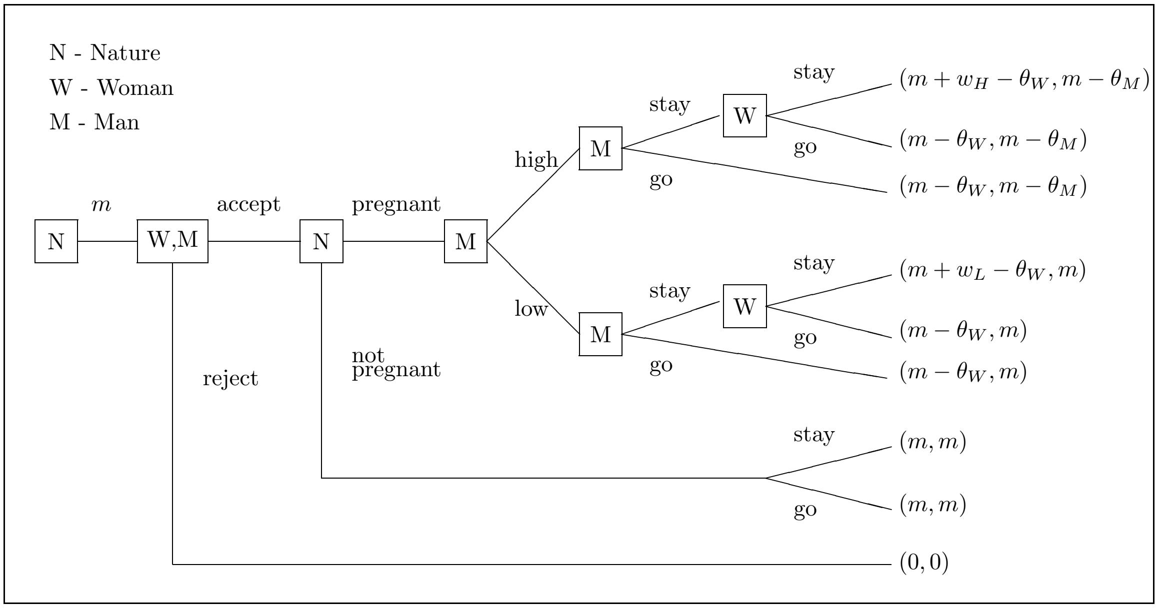


Figure 1: Decisions and *one* period payoffs (woman, man)

of a new match is lower than the value of her current, imperfect, match. When m is below the marriage threshold $\underline{m}$, she never stays, even if her partner chooses high effort, because the expected value of a new match is better than even a good father with the current match quality. For $m \in [\underline{m}, \overline{m})$, she stays only if her partner chooses high effort. For men, the optimal strategy (in period one, when a partner is pregnant) is described by an effort cutoff (E): a man will only choose high effort (and stay) when match quality is E or larger (although a high cost man will stay and choose *low* effort when match quality is above the divorce threshold $\overline{m}$, since his partner will never leave; see below).

There are two types of men, each with a different cost of investing in children. Proportion $0 < \mu' < 1$ of men are low cost types who enjoy children (i.e. they are "loving" fathers) and have investment cost $\theta_M^L < 0$; proportion $1 - \mu'$ are high cost types who find children costly and have cost $\theta_M^H > 0$. The cutoff for high cost type men, E^H (the match quality needed to induce high cost me to invest in their children), is the main interest. Assume for convenience that $\underline{m} < E^H < \overline{m}$.

When women can propose marriage (men can still reject the match), the model features, in equilibrium, marriage, divorce, and illegitimacy. Women reject matches (remain celibate) when m is less than a rejection level R (assume $R < \underline{m}$). They choose to make matches informal (non-marital) when $R \leq m < \underline{m}$ (the woman's expected value is positive, but these matches break up with probability 1; by assumption women choose informal relationships if a breakup is certain) and they make matches formal (marry) when $m \geq \underline{m}$. An institution of marriage, relative to a regime in which first period status is never observable, shifts the high cost type's effort threshold, E^H down and the woman's divorce threshold, $\overline{m}$, up. This increases, at both margins, the number of matches for which high-cost type men choose high effort. The marriage regime also shifts $\underline{m}$ up; at the lower margin, fewer matches involve fathers choosing high effort *and* mothers choosing to stay.

Figure 2: *Decision cutoffs and match outcomes* provides an overview of equilibrium decisions and outcomes. The horizontal axis shows values of m, the match quality. The top line shows, qualitatively, decision cutoffs when unattached men's prior match status is unobservable in period two. The second line shows the effects of introducing marriage: the high cost type man's effort threshold falls ($E^H \downarrow$); the woman's ability

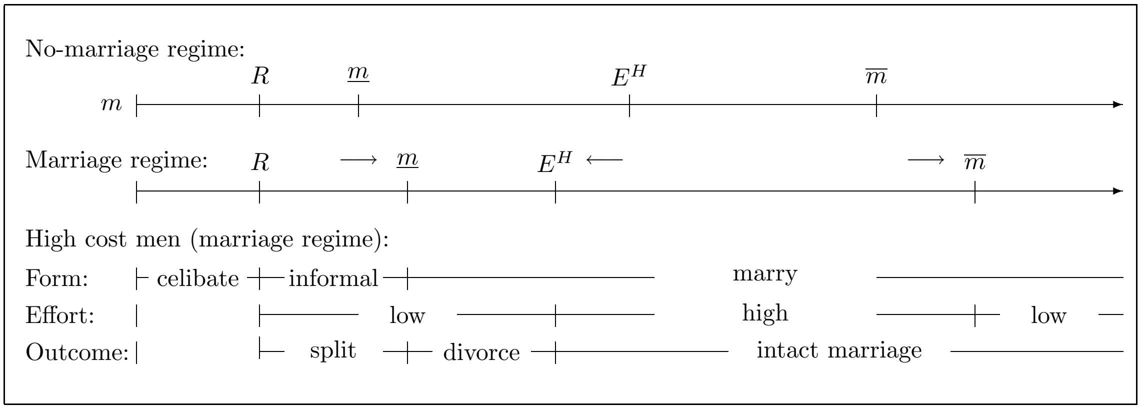


Figure 2: Decision cutoffs and match outcomes

to threaten divorce improves ($\overline{m} \uparrow$) and her marriage threshold also increases ($\underline{m} \uparrow$). The final three lines show (for a high cost man when his partner becomes pregnant) the chosen form, his effort choice, and the final match status by match quality when marriage is available. A "tag", divorce, for men whose marriages fail creates exit costs for men and discourages ex-post abandonment. The costs arise because the tag allows women to discriminate in period two against men who are less likely to invest effort in their children. Other (negative) effects of being identified as divorced (e.g. employers preferring to not hire divorced men, changes in the allocation of property rights over future income) would also contribute to the exit cost, so the likelihood of being a "high cost type" is not the only avenue through which this effect might arise.

III Aggregate U.S. behavior

Consider the United States in the context of the model: populations of size 1 of men, women, and couples. Each couple receives a draw m from $f(m)$. If one knew $f(\cdot)$ and the values of various parameters (e.g. $\theta_W, \theta_M^H, \theta_M^L, p, \ldots$), one could calculate the decision rules $R, \underline{m}, E^H$, and $\overline{m}$. Unfortunately, I know neither $f(m)$ nor the parameter values. But, by assuming a distribution for $f(m)$ and making note of the observed distribution of marital status choices in the U.S. population, I can estimate $R, \underline{m}$, and E^H (though not $\overline{m}$, the divorce threshold). I assume a distribution for m (i.e. $m \sim N(3,1)$) and count the number of couples in different marital status categories. With the assumed distribution and the observed frequencies, I can calculate the implied values of the cutoffs.

The population proportions in observed marital states are:

- single (celibate) - rejected matches: $F(R)$
- informal matches - all accepted matches with match quality below the marriage threshold: $F(\underline{m})-F(R)$
- divorce - marriages with high cost type men and match quality below E^H: $(1-\mu')(F(E^H)-F(\underline{m}))$
- married - marriages with low cost type men and match quality below E^H plus all marriages with match quality above $\overline{m}$: $\mu'(F(E^H)-F(\underline{m}))+1-F(E^H)$

The proportions married and divorced are taken from the appropriate issues of *Current Population Reports* (Census 1982)[2]. The proportion in informal matches is based on the birth[3] (U.S. Dept. of Health and

[2]The proportion divorced is underestimated, and the proportion married overstated, since some of those who divorce have remarried by the time marital status is observed.

[3]Unless otherwise noted, all calculations and data are for women aged 15-44, those in prime child-bearing years.

Human Services 1995) and abortion rates for unmarried women[4]. The sum of the birth and abortion rates is divided by an estimate of the pregnancy rate p and then multiplied by the proportion of available women (neither divorced nor married[5]). So, if i is the proportion of 15-44 year old women in informal matches, I estimate it as follows:

$$i = F(\underline{m}) - F(R) = \frac{\text{unmarried birth rate} + \text{unmarried abortion rate}}{p} \times \text{available women} \tag{1}$$

As a rough estimate, p is the birth rate[6] of 25-29 year old *married* women[7]; this group is almost certainly sexually active and likely to be homogeneous in terms of fertility - unlike younger or older groups. Note that lower values of p will *increase* the estimated size of the informal group; using 25-29 year old married women probably overstates p and so i is likely to be understated. This understatement will be less pronounced when sexually active women have less access to contraception and abortion (i.e. through perhaps 1970). Figure 3: *Live births per married woman aged 25-29*, shows the annual pregnancy rates for 1960-1993.

The proportion single (not sexually active) is the residual:

$$s = 1 - [\text{married} + \text{divorced}] - [\text{informal}] \tag{2}$$

The estimates are shown in Table 1: *Marital and fertility behavior, women 15-44: 1960-1990*, which includes speculative estimates of $F(E^H)$ and of the cutoff values themselves. To estimate $F(E^H)$, I need to make an assumption about μ' (the likelihood that a man is a low cost type). I choose $\mu' = 0.30$ because the resulting values of $F(E^H)$ seem plausible for all years and because, in this range, the values of $F(E^H)$ are not particularly sensitive to the particular value of μ' chosen. The cutoffs at the bottom of Table 1 are generated by assuming that $m \sim N(3, 1)$. For 1980 I can check the estimate of the sexually active population. My estimate for 1980 is $1 - F(R) = 0.849$; the estimate from the Centers for Disease Control (U.S. Dept. of Health and Human Services 1995), is between 0.806 and 0.845[8], which is reasonably consistent.

Several characteristics of these estimates bear noting. First, the match quality threshold for entering into a sexual relationship (R) dropped substantially: from 2.28 in 1960 to 1.90 in 1990. At the same time, the threshold for marriage ($\underline{m}$) increased: from 2.36 to 2.65. While the rejection cutoff (R), the marriage threshold ($\underline{m}$), and the high-cost man's effort threshold (E^H) were fairly close in 1960 ($E^H_{1960} - R_{1960} = 0.19$), by 1990 they had diverged significantly ($E^H_{1990} - R_{1990} = 1.10$). In 1960, almost all accepted matches were formal ones (the rejection and marriage levels were close) and the match quality required to assure that a

[4] The two primary sources for abortion data in the United States are the Centers for Disease Control and the Alan Guttmacher Institute; neither has a complete authoritative dataset. The abortion rate is calculated from figures reported in (Centers for Disease Control 1996). Total abortions are calculated from the Alan Guttmacher Institute abortion rate for 15-44 year olds multiplied by total women 15-44. The total abortions for 15-44 year olds is multiplied by the Centers for Disease Control estimate of the proportion of abortions had by unmarried women and then divided by the 15-44 unmarried population. For 1960, I assume the abortion rate is zero; for 1970, I *assume* that the rate is 0.030. Akerlof, Yellen, and Katz (1996) present an estimate of 0.035 for 1970-1974; *Roe v. Wade* was decided in January, 1973 and several states, including California and New York, had already liberalized their abortion laws.

[5] Since the available woman group is neither divorced nor married, then the pregnancy and abortion rates should be defined similarly. The rates used instead include divorced women in the unmarried group. The divorced group is small relative to the neither divorced nor married group, so this will not change the estimates much. If divorced women are more likely to be sexually active than never married women, then the sexually active population is over-estimated.

[6] The *fertility rate* is the number of births per 1,000 women aged 15-44; the *birth rate* is ordinarily births per 1,000 women in some other (i.e. not the 15-44 reference group) population. Here, the birth rate for married women 25-29 is expressed as births per woman.

[7] The exception is 1990, for which I use the 1980 birth rate. The birth rates for married 25-29 year olds increased from the late 1970s to the early 1990s. Using the higher 1990 rate (16.2% vs. 14.5% in 1980) as a risk of pregnancy seems implausible.

[8] Depending on whether one classifies those reporting no intercourse within the last three months as sexually active or not; see Table IV-3, p. 123, (U.S. Dept. of Health and Human Services 1995).

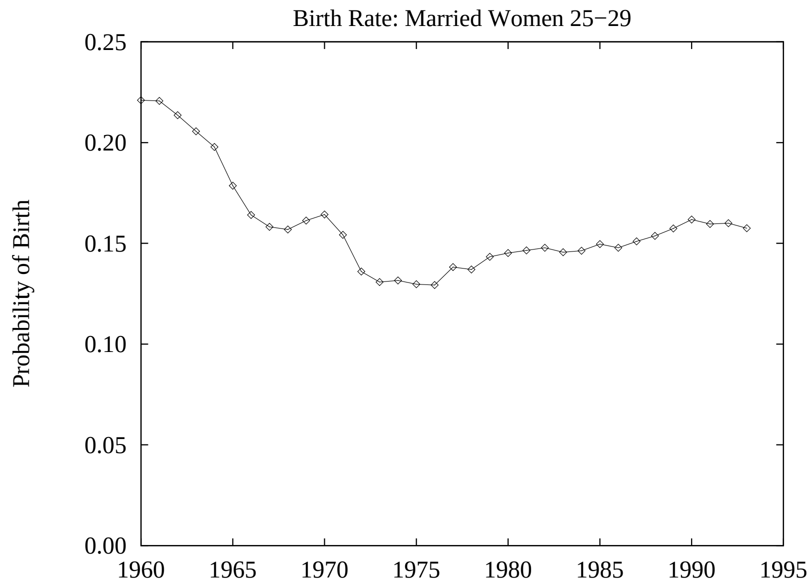


Figure 3: Live births per married woman aged 25-29

Measure		*1960*	*1970*	*1980*	*1990*
Unmarried birth rate		.022	.026	.029	.044
Unmarried abortion rate		*.000*	*.030*	.055	.048
Probability of pregnancy (p)		.221	.164	.145	*.145*
Neither divorced nor married		.259	.351	.359	.364
Divorced		.026	.032	.069	.086
Married		.715	.617	.572	.550
Celibate	$F(R)$	.234	.231	.152	.135
Informal	$F(\underline{m}) - F(R)$	.025	.121	.207	.230
Marry and divorce	$(1-\mu')(F(E^H) - F(\underline{m}))$	.029	.032	.069	.086
Remain married	$\mu'(F(E^H) - F(\underline{m})) + 1 - F(E^H)$	.715	.617	.572	.550
Cutoffs ($f(m) \sim N(3,1)$):					
Reject	(R)	2.28	2.26	1.97	1.90
Marriage	$(\underline{m})$	2.36	2.62	2.64	2.65
High effort	(E^H)	2.47	2.74	2.99	3.00

Table 1: Marital behavior and fertility, women 15-44: 1960-1990

match survived was fairly low (the marriage threshold and the high cost type's effort threshold were close). Informal matches have become more acceptable to women at both margins (R lower and $\underline{m}$ higher). Men are less willing to invest high effort within a marriage (E^H is higher); presumably the penalties they face from divorce have fallen. Figure 4: *Decision cutoff rules: 1960-1990* shows the estimated decision cutoffs for each year.

In 1960, almost a quarter of women of child-bearing age were celibate and another 71% were in intact marriages. Less than 6% were divorced or in informal relationships. By 1990, less than 15% were celibate and only 55% were married. Close to a third were divorced or in informal relationships. These changes are large, which suggests that the composition of the pool of married couples may have changed significantly.

IV Causes?

A natural question concerning these changes in observed behavior is "Why?" Answering this question is not the primary objective here, but I suggest several plausible candidates. The change in women's economic prospects is a likely contributory factor, but I have abstracted away from income (for reasons discussed earlier), so it is ignored here[9]. Two other likely candidates are changes in the probability of pregnancy (p) and changes in (exogenous) divorce and illegitimacy penalties (see (Murphy 1999)). Below, I briefly consider what effects changes in p and changes in exogenous penalties should have on marital behavior. The purpose of the discussion is not to identify the causes of changes in American marital behavior of the past few decades but to consider how these factors would affect decision rules and observed behavior. The implied changes in the decision rules then suggest a counter-intuitive prediction about changes in the association between marriage and the welfare of children.

A Statics: probability of pregnancy and contraceptives

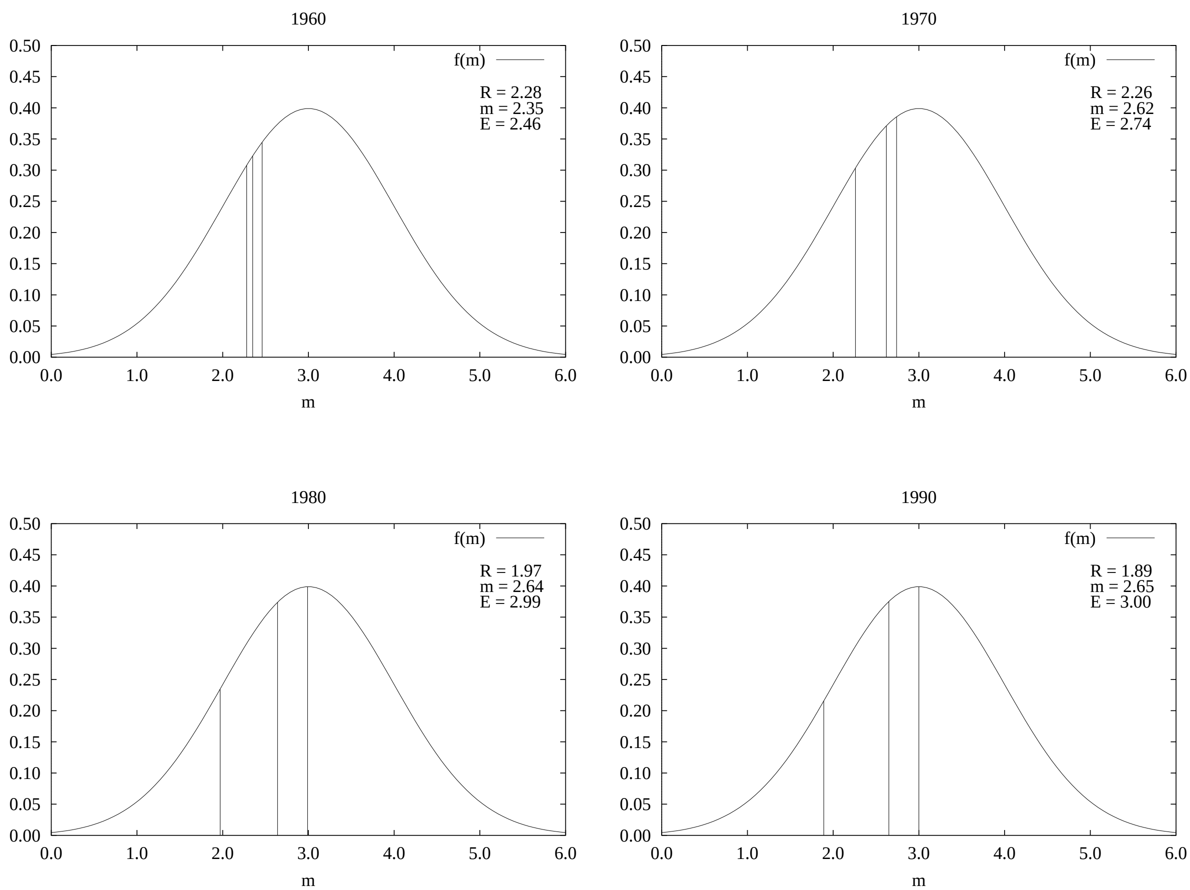
It is not entirely satisfying to treat p as an exogenous parameter, but I do so for analytical convenience[10]. The period one rejection cutoff is $R_1 = p\theta_W$ (expressions for all decision rules are in Appendix A) which increases with p; or, R_1 falls as p does. So, as contraceptive technology improves, a costly pregnancy becomes less likely and a woman is more willing to accept a low value (informal) match. The marriage threshold ($\underline{m}$) increases with p (see Appendix B); again, the more likely a costly pregnancy, the higher the required match payoff is. The high cost type man's effort threshold falls as p increases. As pregnancy becomes more likely, men are more likely to be rejected in subsequent matches, so they are less inclined to take their chances on a new draw. Conversely, as p falls, the effort threshold E^H rises; high effort becomes less attractive from men. The woman's divorce cutoff, $\overline{m}$, also increases as p increases. The cost of staying with a high cost type man increases when a future pregnancy is more likely. Conversely, if p falls, it matters less that he will choose low effort in the future, so $\overline{m}$ falls as p does.

If we believe that p fell over the 1960-90 period, we should expect (Figure 5: *Decreasing p and best response regions*):

- a smaller R (more individuals are sexually active)

[9]Incorporating income is not a simple exercise since we have to worry about the distributions of both men's and women's wages, how wages are shared in matches of different forms, and how they are affected by child bearing. Decision rules for match formation will be conditional on individuals' wage types. The intra-household allocation literature (e.g. (McElroy and Horney 1981), (Chiappori 1992)) highlights the difficulty of addressing these issues even when largely ignoring marital formation and dissolution.

[10]This assumption, however, is unlikely to substantially change the results; personal characteristics, technology, and pecuniary costs together produce an equilibrium value of p for each woman. The p of the model then represents that equilibrium choice for the representative agent. Exogenous changes in p are then changes in the equilibrium choice that would arise from changes in, for instance, technology or costs.



		1960	1970	1980	1990
Rejection cutoff	(R)	2.28	2.26	1.97	1.89
Marriage threshold	$(\underline{m})$	2.35	2.62	2.64	2.65
Effort threshold	(E^H)	2.46	2.74	2.99	3.00

Figure 4: Decision cutoff rules: 1960-1990

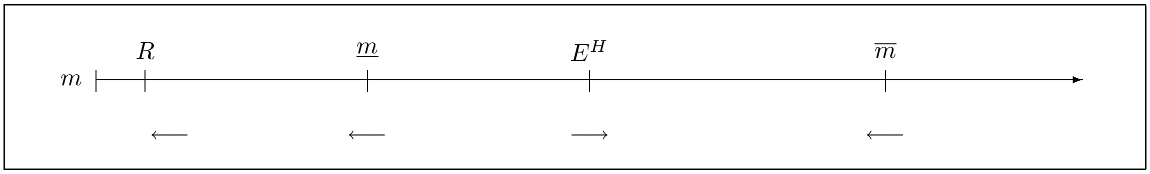


Figure 5: Decreasing p and best response regions

- a smaller $\underline{m}$ (marriages form more readily)
- an increase in E^H (men are less willing to invest high effort)
- a reduction in $\overline{m}$ (divorce threats are less credible)

The range over which informal relationships form ($|\underline{m} - R_1|$) increases[11]; even though $\underline{m}$ falls, it does not fall too much.

Reducing the likelihood of pregnancy may *increase* births out-of-wedlock as a fraction of all births. Define the illegitimacy ratio i_r as the ratio of non-marital to total (marital plus illegitimate) births:

$$i_r \equiv \frac{F(\underline{m}) - F(R_1)}{1 - F(R_1)} \tag{3}$$

Define the "hazard rate" of decision rule x as

$$h(x) \equiv \frac{f(x)}{1 - F(x)} \tag{4}$$

The hazard rate here is in terms of switching *out of* marriage ($\underline{m}$) or sexual activity (R_1). The illegitimacy ratio i_r increases as p falls if:

$$\frac{h(\underline{m})}{h(R_1)} < \frac{1}{1 - \frac{w_H}{\theta_W}} \tag{5}$$

If the hazard rate is constant (across decision rules), then this is always true (note that $w_H < \theta_W$ by assumption). If $f()$ is Gaussian, $R_1 < \underline{m}$, and $f(R_1) < f(\underline{m})$, then it depends on how much high effort is worth relative to the cost of a child (θ_W). The more valuable high effort (w_H) is, the more likely it is that the illegitimacy ratio will *increase* as pregnancy becomes *less* likely.

Matches between $\underline{m}$ and E^H are formal but may end in divorce. Since $\underline{m}$ decreases while E^H increases, increasing divorce would be consistent with a reduction in p. In addition, $|E^H - R|$ is increasing, so we would expect an increase in the likelihood of children experiencing either divorce *or* illegitimacy; increasing illegitimacy would *not* be offset by a reduction in divorce. These results are interesting, because, in a sense, children are "bads" in the model and a reduction in p makes them less likely. One of the most frequently voiced arguments in favor of increasing the availability of contraception is that it gives couples greater control over their fertility. Hence, the argument goes, "every child will be a wanted child." However, behavior changes on multiple margins; as p falls, the incentives men face to avoid divorce fall as do the incentives for single women to reject marginal matches.

The aggregate estimates qualitatively support some of these predictions (see Table 1). Fewer women are celibate ($F(R)$ declines from 0.23 in 1960 to 0.13 in 1990) and more are in informal matches ($F(\underline{m}) - F(R)$

[11] Since $\frac{d(\underline{m}-R_1)}{dp} = -w_H$, $|\underline{m} - R_1|$ increases as p falls.

increases from 0.03 in 1960 to 0.23 in 1990). The divorced group has also increased; the increase in informal relationships was not offset by a decline in divorce. The category of informal *or* divorced grew dramatically (from 0.05 in 1960 to 0.32 in 1990). However, contrary to the prediction, the marriage threshold $\underline{m}$ inferred from the aggregate data *increased*.

B "Easy" divorce, less "illegitimacy" stigma

Another factor which may have contributed to the observed changes is a reduction in exogenous penalties (taxes, stigma) incurred as a consequence of divorce or illegitimate birth. Murphy (1999) shows that imposing these penalties can increase the average welfare of children; such penalties may not be unreasonable in a society of rational agents. Let Ω_A be the proportion of adults who are sexually active and Ω_M be the proportion married (and not divorced).

$$\Omega_A = 1 - F(R) \tag{6}$$

$$\Omega_M = 1 - F(E^H) + \mu'(F(E^H) - F(\underline{m})) \tag{7}$$

Then average children's welfare (the proportion who get high investment) is: (father chooses high effort *and* couple remains together) is:

$$\gamma = \frac{\Omega_M - (1-\mu')(1-F(\overline{m}))}{\Omega_A} \tag{8}$$

Exogenous penalties act as taxes on out-of-wedlock child bearing and divorce. They increase R, discouraging acceptance of low quality informal matches and increase $\underline{m}$, discouraging low quality marriages. The penalties reduce E^H, encouraging high cost type men to invest in their children even in matches with lower match qualities. However, the penalties also make less credible a woman's divorce threat and so reduce $\overline{m}$; more women (and children) will be "stuck" in bad marriages. Under plausible circumstances, γ, a measure of average welfare, increases as penalties are imposed (Murphy 1999).

Whether these penalties (or their levels) are or were optimal, they are certainly lower today than they were 30-40 years ago[12]. Reductions would reinforce the changes in the decision rules induced by the decline in the probability of pregnancy, except for changes in $\underline{m}$ and $\overline{m}$. Reducing exogenous penalties for divorce *raises* both $\underline{m}$ and $\overline{m}$, while reducing p *lowers* them. Since the aggregate data suggest that $\underline{m}$ increased, it seems likely that the "stigma" effects on the marriage and divorce thresholds are larger than the "contraceptive" effects; $\overline{m}$ probably increased along with $\underline{m}$.

V Effects on children?

Whether because of declines in p, reductions in social penalties, or other reasons, behavior *has* changed. Decision cutoffs have shifted and the estimates of Table 1 suggest they have changed substantially. The mix of couples across different marital states has probably changed as well. Conditioning on an intact marriage in 1960 is different than conditioning on an intact one in 1980 or 1990; this has important implications for the interpretation of empirical estimates of the effect of marital status on children's outcomes.

[12]Akerlof, Yellen, and Katz (1996) endogenize a cost of out-of-wedlock births and consider reductions in it as part of their effort to understand to the decline in "shotgun" weddings.

A Measured effects: theory

Suppose we have data on individual children's outcomes and their parents' marital status, but not on parents' investment of effort in their children. A typical estimate of the benefit of marriage to children would be the difference between the outcome, on average, for those whose parents remain married and the outcome, on average, for those children whose parents do not remain married (i.e. the coefficient on a `unmarried` -or `married-` dummy variable). For instance, child i's outcome y_i might be related to a a vector of inputs $\mathbf{X}_i$ and his parents' marital status d_i ($d_i = 1$ if i's parents are not married, 0 otherwise).

$$y_i = \mathbf{X_i}\boldsymbol{\beta} - \gamma d_i + \epsilon_i \tag{9}$$

The estimated OLS coefficient $\hat{\gamma}$ would be interpreted as an estimate of costs faced by children of unmarried parents.

However, the marriage model suggests that the set of intact marriages includes both good and bad ones. Therefore, the estimated effect of marriage is the average outcome of children conditional on their parents remaining married:

$$E(\hat{\gamma}) = \frac{\Omega_M - (1-\mu')(1-F(\overline{m}))}{\Omega_M} \tag{10}$$

As a measure of children's welfare, $E(\hat{\gamma})$ is confounded by the effects of who marries, who does not, and who divorces.

To see why this matters, suppose that high penalties for divorce and illegitimacy were prevalent in the U.S. circa 1960 and that the penalties fell, exogenously. Further suppose the divorce threshold $\overline{m}$ increases (the stigma effect exceeds the contraceptive effect, see Sections III and IV); then the measured penalty of not having married parents may *increase* as fewer couples marry and more divorce. Intuitively, $E(\hat{\gamma})$ increases as fewer marriages form and more break up because the measured benefit of marriage is the proportion of *surviving* marriages with the good outcome. As penalties fall, the woman's upper cutoff, $\overline{m}$, increases. Her threat to leave is more credible, so men in high m matches are more willing to choose high effort to remain in the match. Therefore, reducing the cost of divorce to parents reduces the number of children whose mothers are "trapped" in bad marriages. If the total number of intact marriages does not fall too quickly, the *measured* benefit of marriage rises. Marriages are easier to get out of, so the ones that survive are better on average. As "family values" weaken (the penalties for divorce and births out-of-wedlock fall and the likelihood of growing up *outside* an intact marriage increases), the relative value of married parents increases. Equivalently, the measured penalty to children of growing up outside of an intact married couple household *increases*.

The appropriate welfare measure is not the measured difference between children in intact marriages and those outside of them ($E(\hat{\gamma})$), but the average investment across all children (i.e. γ). This reflects the general equilibrium effects of marital status choices. There is a simple relationship between the measured benefit and aggregate welfare: $E(\hat{\gamma}) = \gamma\frac{\Omega_A}{\Omega_M}$. An increase in accepted matches and a reduction in intact marriages will drive the bias factor $\frac{\Omega_A}{\Omega_M}$ up and if it increases quickly enough, $E(\hat{\gamma})$ increases even while γ falls. From the earlier estimates (Table 1), this ratio was 1.1 in 1960, 1.3 in 1970, 1.5 in 1980, and 1.6 in 1990. Given this increase, even if one thinks that children are worse off on average (i.e. that γ is lower) or alternatively, even if one thinks that marriage is less "important" (perhaps because it has become more acceptable to be unmarried), it is likely that the measured effect of marriage, $E(\hat{\gamma})$, has *increased*.

B Measured effects: evidence

The following exercise is *not* intended to generate estimates of the causal relationship between marital status and school attendance at age 17. Instead, the estimates reflect the incentives to choose marital status as well as the incentives which result *from* marital status. I claim that, given the incentives faced by prospective parents and the choices parents make in equilibrium, growing up outside an intact marriage should be associated with a larger apparent penalty when divorce and illegitimacy are *more acceptable* for parents.

Measuring child outcomes over time is problematic because of changes in real standards of living and in women's wages. A good measure is education, which is strongly correlated with labor force participation, earnings, and other positive outcomes. I take whether or not a child is in school at age 17 as the indicator of child welfare (in school is good and out of school is bad). Of course, if education is the indicator of child welfare, why not estimate γ_t as the proportion, in year t, of 17 year olds still in school? There are two problems with this approach. I want to consider two periods with different constraints on prospective parents and so want two periods separated in time. This, and a desire to compare periods before and after significant changes in marital behavior were observed lead me to use Census samples and particularly the years 1960 and 1980[13].

Unfortunately (in this context), school enrollment has increased over time; arguably, the increase is due to factors exogenous to family organization such as labor market returns, mandatory attendance laws, and declining costs of education. Therefore, the levels are apt to be misleading indicators of the effects of changing family arrangements. One way around this is take the approach used in the model: normalizing the non-married couple outcome at zero. This is fine for a single year, but across time is inappropriate if "society" is better at raising children of divorced and non-married parents. If so, the not-married level is rising over time and the normalization would erroneously cancel out the gain.

The measured effect, the benefit of marriage (or conversely, the penalty conditional on having not-married parents), is more appropriate since it should have declined. General social improvements for children without married parents should have reduced the difference between those whose parents are married and those whose parents are not. Labor market returns and other factors that might be increase enrollment levels should not affect the measured differential unless these factors affect children in intact marriages differently than those outside of them. And, one might think that differential impacts of the past few decades would favor the more disadvantaged.

1 Data

I use samples from the U.S. decennial censuses: the 1960 and 1980 Integrated Public Use Microdata Sample (IPUMS) files (Ruggles and Sobek 1995). The data are far from perfect; there are no longitudinal observation of households (for instance, I observe only parental status at the time of the census) and only a crude measure of child outcome: in or out of school. Nonetheless, dropping out of school is a good indicator of trouble, I know current household status[14], and know several important parental and household characteristics. Standard panel data sets (e.g. the PSID or the NLS samples) do not reach back far enough. The IPUMS

[13] The first is before substantial changes, the second is after (see Table 1 and the accompanying discussion). The intermediate census, 1970, is right in the midst of changes such as abortion legalization and changes in the costs of divorce and illegitimacy. 1980 is more appropriate than 1990 because the estimates in Table 1 suggest that decision cutoffs did *not* change much between 1980 and 1990, but other aggregate data indicate that enrollment rates did. Work in progress includes checking the 1960 and 1980 results with estimates from other years.

[14] Observing only current marital status may not be too bad. Census figures (U.S. Dept. of Commerce 1975), indicate that the median duration of marriages ending in divorce was 7.2 years in 1960. Between 1950 and 1970, the median was no lower than 5.8 and no higher than 7.5 years. Few parental marital status decisions are probably being made when children were 16 years old.

x		1960 $\overline{x}$	1960 s_x	1980 $\overline{x}$	1980 s_x
`school`	In school	.754	.4303	.843	.3636
`female`	Female	.495	.4999	.485	.4997
`white`	White	.879	.3256	.828	.3767
`black`	Black	.111	.3143	.145	.3522
`marry`	Parents married	.731	.4433	.693	.4611
`divorce`	Parents divorced	.026	.1610	.090	.2865
`widow`	Widowed	.051	.2208	.045	.2085
`step`	Step-parent(s)	.014	.1189	.012	.1094
`single`	Single mother	.001	.0334	.010	.1008
`par_oth`	Other HH type	.175	.3801	.148	.3556
`notmarry`	Neither married, nor other	.093	.2912	.158	.3649
`med_med`	Missing mother's educ.	.158	.3650	.134	.3416
`mom_a`	Mother: grade 4 or less	.042	.2027	.015	.1237
`mom_b`	Mother: grades 5-10	.389	.4876	.197	.3979
`mom_11`	Mother: grade 11	.058	.2341	.056	.2310
`mom_12`	Mother: HS graduate	.248	.4320	.390	.4878
`mom_c3`	Mother: 1-3 college	.067	.2509	.126	.3322
`mom_c4p`	Mother: at least college graduate	.035	.1844	.078	.2695
N			14,292		21,026

Table 2: Sample means and standard deviations (17 year olds)

files used, `ip19601` and `ip19802`, are 1% samples from which I select one half of the households[15]. The resulting files are 0.5% population samples for 1960 and 1980; from them, I select all individuals aged 17 years. Sample characteristics (for the 17 year olds) are shown in Table 2.

2 Apparent marriage benefit?

I estimate the apparent penalty to children of not growing up in an intact married couple household with two probit equations: one for 1960 and the other for 1980. The regressions are (naive) estimates of the probability of being in school at age 17 as a function of sex, race, mother's education, and parents' marital status. The estimates are shown in Table 3. The coefficients (probit coefficients translated into probability changes: $\Delta G(\mathbf{X}\hat{\boldsymbol{\beta}})$) are tightly estimated: most are several times larger than their corresponding standard errors. The `notmarry` variable is a dummy variable that is 0 if the individual's parents are married and 1 if his or her mother is divorced, widowed, remarried, or a single mother[16]. Other parental household arrangements (for instance, living with grandparents, not living with either parent) are captured by `par_oth`. The `mom_x` variables are dummy variables indicating the individual's mother's highest level of education (see Table 2).

The coefficients of particular interest are those on `notmarry`. The prediction was that as divorce and illegitimacy became less stigmatized, the measured penalty for not being in a married couple home would be *larger*; that is, the effect would be larger in 1980 than in 1960. From Table 3, we can see that in 1960, the estimated penalty for not growing up in (not being in at the date of the Census) a married couple household is -3.5%; the estimated penalty in 1980 is -6.8% (the figures should be interpreted as the measured change

[15]Each half sample was selected using the provided IPUMS sub-sample numbers; these give access to 100 random sub-samples of each file of which I select the first 50.

[16]Estimates for each marital status separately (not shown) are consistent with the combined estimate.

"Naive" estimates of enrollment status
(dependent variable: `school`)

	1960					1980				
	Full Sample	Mother's Education Known	Parents' Status Known	White Only	Non-white Only	Full Sample	Mother's Education Known	Parents' Status Known	White Only	Non-white Only
`notmarry`	−.0349	−.0445	−.0435	−.0309	−.0460	−.0683	−.0695	−.0647	−.0706	−.0370
	(.0130)	(.0120)	(.0118)	(.0140)	(.0355)	(.0073)	(.0067)	(.0064)	(.0083)	(.0163)
`female`	.0241	−.0264	−.0278	.0265	.0129	.0161	.0251	.0264	.0139	.0311
	(.0072)	(.0115)	(.0119)	(.0075)	(.0230)	(.0047)	(.0045)	(.0045)	(.0051)	(.0125)
`black`	−.0048	.0316	.0319			.0353	.0095	.0111		
	(.0110)	(.0069)	(.0069)			(.0063)	(.0063)	(.0066)		
`par_oth`	−.1632	−.0975		−.1777	−.1263	−.1621	−.0814		−.1847	−.1059
	(.0189)	(.0196)		(.0226)	(.0377)	(.0113)	(.0119)		(.0141)	(.0199)
`mom_a`	−.3366	−.3032	−.3095	−.3672	−.2322	−.2073	−.1871	−.2197	−.2355	−.0865
	(.0213)	(.0198)	(.0202)	(.0246)	(.0563)	(.0224)	(.0200)	(.0216)	(.0251)	(.0499)
`mom_b`	−.1345	−.1119	−.1112	−.1311	−.1185	−.1350	−.1128	−.1150	−.1446	−.0822
	(.0101)	(.0085)	(.0086)	(.0102)	(.0447)	(.0075)	(.0064)	(.0066)	(.0083)	(.0182)
`mom_11`	−.0438	−.0377	−.0296	−.0310	−.1402	−.0805	−.0668	−.0635	−.0772	−.0590
	(.0184)	(.0159)	(.0160)	(.0186)	(.0738)	(.0120)	(.0103)	(.0107)	(.0135)	(.0268)
`mom_c3`	.1059	.0894	.0870	.1078	−.0389	.0342	.0291	.0262	.0383	.0071
	(.0179)	(.0150)	(.0149)	(.0177)	(.1036)	(.0085)	(.0071)	(.0072)	(.0088)	(.0264)
`mom_c4p`	.1238	.1024	.1124	.1238	.0618	.0675	.0554	.0546	.0664	.0587
	(.0239)	(.0197)	(.0200)	(.0237)	(.1197)	(.0104)	(.0087)	(.0088)	(.0107)	(.0361)
`med_miss`	−.3080		−.1572	−.3059	−.2554	−.1911		−.0873	−.1992	−.1173
	(.0219)		(.0336)	(.0255)	(.0564)	(.0122)		(.0147)	(.0145)	(.0245)
observed P	.7545	.8144	.8171	.7659	.6715	.8432	.8843	.8852	.8471	.8245
predicted P	.7813	.8298	.8327	.7945	.6799	.8697	.8977	.8983	.8768	.8357
n	14,292	12,030	11,788	12,569	1,723	21,026	18,189	17,903	17,425	3,601
Pseudo R^2	0.1246	0.0588	0.0585	0.1349	0.0492	0.1236	0.0628	0.0616	0.1456	0.0460

Table 3: Coefficients (dG/dX for unit change in variable) with s.e. in ().

in the probability of being enrolled in school at age 17 for a child whose parents are not married relative to one whose parents are married, conditional on other characteristics). The effects may seem small, but recall that the overall dropout rates are small as well: 24.6%6 in 1960 and 15.7% in 1980. The estimated penalty in 1980 is enough to offset the benefit of having a mother with 4-plus years of college relative to one with only 12 years of education.

The estimates are not sensitive to specification of mother's education (e.g. dummies for each year of education or alternative groupings) or to including additional racial dummies (e.g. each individual race or white, black, all other). The estimates do vary if I construct separate racial sub-samples (white and non-white). Estimates for whites only are similar to those for the full sample, with the change in the `notmarry` coefficient more pronounced than in the combined samples (from -3.1% in 1960 to -7.1% in 1980). For non-whites, the measured 1960 penalty is -4.6%, but with s.e. of 3.6; however, the measured 1980 penalty is smaller, only -3.7%, but the s.e. of 1.6.

This increase in the apparent penalty associated with not having married parents is surprising computationally: the mean probability of being in school at age 17 increased from 0.75 in 1960 to 0.84 in 1980. Ordinarily, as $G(\mathbf{X}\boldsymbol{\beta}) \to 1$, we expect to have a more difficult time detecting significant effects of independent variables. Since $G(\cdot)$ is most sensitive to changes around 0.5, we would expect the estimated coefficient to fall as the observed probability moves away from that. Here we find the opposite; the estimated effect is larger when the observed probability is 0.84 than when it is 0.75.

The results do not change substantially when the sample is restricted to only individuals for whom parental marital status is identified (i.e. `par_oth` $\neq$ `1`) or if restricted to those individuals for whom mother's education level is known (i.e. `med_miss` $\neq$ `1`). The estimates are also quite similar when controls for state of residence are included (see Tables 8 and 9 in Appendix C).

3 High and Low Cost Groups

The prediction concerning increases in the apparent "divorce penalty" was presented in terms of time: the apparent penalty associated with having unmarried parents should be higher in period $t + 1$ when parents face lower costs of divorce and out-of-wedlock child bearing than they do in period t. However, it can also be interpreted cross-sectionally: children in groups in which parents face higher costs of divorce should exhibit a *lower* apparent penalty associated with their parents being unmarried, while those whose parents face lower costs should exhibit *higher* measured penalties.

Two groups likely to have different costs (on average) are residents of states with more Roman Catholics and residents of states with high divorce rates. For Catholics, divorce is unavailable within the Church; a failed marriage means either:

- re-marriage outside of the Church and exclusion from sacraments within the Church
- applying for a declaration of nullity (an "annulment")
- never re-marrying

Annulments are not uncommon now.[17] Prior to the Second Vatican Council (1962-1965) never re-marrying was probably more likely for Catholics whose marriages failed; re-marriage outside of the Church is more common today. In the aftermath of Vatican II, which led the Church to stress social teachings more and adherence to rules less, the differences in divorce costs between Catholics and non-Catholics have probably fallen.

States with higher divorce rates are, arguably, those with lower costs of divorce. Again, one might suppose that these differences were more pronounced in 1960 than in 1980.

[17] There were between 55 and 60 thousand declarations annually 1987-1994 in the U.S., while marriages were roughly 300 thousand annually (Catholic Church. Rationarium Generale Ecclesiae (various years)).

Table 4: Measured dropout penalty: Catholic and divorce effects (`unmarry`)

		1960[a]	1980
Base	`unmarry`[b]	−.0740	−.0887
		(.0135)	(.0082)
% Catholic	`unmarry`	−.1138	−.0856
		(.0191)	(.0140)
	`unmarry * % Catholic`[c]	+.1733	−.0130
		(.0836)	(.0553)
Divorce rate	`unmarry`	−.0457	−.0879
		(.0255)	(.0210)
	`unmarry * Divorce rate`[d]	−.0121	−.0001
		(.0084)	(.0026)
Both	`unmarry`	−.0858	−.0791
		(.0285)	(.0270)
	`unmarry * % Catholic`	+.1386	−.0198
		(.0825)	(.0571)
	`unmarry * Divorce rate`	−.0083	−.0007
		(.0065)	(.0026)

[a] All models include controls for mother's education and state of residence; robust standard errors in ()s.

[b] Dummy variable = 0 if parents are together and currently married; 1 otherwise.

[c] Catholic population from (P. J. Kennedy and Sons (various years)).

[d] Divorce rates from (U.S. Department of Health, Education, and Welfare, Public Health Service 1964), (National Center for Health Statistics; Public Health Service 1985).

Given data on the Catholic population of states and state divorce rates, consider models such as:

$$y_i = \beta_0 \mathbf{X}_i + \beta_1 d_i + \beta_2 (d_i * RC_j) + \epsilon_i \tag{11}$$
$$y_i = \alpha_0 \mathbf{X}_i + \alpha_1 d_i + \alpha_2 (d_i * DIV_j) + \eta_i \tag{12}$$

for individual i in state j ($d_i = 1$ if i's parents are *unmarried*). The prediction is that $\beta_1 < 0$ (children of unmarried parents are more likely to drop out) and $\beta_2 > 0$ (states with more Catholics should exhibit a *smaller* average difference between children of married and unmarried parents. Furthermore, if the excess divorce costs faced by Catholic parents fell over time, then we should expect $\beta_{2,1960} > \beta_{2,1980} > 0$, with $||\beta_{2,1980}||$ depending on the extent to which Catholics still face relatively higher costs of divorce and out-of-wedlock child bearing. Similarly, we expect $\alpha_1 < 0$ and $\alpha_2 < 0$; where divorce is less costly and more prevalent, the difference between children in married families and others should appear to be *larger*.

Estimates from 1960 and 1980 are generally consistent with these predictions (Table 4). The coefficient on `unmarry` (i.e. `unmarry` = 1 if either `notmarry` or `par_oth` is 1) is negative in both years. The interaction with % Catholic is positive in 1960, as expected (i.e., in 1960, the apparent penalty for not having married parents was *lower* in states with more Catholics) and insignificant (but negative) in 1980. The divorce rate interaction is negative in both 1960 and 1980, as expected, although the coefficients are not much larger than their standard errors. These patterns persist when both interactions are included simultaneously.

Less aggregated estimates (i.e. separate `notmarry` and `par_oth` coefficients and separate interactions with the Catholic and divorce regressors were generally similar (see Tables 8 and 9 in Appendix C). The 1960 % Catholic interactions were positive (one significant, one not) while the 1980 ones were negative and about as large as the standard errors). The divorce rate interactions were inconclusive. The estimates are not entirely consistent with the model predictions. However, given the measurement error inherent in variables such as "proportion Catholic" (relative to *being Catholic*) and "divorce rate" (relative to *low cost of divorce*), the overall results are supportive of the model predictions. The *smaller* penalty for children in states with *more* Catholics in 1960 is particularly suggestive.

VI Child Outcomes and Selection into Marriage

The model suggests that the difference in estimated risks reflects measurement issues rather than differences across time in the benefit of marriage. The distinction between the measured and true effects of marital status arises from two specification problems in the simple regression model of the relationship between child outcomes and parents' marital status. One problem is the endogeneity of marital status; parents' choices of marital status are correlated with decisions affecting child outcomes, leading to inconsistent estimates of the effect of marital status. The second problem arises because of an omitted variable: intact "bad" marriages. We cannot distinguish between surviving good and bad marriages and hence our estimated marriage benefit reflects a blend of the benefits of good marriages and the costs of bad ones.

Using two estimation techniques (joint maximum likelihood estimation of child dropout and parental marital status equations and instrumental variables estimation of the child dropout equation), I show that the estimated adverse effect of having unmarried parents persists after accounting for the correlation between marital status and the error term in the dropout regression model, while the difference over time does not. Children in households with unmarried parents are still, on average, worse off than children in households with married parents. However, accounting for the selection of couples into married or unmarried states suggests that, after selection effects are accounted for, the benefits of marriage (to children) are quite similar in 1960 and 1980 (and larger than either of the naive 1960 or 1980 estimates).

A Endogeneity of marital status

Since father's effort affects both child outcomes and mothers' payoffs, marital status is determined endogenously. Let m_i^* be value of the current match (i.e. match quality) and w^* be the value of the father's effort choice (high or low) to the child's mother[18]. From Section II, the value of low effort is less than the value of high effort: $w_L^* < 0 < w_H^*$. Marital status, s_i, is determined by:

$$s_i = \begin{cases} 1 & \text{(the parents are married) if: } m_i^* + w^* + \eta_i > M \\ 0 & \text{otherwise} \end{cases} \tag{13}$$

for a constant M and a suitably behaved error term η_i.

The intact marriages that we observe are those that are valuable enough, in terms of match quality and father's effort, that women choose not to leave. On average, intact marriages will have large ηs. Note that M may, in fact, vary (e.g. across time): one can interpret the results of Section V as suggesting that the threshold for marital "success" (i.e. no divorce) in 1960 was lower than the threshold for success in 1980, or that $M_{1960} < M_{1980}$. For the discussion that follows, however, I consider one year at a time and normalize M to 0.

[18]Notation: the * variables are unobservable.

We can then write this as a (simplified) switching model with varying coefficients; children are produced in either a married regime or in an unmarried one. Although the outcomes are potentially well defined for children under both regimes, we observe outcomes under only one regime for each child. Since we expect parents (only mothers in the model) to weigh the effect of marital status on children in *choosing* marital status, coefficient estimates will be inconsistent.

Let unobserved (potential) child outcomes in the two regimes be given by:

$$y^*_{mi} = \mathbf{X}_i\boldsymbol{\beta} + \alpha_i + u_i \quad \text{If } i\text{'s parents are married} \tag{14}$$
$$y^*_{ni} = \mathbf{X}_i\boldsymbol{\beta} + v_i \qquad \text{if } i\text{'s parents are unmarried} \tag{15}$$

or

$$y^*_i = s_i(\mathbf{X}_i\boldsymbol{\beta} + \alpha_i + u_i) + (1 - s_i)(\mathbf{X}_i\boldsymbol{\beta} + v_i) \tag{16}$$

Here X_i is a vector of exogenous characteristics. To simplify the estimation, I assume that the regimes vary only in intercept terms ($\beta_0 + s_i\alpha_i$) and that disturbances are equivalent in the two regimes ($u_i = v_i$). Both assumptions are restrictive and could be relaxed. The current formulation, however, is consistent with both the existing literature and previous sections. It should be sufficient to provide indications of at least simple differences in the production of children. The net effect of marital status, α_i, may vary across matches:

$$\alpha_i = \overline{\alpha} + \epsilon^*_i \tag{17}$$

where $E(\epsilon^*_i) = 0$.

Parents (mothers) choose marital status while taking into account individual characteristics ($\mathbf{Z}_i$) and the effect of marital status on child outcomes. They value these differences according to a weighting function $W()$. So status, s_i, is determined as:

$$s_i = \begin{cases} 1 & \text{if } \mathbf{Z_i}\boldsymbol{\delta} + W(y^*_{mi} - y^*_{ni}) > 0 \\ 0 & \text{otherwise} \end{cases} \tag{18}$$

In terms of the earlier notation (13), $m^*_i = \mathbf{Z_i}\boldsymbol{\delta}$, $w^* = W(\overline{\alpha}\ )$, and $\eta_i = W(\epsilon^*_i) = \epsilon_i$. The function $W()$ is such that $E(\epsilon_i) = E(\epsilon^*_i) = 0$. Marital status is chosen on the basis of match specific characteristics $\mathbf{Z}_i\boldsymbol{\delta}$, the "true" (average) benefit from an intact marriage $W(\overline{\alpha})$, and the idiosyncratic consequences of *this* marriage, if it should remain intact, ϵ_i. The vector $\mathbf{Z}_i$ of exogenous regressors affects parents' (mothers') marital status, but not the child outcomes y^*_{mi} and y^*_{ni}. Then

$$y^*_i = s_i(\mathbf{X}_i\boldsymbol{\beta} + \overline{\alpha} + \epsilon_i + u_i) + (1 - s_i)(\mathbf{X}_i\boldsymbol{\beta} + u_i) \tag{19}$$

and marital status is:

$$s_i = \begin{cases} 1 & \text{if } \epsilon_i > -\mathbf{Z_i}\boldsymbol{\delta} \\ 0 & \text{otherwise} \end{cases} \tag{20}$$

where the value of the unknown marital status effect ($w(\overline{\alpha})$) has been incorporated into the marital status *regressors* Z_i (i.e. as part of the constant). We observe outcomes for children only under one regime:

$$y^*_i = \begin{cases} y^*_{mi} & \text{if } s_i = 1 \\ y^*_{ni} & \text{if } s_i = 0 \end{cases} \tag{21}$$

For estimation purposes, the data are such that we observe only a binary outcome; y_i^* is a latent, rather than observed, outcome:

$$y_i = \begin{cases} 1 & \text{if } y_i^* > 0 \\ 0 & \text{otherwise} \end{cases} \tag{22}$$

Estimation of (19) by OLS (or ML given (22)) will be inconsistent since the error term $s_i\epsilon_i + u_i$ is correlated with s_i. To generate consistent estimates of $\overline{\alpha}$, the parameter of interest in (19), we need to deal with (or avoid) the endogeneity of marital status in the child outcome equation.

B Father's effort

There is another specification issue as well. Ignore, for now, the endogeneity of marital status and suppose that status is an observable, exogenous regressor which we believe affects child outcomes.

The model predicts a qualitative difference between the outcomes of children whose parents' match quality is sufficiently high (above the high cost man's effort threshold E^H) and those whose parents' match quality is below it. For matches above this effort threshold, but below the woman's credible divorce threat cutoff ($\overline{m}$), fathers choose to invest high effort even when high effort is costly. They do so to avoid losing the match and, at the margin, to avoid the costs associated with being identified as divorced.

In matches *above* the mother's credible divorce threat level ($\overline{m}$), fathers know they can choose high *or* low effort without losing the match. The fathers in these matches who choose low effort are never observed. Therefore, intact marriages are not all alike; in "bad" marriages, the father chooses low effort, but the marriage survives anyway (which is worse for the child than a breakup).

Couples are characterized by:

1. unobservable match quality ($m_i^* \in \mathbb{R}^1$)
2. unobservable father's effort: high or low ($e_i^* \in \{H, L\}$)
3. observable marital status: married or not ($s_i \in \{1, 0\}$)

Implicitly, child outcomes in the model are good, bad, or indifferent; the observed outcome, y, would be:

$$y_i = \begin{cases} 1 & \text{if } s_i = 1 \text{ and } e_i^* = H \\ -1 & \text{if } s_i = 1 \text{ and } e_i^* = L \\ 0 & \text{otherwise} \end{cases} \tag{23}$$

If we observe father's effort choice and marital status then we can consistently estimate the effects of marital status and effort choice (let $e_i = 1$ designate *low* effort: $e_i = 1$ if $e_i^* = L$ and 0 otherwise). Suppose, however, that we do not observe father's choice of (low) effort in intact marriages (s_ie_i). Instead of:

$$y_i = \beta_0 + \beta_1 s_i + \beta_2 s_i e_i + u_i \qquad u_i \sim N(0, \sigma^2) \tag{24}$$

we estimate the incorrectly specified model:

$$y_i = \alpha_0 + \alpha_1 s_i + u_i \tag{25}$$

Then $E(\hat{\alpha}_1)$ has the intuitive form of

$$E(\hat{\alpha}_1) = \beta_1 + \frac{\Sigma s_i e_i}{\Sigma s_i}\beta_2 \tag{26}$$

Since $\beta_2 < 0$ (β_2 is the effect on a child's welfare of his parents remaining together even though his father chooses low effort), the estimate of β_1 (the "good effect" of marriage when marital status is exogenous) is biased downwards. The extent of the bias depends on the cost (to the child) of low effort, β_2, and on the proportion of intact marriages in which the father chooses low effort, $\frac{\Sigma s_i e_i}{\Sigma s_i}$. The more often intact marriages involve low effort by fathers, the larger the bias. As long as fathers' effort is inherently unobservable, none of our estimation procedures will allow us to generate consistent estimates of β_1. Given a proxy for father's effort, one possibility is to simply use the proxy and accept the resulting (possibly poor) estimates as the best possible.

C Two problems or one?

Conceptually, we face two estimation problems:

1. If parents (mothers) choose marital status on the basis of the effects of their spouse's actions on their own welfare as well as their child's, then marital status will not be independent of the child outcome (i.e. enrollment status) equation error term.
2. Because father's effort within intact marriages is unobservable, estimates of the effect of marital status on child outcomes will suffer from an omitted variable problem.

For several reasons, however, I argue that they should be treated as one problem.

First we are interested in the *net* effect of marital status. The incentives arising from the institution of marriage indicate that there is neither *a* "good" nor *a* "bad" effect of marrying. The theory makes clear that surviving *bad* marriages and the incentives provided within *good* marriages are two sides of similar coins: exit costs within formal matches. Hence, the omitted variable (father's effort) "problem" is inherent in the institution; it neither can nor should be entirely corrected.

Second, the bias arising from not observing father's effort choice within an intact marriage is proportional to the number of surviving bad marriages (26). Provided that the estimation procedures do a reasonable job of accounting for the endogeneity of marital status, I will have controlled (somewhat) for selectivity into (and out of) marriage; the estimated effects of marital status ought to be comparable across years. That is, the estimated effects will still reflect the omitted variable problem, but dealing with the selectivity effects should make the magnitude of the omitted variable effects similar in different periods.

Third, although the problems of father's effort and endogeneity of marital status are potentially separable in regression models such as (24), both practical and theoretical considerations suggest that factors which influence one will also affect the other. In particular, both effects are driven by the costs of choosing one marital status over another.

Finally, the data are limited; I have few plausible regressors which may be used to induce exogenous variation in marital status for estimation purposes. Given the inherent limitations of the data, prudence is appropriate.

D Estimators

There are several alternative methods for consistently estimating the effects of marriage while dealing with the likely selection of couples into intact marriages. I focus on the two methods most likely to generate consistent estimates of the effect of marital status in the current context: bivariate maximum likelihood and instrumental variables. Maximum likelihood requires specification of a joint structure on u_i and ϵ_i (the disturbance terms in the outcome and marital status equations, respectively), but once the structure is specified (and assuming that the specification is correct), the likelihood function for the sample allows iteratively solving for the appropriate estimates. Alternatively, with suitable instruments, I can bypass the

Variable	Description
`mnchild`	Number of children of mother (or mother surrogate).
`ftotinc`	Reported total family income.
`mage`	Mother's age.
`dfborn`	Dad is foreign born (including US possessions such as Puerto Rico).
`mfborn`	Mother is foreign born (including US possessions such as Puerto Rico).
`mmig5`	Mother did not live in this county 5 years ago.
`moccscr`	Average income, in 1950, of mother's occupation.
`unmarry`	Parents are *not* married and living together with the child.

Table 5: Bivariate probit: variables

dual-equation problem and handle the endogeneity problem in a computationally cheaper framework[19]. Both the bivariate ML and IV approaches solve (conceptually) the endogeneity problem, but accommodate the omitted variable problem only in the rough, indirect manner discussed above.

E Results

Ideally we would like variables that provide exogenous variation in marital status without affecting school dropout propensity. Unfortunately, the options are somewhat limited.

Explain IV choices.

As an alternative, I estimate several different specifications of the *not married* equation and check to see how the different specifications affect the estimated effect of marital status.

I also make use of two additional pieces of information from the estimation procedures. From the first stages of the IV and the bivariate ML procedures, I have a sense of how good the instruments are in explaining variations in marital status. With the IV estimates, by restricting attention to over-identified models (multiple instruments for marital status), I can test the over-identifying restrictions[20]. This provides a test of the joint hypothesis that the instruments do not belong in the school dropout equation *and* that they are uncorrelated with the error term in the dropout equation.

The variables are described in Table 5. All of the information about parents is collected for biological parents *or* for the equivalent within the child's household (e.g. step-mother). Dropouts are too heavily concentrated among those children for whom we have limited parental information for samples of only children with complete biological parent information to be representative.

The most promising instruments are those related to mobility. Arguably, living apart from family and friends lowers the costs of ending marriages. I have two sets of measures of mobility: mother's movement between counties within the past five years (`mmig5`) and child's parents being born in or outside of the United States[21] (`mfborn` and `dfborn`). One might believe that these characteristics would also affect the likelihood of child enrollment in school. I argue nonetheless for their use for two reasons. First is the matter

[19] Since the outcome variable is binary, I use an efficient instrumental variables estimator for limited dependent variables developed by Newey (1987).

[20] The test is based on an OLS regression of residuals from the probit IV model on all of the instruments ($\mathbf{X_1}$ and $\mathbf{X_2}$). The test statistic, which is (for non-linear least squares models) distributed $\chi^2(l)$ (l the degree of over-identification) is calculated as n times the ratio of the explained sum of squares to the total sum of squares of this regression (Davidson and MacKinnon 1992). On the basis of a limited amount of Monte Carlo testing of models similar to the ones estimated here, this statistic, used in the probit IV case, appears to successfully reject poorly specified models, but to too often reject the null hypothesis that the over-identifying restrictions are valid. Critical values for the simulated model with valid instruments (two degrees of freedom) are 12.55 (p=0.05) and 19.88 (p=0.01) as opposed to 5.99 and 9.21 for the χ^2 distribution.

[21] I have classified those born outside of the 50 states and the District of Columbia (e.g. Puerto Rico) as foreign born.

of practicality: there are few reasonable possibilities. Second, I can test (somewhat) for their suitability (i.e. tests of the over-identifying restrictions); as an empirical matter, they *do* seem to be reasonable instruments. I also present estimates augmenting them with other possible variables (`ftotinc, chborn, moccscr`, and `mage`). Results are not presented for other specifications, which:

- fare poorly on estimation quality criteria; for instance: convergence was difficult, the instruments are clearly not exogenous, or the instruments were unrelated to marital status
- do not, on balance, provide widely divergent evidence on the effects of marital status (although naturally the estimates are quite noisy)

These alternative specifications included IPUMS (i.e. Census) variables such as ages of oldest and youngest children and number of children ever born and also local environment measures such as state divorce rate and proportion Catholic.

The main instruments, parents' foreign born status (`mfborn, dfborn`) and mother's migration status (`mmig5`) are strongly related to parents' marital status (Table 6). The first stage models produce R^2 figures in the range of 0.45-0.65. The coefficients are individually tightly estimated with the signs generally as expected. That is, mother's migration status and parents' foreign birth are associated with a greater likelihood of one's parents being unmarried. The one exception to this is mother's foreign birth is positively related to parents being unmarried in 1960, but negatively related in 1980. The sign of the 1980 estimate seems to be related to interactions with race. White-only models have similar results to the overall models (i.e. the `school` equation coefficients are similar), and the sign on foreign born mothers is consistent across 1960 and 1980.

It is unlikely that parents' foreign born status arises *from* parents' marital status, unless divorced and unmarried parents are disproportionately likely to move to the United States. Mother's migration status is, arguably, more likely to arise *from* marital status rather than affecting the costs of different marital states. However, my measure of migration status is movement from another county within the last five years. To the extent that moves are caused by changes in marital status, this variable is a reasonable instrument as long as these status induced changes are unlikely to occur across counties. The moves most likely to be a worry, those of mothers and children post-divorce, are likely to be to less expensive housing within the same area. Moves arising *from* marital status changes are more likely to be local when women want to remain near existing jobs and social networks. The tests of overidentifying restrictions support these conclusions; there is little evidence that these instruments belong in the basic model or that they are correlated with the error term. Separate results (not reported) are qualitatively similar for parents' foreign born status alone as instruments (i.e. without mother's migration status).

Summary results, for the estimated effects of the reduction in the probability of being in school at age 17 for children of unmarried parents[22], are presented in Table 7. Since the pattern of interest is the same, but the data is a little cleaner for the sub-sample for which parental marital status is known (i.e. `notmarry` $\in \{0, 1\}$ and `par_oth` $\neq 1$, see Table 3), the estimates here all pertain to this group. Results for the full sample are qualitatively similar. The main finding of this section is summarized by the estimates on the first three lines of Table 7 for the Naive probit, Base BV, and Base IV estimators. The naive estimates (i.e. Section V) show the clear difference between 1960 and 1980 estimates. The endogenous selection models narrow the gap between these figures considerably. The new 1960 estimates average 7.6% while the 1980 ones average 7.9%; both sets of estimates are in a fairly narrow range. The limited diagnostics suggest that the instruments used in at least the first model are plausibly correlated with marital status and exogenous to dropout status. The models are well behaved; the other coefficient estimates (e.g. on mother's education) are also quite similar from one specification to another (see Tables 10, 11, 12, and 13 in Appendix C).

[22]The probability effects are calculated on the basis on single equation (i.e. dropout equation) coefficients and sample means of the regressors. The standard errors for the probability effects are the index coefficient standard errors scaled by the ratio of the probability effect to the index coefficient value.

(dependent variable: `notmarry`)

	1960			1980		
	Base	I	II	Base	I	II
`female`	−.0024	−.0021	−.0023	.0013	.0029	.0016
	(.0040)	(.0039)	(.0040)	(.0035)	(.0033)	(.0039)
`black`	.1006	.0963	.0859	.0976	.0902	.0876
	(.0072)	(.0071)	(.0075)	(.0054)	(.0052)	(.0056)
`mom_a`	−.0490	−.0340	−.0735	−.1472	−.1078	−.1624
	(.0101)	(.0098)	(.0102)	(.0141)	(.0132)	(.0140)
`mom_b`	−.0091	−.0050	−.0234	−.0110	−.0131	−.0342
	(.0048)	(.0047)	(.0049)	(.0046)	(.0044)	(.0047)
`mom_11`	.0013	.0008	−.0029	.0005	−.0023	−.0085
	(.0086)	(.0082)	(.0085)	(.0075)	(.0070)	(.0074)
`mom_c3`	.0037	.0060	.0188	−.0003	.0008	.0129
	(.0080)	(.0077)	(.0080)	(.0052)	(.0049)	(.0052)
`mom_c4p`	−.0079	−.0105	.0204	−.0169	−.0103	.0109
	(.0105)	(.0102)	(.0106)	(.0063)	(.0060)	(.0064)
`med_miss`	.7903	(dropped)	(dropped)	1.0710	(dropped)	(dropped)
	(.0190)			(.0131)		
`mfborn`	.0208	−.0525	.0269	−.1720	−.2542	−.1637
	(.0083)	(.0083)	(.0083)	(.0064)	(.0062)	(.0063)
`dfborn`	.5552	.5767	.5464	.7200	.7524	.6996
	(.0060)	(.0059)	(.0061)	(.0046)	(.0044)	(.0048)
`mmig5`	.1261	.0965	.1251	.0266	.0167	.0253
	(.0060)	(.0059)	(.0060)	(.0035)	(.0033)	(.0035)
`moccscr`		.0013			.0006	
		(.0002)			(.0001)	
`mage`		−.0007			−.0017	
		(.0003)			(.0003)	
`ftotinc`			−.0077			−.0025
			(.0005)			(.0001)
`mnchild`			−.0038			−.0066
			(.0012)			(.0012)
`constant`	.0000	.0149	.0755	.0096	.0741	.1026
	(.0043)	(.0150)	(.0067)	(.0037)	(.0126)	(.0061)
N	11,773	11,533	11,597	17,398	16,895	16,988
R^2	0.5225	0.4865	0.4742	0.6527	0.6620	0.6278

Table 6: First stage (OLS) Newey IV estimates: coefficients (s.e.)

		1960			1980		
Model	Estimator	Δ Risk	1st stage (Pseudo) R^2	χ^2 (Over-ID)	Δ Risk	1st stage (Pseudo) R^2	χ^2 (Over-ID)
Naive	probit	−.0444	–	–	−.0725	–	–
		(.0126)			(.0075)		
Base	BV	−.0757	(0.51)	–	−.0743	(0.46)	–
		(.0179)			(.0091)		
	IV	−.0744	0.52	9.7	−.0843	0.65	8.4
		(.0184)			(.0090)		
Base plus	BV	−.0758	(0.55)	–	−.0785	(0.63)	–
`moccscr,`		(.0184)			(.0084)		
`mage`	IV	−.0901	0.49	198.9	−.0872	0.66	13.3
		(.0211)			(.0087)		
Base plus	BV	−.0706	(0.26)	–	−.0816	(0.36)	–
`ftotinc,`		(.0172)			(.0085)		
`mnchild`	IV	−.0680	0.47	159.9	−.0674	0.63	164.6
		(.0167)			(.0099)		

Table 7: Change in enrollment probability for children of unmarried parents

The general pattern, negative (and significant) coefficients for not married parents and similar effects in 1960 and 1980, is fairly robust to alternative specifications such as samples limited to only whites or only those whose mother's education is known (although, as in Section V, the estimates for non-whites only are less precise and often significantly different than those for whites). These results are presented in Appendix C[23]. Note that, consistent with the earlier discussion of increasing marriage thresholds (k_{1960} and k_{1980}), the marital status equation constant terms are larger in 1980 than in 1960.

VII Conclusion

Couples face a commitment problem in child rearing: intimacy precedes investment in children. A man who promises to care for his partner (and child) if she becomes pregnant must fulfill the promise after they have already been intimate. This view of marriage seems more natural than one based on explicit, enforceable contracts. Marriage does not simply provide outcomes for children, but does provide incentives for parents. Divorce, and the costs associated with divorce, directly affects these incentives. This has implications for empirical efforts to study the effects of marital status on children.

Empirically, marital status is easy to examine. Large public datasets capture marital status and other individual characteristics and are readily available at decreasing cost. Estimating, with one of these datasets, "the effect of marriage on children," "the effect of divorce on children," or other marital "effects" is easy to do. Analysts typically regress an outcome measure against individual characteristics and interpret the coefficient on the marital status dummy variable as the marginal, or average marginal, effect of that marital state. But this interpretation is inappropriate because it fails to account for the incentives for choosing and remaining in different marital states. The estimated increase between 1960 and 1980 in the (naive) estimated dropout risk associated with unmarried parents suggests that this problem is not small.

Another empirical implication (although not one explored here) is that efforts to use longitudinal data to identify the effects on children of the *event* of divorce may be misguided. Divorce results from choices made

[23]Note that Tables 10-13 of Appendix C report index coefficients and *not* probability effects. That is, Table 7 presents $\Delta G(\mathbf{X}\hat{\boldsymbol{\beta}})$ while Tables 10-13 present $\hat{\boldsymbol{\beta}}$. In Tables 10-13, the reported *Pseudo* R^2 is for the first stage marital status equation estimates of the bivariate probit estimate (i.e. from the single equation results used to generate starting values for the two equation estimates).

within marriage; these choices influence child outcomes - and lead to divorce. Longitudinal studies which find little effect of divorce (i.e. they find bad outcomes pre- as well as post-divorce among children who have experienced divorce; see (Cherlin et al. 1991), (Amato and Booth 1996)) do not indicate that family status does not matter. Instead, they tell us about the effect of divorce *in families whose choices lead them to experience divorce.*

Accounting for exit costs, which the theory indicates is responsible for the estimated differences between 1960 and 1980, suggests that the net (i.e. average) benefit of marriage is comparable across time. On average, the additional risk of dropping out of school at age 17 appears to be roughly 7.0%-8.0% for children whose parents are not married. These results seem consistent across various samples for which information is more or less complete, with the notable exception of non-white 17 year olds. Estimates for these children are much noisier.

These results suggest that empirical work on the effects of marital status should account for the incentives within different marital states and the incentives for choosing different states. And, more generally, policy makers should consider incentive effects of marital status when evaluating policies which change the costs of entering and leaving different marital states. Increasing the costs of “bad” states such as divorce and single motherhood may increase the average welfare of children - but will make some women and children strictly worse off. Conversely, welfare programs such as AFDC may encourage child bearing among disadvantaged couples (those in which the father will not, in equilibrium, invest high effort in his children), but will also make leaving bad matches less costly.

A Decision rules

The decision rules of interest are:

1. The woman’s period 1 rejection cutoff is $R_1 = p\theta_W$; her period 2 rejection threshold when matched with probability x to a man who is a low cost type:

$$R(x) = p\theta_W - pxw_H \tag{27}$$

2. The expected value to a woman of a new draw when x is the probability that a new man will be a low cost type is:

$$v(x) = \int_{R(x)}^{\infty} (m - p\theta_W + pxw_H)\, dF(m) \tag{28}$$

3. The marriage ($\underline{m}$) and divorce ($\overline{m}$) thresholds have three terms: the expected value of a new draw ($\alpha v1 + (1-\alpha)v2$), the cost (benefit) of avoiding the effects of this man’s actions (w_H or w_L), and the expected cost of a child next period ($R(x)$):

$$\underline{m} = \alpha v(\mu^S) + (1-\alpha)v(\mu^D) - \frac{w_H}{\beta} + R(1.0) \tag{29}$$

$$\overline{m} = \alpha v(\mu^S) + (1-\alpha)v(\mu^D) - \frac{w_L}{\beta} + R(0.0) \tag{30}$$

The probability of being matched with a *single* man in period 2 is α. The probability that a divorced man is a low cost type is μ^D; α and μ^D are determined endogenously.

4. The man's (high) effort threshold is:

$$E^H = \int_{R(\mu^D)}^{\infty} m dF(m) + \frac{\theta_M^H}{\beta} \tag{31}$$

B Changes in the probability of pregnancy

The *first* period rejection cutoff is $R_1 = p\theta_W$; it increases with p:

$$\frac{dR_1}{dp} = \theta_W > 0 \tag{32}$$

The *second* period rejection cutoff also increases with p:

$$\frac{dR(x)}{dp} = \theta_W - x w_H \tag{33}$$

Since $v(x)$ is defined by indifference, matches at the margin have no value; $\frac{dv(x)}{dp} = 0$.
The marriage threshold ($\underline{m}$) increases with p:

$$\frac{d\underline{m}}{dp} = (\theta_W - w_H) > 0 \tag{34}$$

The woman's divorce cutoff, $\overline{m}$, also increases as p increases

$$\frac{d\overline{m}}{dp} = \theta_W > 0 \tag{35}$$

When children are more likely, the expected costs of a man's bad behavior increases.
The high cost type man's effort threshold falls as p increases:

$$\frac{dE^H}{dp} = -\frac{[R(\mu^D)]^2}{p} f(R(\mu^D)) < 0 \tag{36}$$

As pregnancy becomes more likely, men are more likely to be rejected in a new match ($R(\mu^D)$ increases), so they are less inclined to take their chances on a new draw.

C Additional estimates

	A	*B*	*C*	*D*	*E*	*F*	*G*	*H*	*I*
notmarry	−0.0363	−0.0344					−0.0415	−0.0372	−0.0490
	(0.0135)	(0.0135)					(0.0245)	(0.0225)	(0.0371)
nmcath							0.0279		0.0368
							(0.0908)		(0.0974)
nmdiv								0.0010	0.0022
								(0.0076)	(0.0081)
par_oth	−0.1744	−0.1720					−0.2352	−0.1239	−0.1864
	(0.0198)	(0.0199)					(0.0270)	(0.0240)	(0.0342)
pocath							0.2456		0.1937
							(0.0672)		(0.0711)
podiv								−0.0183	−0.0129
								(0.0056)	(0.0058)
unmarry			−0.0740	−0.1138	−0.0457	−0.0858			
			(0.0121)	(0.0184)	(0.0161)	(0.0247)			
uncath				0.1733		0.1386			
				(0.0583)		(0.0619)			
undiv					−0.0121	−0.0083			
					(0.0048)	(0.0050)			
female	0.0239	0.0227	0.0215	0.0217	0.0217	0.0218	0.0232	0.0230	0.0233
	(0.0072)	(0.0072)	(0.0072)	(0.0072)	(0.0072)	(0.0072)	(0.0072)	(0.0072)	(0.0072)
black	−0.0028	0.0054	0.0005	0.0014	−0.0002	0.0007	0.0060	0.0044	0.0052
	(0.0110)	(0.0114)	(0.0115)	(0.0115)	(0.0115)	(0.0115)	(0.0114)	(0.0114)	(0.0114)
mom_a	−0.3356	−0.3226	−0.3205	−0.3237	−0.3208	−0.3233	−0.3275	−0.3227	−0.3265
	(0.0229)	(0.0236)	(0.0236)	(0.0236)	(0.0236)	(0.0236)	(0.0236)	(0.0236)	(0.0236)
mom_b	−0.1350	−0.1304	−0.1318	−0.1331	−0.1317	−0.1328	−0.1320	−0.1300	−0.1313
	(0.0103)	(0.0104)	(0.0104)	(0.0104)	(0.0104)	(0.0104)	(0.0104)	(0.0104)	(0.0104)
mom_11	−0.0479	−0.0461	−0.0480	−0.0490	−0.0482	−0.0489	−0.0470	−0.0457	−0.0465
	(0.0192)	(0.0191)	(0.0192)	(0.0192)	(0.0192)	(0.0192)	(0.0192)	(0.0191)	(0.0192)
mom_c3	0.1059	0.1024	0.1041	0.1035	0.1039	0.1035	0.1021	0.1023	0.1021
	(0.0147)	(0.0148)	(0.0147)	(0.0148)	(0.0147)	(0.0147)	(0.0148)	(0.0148)	(0.0148)
mom_c4p	0.1243	0.1247	0.1239	0.1239	0.1241	0.1241	0.1247	0.1250	0.1249
	(0.0179)	(0.0177)	(0.0177)	(0.0177)	(0.0177)	(0.0177)	(0.0177)	(0.0176)	(0.0177)
medmiss	−0.2976	−0.2984	−0.3923	−0.3921	−0.3908	−0.3911	−0.2978	−0.2957	−0.2960
	(0.0232)	(0.0234)	(0.0188)	(0.0188)	(0.0189)	(0.0189)	(0.0234)	(0.0234)	(0.0235)
State dummies	*No*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*
N	14,292	14,292	14,292	14,292	14,292	14,292	14,292	14,292	14,292
Pseudo R^2	0.1241	0.1336	0.1306	0.1312	0.1310	0.1314	0.1345	0.1344	0.1348

Table 8: Alternative ("naive") estimates: 1960; probability effects

	A	*B*	*C*	*D*	*E*	*F*	*G*	*H*	*I*
notmarry	−0.0716	−0.0723					−0.0634	−0.0737	−0.0519
	(0.0082)	(0.0082)					(0.0159)	(0.0240)	(0.0371)
nmcath							−0.0355		−0.0475
							(0.0554)		(0.0656)
nmdiv								0.0002	−0.0013
								(0.0032)	(0.0038)
par_oth	−0.1613	−0.1646					−0.1644	−0.1567	−0.1487
	(0.0134)	(0.0135)					(0.0207)	(0.0278)	(0.0427)
pocath							−0.0012		−0.0142
							(0.0516)		(0.0605)
podiv								−0.0010	−0.0014
								(0.0030)	(0.0035)
unmarry			−0.0887	−0.0856	−0.0879	−0.0791			
			(0.0072)	(0.0124)	(0.0178)	(0.0282)			
uncath				−0.0130		−0.0198			
				(0.0424)		(0.0501)			
undiv					−0.0001	−0.0007			
					(0.0025)	(0.0029)			
female	0.0160	0.0154	0.0140	0.0140	0.0140	0.0140	0.0154	0.0154	0.0155
	(0.0048)	(0.0048)	(0.0048)	(0.0048)	(0.0048)	(0.0048)	(0.0048)	(0.0048)	(0.0048)
black	0.0359	0.0408	0.0369	0.0369	0.0369	0.0369	0.0408	0.0408	0.0408
	(0.0059)	(0.0059)	(0.0060)	(0.0060)	(0.0060)	(0.0060)	(0.0059)	(0.0059)	(0.0059)
mom_a	−0.2129	−0.2020	−0.2072	−0.2073	−0.2072	−0.2074	−0.2023	−0.2021	−0.2024
	(0.0278)	(0.0277)	(0.0279)	(0.0279)	(0.0279)	(0.0279)	(0.0277)	(0.0277)	(0.0278)
mom_b	−0.1350	−0.1293	−0.1304	−0.1303	−0.1304	−0.1303	−0.1293	−0.1293	−0.1293
	(0.0086)	(0.0086)	(0.0086)	(0.0086)	(0.0086)	(0.0086)	(0.0086)	(0.0086)	(0.0086)
mom_11	−0.0819	−0.0771	−0.0775	−0.0774	−0.0775	−0.0774	−0.0771	−0.0772	−0.0771
	(0.0139)	(0.0138)	(0.0138)	(0.0138)	(0.0138)	(0.0138)	(0.0138)	(0.0138)	(0.0138)
mom_c3	0.0342	0.0345	0.0352	0.0353	0.0352	0.0352	0.0345	0.0344	0.0344
	(0.0079)	(0.0078)	(0.0078)	(0.0078)	(0.0078)	(0.0078)	(0.0078)	(0.0078)	(0.0078)
mom_c4p	0.0669	0.0658	0.0660	0.0660	0.0660	0.0660	0.0658	0.0658	0.0658
	(0.0084)	(0.0083)	(0.0083)	(0.0083)	(0.0083)	(0.0083)	(0.0083)	(0.0083)	(0.0083)
medmiss	−0.1884	−0.1782	−0.2400	−0.2400	−0.2400	−0.2400	−0.1779	−0.1778	−0.1777
	(0.0148)	(0.0147)	(0.0131)	(0.0131)	(0.0131)	(0.0131)	(0.0147)	(0.0147)	(0.0147)
State dummies	*No*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*
N	20,475	20,475	20,475	20,475	20,475	20,475	20,475	20,475	20,475
Pseudo R^2	0.1219	0.1306	0.1276	0.1276	0.1276	0.1276	0.1307	0.1307	0.1307

Table 9: Alternative ("naive") estimates: 1980; probability effects

Sample: All (with known parents' marital status)

	1960				1980			
	Means	Naive	BV	IV	Means	Naive	BV	IV
notmarry	.1123	−.1670	−.2746	−.2719	.1845	−.3538	−.3618	−.4072
		(.0447)	(.0651)	(.0673)		(.0326)	(.0445)	(.0435)
female	.4808	.1294	.1309	.1288	.4732	.1515	.1516	.1527
		(.0278)	(.0278)	(.0278)		(.0259)	(.0259)	(.0260)
black	.0891	−.1110	−.1229	−.0991	.1230	.0725	.0722	.08551
		(.0457)	(.0459)	(.0460)		(.0387)	(.0387)	(.0394)
mom_a	.0490	−.9264	−.9054	−.9374	.0164	−.7897	−.7852	−.8232
		(.0605)	(.0611)	(.0606)		(.0811)	(.0829)	(.0812)
mom_b	.4508	−.4346	−.4302	−.4403	.2142	−.5163	−.5155	−.5204
		(.0337)	(.0338)	(.0338)		(.0312)	(.0313)	(.0312)
mom_11	.0683	−.1134	−.1123	−.1152	.0618	−.2890	−.2888	−.2876
		(.0611)	(.0611)	(.0612)		(.0516)	(.0516)	(.0517)
mom_c3	.0810	.4168	.4190	.4189	.1446	.1768	.1767	.1772
		(.0715)	(.0715)	(.0715)		(.0442)	(.0442)	(.0442)
mom_c4p	.0418	.6046	.6101	.6084	.0912	.3883	.3886	.3874
		(.1076)	(.1075)	(.1077)		(.0609)	(.0609)	(.0609)
med_miss	.0149	−.4797	−.5040	−.3673	.0236	−.2312	−.2407	−.2347
		(.1109)	(.1110)	(.1200)		(.0778)	(.0857)	(.0814)
constant	1.0000	1.1293	1.1382	1.1487	1.0000	1.3382	1.3397	1.3558
		(.0307)	(.0309)	(.0310)		(.0239)	(.0245)	(.0242)
mmig5	.1487		1.4216		.5717		.4274	
			(.0590)				(.0324)	
mfborn	.0850		.6183		.1183		.3961	
			(.0513)				(.0332)	
dfborn	.1415		2.4801		.2090		2.2809	
			(.0552)				(.0305)	
constant	1.0000		−2.5872		1.0000		−2.0872	
			(.0497)				(.0318)	
$\hat{\rho}$			0.11				0.01	
Pseudo R^2			0.51				0.46	
χ^2 (over-ID)				9.7				8.4
N	11,773	11,773	11,773	11,773	17,398	17,398	17,398	17,398

Table 10: Coefficients (for $\mathbf{X}\hat{\boldsymbol{\beta}}$ index, *not* probability effects) and s.e. in ().

Sample: Mother's education known

	1960				1980			
	Means	Naive	BV	IV	Means	Naive	BV	IV
notmarry	.0988	−.1680	−.2678	−.2644	.1648	−.3512	−.3733	−.4088
		(.0447)	(.0671)	(.0671)		(.0327)	(.0418)	(.0433)
female	.4811	.1276	.1289	.1269	.4740	.1437	.1442	.1450
		(.0281)	(.0281)	(.0281)		(.0264)	(.0264)	(.0264)
black	.0880	−.0996	−.1089	−.0883	.1227	.0599	.0602	.0742
		(.0465)	(.0467)	(.0468)		(.0392)	(.0392)	(.0399)
mom_a	.0498	−.9291	−.9122	−.9407	.0168	−.7891	−.7797	−.8219
		(.0605)	(.0610)	(.0607)		(.0811)	(.0819)	(.0812)
mom_b	.4576	−.4355	−.4316	−.4414	.2194	−.5150	−.5133	−.5190
		(.0337)	(.0338)	(.0338)		(.0312)	(.0312)	(.0312)
mom_11	.0693	−.1138	−.1125	−.1157	.0633	−.2873	−.2867	−.2859
		(.0611)	(.0611)	(.0612)		(.0516)	(.0517)	(.0517)
mom_c3	.0823	.4168	.4184	.4186	.1480	.1769	.1768	.1773
		(.0715)	(.0715)	(.0715)		(.0442)	(.0442)	(.0442)
mom_c4p	.0424	.6044	.6090	.6079	.0934	.3882	.3886	.3873
		(.1076)	(.1075)	(.1076)		(.0609)	(.0609)	(.0609)
constant	1.0000	1.1297	1.1364	1.1486	1.0000	1.3422	1.3450	1.3605
		(.0308)	(.0309)	(.0310)		(.0240)	(.0242)	(.0243)
mmig5	.1358		1.2540		.5614		.3162	
			(.0630)				(.0361)	
mfborn	.0711		.2129		.0970		−.4371	
			(.0568)				(.0402)	
dfborn	.1424		2.63776		.2123		2.7766	
			(.0581)				(.0358)	
constant	1.0000		−2.6277		1.0000		−2.2492	
			(.0524)				(.0363)	
$\hat{\rho}$			0.10				0.03	
Pseudo R^2			0.52				0.57	
χ^2 (over-ID)				9.7				8.0
N	11,597	11,597	11,597	11,597	16,988	16,988	16,988	16,988

Table 11: Coefficients (for $\mathbf{X}\hat{\boldsymbol{\beta}}$ index, *not* probability effects) and s.e. in ().

Sample: White only

	1960				1980			
	Means	Naive	BV	IV	Means	Naive	BV	IV
notmarry	.1019	−.1743	−.2714	−.2705	.1527	−.4084	−.4093	−.4504
		(.0496)	(.0741)	(.0773)		(.0374)	(.0530)	(.0519)
female	.4813	.1346	.1358	.1337	.4690	.1445	.1445	.1457
		(.0296)	(.0296)	(.0297)		(.0285)	(.0285)	(.0285)
mom_a	.0377	−1.0235	−1.0021	−1.0383	.0153	−.8643	−.8638	−.8996
		(.0690)	(.0700)	(.0692)		(.0900)	(.0929)	(.0900)
mom_b	.4351	−.4421	−.4387	−.4465	.1925	−.5628	−.5627	−.5673
		(.0349)	(.0349)	(.0349)		(.0341)	(.0343)	(.0342)
mom_11	.0699	−.0949	−.0941	−.0963	.0548	−.2870	−.2870	−.2858
		(.0640)	(.0640)	(.0640)		(.0588)	(.0588)	(.0588)
mom_c3	.0869	.4494	.4514	.4508	.1493	.2178	.2178	.2180
		(.0744)	(.0744)	(.0744)		(.0484)	(.0484)	(.0484)
mom_c4p	.0441	.6260	.6295	.6299	.0968	.4104	.4104	.4100
		(.1125)	(.1124)	(.1125)		(.0655)	(.0655)	(.0655)
medmiss	.0136	−.4305	−.4510	−.3539	.0233	−.2611	−.2618	−.2703
		(.1235)	(.1238)	(.1352)		(.0840)	(.0906)	(.0894)
constant	1.0000	1.1313	1.1379	1.1461	1.0000	1.3627	1.3628	1.3778
		(.0316)	(.0318)	(.0320)		(.0255)	(.0261)	(.0259)
mmig5	.1511		1.3595		.5760		.4824	
			(.0615)				(.0371)	
mfborn	.0866		.5806		.1142		.5656	
			(.0541)				(.0369)	
dfborn	.1379		2.3841		.1751		2.1789	
			(.0580)				(.0347)	
constant	1.0000		−2.5964		1.0000		−2.1996	
			(.0523)				(.0368)	
$\hat{\rho}$			0.10				0.00	
Pseudo R^2			0.49				0.44	
χ^2 (over-ID)				7.3				6.0
N	10,627	10,627	10,627	10,627	14,860	14,860	14,860	14,860

Table 12: Coefficients (for $\mathbf{X}\hat{\boldsymbol{\beta}}$ index, *not* probability effects) and s.e. in ().

Sample: Non-white only

	1960				1980			
	Means	Naive	BV	IV	Means	Naive	BV	IV
notmarry	.2086	−.1308	−.2548	−.1807	.3708	−.1841	−.2129	−.3079
		(.1039)	(.1220)	(.1196)		(.0662)	(.0871)	(.0794)
female	.4756	.0970	.0990	.0969	.4976	.1976	.1969	.1991
		(.0814)	(.0811)	(.0819)		(.0635)	(.0635)	(.0636)
black	.9154	−.2786	−.3948	−.2456	.8432	.0126	−.0033	.1416
		(.1591)	(.1703)	(.1593)		(.0904)	(.0956)	(.0916)
mom_a	.1536	−.6237	−.6239	−.6606	.0232	−.4588	−.4590	−.4697
		(.1598)	(.1590)	(.1608)		(.1927)	(.1926)	(.1943)
mom_b	.5960	−.3214	−.3213	−.3577	.3412	−.3082	−.3080	−.3018
		(.1376)	(.1369)	(.1385)		(.0770)	(.0770)	(.0772)
mom_11	.0532	−.2358	−.2340	−.2490	.1024	−.2271	−.2272	−.2151
		(.2186)	(.2179)	(.2202)		(.1111)	(.1111)	(.1116)
mom_c3	.0262	−.1034	−.1109	−.1231	.1170	−.0327	−.0337	−.0230
		(.2911)	(.2898)	(.2936)		(.1128)	(.1127)	(.1131)
mom_c4p	.0201	.3677	.4711	.3242	.0579	.2281	.2295	.2082
		(.3921)	(.3952)	(.3963)		(.1697)	(.1697)	(.1700)
medmiss	.0279	−.6173	−.6475	−.4736	.0252	.2152	.1468	.0930
		(.2705)	(.2672)	(.2751)		(.2304)	(.2669)	(.2315)
constant	1.0000	1.2050	1.3344	1.2516	1.0000	1.2090	1.2349	1.1118
		(.1923)	(.2030)	(.1930)		(.0996)	(.1115)	(.1003)
mmig5	.1265		5.8713		.5465		.3223	
			(588.2801)				(.0726)	
mfborn	.0707		1.2605		.1422		−.2703	
			(.2025)				(.0793)	
dfborn	.1754		7.3422		.4074		2.4635	
			(588.2801)				(.0693)	
constant	1.0000		−6.4764		1.0000		−1.6599	
			(588.2801)				(.0693)	
$\hat{\rho}$			0.30				0.05	
Pseudo R^2			0.74				0.50	
χ^2 (over-ID)				8.3				12.1
N	1,146	1,146	1,146	1,146	2,538	2,538	2,538	2,538

Table 13: Coefficients (for $\mathbf{X}\hat{\boldsymbol{\beta}}$ index, *not* probability effects) and s.e. in ().

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Revision: 1.13
Date: July 19, 1999, 13:53:40

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