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EDUCATION AND CONTRACEPTIVE CHOICE: CONDITIONAL LOGIT

AND STRUCTURAL MODELS

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I. Introduction

In recent years economists have been devoting attention to measuring the effects of schooling on skills which are not traded in the market and thus not directly reflected in wage earnings (Welch, 1970; Huffman, 1975; Leibowitz, 1974; Rosenzweig, 1977). The more rapid adoption of a new contraceptive technology, the pill, by more educated women in the early 1960's, documented by Ryder (1972), has been cited as one example of the role of schooling in improving allocative abilities in the household sector (Huffman, 1976). Examination of the available evidence on the relationship between schooling and contraceptive choice, however, suggests that the hypothesized advantages of schooling in lowering the costs associated with contraceptive decisions is not well established. Unless it is assumed that the pill dominated all other contraceptive techniques, that it was efficient for all women to adopt the pill independently of family size desires, one cannot interpret the correlations between female education and pill use unambiguously. For instance, if less educated women tend to want larger families because, say, their time costs are lower, adoption of a contraceptive device such as the pill may not be optimal for them; the overall positive association between schooling and pill adoption may thus simply reflect the underlying determinants of the demand for children.

Michael (1973) and Michael and Willis (1975), based on an economic model of contraceptive choice, have estimated the determinants of contraceptive use by women within parity groups and conditional on measures of birth intentions. The estimated associations between the schooling levels of women and the use of the pill or other efficient techniques in these studies are generally much weaker than the uncondi-

tional (on parity or intentions) or "reduced form" correlations, casting
some doubt on the existence of true cost or allocative skill effects of
schooling. However, both parity and total birth intentions are treated
as if they were exogenous variables, which may mean that these results
are not unbiased; moreover, only women in closed birth intervals are
included in the samples, resulting in selectivity bias.[1] Finally, no
studies have taken into account the differential time costs associated
with using contraceptive devices, which may be strongly correlated with
schooling.

As a consequence of the lack of knowledge concerning the magnitude
of the ceteris paribus effect of education on contraceptive choice, a
parallel ambiguity characterizes empirical applications of house-
hold models of fertility choice in which the level of the mother's
schooling is utilized as a proxy for the value of her time (Willis,
1973; Ben-Porath, 1973). If education lowers the (marginal) cost of
controlling family size due to allocation efficiency or informational
effects, the negative coefficient on the wife's schooling in regressions
on parity would be an overestimate of the substitution effect arising
from the presumed time-intensity of children implied by this framework
and indeed may wholly reflect differences in contraceptive costs.

In this paper we utilize the theory of contraceptive choice to
formulate tests for the importance of schooling net of wage costs in
contraceptive decisions using birth intentions information and apply
them to data from the 1970 NFS, which reflect the availability of
both the pill and more recent innovations in contraceptive technology.
We demonstrate that knowledge of the schooling-contraceptive choice

relationship conditional on future birth intentions in the open interval
is required to ascertain the magnitude of true schooling effect on
adoptive efficiency but that least squares cannot be used to obtain
consistent estimates of this structural relationship. However, it is
shown that the direction of the least squares bias embedded in the birth
intentions coefficient in the contraception structural equation can
provide information on the extent to which the reported birth intentions
of women reflect contraceptive costs, without the need for estimates
from a structural equation determining birth intentions. We also show
that the relationship between reduced form (unconditional) and structural
estimates of the effect of schooling on contraceptive choice, which
depends crucially on the theoretically unsigned relationship between
birth intentions and education, can be ascertained from the direction
of the least squares bias in the structural schooling coefficients.
Finally, we demonstrate why estimates of the relationship between school-
ing and contraceptive choice condtional on parity, even if consistent,
cannot resolve the issue of whether education augments the ability to
adopt new contraceptive technologies or lowers costs.

The empirical results support the hypothesis that women with higher
levels of schooling tended to adopt more readily the newer contraceptive
technologies, given their demand for children, and suggest that both
reduced form and (inconsistent) conditional estimates of the association
between female schooling and pill use in prior studies understate the
"true" structural relation between schooling and contraceptive efficiency,
whether or not time costs associated with the wage rates of women are
accounted for. Tests applied to residuals indicate that birth intentions
cannot be treated as an exogenous determinant of contraceptive choice

but the estimated sign of the bias in the OLS estimators suggests that
this is due more to the presence of serially correlated variables re-
flecting unobserved factors associated with contraceptive choice than
to the influence of anticipated contraceptive costs on the reported
intentions of women. Categorization of contraceptive methods on the
basis of cost characteristics and use of the log-linear model in con-
junction with the structural estimates also suggests that women with
higher levels of schooling who do not choose the newer birth control
methods tend to use the traditional contraceptive techniques more effi-
ciently but that women with higher values of time are less likely to be
using the newer techniques. A decomposition of the gross relationship
between completed family size and female schooling based on the struc-
tural parameter estimates thus suggests that differences in contra-
ceptive methods associated with the schooling attainment of the wife
account for approximately 50 to 90 percent of the estimated fertility
effects of female education reported in a number of fertility studies
utilizing U. S. data from the mid 60's to 1970.

In section 2 we briefly review the theory of optimal contra-
ceptive choice. Section 3 contains the discussion of the econometric
tests which are then applied to the data in section 4. In the final
section we review the results obtained in terms of their quantitative
significance in determining fertility differentials and their implications
for intergenerational changes in the distribution of earnings.

II. Theory

The main assumption of the theory of contraceptive choice in its deterministic form (Michael, 1973; Michael and Willis, 1975)[2] is that the achievement of a given family size goal, however formed, necessitates behavior aimed at reducing the flow of births that would occur, given levels of coition, in the absence of fertility control. Consider a married woman aged a who has P_a children. The number of additional births n for a desired completed family size of P_T with costless fertility control is P_T-P_a Let the probability of a birth at each age x conditional on coital frequency be given by n(x). Then the total number of additional births to be averted, n_A, is

$$(1) \quad n_A = \int_a^\beta n(x)dx - n = \bar{e} \int_a^\beta n(x)dx = \bar{e} \, N_a^*$$

where $\bar{e}$ is the average contraceptive efficiency required to achieve n_A $(o \leq \bar{e} \leq 1)$ and n(x) = o for x $\geq$ β. The choice problem is thus to select a contraceptive technique which minimizes the total cost over β-a years of achieving $\bar{e}$(or n).

To bring out the major implications of the analysis, assume that contraceptive techniques can be divided into two classes according to their cost characteristics--those with relatively high fixed costs whose use is unrelated, at the margin, to coition (pill, IUD, sterilization) and those with costs (psychic and pecuniary) which depend on the frequency of coition and the number of births to be averted (condom, diagraphm). The fixed costs associated with the first techniques include the cost of acquiring and decoding the available information about these devices--the value of time

and goods used in the search for relevant information, including consultations with medical personnel--and costs of acquisition, with part of the cost of information being a positive function of the date of introduction of the technique. The acquisition costs also contain time and direct expenditure components--visits to doctors or clinics for prescriptions and/or installation and the monitoring of both contraceptive performance (IUD) and possible health side-effects (pill, sterilization). Variable costs associated with coition and/or the length of use would include both psychic costs and those costs associated with perceived health hazards.

The relative importance of fixed and variable costs characterizing contraceptive methods will, in general, depend both upon societal institutional structures and technological characteristics of the methods. In the United States in 1970, to which the data we shall use refer, adoption of the pill, for example, necessitated at least one visit to a doctor or clinic and the health side-effects of prolonged pill consumption (a variable cost) were not yet well-established so that fixed costs of the pill clearly dominated. In other contexts, such as in some developing countries where the pill is available without prescription or in the U.S. when pill health hazards became known, the rankings of methods by fixed relative to variable cost as well as the correspondence between contraceptive efficiency ($\bar{e}$) and fixed costs may differ. Without imposing some cost structure on the techniques, however, little can be said about contraceptive behavior.[3]

Figure 1 graphs the total cost curves of two techniques against n_A, for women aged a, where e_1 is a coition-dependent and e_2 a high fixed cost technique. As long as the fixed costs of e_2 are less than $\bar{e}$ and the slope
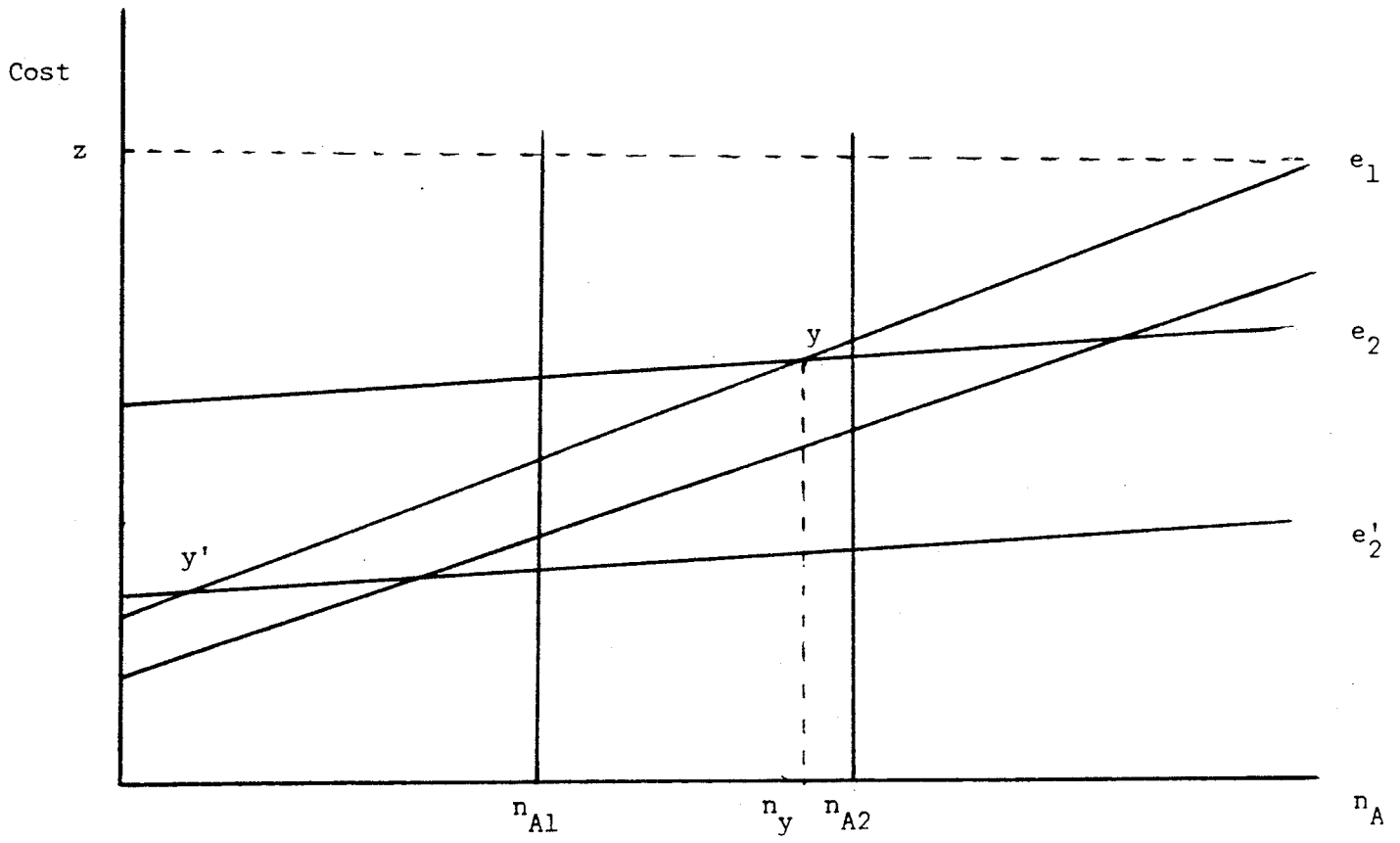
Figure 1.

of the e_2 curve is less than that of the e_1 curve there will be a switching point; such as at y, associated with a particular n_A (n_y) above which it is cost-efficient to incur the larger fixed costs associated with e_2. Thus a household which wishes to avert n_{A1} births selects technique e_1; a household choosing to avert n_{A2} births (a lower n), given the same N_a^*, utilizes technique e_2.

The actual number of births the household will be observed to avoid (actual additional births) for given n will then depend on the control technique selected. If e_1 is the least-cost choice, then the cost of averting a birth at the margin is now higher, by the slope of the e_1 curve at n_{A1}, leading to a greater number of additional births than would occur if fertility control were costless. Thus, among households with the same n, users of the e_1 techniques will display greater fertility rates.[4]

The cost-minimizing framework suggests at least three channels through which schooling, which will be assumed to be that of the wife, may affect the choice of contraceptives. If education augments skills in decoding available information and adopting new techniques, (Welch, 1970; Schultz, 1974) or is associated with greater access to sources of information then the 'fixed' costs of the newer devices will be lower for the more educated, as depicted in Figure 1 by the shift in curve e_2 to e_2^1. The switching point will thus be lowered (to y') and given that the level of desired additions to family size remains the same, the household which formerly used e_1, given a higher level of schooling, will choose technique e_2.

However, time costs embedded in the fixed costs of e_2 will also increase with schooling attainment, thus raising the relative price of the time-intensives effective techniques, if education increases female wages. Indeed, consistent with the price of time hypothesis, Acton (1975) has found that high-wage women purchase less physician and clinical health services than low-wage women. This effect of schooling thus may partly or wholly obscure the informational or allocative ability effects of schooling.

The third possible effect of schooling is that of increasing the efficiency with which a given technique is used; i.e., schooling may lower the <u>marginal</u> cost of the e_1 techniques, as depicted in Figure 1 by a downward rotation of the e_1 cost curve. Controlling for the price of time, women with higher levels of schooling and the same n_A will be more likely to use an e_1 technique relative to none if education augments use-effectiveness, and given that an e_2 technique is not being used, and will be more likely to use an e_2 technique if education lowers information costs or improves dynamic allocative skills.

III. Estimation Framework

It is clear from Figure 1 that the 'efficiency' or structural effects of schooling can only be identified if desired additions to the existing stock of children are known and held constant when e_2 techniques are not cost efficient over all n_A. As long as there exists a switching point, if schooling shifts both the cost curves and n_A the observed unconditional or 'reduced form' relationship between schooling and adoption of, say e_2, can be of any sign unless further restrictions are imposed on the determination of expected fertility. While the theory of contraceptive choice, even in its most rudimentary form, thus clearly calls for estimates from a simultaneous equation system, no prior studies of the demand for contraceptives provide estimates of the relevant structural equations. We now construct a framework for testing for the allocative effects of schooling in contraceptive choice based on the theoretical model and use it to discuss evidence on the schooling-contraception adoption association appearing in the literature.

Consider the following two linearized stochastic structural equations as representing estimating equations suggested by the theory:[5]

$$(2) \quad n^* = \alpha_0 + \alpha_1 E_f + \alpha_2 e_i + \alpha_3 E_m + \alpha_4 F + \Sigma \alpha_j z_j + \epsilon_1$$

$$(3) \quad e = \beta_0 + \beta_1 E_f + \beta_2 n^* + \beta_3 E_m + \beta_4 F + \beta_5 \pi_e + \epsilon_2$$

where n^* is the desired number of additional children, obtained, say, from a survey question on expected or intended births, e is a variable representing the adoption of the newer contraceptive techniques, E_f = female schooling attainment, E_m = male schooling, F = husband's income,

π_e = money price index of contraceptive devices and the z_j are variables which are assumed to influence fertility goals but not the choice of contraceptives. Each equation contains explanatory variables, in addition to the unspecified z_j, which are meant to be representative of those used in prior analyses of contraceptive choice, and which are assumed to be uncorrelated with the two error terms. The latter may, however, be themselves correlated. As specified, equation (2) is just identified if data on market"prices" of contraceptives are available; (3) is overidentified.

The theoretical framework suggests that $\alpha_2 \neq 0$ if n* is the intended number of additional children given the costs of control, $\alpha_2 = 0$ if n* is the number of additional children desired conditional on costless control, $\beta_2 > 0$ if there is a switching point, and $\beta_1 > 0$ if the educational attainment of women lowers contraceptive costs associated with information (we will concentrate on the effects of women's education, on the presumption that women bear more of the cost of contraceptive decisions than men, but the same considerations apply to the husband's schooling as well. The reduced form equations,-corresponding to (2) and (3), are:

$$(4) \quad n^* = \gamma_{10} + \gamma_{11}E_f + \gamma_{12}E_m + \gamma_{13}F + \gamma_{14}\pi_e + \Sigma\gamma_{1j}z_j + u_1$$

$$(5) \quad e_i = \gamma_{20} + \gamma_{21}E_f + \gamma_{22}E_m + \gamma_{23}F + \gamma_{24}\pi_e + \Sigma\gamma_{2j}z_j + u_2$$

Most of the evidence on the schooling-contraceptive use association is based on zero-order correlations or reduced form equations such as (5)

in which other exogenous household characteristics are netted out (Ryder, 1972).
To see that the estimation of (5), without (4) or (3), does not provide
any evidence on the role of schooling in fostering the adoption of the
new effective techniques, note that the coefficients γ_{11} and γ_{21}, the
reduced-form effects of female schooling on birth intentions and contra-
ception, in terms of the structural parameters, are:

$$\gamma_{11} = (\alpha_1 + \alpha_2\beta_1) / \Psi$$

$$\gamma_{21} = (\beta_1 + \beta_2\alpha_1) / \Psi$$

where $\Psi = (1 - \beta_2\alpha_2)$

Even with the imposition of the sign restrictions implied by the stability
conditions (which imply that $\psi > 0$) and the theoretical analysis, the sign
of γ_{21} is unknown, depending on the sign of α_1, the effect of education on
the demand for additional children (birth intentions) in structural equa-
tion (2), which is not indicated by the theory of contraceptive choice.

It has usually been implicitly assumed that n* and E_f are negatively
correlated ($\alpha_1 < 0$), perhaps based on the available evidence on female edu-
cation and completed family size (or parity) which shows these variables
to be negatively correlated (see Ben-Porath, 1973; Willis, 1973; for ex-
amples). Thus the positive sign of γ_{21}, observed in all studies of pill
use (for example Ryder, 1972), is assumed to be an _overestimate_ of the
structural (or efficiency) effect of schooling (Westoff and Ryder, 1977;
Vaughan et al., 1977). Indeed, with $\alpha_1 < 0$, β_1 may be zero while $\gamma_{21} > 0$.
None of the studies of contraceptive choice, however, report structural

estimates of the determinants of birth intentions to provide evidence

on this assumption. Moreover, it is not obvious that birth intentions

and female schooling are inversely correlated just because P_T and E_f

are negatively related. To see this, rewrite (1) as:

$$(6) \quad n = P_T - (1-\bar{e}') \int_m^a n(x)dx$$

$$= P_T - (1-\bar{e}')[N(a) - N(m)]$$

where $\bar{e}'$ is the average level of contraceptive effectiveness up to age a,

m is the age of marriage and N(x) is cumulative fertility at age x in the

absence of control. The total effect on birth intentions of a change in

E_f is thus:

$$(7) \quad \frac{\partial n}{\partial E_f} = \frac{\partial P_T}{\partial E_f} + [N(a) - N(m)] \frac{\partial \bar{e}'}{\partial E_f} - \frac{\partial N(m)}{\partial E_f} \bar{e}' = \alpha_1$$

While the first term in (7), the completed family size effect, is negative

(by assumption), the other terms, the contraceptive efficiency and marriage

effects, are most likely positive since education is presumed to increase

contraceptive efficiency while age at marriage and female schooling tend

to be positively correlated (for evidence, see Keeley, 1976).[6] Thus among

women of the same age, the more educated, even if their ultimate family

size goals are lower, may desire a greater number of additional children

than their less educated counterparts since their existing stock of children

(P_a) may be lower due to both delayed marriage and more efficient

contraception. If these latter effects dominate, the reduced-form school-
ing coefficient γ_{21} is then an underestimate rather than overestimate of
the "true" or structural effect of schooling on contraceptive efficiency.

While the joint reduced form estimates can be used as a rough test
of the theory, since the theoretical restrictions rule out $\gamma_{11} < 0$ and
$\gamma_{21} > 0$ as a joint result, estimation of β_1 in structural equation (3),
which makes use of birth expectations data, is a more direct test. One
potential problem, however, is that it is not clear whether women who
report birth intentions take into account all the costs, particularly
contraceptive costs, pertaining to children in their responses. Estimation
of α_2 in (2) would provide direct evidence on this question; however,
π_e, the price variable which identifies the birth intentions equation, is
not likely to vary significantly in the cross-section and/or is difficult
to measure. We will thus assume that (2) is under-identified ($\text{var}(\pi_e) = 0$)
but will show that the estimation of equation (3) also sheds some light
on this issue.

Perhaps because of the questionable nature of expectations data,
however, many estimates of the correlation between schooling and con-
tractive usage have been reported conditional on parity, (as in Michael
and Willis, 1975). Expression (7), demonstrates that such estimates,
even if obtained using appropriate techniques, are weak evidence
of the allocative ability effect of schooling. The assumption that P_T
and schooling are negatively correlated implies that within parity groups
n and schooling will be negatively associated. Thus, if it is cost-
efficient to use the coition-dependent techniques over lower values of

n_{al}, i.e., $\beta_2 < 0$, schooling and contraceptive efficiency will be positively

related even if schooling does not affect contraceptive choice directly.

The 'structural' estimates obtained by controlling for parity thus

must overstate the allocative effect of education on contraceptive

costs; moreover, the effect of parity itself on contraception is not

'predicted' by the theory of contraceptive choice unless the sign of

the relationship between parity and n, given by (6) is specified.

The relationships between parity, birth intentions and contra-

ceptive choice, given by equations (2), (3) and (6), also imply that the

use of (ordinary) least squares to estimate the effect of schooling on

contraceptive use conditional on birth intentions, as in (Michael, 1975), will

result in inconsistent parameter estimates. Moreover, the source of the

estimation problem, the correlation between n* and E_2, which will bias

all the coefficients in (3), has two components and cannot be signed a priori:

First, birth intentions, as reported by women, may reflect the cost of

the contraceptive being used. If women who choose the techniques with

relatively large marginal costs consequently anticipate higher numbers

of births and report these as intentions, the relationship between n*

and e in (3) will be negatively biased.

However, the disturbances in (2) and (3) may also be correlated.

Indeed, expression (6) suggests that the sign of the error covariance

is likely to be positive--among observationally identical women, those

who face lower contraceptive costs throughout their childbearing years will be

more likely to have married later (if contraceptive "failure" leads to

marriage) and to have contracepted more efficiently once married; "permanently"

more efficient contraceptors will thus have lower parities at every

age and be more likely to be expecting more children than otherwise

identical women facing higher contraceptive costs or less able to adopt

the newer, more efficient techniques. The positive covariance of ϵ_1 and

ϵ_2 will thus bias upward the coefficent of n* in (3).

More rigorously, it can be shown that

$$(8) \quad \underset{N\to\infty}{\text{plim}} \; (\hat{\beta}_2 - \beta_2) = \frac{(1 - b_{EH}\beta_{HE.F} - b_{EF}\beta_{FE.H})(1 - \rho^2_{HF})}{\psi\phi} \cdot \frac{[\alpha_2 \text{var}(\epsilon_1) + \text{cov}(\epsilon_1,\epsilon_2)]}{\text{var}(E)}.$$

where ϕ is the determinant of a positive definite matrix ($\phi > 0$), ρ^2_{HF} is
the squared simple correlation coefficient between husband's schooling

and income, b_{EH} and b_{EF} are the simple regression coefficients of female

schooling on male schooling and income respectively, and $\beta_{HE.F}$ and $\beta_{FE.H}$

are the partial regression coefficients of male schooling on female schooling

given income, and income on female schooling holding constant male education.

Since the first term in parentheses is a principal minor of the positive

definite matrix ϕ , and must therefore be positive, we see that the sign

of (8) depends on the sign of the expression in brackets, which in turn

depends on (i) the sign of α_2, which is negative if intentions take into

account contraceptive costs, and (ii) the sign of the error covariance

term, which we have shown is likely to be positive.

The existence of offsetting components in the bias characterizing

the OLS estimator of β_2, however, is fortuitous since the sign of the

observed bias in β_2 will provide information on the relative quantita-

tive importance of the existence of differential contraceptive costs in

influencing the reported birth intentions of married women (the magnitude

of α_2), estimates of which would be difficult to obtain directly given
that exogenous instruments which would identify structural equation (2)
are unlikely to be found in most cross-sectional data sets. Thus if the
bias in β_2 is negative, this would be evidence that women do take into
account contraceptive costs when reporting their fertility expectations--
birth intentions, barring changes in plans, could therefore be used to
accurately predict subsequent fertility. A positive bias in β_2
would suggest that women <u>may</u> tend to ignore or underestimate such costs
in their reported intentions; if so, actual fertility will exceed "predicted"
fertility, particularly for (less educated) women facing higher contra-
ceptive costs.[7]

The bias in the OLS schooling coefficient β_1 in (3), given by (9),
however, depends not only on the relative magnitudes of α_2 and the error
covariance term but on the sign of the relationship between female schooling
and birth intentions, the important determinant of whether the reduced-form
schooling effect on contraceptive choice was an over or under-estimate of
the structural relationship.

$$(9) \quad \plim_{N \to \infty} (\hat{\beta}_1 - \beta_1) = \frac{-(b_{En} - b_{EH}\beta_{Hn.F} - b_{EF}\beta_{Fn.H})(1-\rho^2_{HF})}{\psi\phi} \ [\ \alpha_2 \text{var}(\varepsilon_1) + \text{Cov}(\varepsilon_1,\varepsilon_2)] \text{var}(E$$

where b_{En} is the simple regression coefficient of female schooling on
birth intentions and $\beta_{Hn.F}$ and $\beta_{Fn.H}$ are the partial regression coefficients
of male schooling and income on intentions, holding constant income and male
schooling respectively. Since, given plausible values of the coefficients,
the sign of the first term in (9) will be dominated by b_{En}, the relationship

between the OLS biases in β_1 and β_2 can provide information on the direction of the association between schooling and birth intentions. For example, if $\hat{\beta}_2 - \beta_2 > 0$ and $\hat{\beta}_1 - \beta_1 < 0$, then $b_{En} > 0$. However, if birth intentions and female schooling are positively associated and $\left| \alpha_2 \, \text{var} \, (E_1) \right| > \text{cov} \, (\epsilon_1 \epsilon_2)$, least squares estimates of (3), as in Michael (1973), would result in biases in the female schooling and intentions coefficients such that they would be more likely to be in accord with theory; i.e., the intentions coefficient would be biased negatively and the schooling coefficient biased upward. Alternatively, if reports on birth expectations ignore contraceptive costs and more educated women tend to postpone childbearing, the least squares estimator β_1 will be less then the true structural effect of schooling on contraceptive choice. In that case OLS estimates of the female schooling coefficient in both reduced form and structural (conditional on intentions) contraceptive choice equations would understate the magnitude of the role of schooling effects in contraceptive decisions.[8]

IV. Empirical Application

In this section we present estimates of reduced form, conditional and structural equations determining birth intentions and contraceptive choice based on data from the 1970 National Fertility Survey (NFS), a national probability sample of 5981 currently-married and 771 post-married women (Westoff and Ryder, 1977). Our sample consists of white women aged 25-39 with husband present, in which both spouses are in their first marriage. Of this group, we excluded women who were not "presumably fecund" (exclusive of contraceptive operations), living on farms and who could not provide information on any of the variables used

in the analysis. The total sample thus comprises 1477 women, divided
into three age groups, 25-29 years, 30-34 years, and 35-39 years of age.

a. Reduced Form and Conditional Logit Estimates

As a first test of the hypothesis that schooling affects the
costs of contraceptives and/or allocative decisions, we categorize the
eleven or more contraceptive methods on which there is survey information
into two groups corresponding to the relatively new (in 1970) fixed-cost and
"traditional" variable-cost techniques of the theoretical analysis.
We let P_E take on the value of one if the woman is currently using either
of two devices--the pill or IUD--of if either the husband or wife has
undergone a sterilization operation for the purpose of contraception.
All of these contraceptive methods (the "Effective" techniques) repre-
the products of recent technological advances, are significantly more
effective than other methods (Westoff and Ryder, 1977), require visits
to doctors or clinics for use or purchase, and are not coition-dependent.
We let P_R = 1 if the women is using one of the less effective ("Risky")
devices and let P_M = 1 if the woman states, in response to a survey
question, that she intends to have at least one more child ("More").[9]

Nerlove and Press (1973) have shown that if the probabilities
of these joint dichotomous outcomes are characterized by a multivariate
logistic distribution, then the log of the probability ("odds") ratio
for each outcome conditional on the occurrence or non-occurrence of the
other events, P_{ci} (i = E,R,M), can be expressed as a linear function of

"main" effects and "interaction" effects involving the other endogenous

variables:

$$i = E, \; j = R, \; k = M$$
$$(10) \quad P_{ci} = \gamma_i' + \pi(i,j) + \pi(i,k) + \eta(i,j,k) \qquad i = R, \; j = E, \; k = M$$
$$i = M, \; j = R, \; k = E$$

If the set of interaction terms is set equal to zero ($\pi(i,j) = \pi(i,k) = \eta(i,j,k) = 0$), it can be easily shown that the conditional equations or

log-linear model in (10) becomes the simple logit model, where:

$$(11) \quad P_i = 1 \; / [1 + e^{-2\gamma_i'}] \qquad i = E, R, M$$

We assume that the γ_i' are linear functions of a set of exogenous explana-

tory variables, such that:

$$(12) \quad \gamma_i' = \gamma_{0i} + \gamma_{1i}EDW + \gamma_{2i}EDH + \gamma_{3i}INCH + \gamma_{4i}LWAGEW + \gamma_{5i}CATH$$
$$+ \; \gamma_{6i}AGEW + \gamma_{7i}COMSIZE + \gamma_{8i}SIBW + \gamma_{9i}SIBH + v_i$$

where EDW = schooling attainment, in years, of the wife, EDH = schooling

attainment of the husband, INCH = husband's annual income, LWAGEW =

predicted log of the weekly wage of the wife based on an auxiliary re-

gression equation estimated over working women,[10] CATH = dummy variable

which equals one if the wife is Catholic, AGEW = wife's age in years,

COMSIZE = size of the community in which the wife currently resides,

SIBW = number of the wife's siblings, and SIBH = number of the husband's

siblings. We assume that all of these variables, including

EDW, are uncorrelated with the disturbance term v_i. Sample means and

standard deviations for the three age groups are given in Table 1, along

with those for the monthly probability of conception based on the average

use-effectiveness of the contraceptive technique being used (Tietze, 1971),

parity, and the number of intended births.

Table 1

Means and Standard Deviations, White Women 25-39[a]
by Age Group

Variable	Age Group = 25-29		Age Group = 30-34		Age Group = 35-39	
	Mean	Standard Deviation	Mean	Standard Deviation	Mean	Standard Deviation
Probability of Conception	.048	.079	.045	.076	.046	.075
Effective (%)	47.5	-	47.2	-	40.1	-
Risky (%)	31.8	-	33.7	-	41.2	-
More (%)	51.0	-	22.5	-	7.4	-
Intend	.650	.863	.286	.594	.106	.498
EDW	12.70	2.07	12.71	2.08	12.33	2.22
EDH	13.28	2.53	13.26	2.75	12.86	2.88
INCH	10211	4223	11486	4966	11806	5372
LWAGEW	4.59	.245	4.60	.247	4.59	.294
CATH	.267	.443	.284	.452	.280	.449
PARITY	2.01	1.29	2.72	1.30	3.29	1.60
AGEW	27.01	1.36	32.04	1.42	36.97	1.38
AGEH	29.46	3.17	34.48	3.32	39.62	3.33
COMSIZE	3.77	1.61	3.72	1.62	3.78	1.57
SIBW	3.00	2.35	3.01	2.45	3.45	2.73
SIBH	3.03	2.56	3.09	2.50	3.87	2.88
n	623		475		379	

[a]Currently married, husband present, once fecund and non-farm.

The first three explanatory variables correspond to those commonly included in empirical studies of contraception. The instrumental wage is used to capture price-of-time effects on both intentions and contraceptive choice.[11] CATH is included to test if religious affiliation impinges upon contraceptive decisions, AGEW captures age differences in fecundability, and the sibling variables are meant to reflect background characteristics which may condition fertility goals. Finally, we include the community size variable to control for differences in informational costs associated with population density.

The equations in (11), consisting only of exogenous variables, are the logistic analogues of the reduced form equations (4) and (5) in section 3 (with the addition of an equation determining the choice of risky techniques). We use the conditional model and the logit reduced-form estimates, which can be estimated using maximum likelihood, to test if the joint contraceptive choice and intentions outcomes are independent of each other, i.e., if we drop the trivariate interaction term ($\eta(ijk) = 0$) and assume that the bivariate terms are constants in each equation (to reduce computational costs), this is equivalent to testing the hypothesis that the interaction coefficients, $\pi(ij)$, $\pi(ik)$, are all zero. We cannot, however, obtain the signs or magnitudes of the structural coefficients, such as β_2 and α_2 in (2) and (3), since the estimates obtained using the log-linear model provide only a description of the probabilistic relationships among the joint outcomes specified, not estimates of structural parameters (Heckman, 1977)). An advantage of the log-linear model, however, is that it allows us to distinguish between the choice of techniques by the characteristics suggested to be important by the theory

and yet to go beyond the estimation of reduced forms.[12]

The joint conditional estimates for the three age groups are reported in Table 2; the corresponding reduced-form estimates are given in Table 3.[13] Comparison of the values of the lnlikelihoods for each age group in Table 2 and 3 indicates that the set of interaction effects are statistically significant--minus twice the differences in likelihood values, distributed asymptotically as χ^2 with three degrees of freedom (the number of independent interaction coefficients), exceeds the relevant χ^2 fractile point at the .01 level in every age group. Thus, as suggested by the model, with other characteristics, including schooling levels, held constant, women who intend to have more children are less likely to use either type of contraceptive technique, with a woman intending additional children being less likely to adopt an efficient than a risky technique. However, as was indicated, this statistical framework does not differentiate between the causes of this mutual dependence--whether effective techniques are cost-efficient when large numbers of births must be averted (structural coefficient $\beta_2 < 0$), whether the use of the efficient methods is associated with lower costs of averting a (marginal birth ($a_2 < 0$), or whether there are omitted characteristics which tend to make women report positive birth intentions and also tend to condition them to adopt less effective technologies; i.e., the cov (ϵ_1, ϵ_2)<0. Since, however, we have argued that it is more likely that the omitted variables would produce a positive error covariance (equation (6)), the negative coefficients of the More and Effective interaction terms suggest that either β_2 or α_2 (or both) must have a negative sign in the structural equations.

Table 2

Full Information Maximum-Likelihood Logit Coefficients: 'Conditional' Estimates,

Contraceptive Choice and Fertility Plans by Age Group of Wife

Independent Variable	Age Group = 25-29			Age Group = 30-34			Age Group = 35-39		
	Effective	Risky	More	Effective	Risky	More	Effective	Risky	More
EDW	.074 (.060)a	.094 (.062)	.124** (.052)	.209** (.061)	.175** (.062)	.090 (.058)	.140** (.060)	.085 (.057)	.069 (.084)
EDH	-.020 (.030	-.022 (.031)	.001 (.025)	.030 (.034)	.011 (.035)	-.003 (.033)	.022 (.036)	.048 (.035)	-.033 (.052)
INCH	.008 (.015)	-.025 (.015)	-.014 (.013)	-.003 (.016)	.021 (.016)	-.007 (.014)	.012 (.017)	.015 (.017)	.010 (.024)
LWAGEW	-.144 (.519)	-.253 (.532)	.923** (.444)	-1.031** (.475)	-.905 (.684)	.746* (.451)	-.366 (.407)	-.553 (.693)	.421 (.618)
CATH	-.230* (.129)	.128 (.129)	.459** (.113)	-.117 (.153)	-.040 (.155)	.324** (.136)	-.622** (.177)	-.127 (.160)	.223 (.223)
AGEW	-.0041 (.048)	.0050 (.048)	-.260** (.041)	-.052 (.054)	-.024 (.434)	-.148** (.049)	-.012 (.062)	.018 (.061)	-.108** (.036)
COMSIZE	-.026 (.044)	-.029 (.045)	-.024 (.038)	-.032 (.049)	-.052 (.051)	.036 (.045)	-.045 (.054)	-.093* (.052)	.038 (.075)
SIBW	.012 (.025)	-.045* (.027)	.054** (.022)	.018 (.030)	.006 (.031)	.014 (.028)	-.004 (.029)	-.021 (.028)	.009 (.039)
SIBH	.016 (.024)	-.014 (.025)	.011 (.021)	.005 (.029)	-.006 (.030)	.005 (.027)	.019 (.029)	.050* (.028)	.060 (.038)
Effective	...	-1.529** (.185)	-.560** (.072)	...	-1.536** (.187)	-.608** (.081)	...	-1.394** (.188)	-.473** (.149)
Risky	-1.529** (.185)	...	-.409** (.074)	-1.536** (.187)	...	-.465** (.080)	-1.394** (.188)	...	-.245** (.117)
More	-.560** (.072)	-.409** (.074)	...	-.608** (.081)	-.465** (.080)	...	-.473** (.149)	-.245** (.117)	...
Constant	-.192	-.393	1.037	2.983	1.016	-.597	-1.346	-.723	3.043

-lnlikelihood = 969.326 (25-29),   691.892 (30-34),   470.167 (35-39)

aAsymptotic standard errors in parentheses
**Significant at .05 level
*Significant at .10 level

Table 3

Full Information Maximum-Likelihood Logit Coefficients: Reduced Form Estimates,

Contraceptive Choice and Fertility Plans by Age Group of Wife

Independent Variable	Age Group = 25-29			Age Group = 30-34			Age Group = 35-39		
	Effective	Risky	More	Effective	Risky	More	Effective	Risky	More
EDW	.0090	.0456	.1026**	.0775*	.0218	.0167	.0739*	.0041	.0361
	(.0435)	.0465	(.0487)	(.0405)[a]	(.0424)	(.0526)	(.0432)	(.0413)	(.0317)
EDH	-.0085	-.0079	.0072	.0232	-.0104	-.0138	-.0079	.0327	-.0387
	(.0217)	(.0233)	(.0238)	(.0241)	(.0256)	(.0302)	(.0265)	(.0263)	(.0509)
INCH	.0261**	-.0302**	-.0161	-.0154	.0232**	-.0075	.0002	.0063	.0076
	(.0108)	(.0117)	(.0118)	(.0109)	(.0114)	(.0128)	(.0124)	(.0121)	(.0240)
LWAGEW	.0795	-.3864	.8871**	-.5634*	-.1987	1.1208**	-.0058	-.3352	.5055
	(.3672)	(.3893)	(.4155)	(.3024)	(.3150)	(.4182)	(.2851)	(.2759)	(.6076)
CATH	-.4116**	.2667**	.5178**	-.1531	.0194	.3569**	-.5504**	.2585*	.3650*
	(.0975)	(.0989)	(.1065)	(.1074)	(.1121)	(.1246)	(.1357)	(.1215)	(.2164)
AGEW	.0181	.0374	-.2494**	-.0095	.0211	-.1300**	-.0179	.0264	-.1057
	(.0342)	(.0356)	(.0387)	(.0371)	(.0385)	(.0450)	(.0437)	(.0431)	(.0869)
COMSIZE	-.0031	0.0121	-.0163	.0167	-.0440	.0422	.0165	-.0603	.0502
	(.0317)	(.0339)	(.0354)	(.0331)	(.0349)	(.0411)	(.0391)	(.0382)	(.0735)
SIBW	.0235	-.0525**	.0519**	.0104	-.0076	.0093	.0078	-.0206	.0105
	(.0186)	(.0213)	(.0209)	(.0209)	(.0222)	(.0261)	(.0216)	(.0215)	(.0386)
SIBH	.0215	-.0250**	.0053**	.0076	-.0099	.0044	-.0197	.0361*	.0559
	(.0175)	(.0193)	(.0193)	(.0260)	(.0212)	(.0248)	(.0206)	(.0203)	(.0376)
Constant	.0021	-.0835	.9443	2.553*	-1.075	-1.786	-.7213	.3177	3.254
	(1.535)	(1.631)	(1.706)	(1.556)	(1.618)	(1.939)	(1.827)	(1.797)	(3.722)
-lnlikelihood[b]	-418.97**	-380.73**	-361.722**	-321.89**	-299.72**	-236.21**	-242.51**	-250.13**	-91.09*

[a]Asymptotic standard errors in parentheses.

[b]From logit equations estimated independently. Joint lnlikelihood = -1161.42 (25-29), -858.28 (30-34), -583.73 (35-39).

**Significant at .05 level

*Significant at .10 level

More importantly, the conditional estimates support the hypothesis
that the level of female schooling, for given birth intentions, is posi-
tively related to the probability that the newer, more effective, contra-
ceptive methods are adopted, with the female schooling coefficients being
statistically significant for the two older age groups. The estimates
also suggest that increases in female schooling attainment raise the pro-
bability of adopting the traditional contraceptive methods if the newer
techniques are not being used and whether or not birth intentions are non-
zero. Thus education (of women) appears to both reduce informational costs
and to increase the effectiveness with which the more familiar contracep-
tive techniques are used, although the latter effects are weaker.

The female wage coefficients are negative in all age groups and
statistically significant in the Effective equation for women in the 30-34
age group, however, suggesting that time costs may be a significant deter-
ent in the use of the more efficient methods. None of the other charac-
teristics of the families in the sample, except religion but including
the husband's education, appear to have significant effects on the con-
traceptive choice decision. The results of Table 2 do suggest, however,
that Catholic women are more likely to report that they intend to bear
additional children whatever their current contraceptive practice and
are less likely to select from among the effective techniques regardless
of their birth intentions.

The joint conditional logit estimates also reveal that female
schooling and birth intentions are positively associated whatever the
type of contraceptive chosen. Consequently, the reduced-form female
schooling coefficients in the contraceptive equations in Table 3 are

significantly lower than the corresponding conditional estimates. The jointly estimated More and Effective equations thus indicate that (i) better educated women are more likely to be expecting additional children among married women of the same age, despite using less costly, on the margin, contraceptives which lower the cost of averting births,[14] and (ii) more educated women have higher probabilities of using one of the recently introduced, highly effective contraceptive methods despite their fertility plans, not because of them. Thus, reduced-form female schooling-contraceptive choice associations are lower-bound estimates of the structural or allocative schooling effect.

b. Structural Estimates and Further Tests

While the joint conditional logit results reported in Table 2 appear to point to the existence of allocative/cost effects associated with the wife's schooling attainment in contraception decisions, additional information concerning the quantitative magnitude of the structural schooling effects on contraceptive efficiency, the degree to which reports on birth intentions anticipate contraceptive costs and the sign of the covariance between residual errors in the contraception and intentions structural equations can be obtained by direct estimation of the structural equation determining contraceptive choice, (4), as was shown in section 3. Estimation of (4) requires, however, imposing a particular set of identifying restrictions on the data which do not come directly from the theoretical analysis. We assume that a set of background variables including childhood farm residence, the number of siblings of both spouses, and the Duncan index corresponding to the occupation of the wife's father

are significantly more important in influencing the family size goals of the household than the wife's decisions to adopt new contraceptive technologies. Such decisions, given fertility plans, are thus only influenced by schooling, wages, income, age, etc.

To estimate the contraceptive adoption structural equation using classical simultaneous equations techniques we utilize the number of intended children (INTEND) as the continuous analogue of the 'More' Variable. We also redefine the contraception variable as minus the monthly probability of conception associated with the method used by the household, based on the computed contraceptive efficiencies reported in Tietze, 1971.[15] The transformation of the qualitative contraceptive variables to one continuous variable based on use-effectiveness requires the assumption that the fixed/variable cost ratios associated with the techniques are positively correlated with contraceptive efficiency. Because, however, the newer methods are the most effective and appeared to have small variable cost components in 1970, the predictions of the model concerning the choice of techniques by cost characteristics should also hold for techniques ranked by effectiveness, i.e., more educated women should select contraceptive methods with low average birth probabilities and intentions and efficiency should be negatively associated if there is a switching point. As was discussed earlier, this fixed-cost/effectiveness association may not be true in other contexts.

The structural equation results are reported in Table 4. The first two columns in each age-group contain the ordinary least squares (OLS) parameter estimates with and without a parity variable in addition to birth intentions; the third and fourth columns report the parameters obtained using two-stage least squares (2SLS) for the same specifications. The

Table 4

Ordinary and Two-Stage Least Squares Coefficients: Structural Estimates and Wu Tests,

Contraceptive Choice (Monthly Probability of Conception) by Age Group of Wife

Estimation Technique Independent Variables	Age Group = 25-29				Age Group = 30-34				Age Group = 35-39			
	OLS	OLS	2SLS	2SLS	OLS	OLS	2SLS	2SLS	OLS	OLS	2SLS	2SLS
EDW	.?04 (.0?31)	.0036 (.0031)	.0055* (.0032)	.0041 (.0040)	.0090** (.0027)	.0071** (.0026)	.0093** (.0028)	.0092** (.0036)	.0075** (.0023)	.0071** (.0028)	.0120** (.0031)	.011?* (.00?2)
EDH	-.0012 (.0015)	-.0007 (.0012)	-.0009 (.0016)	-.0001 (.0022)	.0003 (.0016)	.0007 (.0015)	-.0007 (.0017)	.0018 (.0019)	.0015 (.0017)	.0019 (.0017)	.0012 (.0018)	.0022 (.0019)
INCH(x10^{-3})	-.1158 (.7712)	-.0255 (.7712)	.2357 (.8002)	.3685 (.9214)	.3331 (.7243)	.4142 (.7012)	.0595 (.7600)	.4095 (.8000)	.3375 (.3301)	.6091 (.8243)	.7107 (.8300)	.?1?? (.99??)
LWAGEW	.0019 (.0265)	.0228 (.0277)	.0150 (.0281)	.0478 (.0809)	-.0463** (.0203)	-.0172 (.0202)	-.0793** (.0290)	-.0803* (.0453)	-.0320** (.0192)	-.0285 (.0188)	-.0351* (.0192)	-.024? (.0213)
CATH	-.0084 (.0069)	-.0096 (.0069)	-.0004 (.0078)	-.0032 (.0125)	-.0037 (.0072)	-.0108 (.0071)	-.0092 (.0083)	-.0161 (.0125)	-.0250** (.00?5)	-.0333** (.0067)	-.0148* (.0090)	-.037?* (.0?1?)
AGEW	.0005 (.0025)	-.0011 (.0025)	-.0033 (.0029)	-.0053 (.0081)	-.0042* (.0025)	-.0058** (.0024)	-.0091** (.0030)	.0017 (.0047)	.00?3 (.0030)	-.0005 (.0029)	-.0040 (.0032)	-.00?? (.0?5?)
COMSIZE(x10^{-2})	-.1567 (.2290)	-.0624 (.2317)	-.1012 (.2430)	.0404 (.3760)	-.1826 (.2210)	.1629 (.2134)	.0066 (2340)	.2469 (.2550)	-.4343* (.2590)	-.4958* (.2563)	.3576 (.2610)	.53?? (.2?7?)
INTEND[b]	-.0324** (.0037)	-.0287** (.0040)	-.0474** (.0085)	-.0404 (.0261)	-.0583** (.0030)	-.0492** (.0059)	-.1127** (.0169)	-.0140 (.0373)	-.0353** (.0031)	-.0320** (.0060)	-.1327** (.0462)	-.2??? (.13?6)
PARITY[b]	-	.0076** (.0030)	-	.0124 (.0235)	-.	.0153** (.0026)	-	-.0004 (.0166)	-	.0086** (.0024)	-	.0?1? (.0?3?)
CONSTANT	-.3130	-.0160	-.0104	-.1346	.1971	.0985	.2037	.1614	.1091	-.0085	.2433	.2415
Wu Statistic[c]	-	-	7.06(t)**	40.36(F)**	-		8.22(t)**	54.09(P)**	-		3.78(t)**	15.37(F)
$\bar{R}^2$	.114	.122			.199	.254			.093	.121		
S.E.E.	.074	.074	.076	.078	.068	.066	.067	.075	.071	.070	.070	.073

[a] Standard error in parentheses.

[b] Endogenous variable. Excluded exogenous variables are SIBHW, SIBH, the occupation of the father of the wife, and farm background, both spouses.

[c] Refers to the "t" statistic on the coefficient of the residual of preducted birth intentions from actual intentions or the joint "F" statistic on the intentions residual and that between predicted and actual parity.

** Significant at .05 level.
 * Significant at .10 level.

signs of both the wife's schooling and INTEND coefficients in all age
groups, estimated with either technique, are again consistent with
the hypotheses that the more effective techniques have relatively greater
fixed-cost components and that there is a switching point. The structural
estimates thus confirm the conditional logit results indicating that
the wife's educational attainment lowers informational costs and/or in-
creases her ability to make allocative decisions when faced with new
technologies, independent of fertility plans. While predicted parity
and intentions are highly collinear, the introduction of actual parity
in addition to intentions significantly lowers the EDW coefficients.

Wu (1973) tests support the hypothesis that both the birth inten-
tions and parity variables are correlated significantly with the struc-
tural equation residuals, confirming the inappropriateness of least
squares estimates. Moreover, the positive bias in the intentions coef-
ficient resulting from the application of OLS is consistent with the
hypothesis that the errors in the two structural equations (2) and (3)
are positively correlated; indeed, the residual covariance term, from (9),
appears to dominate the effect of contraceptive costs on the reports of
birth intentions by women. This result, in addition to the fact that
approximately 8 percent of all women in our sample who indicated that
they intended to have no more children also reported that they used
either no or highly ineffective contraceptive methods, suggests that birth
intentions information appears to be provided as if contraception were
costless and thus underestimates subsequent fertility, particularly for
women with low levels of schooling.[16]

Consistent with the positive correlation between errors in (2) and (3) implied by the sign of the least squares bias in the structural birth intentions coefficient and the positive correlation between female schooling and birth intentions indicated by the conditional logit estimates, the structural schooling coefficients are biased downward in all age groups when ordinary least squares is used. Thus the specification corresponding to structural equation (3) in part 3, estimated using 2SLS, provides the highest estimate of the female schooling effect on the monthly birth pro-babilities in all three age groups, with that coefficient also statistically significant in all age groups.[17]

The other coefficients in the 2SLS equations excluding parity are qualitatively similar to those obtained from the log-linear model: The value of the wife's predicted wage appears to be significantly and negatively correlated with the choice of the highly effective contraceptive devices in the two older age groups, consistent with the hypothesis that working women purchase less medical services from clinics or doctors, given schooling levels and fertility plans. However, while religion appears to have a significant independent effect on the choice of contraceptives among women over 35, neither community size nor the levels of the husband's schooling attainment or income appear to be important determinants of contraceptive adoption.

Table 5 summarizes the quantitative effects of female education on contraceptive choice and efficiency obtained from the reduced form, conditional and structural equations. The computations reported in the upper part of the table indicate that with educational effects estimated conditional on birth intentions, an additional year of the wife's schooling attainment is associated with a 20 percent higher probability of adopting the newer, more effective contraceptive methods among women 30-39, more than two and one-half times the corresponding reduced-form effects, as a consequence of the positive association between female schooling and incremental birth intentions. The probability that a woman has adopted the traditional methods, if the more effective techniques are not used, similarly rises by 16 percent with a one year increase in female educational attainment, a result which is almost totally obscured in the reduced form.

In terms of the estimated changes in average monthly birth probabilities as a consequence of technique choice (exclusive of technique use-efficiency), given in the lower part of Table 6, the 2SLS structural estimates indicate that a one-year increase in female schooling lowers the monthly probability of a birth by approximately 23 percent for women 30-39. These effects are more than twice as large as the reduced-form estimates and almost three times the magnitude of the effects computed from the estimates obtained by Michael from the 1965 National Fertility Survey for the same age groups. It is not clear, however, whether the selectivity bias which characterizes Michael's estimates accounts for all the difference between the OLS structural estimates obtained from the 1970 data and those of Michael, since his data reflect another time period, although one in which the pill was most recently introduced and thus when schooling should also have played a

Table 5

Effects on Contraceptive Choice and Conception Probabilities of a _Ceteris Paribus_
Increase in Female Schooling by One Year

Age Group	% Increase in Probability of Selecting Either Pill IUD, Sterilization		% Increase in Probability of Selecting Other Methods		% Increase in Probability of Intending Additional Children	
	Reduced Form	Conditional	Reduced Form	Conditional	Reduced Form	Conditional
25-29	1.3[a]	8.4	6.5	13.3	10.0	12.5
30-34	8.1	22.3	2.8	23.1	3.4	13.9
35-39	9.4	17.6	0.5	9.8	6.7	13.2

% Decrease in Monthly Birth Probability

	Michael(1973) OLS	Reduced Form	Structural OLS	Structural 2SLS
25-29	6.2 [c]	2.7[b]	7.5	11.5
30-34	9.1	14.0	15.8	20.7
35-39	6.5	12.2	15.4	26.1

[a] $2\gamma'_{11}(1-\bar{P}_i)$ where $\bar{P}_i$ is the mean probability of selecting technique set i or intending more children.

[b] $\beta_1(\overline{MBP}_i)^{-1}$ where $\overline{MBP}_i$ is the mean monthly probability of conception based on the contraceptive methods chosen.

[c] Weighted average of effects for three closed birth-intervals. The results are based on white, married, non-Catholic women in the 1965 NFS.

significant role.

V. Implications for Fertility Differentials and School Investment

While the structural estimates appear to indicate that the schooling of the wife plays a quantitatively significant role in the choice of contraceptive methods, an important aspect of the schooling-contraception relationship is its significance in ultimately determining differentials in completed family size and schooling investment. We can utilize the structural estimates to partition the gross ceteris paribus relationship between female schooling and completed fertility into contraceptive-choice and non-contraceptive effects by translating the predicted changes in average monthly birth probabilities into changes in family size. If $\bar{\lambda}_i$ is the mean monthly birth probability associated with the contraceptive method used by the "average" household in age group i, the expected number of additional children over a period of 10 years, $E(N)_{10}$, is given by :

$$(13) \quad E(N)_{10} = 1 - (1 - \bar{\lambda}_i)^{120}$$

The difference in expected additional children between the average household and a household where the woman has four additional years of schooling, given the same level of birth intentions, $\Delta E(N)_{10}^4$, is thus:

$$(14) \quad \Delta E(N)_{10}^4 = (1 - \lambda_i \cdot (1 - 4\hat{\beta}_{1i}))^{120} - (1 - \bar{\lambda}_i)^{120}$$

where $\hat{\beta}_{1i}$ is the structural female education coefficient on $\bar{\lambda}_i$. Based on the estimates in Table 4, this differential is .214 children for women 30-34.[18] That is, among women in the 30-34 age group with the same desires for additional children under a costless control regime, those with four years of

additional schooling attainment (approximately 16 years) will by age 40-44
have borne, on average, one fifth of a child _less_ than the average family (in
which the woman has 12 years of schooling) due only to differences in method
choice.

Table 6 presents a decomposition of cross-sectional female schooling
effects on completed family size based on estimates taken from U.S. studies
of fertility for the 1960-70 period computed using the .214 figure as the fer-
tility differential attributable to contraceptive choice. Since completed
family size in 1960 would not reflect the availability of the newer contra-
ceptive methods, the estimated gross schooling effects for that year are
displayed for comparative purposes only. If the 1970 contraception results
are a reasonable approximation for the 1965-1970 period, then the estimates
suggest that differences in allocative ability (or access to contraceptive
information) accounted for a substantial part of completed fertility dif-
ferentials associated with female education in those years, with almost
half of the educational differential in completed family size in the mid-
60's and more than 90 percent of this difference in 1970 being attributable solely
to differential contraceptive choice.

Household models of fertility (Willis, 1973) suggest that differential
contraceptive costs may also be reflected in the levels of resources devoted
to individual children. Based on the .212 contraceptive fertility differen-
tial, we can   approximate the effects of differential contraceptive adop-
tion behavior associated with female education on the distribution of the
educational attainment of children. As a first-order approximation of the
effects of family size on schooling, we regressed the educational level

Table 6

Decomposition of Completed Fertility Differentials by Female Schooling

Four-Year Increments, White Females

Data	Total Effect	Effect Net of Contraceptive Choice	Proportion Due to Contraceptive Choice
1960 Census[a]	-.355, -.401	-	-
1965 NFS[b]	-.452	-.238	.47
1967 SEO[c]	-.468	-.254	.46
1970 NFS[d]	-.236	-.022	.91

[a]Willis (1973). Urban Women 35-64. Other variables: husband's schooling, (1) husband's predicted income or (2) actual income, an interaction variable involving (1) or (2) and the wife's schooling, SMSA size, cohort trend.

[b]Boulier and Rosenzweig (1978). Non-Farm women 45-54. Other variables: husband's schooling, husband's predicted income.

[c]T. P. Schultz (personal communication). Women 35-44. Other variables: husband's income, nonemployment income, wife's age.

[d]Women 35-44. Other variables: husband's schooling, husband's income, wife's age, community size.

(years) of the wife (husband) against SIBW(H), CATH, the Duncan index of her (his) father's occupation, and a variable representing whether or not she (he) was born on a farm. The coefficients on SIBW and SIBH, while probably not unbiased estimates of structural sibling effects, were statistically significant and indicated that the presence of an additional child reduced schooling attainment by approximately one-third of a year for either sex.[19] These results suggest that the educational attainment of children from families where the mother has 12 years of schooling differs by less than one-tenth of a year from that of children in households in which the wife has 16 years of schooling as a consequence of differences in ability to adopt the newer contraceptive methods or in contraceptive efficiency. Thus, despite the importance of female education in contraceptive adoption, it does not appear that the existence of differences in family planning skills or information access associated with the schooling of women will play a large role in the distribution of earnings, as transmitted through educational investment, in subsequent generations.

VI. Conclusion

In this paper tests were formulated to identify the significance of the allocative-efficiency role of schooling in the adoption by households of newly-introduced contraceptive technologies. Use of appropriate estimation procedures in a simultaneous equations framework revealed that the education of the wife, among households with identical fertility plans, was significantly and positively associated with the adoption of the newer birth control methods in 1970, with quantitative effects significantly greater than those indicated in reduced-form and inconsistently estimated structural equations appearing in prior studies. Decomposition of the OLS

biases in the structural equation coefficients indicated that the estimates obtained from the latter <u>understate</u> the "true", i.e., conditional on fertility plans, allocative-efficiency or informational effects of female schooling in contraceptive adoption because of (1) the tendency of more educated women to delay their child-bearing and (2) the existence of a strong serial correlation of unobserved characteristics determining contraceptive choice.

The results also support the hypothesis that the relative magnitudes of fixed and variable contraceptive costs influence contraceptive decisions, with the value of the wife's time an important component of the former, and suggest that women with higher levels of schooling tend also to utilize traditional contraceptive techniques more efficiently. Computations based on the structural parameter estimates suggest that more than half of the differences in completed family size associated with the schooling of the wife in the latter part of the 1960s, often cited as evidence of value-of-time effects on fertility, can be attributed solely to differences in method choice. Both allocative efficiency and 'worker' effects of schooling thus appear to have played important roles in determining fertility differentials in that period.

FOOTNOTES

[1] The criterion for the inclusion in the sample of a woman with given parity was that she had to have an additional birth. It is obvious that the more effective contraceptors were thus selected out, particularly in the higher closed-interval sub-samples. Indeed, none of the sub-samples can contain households in which one of the spouses chose sterilization as a control method. This selectivity is likely to bias down-ward the estimated coefficients of schooling on the choice of effective contraceptive techniques.

[2] Michael and Willis emphasize the stochastic aspect of fecundity. One implication is that more efficient contraceptive techniques are associated with both lower expected birth probabilities and lower variances in expected parity for given durations of use. It is not clear how this additional attribute of the theory relates to the question of the role of education in the contraceptive choice decision unless education, for a given value of time and desired number of children, is positively correlated with the cost of "timing failures."

[3] It is not necessary that these costs be a significant part of the total household budget for decisions concerning the choice of contraceptive techniques to be importantly influenced by their size. What matters is the distribution of fixed and variable costs among the techniques.

[4] The existence of differential (marginal) costs to fertility control means that fertility is "subsidized" differentially across households according to the characteristics which determine contraceptive

choice. The analysis thus suggests that since women (of the same age) who _want_ relatively more births are more likely to use techniques with positive marginal costs, _cet.par._, a reduction in the cost of contraceptives of either type would lower actual fertility relatively more for high birth-rate than low birth-rate women and thus would narrow fertility differentials.

[5]A more complete system would include equations determining parity, age at marriage and coital frequency. The latter is the most directly relevant to the contraceptive choice problem; however, attempts to integrate coition into the empirical analysis reported in the next section met with no success: we could not reject the hypothesis (F-test, 10 percent level) that the set of twelve economic variables, described below, which were important correlates of birth intentions, parity and contraceptive choice, explained none of the variation in the monthly frequency of intercourse reported by women in the sample survey.

[6]Potential or "natural" fertility is assumed to be uninfluenced by schooling attainment. However, if more fecund women marry earlier as a consequence of pre-marital sexual activity and lack of information on individual fecundity, cut short their schooling, and, in response to the information on their individual stochastic component of natural fertility, choose more effective contraceptive techniques after their first pregnancy, the positive association between schooling and contraceptive efficiency will be less strong.

[7]Alternatively, the coefficient on intentions will be biased towards zero if reported birth intentions randomly differ from the actual fertility plans which influence contraceptive behavior. Specifically, if v is the value of the measurement error in n* and the v are uncorrelated (in the limit) with both ϵ_2 and the exogenous variables, then:

$$\operatorname*{plim}_{N\to\infty}\hat{\beta}_2 = \beta_2 \left[1 + \frac{\operatorname{var}(v)(1 - b_{EH}\beta_{HE.F} - b_{EF}\beta_{FE.H})(1 - \rho_{HF}^2)}{\phi} \right]^{-1}$$

However, tests are available, given that (3) is identified, to ascertain if n* and ϵ_2 are correlated (Wu, 1973). These are applied in the empirical section. The existence of a correlation between the residuals and birth intentions in the contraception equation does not, of course, rule out measurement error.

[8]The presence of serially correlated unobservable factors which influence contraceptive choice (Heckman and Willis, 1975) implies that if parity is used inappropriately in place of birth intentions in (3), application of ordinary least squares would also produce inconsistent parameter estimates. Since more efficient contraceptors would be observed to have lower parities at every age, cet.par., the covariance of the disturbances in the contraceptive choice and (implicit) parity equations would be negative and the least squares parity coefficient would be biased downward. The presumed negative effect of education on parity would thus imply that the schooling coefficient conditional on parity would also be biased downward. As was discussed above, however, even consistent estimates of the effect of schooling on contraceptive choice **within parity groups do not provide much information on the role of education in enhancing allocative performance or reducing information costs associated with contraceptives.**

[9]Eight percent of the women in the sample were pregnant at the time of the interviews. Rather than omit these women, which would result in selectivity bias, we assigned to them the contraceptive technique (if any) in use prior to the current pregnancy and computed the P_M variable as if the birth were intended. Because, in fact, some of these pregnancies may have been unintended, we also experimented with regressions run with all pregnant women assigned the value of 0 for the P_M variable. The results were not significantly different with this classification scheme.

[10]The explanatory variables in the instrumental wage equation, estimated for women working 40 weeks or more in 1970, were EDW, AGEW, COMSIZE, and linear and quadratic work experience variables, computed from actual work histories.

[11]The usual caveat concerning selectivity bias (Gronau, 1973) is invoked here. However, most of the evidence from labor supply studies suggests that the quantitative significance of this bias is slight (see, for example, Heckman (1974).

[12]It is obvious that estimates of <u>structural</u> equations involving qualitative variables would be the most useful in this context. However, the properties of the estimates obtained from discrete variable simultaneous equations models which have been proposed (Schmidt and Strauss, 1975; Heckman, 1977) are not well-established. In the next section we redefine the discrete contraceptive choice and intentions variables so that classical simultaneous equations estimators and statistical tests can be applied.

[13] To achieve convergence it was necessary to create two fictitious observations to fill the empty cells corresponding to $(P_R = 1, P_E = 1, P_M = 1)$ and $(P_R = 1, P_E = 1, P_M = 0)$. The fictitious individuals were assigned sample mean values of the personal characteristics. Nerlove and Press (1973) report on an experiment with this missing cell technique in which the resulting estimates were insignificantly changed. However, as they note, very little is known about the sensitivity of the log-linear estimates to imputed observations when their use is required to obtain results.

[14] The positive relationship between the female wage and birth intentions is not inconsistent with the negative wage effect on completed family size predicted by household models (Willis, 1973). Computations from reduced form equations, not reported, indicate that the net effect of the female wage rate on intended completed family size $(CFS = P_a + n)$ is negative for all three age groups, as is the total female schooling effect, the sum of the indirect wage and direct schooling effects on both parity and intentions.

[15] This is the same definition employed in Michael (1973). The change in sign is to facilitate qualitative comparisons of the structural and logit equations. The measure of contraceptive effectiveness is fraught with problems (Tietze, 1962); however, it is unlikely that the qualitative characteristics of the contraceptives suggested by Tietze's measures--i.e., the significantly higher effectiveness of the pill and IUD--are sensitive to estimation techniques. We assume, following convention, that 'natural' fecundability, is 0.20 and that the monthly probability of conception associated with sterilization is 0.0.

[16] A follow-up to the 1970 NFS, in which women in that sample were reinterviewed in 1975, indicates that approximately nine percent of the women reporting that they intended no more children in 1970 and who did not change their intentions had a birth in the subsequent 5-year interval (Westoff and Ryder, 1977), a result which could have been anticipated on the basis of the joint intentions and contraceptive use information in the 1970 survey. The follow-up study by Ryder and Westoff also indicates, consistent with the results obtained here, that more educated women were significantly less likely to experience a "contraceptive failure."

[17] One reason that the educational attainment of women may appear to be positively associated with the adoption of the newer contraceptive technologies in the structural equations is that schooling may reflect pre-existent abilities in allocative decision-making or in extracting information. If these innate skills are correlated with abilities related to chooling investment or if there is an unobserved homogeneous skill variable which leads to both higher levels of schooling and more efficient adoption decisions, the coefficient of EDW will be biased. As one means of testing for an "ability bias" without information on pre-school skills, we treated schooling as an endogenous variable (and excluded the predicted wage variable) enploying the same instruments used to predict birth intentions. Wu tests (not reported) indicated that we could not reject the hypothesis that EDW and the residual were orthogonal in the oldest age group but the female schooling variable and the error term did appear to be correlated in the two younger sub-samples. Differences between the coefficients

obtained with EDW exogenous and the wage variable excluded and those obtained with EDW endogenous   were small, on the order of five percent, a difference less than that caused solely by the exclusion of the female wage. The application of the Wu test is only valid, however, if the background variables associated with the wife--number of siblings and the occupational status of her father--are uncorrelated with the unobserved ability component impounded in the residual.

[18]The same computations applied to Michael's results yields an estimated differential of .004 children over the decadal period.

[19]Lindert (1977) reports regression results in which an additional child, on average, reduced the time devoted to the care of every child by from 20 to 35 percent, depending on spacing and birth order, and shows that child care time is significantly and positively related to schooling attainment; the resultant impact of family size on schooling, however, is quantitatively small.

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