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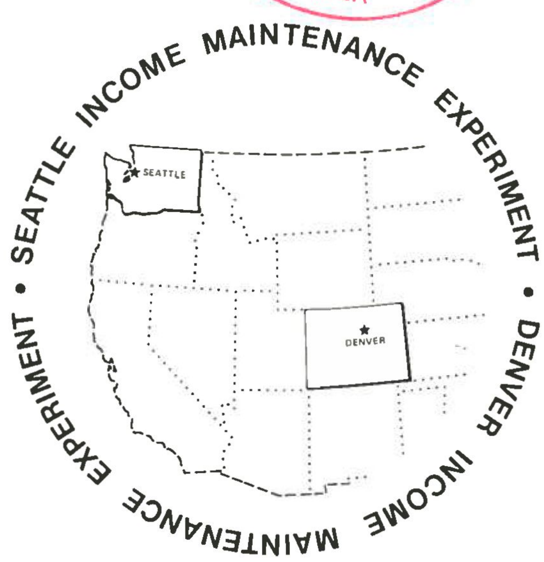

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SRI Project No. URD 8750/1190

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I INTRODUCTION

There are several reasons to expect the impact of the Seattle and Denver Income Maintenance Experiments (SIME/DIME) on changes in marital status to vary over experimental time. Most of these reasons should also apply to a permanent national program. Consequently, the effects of a national program on rates of marital formation and dissolution should depend on the length of time that the program has been in operation.

Our earlier analyses (Hannan et al., 1976; Tuma et al., 1976) revealed significant impacts of income maintenance (IM) on marital status changes, particularly on marital dissolution rates of White, Black, and Chicana women and on remarriage rates of Chicana women. With one exception, our analyses in these reports assumed that IM impacts did not vary over experimental time. Thus, our previous estimates essentially averaged the effects of IM over the total time period observed (either 18 or 24 months).

The present report has three main objectives. First, we are concerned with explicating the reasons for expecting time-variation in IM impacts on rates of change in marital status. Although we are presently unable to develop a structural model of time-variation in impacts, this discussion may suggest directions for the future development of a structural model of time-variation. This discussion appears in Section II.

Second, we want to examine the degree and kinds of variation in IM impacts during the first 24 months of SIME/DIME. Our data and analytic methods are described in Section III. Section IV contains the presentation and discussion of the results of our analyses.

Third, we wish to consider the policy implications of time-variation in IM impacts. Our analyses provide estimates of IM impacts on rates of marital formation and dissolution. Since these variables are not directly observable, it is extremely important to relate our empirical findings to an observable variable of policy interest. Our model has implications for an outcome of particular policy interest: the proportion of unmarried women in the population under an IM program. In Section V we explore implications of our findings and of different assumptions about time-variation in IM impacts for this outcome in a population like our sample.

II REASONS FOR TIME-VARIATION IN IMPACTS

There are at least three main reasons for expecting time-variation in impacts of an IM program (whether SIME/DIME or a national program) on rates of marital status change. A fourth reason applies to SIME/DIME, but not to a national program. We discuss each of these in turn below. In each case we indicate whether it leads us to expect initial impacts to overstate or to understate the long-term impact of IM on rates of marital status change.

Distribution Effects

Imagine that the marital dissolution rate varies directly with marital satisfaction and that the marriage rate varies with satisfaction with being unmarried. We suppose that satisfaction is a continuous variable and that events causing an individual's satisfaction to increase or decrease occur randomly over time. If these events are governed by a stationary stochastic process, there will eventually be an equilibrium distribution of satisfaction among married and unmarried people.

Now suppose that the introduction of an IM program immediately alters the satisfaction of people with their current marital status.* This causes an immediate change in the distribution of satisfaction and consequently changes the distribution of rates. It may also affect the parameters of the stochastic process governing changes in satisfaction over time. Eventually there will be a new equilibrium distribution of satisfaction among married and unmarried people, and thus a new equilibrium distribution of rates of change in marital status. However, the short-run distribution of satisfaction (and of rates) may differ from the old <u>and</u> the new equilibrium distributions of satisfaction (and of

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*Satisfaction may change due to change in financial opportunities in different marital statuses.

rates). This means that marriage and dissolution rates shortly after the treatment begins may differ from their ultimate values.

Understanding Treatments

Financial treatments were carefully explained to SIME/DIME participants at enrollment. However, not everyone may understand (or believe) them, and hence not respond. Some may understand an IM program through experiencing fluctuations in monthly payments as the earnings of family members change over time. Once people understand the effects of IM, they still may not change their marital status until they have made certain preparations. For example, unmarried persons may need to search for a spouse; married persons may need to find other housing arrangements. Both types of adjustment should delay change in marital status due to an IM program. Both should also apply in a national program, though the length of delay may differ from that in SIME/DIME due to differences in the distribution and acquisition of information about IM.

Indirect Treatment Effects

If IM has indirect effects that take time to occur, then its long-term effect on marriage and dissolution rates may not be immediately apparent. For example, family income affects the dissolution rate (e.g., Glick and Norton, 1971). Thus, for a given level of the wife's earnings, a decrease in the husband's earnings should decrease family income and increase the marital dissolution rate. Suppose that under an IM program the husband has the same gross wage rate but works fewer hours per year; this reduces his earnings. Due to institutional constraints on hours worked per week, reductions in annual hours worked may occur by people spending a shorter time in any particular job, a longer time between jobs, or both. Within a population, work adjustments of this type may be spread over a period of time; consequently, their effect on the dissolution rate might not be fully manifested for some time after an IM program begins. Whether initial IM impacts overstate or understate long-term impacts depends on the effects of IM on such intervening outcomes and also on the relationship between these outcomes and rates of change in marital status.

Termination Effects

Unlike a national IM program, SIME/DIME treatments have a definite known termination date (3 or 5 years after enrollment). Since most persons change marital status rather infrequently, people who are contemplating a change in marital status probably consider income for the whole period that they expect to be in the new status. The shorter the time remaining until disenrollment from SIME/DIME, the smaller the impact of IM on their average income in the new status. Consequently, the size of treatment effects may decline as disenrollment approaches, i.e., as the length of time from enrollment increases.

III DATA AND METHOD

Data

The analyses reported here use data on all changes in marital status of originally enrolled female heads of families that occurred in the first 24 months of SIME/DIME. Details are given in Tuma et al., 1976.

In our previous investigations of IM impacts on changes in marital status we found that dummy variables representing the eleven financial treatments do not improve significantly upon a representation with dummy variables for the three support levels ($3,800, $4,800, and $5,600 per year for a family of four in 1971). Consequently, in this study we continue to represent experimental treatments by dummy variables indicating the three support levels and the program length (one if three years, zero otherwise).

Assignment to experimental treatments was random within combinations of the four stratification variables: race-ethnicity (Black, White, and Chicano), site, level of normal income (seven categories), and marital status. Any analysis must take these variables into consideration to prevent erroneous inferences about IM impacts. We analyze observations on the three race-ethnic groups separately. However, we combine observations on the two sites and different levels of normal income, and represent them in the analyses by dummy variables. Tests of statistical significance in our previous research indicated that data from the two sites can be pooled but that data for the three race-ethnic groups cannot. Normal income categories are too small to permit separate analysis. Dissolution and remarriage are, of course, analyzed separately.

Our analyses include a variety of other causal variables in addition to the experimental treatment and assignment variables. They are: a dummy variable for being on AFDC before enrollment, a dummy variable for having children younger than six years, number of children, the woman's age, her years of schooling, and her wage. The inclusion of these variables improves the efficiency of the estimates of IM impacts.

The means of all variables included in our analyses are reported in Appendix A.

Method

The method of analysis is an adaptation of the log-linear rate model (Hannan et al., 1976; Tuma et al., 1976). This model assumes that the instantaneous rate of change from one marital status to another, r, is a log-linear function of a vector of experimental treatment and other causal variables $\underline{X}$ describing a woman and her family at the time of enrollment and of a parameter vector $\underline{B}$:

$$r = \exp(\underline{B} \cdot \underline{X}) \tag{1}$$

or

$$\ln r = \underline{B} \cdot \underline{X} = b_0 + b_1 x_1 + b_2 x_2 + \ldots \tag{2}$$

where the rate of change r is defined as:

$$r = \frac{dF(t|t', \underline{X})/dt}{[1 - F(t|t', \underline{X})]} \tag{3}$$

and $F(t|t',\underline{X})$ is the probability that a woman with characteristics $\underline{X}$ changes marital status before time t, given that t' is the time of either her last marital status change or her enrollment on SIME/DIME, whichever is more recent. The rate of change must be positive. It is not a probability and can be greater than one.

In Section II we gave four arguments for expecting the impact of SIME/DIME to vary with experimental time t. The first and third arguments could imply either large or small initial impacts. The second implies that the initial impact would be smaller than the later one. The fourth implies that the impact declines as experimental time increases. Considered jointly, we might expect the impact of IM over experimental time to increase monotonically, to decrease monotonically, or to increase and then later to decrease, depending on which time-relevant factors affecting response to IM are dominant at any particular time.

We have no a priori prediction concerning the above patterns. Thus, we could allow the effects of the exogenous variables on the rate of a marital status change to vary freely from one experimental period to another. Mathematically stated, this model is:

$$r_p = \exp(\underline{B}_p \cdot \underline{X}) \qquad t_{p-1} \leq t < t_p \tag{4}$$

where $p = 1, \ldots, P$; P is the total number of time periods; t_p is the last moment in period p, and $t_0 = 0$, i.e., the time of enrollment.

While both the experimental treatment and other causal variables may affect rates differently in different time periods, we are interested primarily in the time-dependence of the impacts of the experimental treatments. Furthermore, equation (4) contains P times as many parameters as equation (3), which greatly increases the cost of the estimation. So rather than the model in equation (4), we assume the following in the present analyses:

$$r_p = \exp\,(\underline{B} \cdot \underline{Z} + \underline{C}_p \cdot \underline{X}) \qquad t_{p-1} \leq t < t_p \tag{5}$$

where $\underline{Z}$ is the vector of other causal variables, $\underline{X}$ is the vector of experimental treatment variables, $\underline{B}$ is the vector of effects of the other causal variables and $\underline{C}_p$ is the vector of period-specific effects of the experimental treatments.*

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*Equations (3) and (5) imply that

$$F_j(t|t',\underline{X},\underline{Z}_p) = 1 - \exp[-\sum_{p=1}^{P} \Delta T_p \; r_p(\underline{X},\underline{Z}_p)] \tag{6}$$

where P is the total number of time periods, $\Delta T_p = T_p - T'_p$, and $T'_p = T_p = 0$ if $t_p < t'$ or $t < t_{p-1}$; $T'_p = t_{p-1}$ if $t' < t_{p-1} < t$; $T_p = t_p$ if $t' < t_p < t$; $T'_p = t'$ if $t_{p-1} \leq t' < t_p$; and $T_p = t$ if $t_{p-1} < t \leq t_p$.

$\underline{B}$ and each $\underline{C}_p$ contain a constant term; however, only as many constants can be identified as there are time periods. Therefore, in the analyses reported below, the constant term in $\underline{B}$ was constrained to be zero to achieve identification. We have estimated a constant term for each time period. Thus, the rates of marital status change of those in the control group, who have the value zero on all experimental treatment variables, are permitted to vary from one time period to another, even though the effects of other causal variables are not.*

Equation (5) says nothing about the number of time periods or about the time points that divide one period from the other. We have no firm theoretical notions to guide us in this decision. We wish to have as many periods as possible to describe the pattern of time-variation in impacts accurately. On the other hand, we wish to have enough observations on changes of marital status in every time period to obtain reasonably precise estimates of effects. Four time periods seemed a suitable compromise between these competing aims. We have used 0.0-0.5, 0.5-1.0, 1.0-1.5 and 1.5-2.0 years after enrollment as the time periods.

As in our earlier work with log-linear rate models, we have used the method of maximum likelihood to estimate parameters. Since our data consist of information on whether or not a change has occurred before T, the end of the observational period (two years except for women who drop out of SIME/DIME before this, in which case T is the length of time until they drop), and the dates of any marital status changes that have occurred (expressed in terms of the length of time from enrollment), the appropriate likelihood equation is:†

$$L = \prod_{i=1}^{N} f(t_i | t_i', \underline{X}_i, \underline{Z}_i)^{s_i} \cdot [1-F(T_i | t_i', \underline{X}_i, \underline{Z}_i)]^{(1-s_i)} \qquad (7)$$

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*The rates of those in the control group may vary over time because of secular trends, aging, etc.

†This equation is an adaptation of Bartholomew's (1957) equation for a constant rate model.

where N is the number of spells* analyzed; s_i is one if the spell terminated before T_i, and is otherwise zero; $F(T_i|t_i',\underline{X}_i,\underline{Z}_i)$ is the probability that the change occurs before T_i given t_i', $\underline{X}_i$, and $\underline{Z}_i$, and is calculated from equation (6); $f(t_i|t_i',\underline{X}_i,\underline{Z}_i)$ is the conditional probability density of a change at t and is found by differentiation of equation (6). The FORTRAN program RATE [Tuma and Crockford (1976)], was adapted to estimate the parameters in equation (5) using the likelihood equation (7).

Results of the parameter estimation could be reported in a variety of different ways. We could, of course, report estimates of the elements of the parameter vectors $\underline{B}$ and $\underline{C}_p$. Each element of these vectors indicates the effect of a unit increase in a variable on the logarithm of the rate of marital status change. The effect of a variable on the rate itself (rather than its log) is of more interest. Consequently, we have chosen to report the antilog of parameter estimates in all tables given in this report. The antilog of a parameter indicates the multiplier of the rate for a unit increase in a variable. For dummy variables, which we have used to represent experimental treatments, the antilog of the coefficient of the variable is the ratio of the rate for those people whose value on the dummy variable is one to the rate for those in the omitted category. For example, if for some experimental treatment $\exp(b_j) = 2$, then the rate for those on this treatment is twice the rate for those in the control group.

Sometimes it is useful to express an experimental effect in terms of the percentage change in the rate relative to the rate for the control group. This can be calculated quite readily. The percentage change in the rate for an experimental treatment relative to the rate for those in the control group is just $100\cdot(\exp(b_j) - 1)$, where b_j is the coefficient of the dummy variable representing the experimental treatment. Thus, if $\exp(b_j) = 2$, the percentage change in the rate for this treatment relative to the rate for the control group is 100%.

Time-dependence of experimental effects on dissolution (or remarriage) rates can be assessed in several ways. In this report we place

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*A spell is defined as a continuous period of time in a marital status.

considerable emphasis on the appearance of consistent, interpretable patterns to the magnitude and direction of effects, which we examine through graphical displays. This emphasis is largely because of our present goal, which includes trying to detect a pattern of time-variation to impacts that might be parameterized explicitly in future work. The model of four six-month periods was selected with this goal in mind and is not a parsimonious way of representing time-variation. Hence we report but do not stress statistical significance of time-dependence of the experimental impacts.

IV RESULTS

Marital Dissolution

Table 1 reports the experimental effects on dissolution rates of the three race-ethnic groups for both the time-independent rate model* [Equation (1)] and for the rate model in which experimental effects and the constant term may vary from one six-month period to another, but the effects of background variables are constrained to be constant over time [Equation (5)]. The effects of the other causal variables are reported in Appendix B.

Time-Independent Model

The results for the time-independent rate model (Column 1) indicate that for Whites and Chicanas the set of experimental treatments significantly improves upon the model that includes only the other causal variables (see the likelihood ratio test for experimental treatments). For Blacks, the experimental treatments just miss statistical significance at the 0.10 level. For both Whites and Blacks, women on each support level have a higher dissolution rate than those in the control group with comparable levels of the other causal variables. The percentage increase in the dissolution rate relative to the controls ranges from 21% for Blacks on the high support to 144% for Whites on the low support. For Chicanas, only those on the $3800 support level have a significantly higher dissolution rate (94% greater) than otherwise comparable controls. Finally, we note that the impact of IM on the dissolution rate is smaller for those on the three-year program than for those on the five-year program. However, the difference between the two program lengths is not statistically significant.

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*The results for the time-independent rate model reported here differ slightly from those reported in our twenty-four month report (Table 1 of Tuma et al., 1976). These minor differences partially result from the use of a slightly different set of causal variables. In addition, here we are analyzing all changes in marital status whereas our twenty-four month report analyzed only the first change.

Table 1

EFFECTS OF TREATMENTS ON RATES OF MARITAL DISSOLUTION BY RACE-ETHNICITY†

	One Period Model	Four Period Model			
	(0-2.0 years)	First Period (0-0.5 year)	Second Period (0.5-1.0 year)	Third Period (1.0-1.5 years)	Fourth Period (1.5-2.0 years)
Whites (N = 1367)					
$3800 Support	2.44***	5.30***	2.09*	2.34**	1.40
$4800 Support	2.13***	3.18**	1.51	2.00*	2.17
$5600 Support	1.64*	3.43**	.82	1.44	1.62
Three-year treatment	.82	.69	.67	.72	1.49
Likelihood ratio test (y^2) for experimental treatments	19.58***		35.93**		
Degrees of freedom	4		19		
Likelihood ratio test (χ^2) for time-dependent effects of experimental treatments			12.69		
Degrees of freedom			12		
Blacks (N = 976)					
$3800 Support	1.69**	.91	1.83	2.59**	1.33
$4800 Support	1.72**	1.28	2.72***	2.22*	.70
$5600 Support	1.21	1.07	1.60	1.52	.52
Three-year treatment	.82	1.27	.45**	.75	1.46
Likelihood ratio test (χ^2) for experimental treatments	7.71		21.15		
Degrees of freedom	4		19		
Likelihood ratio test (χ^2) for time-dependent effects of experimental treatments			11.53		
Degrees of freedom			12		
Chicanas (N = 601)					
$3800 Support	1.94**	1.45	2.74*	2.23	1.16
$4800 Support	1.03	.81	.40	1.29	1.80
$5600 Support	.86	.28	.26	.90	2.38
Three-year treatment	.94	.59	.91	1.42	.77
Likelihood ratio test (χ^2) for experimental treatments	8.44*		34.03**		
Degrees of freedom	4		19		
Likelihood ratio test (χ^2) for time-dependent effects of experimental treatments			19.46*		
Degrees of freedom			12		

† All equations contain the other causal variables given in Appendix B. Coefficients are exp($\hat{b}_j$) and indicate the multipliers of the rate. A coefficient of 1.0 means "no effect" of that variable.

* $0.10 \geq p > 0.05$

** $0.05 \geq p > 0.01$

*** $0.01 \geq p$

Time-Dependent Model

Columns 2 through 5 of Table 1 give the multipliers of the dissolution rate for the different experimental treatments for the first through fourth six-month periods, respectively. Again we see that for both Whites and Chicanas the set of experimental treatments significantly improves upon a model with just the other causal variables. For Whites, all but one of the twelve support level multipliers for the four periods indicate increases in the dissolution rate relative to controls; for Blacks, all but three indicate increases in the rate. For Chicanas, for whom the time-independent model shows a significantly positive impact only for those on the low support level, the low support multipliers imply an increased dissolution rate in each of the four periods.

For each race-ethnic group, the second row of likelihood ratio tests in Table 1 provides a test of the model in which experimental impacts on marital dissolution rates may vary over the four six-month periods against a model with identical variables in which only the constant term may vary over the four periods. Based on this test, only Chicanas have significant variation (at the 0.01 level) in experimental impacts over the first two years of SIME/DIME. For the purposes of this paper, this test is less important than ascertaining whether there are any consistent, interpretable patterns of time-variation that we might explicitly parameterize in future work.

We do, in fact, find some consistent patterns of time-variation in impacts. For Whites (see Figure 1), all three support levels produce a very large increase in the dissolution rate during the first six-month period. The support level effects in the remaining three periods are positive in all but one instance (the $5600 support in the second period), but are substantially below the enormous impacts in the first period. There is no clear pattern of time variation in the final three periods. Thus, for Whites, there seems to have been an initial burst of marital dissolutions in the first six months, along with a smaller positive increase relative to the control group thereafter. This is consistent with the argument that women who are dissatisfied with their marriage

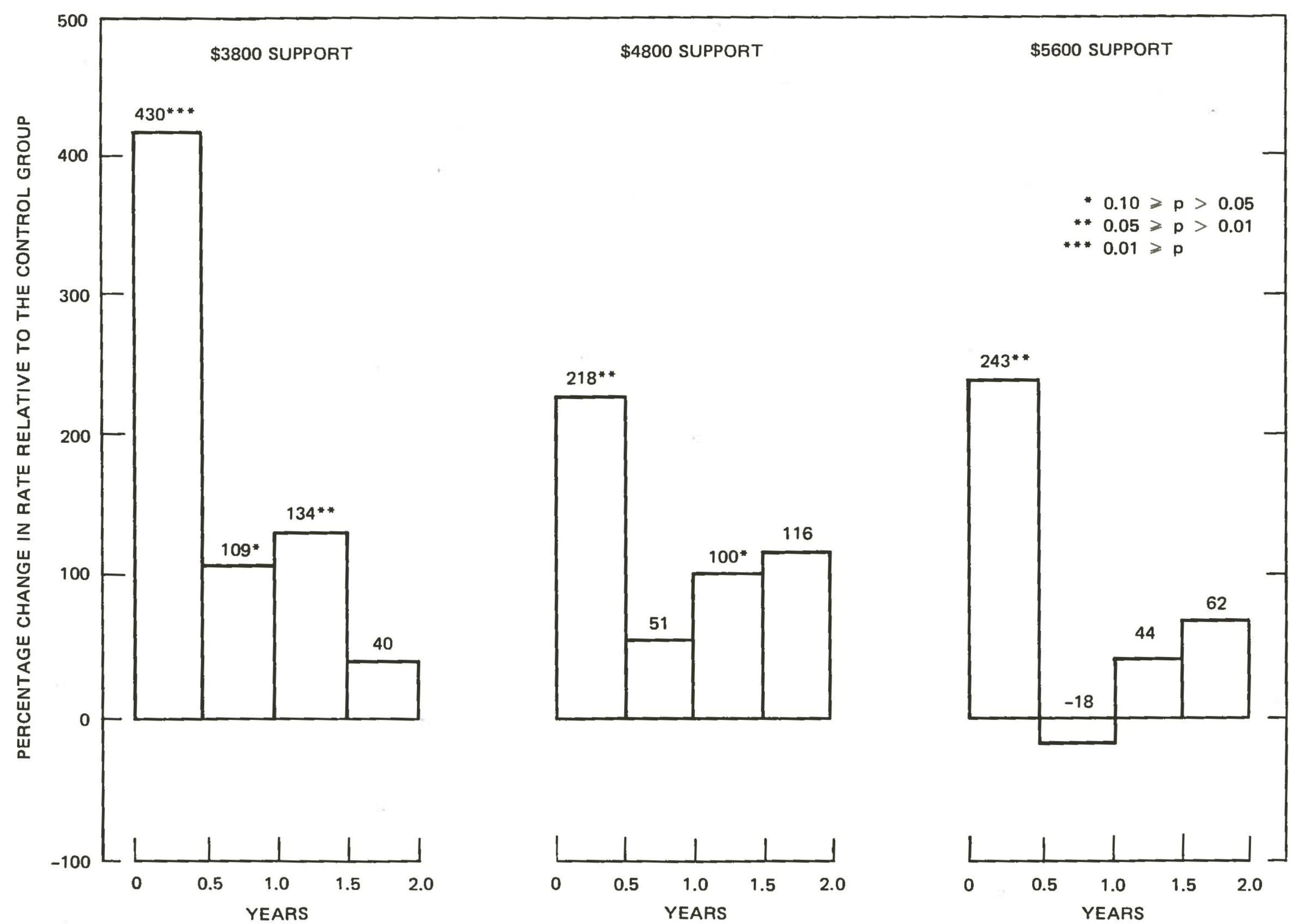

FIGURE 1 RELATIONSHIP OF IM SUPPORT LEVELS TO MARITAL DISSOLUTION RATES OF WHITE WOMEN OVER TIME

prior to IM find singleness preferable after IM begins and quickly respond to their altered financial opportunities under IM by changing their marital status.

There is also a clear pattern of time-variation in the effects of the support levels on the dissolution rates of Blacks (see Figure 2). However, the pattern for Blacks differs from that for Whites. For Blacks, the effects of all three support levels are exceedingly small relative to the effects for Whites in the first six-month period. In the next two periods, there are substantial positive effects of all three support levels on the dissolution rate of Blacks. As we have found previously, these increases are smallest for those on the high support. Support level effects are smaller in the fourth period, and actually indicate lower rates than the controls for the \$4800 and \$5600 treatments.

The patterns of time-variation in IM impacts on dissolution rates of Chicanas are not consistent across support levels (see Figure 3). For the low support level, which is the only treatment significantly affecting the Chicana dissolution rate, the pattern resembles that of Blacks: a fairly small effect in the first period, large positive effects in the second and third periods, and a very small positive effect in the fourth period. The patterns for both Blacks and Chicanas suggest that members of these groups needed time to adjust to IM.

Note that for Whites and Blacks on all three support levels and for Chicanas on the low support level, the largest support level impacts on marital dissolution rates never occur in the fourth period. One might argue that effects in the fourth period, which is near the middle of the experiment for most persons, would be closer to the long-term impacts of a permanent national IM program than the effects in earlier or later periods. If this argument is valid, then IM impacts on marital dissolution, though mainly positive, are smaller than the time-independent model suggests. The final section of this report explores the implications of this possibility.

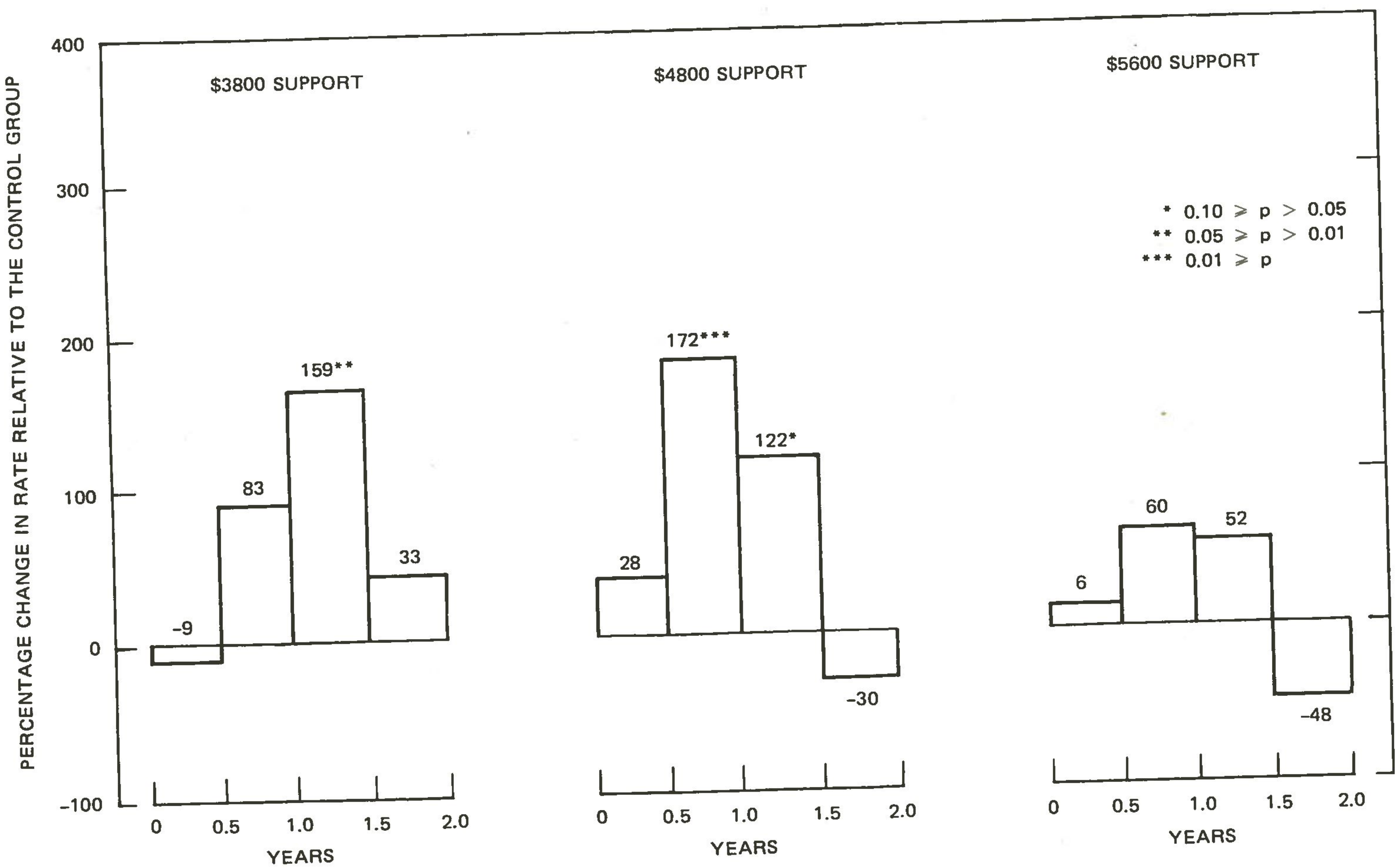

FIGURE 2 RELATIONSHIP OF IM SUPPORT LEVELS TO MARITAL DISSOLUTION RATES OF BLACK WOMEN OVER TIME

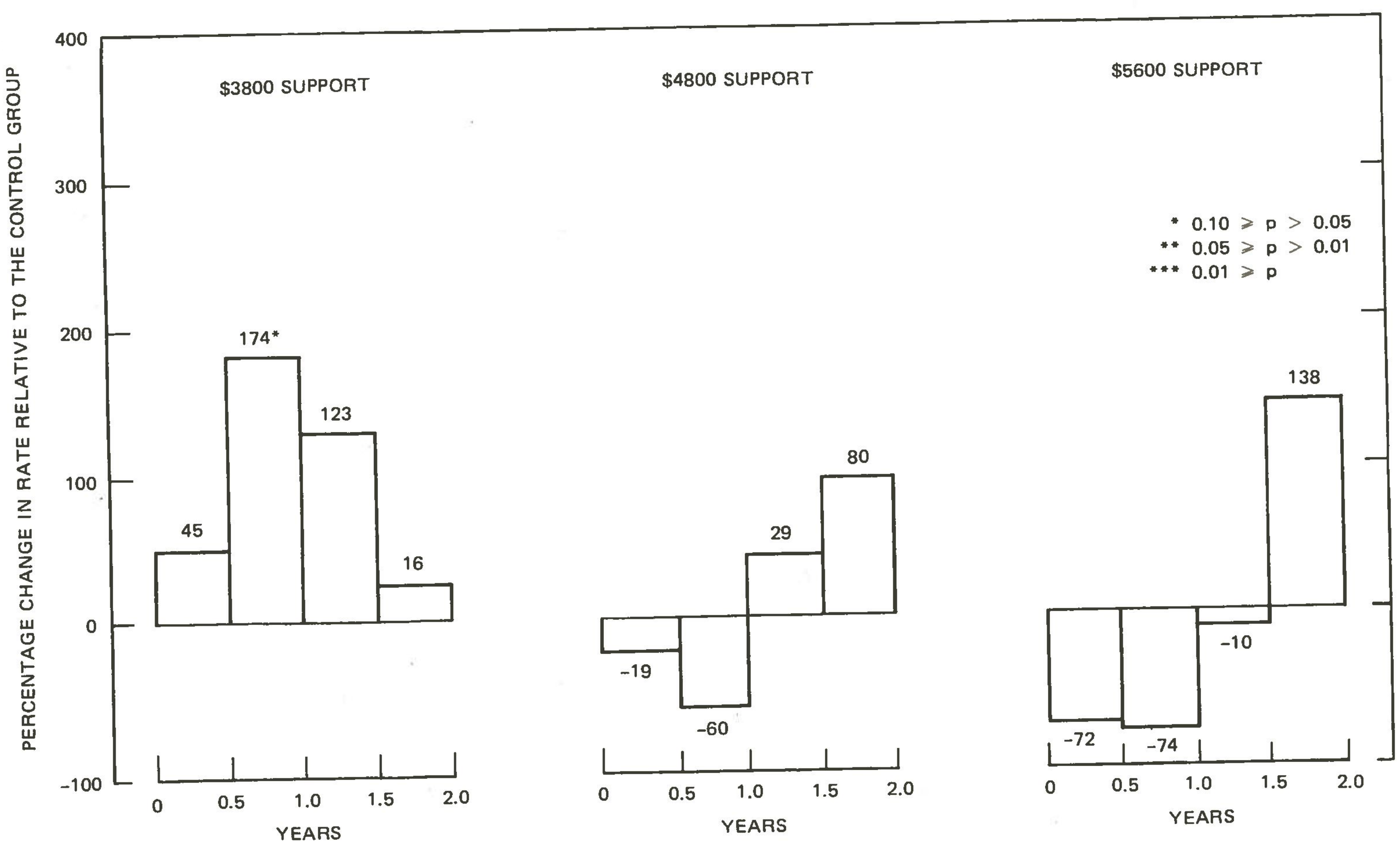

FIGURE 3 RELATIONSHIP OF IM SUPPORT LEVELS TO MARITAL DISSOLUTION RATES OF CHICANA WOMEN OVER TIME

Remarriage

Table 2 contains the estimated experimental effects on remarriage rates of the three race-ethnic groups for the time-independent rate model and for the rate model in which effects of experimental treatments and the constant term may vary with the time period. The effects of other causal variables are reported in Appendix C.

For both the time-independent and time-dependent models, significant experimental effects occur only for Chicanas. (See the likelihood ratio tests for the set of experimental effects.) For these models, the remarriage rate of Chicanas falls as the support level rises, except for the fourth six-month period. And, as in the case of dissolution rates, we find that women with significant experimental impacts (in this case, the Chicanas) tend to show larger responses if enrolled on the five-year program rather than on the three-year program. The two program lengths do not have significantly different effects, however.

For each race-ethnic group, the second row of likelihood ratio tests in Table 2 provides a test of the model in which experimental impacts on remarriage rates may vary over the four six-month periods against a model with the same variables, in which only the constant term may vary over time. This test indicates significant time variability in impacts for Whites (at the 0.01 level) but not for Blacks or Chicanas. Since the experimental impacts on remarriage rates of Whites are never significant and are almost always smaller in magnitude than support level impacts for other race-ethnic groups, we do not attribute much importance to the statistical significance of the time-variation in the impacts for Whites.

To look for possible patterns of variation in impacts over time, we again examine graphical displays of the support level multipliers of remarriage rates over time for each of the three race-ethnic groups. Figure 4 shows no clear pattern of impacts for Whites, while Figure 5 indicates that most support level effects for Blacks are positive but do not vary in any consistent fashion over time. As we indicated above, the support level effects for Chicanas (see Figure 6) are all negative and usually large in magnitude. Note that the differences in the remarriage

Table 2

EFFECTS OF TREATMENTS ON RATE OF REMARRIAGE BY RACE-ETHNICITY†

	One Period Model	Four Period Model			
	(0-2.0 years)	First Period (0-0.5 year)	Second Period (0.5-1.0 year)	Third Period (1.0-1.5 years)	Fourth Period (1.5-2.0 years)
Whites (N = 902)					
$3800 Support	1.27	1.58	.62	1.22	1.93
$4800 Support	1.09	1.09	.81	1.08	1.42
$5600 Support	.79	.62	1.36	.00	.68
Three-year treatment	1.02	1.35	1.46	1.02	.52
Likelihood ratio test (χ^2) for experimental treatments	2.99		23.62		
Degrees of freedom	4		19		
Likelihood ratio test (χ^2) for time-dependent effects of experimental treatments			19.14*		
Degrees of freedom			12		
Blacks (N = 1046)					
$3800 Support	1.30	.88	1.75	1.23	2.03
$4800 Support	1.73*	1.59	1.85	1.91	1.72
$5600 Support	1.17	1.78	1.47	.70	.72
Three-year treatment	.91	.82	.93	1.03	.81
Likelihood ratio test (χ^2) for experimental treatments	4.40		18.63		
Degrees of freedom	4		19		
Likelihood ratio test (χ^2) for time-dependent effects			4.64		
Degrees of freedom			12		
Chicanas (N = 407)					
$3800 Support	.47*	.36	.39	.68	.41
$4800 Support	.35**	.18*	.10	.40	.88
$5600 Support	.20***	.01**	.00	.18	.61
Three-year treatment	2.16**	2.73	2.33	3.19	1.32
Likelihood ratio test (χ^2) for experimental treatments	11.05**		28.76*		
Degrees of freedom	4		19		
Likelihood ratio test (χ^2) for time-dependent effects of experimental treatments			10.72		
Degrees of freedom			12		

† All equations contain the other causal variables given in Appendix C. Coefficients are $\exp(\hat{b}_j)$ and indicate the multipliers of the rate. A coefficient of 1.0 means "no effect" of that variable.

* $0.10 \geq p > 0.05$

** $0.05 \geq p > 0.01$

*** $0.01 \geq p$

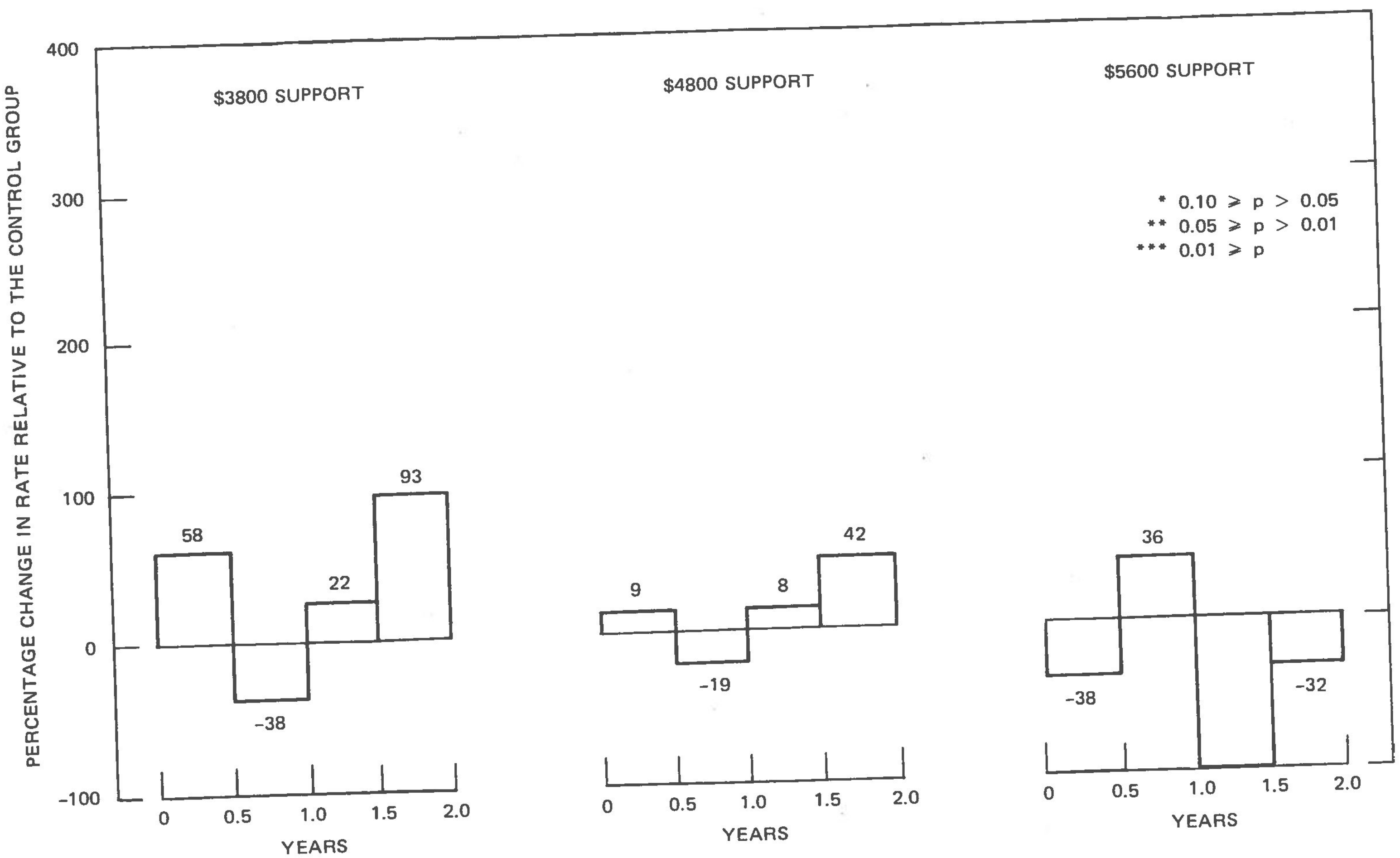

FIGURE 4 RALATIONSHIP OF IM SUPPORT LEVELS TO REMARRIAGE RATES OF WHITE WOMEN OVER TIME

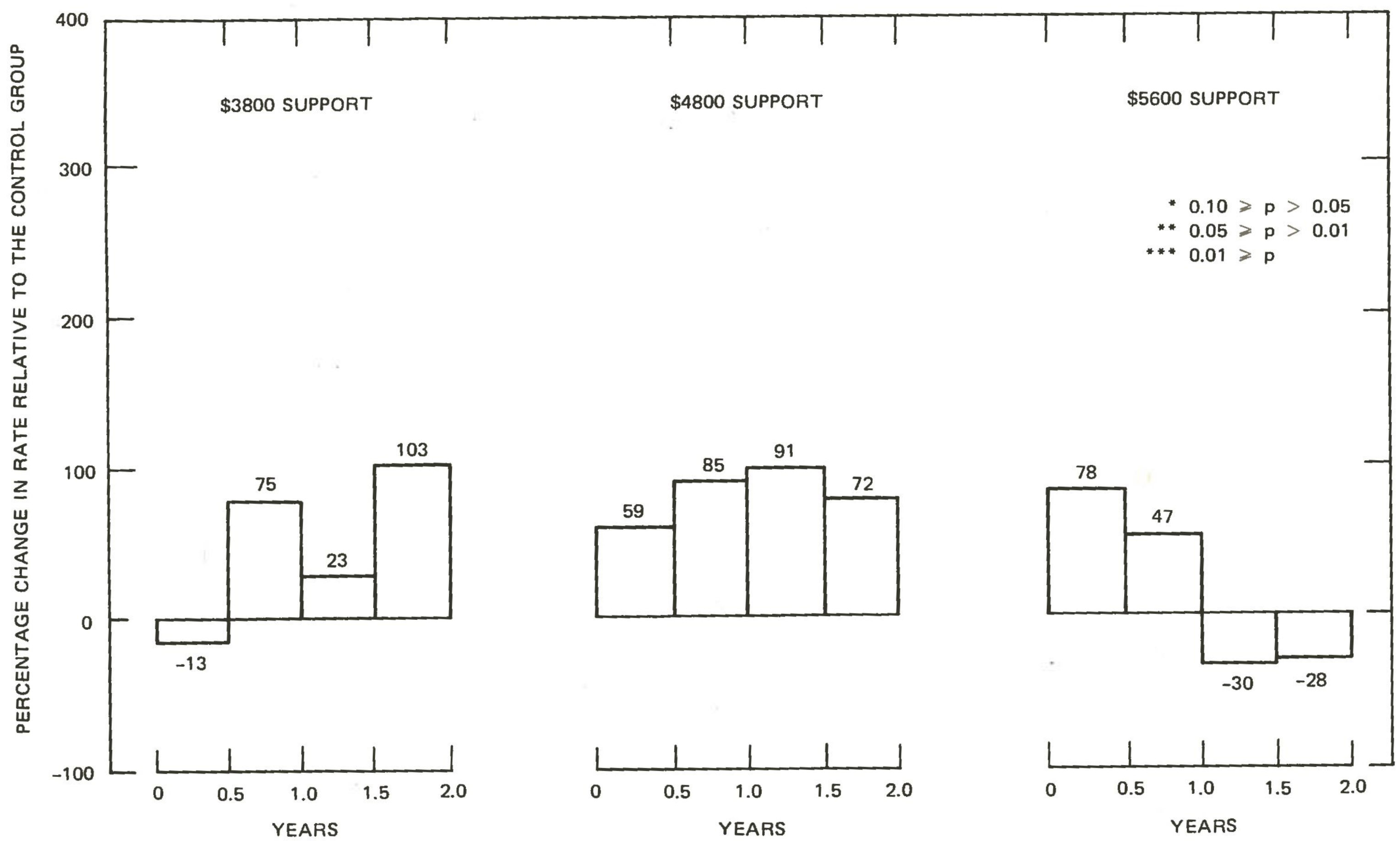

FIGURE 5 RELATIONSHIP OF IM SUPPORT LEVELS TO REMARRIAGE RATES OF BLACK WOMEN OVER TIME

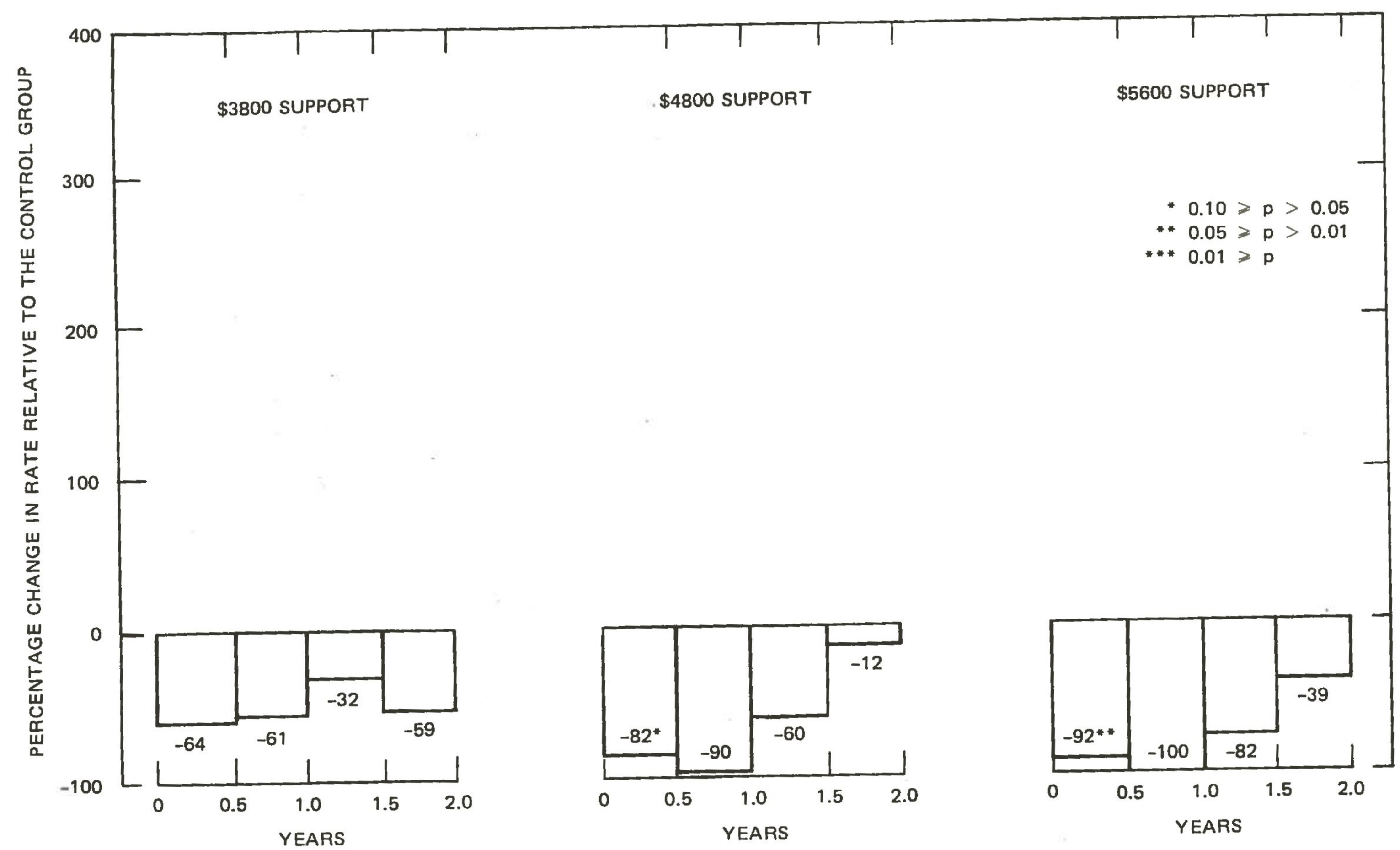

FIGURE 6 RELATIONSHIP OF IM SUPPORT LEVELS TO REMARRIAGE RATES OF CHICANA WOMEN OVER TIME

rates of Chicana controls and those on the $4800 and $5600 supports (which generally have the largest impacts) are much smaller in the fourth six-month period than in the first year of SIME/DIME. If the fourth period impacts are most similar to the long-term impacts of a permanent IM program, then the display in Figure 6 suggests that the time-independent rate model exaggerates the dampening effect of IM on remarriage of Chicanas.

Summary of Results

The results presented in this section agree with our previous findings of significant support level effects on the dissolution rates of Whites and Chicanas and on the remarriage rates of Chicanas. In addition, the impact of the support levels on the dissolution rates of Black women just misses statistical significance at the 0.10 level. These findings also indicate that a rate model in which experimental impact may vary from one six-month period of the experiment to another did not significantly improve upon the explanatory power of a time-independent model, except in the case of dissolution rates of Chicanas and of remarriage rates of Whites. We do not put much importance on experimental impacts on remarriage of Whites as these impacts are never significant.

Several consistent patterns emerged in the time-dependent analyses. We found that for Whites, all three support levels led to huge increases (relative to controls) in the dissolution rate during the first six-month period and to smaller increases in subsequent periods. For both Blacks and Chicanas, the support level effects on dissolution rates were negligible in the first six-month period, but increased thereafter. We also found that whenever there are substantial support level effects on rates of marital status change, these effects are never largest in the fourth six-month period, which might be expected to have effects most like the long-term effects of a permanent IM program.

V IMPLICATIONS OF TIME-VARIATION

At present it is unclear to what extent we should rely on the statistical tests indicating insignificant time-variation in experimental impacts, and to what extent we should be influenced by the nonsignificant, but often consistent, patterns of time-variation in support level effects. When experimental treatments are parameterized in terms of income and independence effects rather than support levels and time is also parameterized, time-variation in impacts of an IM program may be better understood.

In the interim, it is useful to consider what different patterns of time-variation in support level effects imply about the predicted proportion of unmarried women at various future points in time. This proportion is of considerable policy importance since program costs tend to vary directly with it.

The examination of these implications has three purposes. First, we wish to learn the consequences of various patterns of time-variation in IM effects on the steady-state or equilibrium* proportion of unmarried women in a population like our sample. If our estimates of time-independent and time-dependent effects of the support levels on rates of change in marital status imply similar steady-state proportions of unmarried women, then there is less reason for concern about time-dependence in IM effects. Second, we want to examine the consequence of different patterns of time-variation in impacts on the time path to equilibrium. Among other things, the time path indicates how rapidly the steady-state is attained. If a national IM program is instituted, it will undoubtedly be evaluated within

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*By steady-state or equilibrium proportion, we mean the proportion that would eventually be obtained assuming IM affects rates of marital status change but does not alter the values of other causal variables used in the analysis. A formal definition is given below.

a few years after its beginning. It is useful to know whether the steady-state will have been reached at the time of such an evaluation. Finally, we would like to know the time path from an IM to non-IM condition if the effects of IM terminate at a particular time. The termination of these effects may be because of either the purely transitory nature of IM impacts or termination of the IM program, coupled with a return to pre-IM welfare programs. The latter not only happens with certainty on SIME/DIME, but could also happen in a national program if it had highly undesirable, unforeseen consequences that led to its abandonment.

The consequences of different patterns of time-variation in IM impacts for the proportion of unmarried women in a population can be predicted through the simple model implicitly underlying our analyses. This model assumes that: (1) there are only two states in which a woman can be (married or not married); (2) for each woman, change in marital status is a first-order Markov process within a given time period p; and (3) a woman's rates of forming and dissolving a marriage within time period p are log-linear functions [see equation (5)] of the IM support level and of her values on the other causal variables used in our analyses (which we assume to be fixed and exogenously determined).*

Let $q_{ij}(t)$ represent the probability that woman i with characteristics $\underline{X}_i$ is unmarried at time t when an IM program with support level j begins at time 0. By the above assumptions,

$$\frac{dq_{ij}(t)}{dt} = -\mu_{ijp}q_{ij}(t) + \delta_{ijp}(1 - q_{ij}(t)) \quad , \tag{8}$$

$$t_{p-1} \leq t < t$$

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*Values of the other causal variables may vary over time either for exogenous reasons (e.g., age), because they are affected by a change in marital status (e.g., normal income), or because they are influenced by IM. The latter two probably cause the most serious errors in drawing implications of our analysis for the proportion of unmarried women in a population. Some type of simultaneous-equations approach is needed to solve these problems.

where μ_{ijp} and δ_{ijp} are her rates of marital formation and dissolution, respectively, in period p. The solution of this differential equation is:*

$$q_{ij}(t) = \frac{\delta_{ijp}}{\delta_{ijp} + \mu_{ijp}} + \left[q_{ij}(t_{p-1}) - \frac{\delta_{ijp}}{\delta_{ijp} + \mu_{ijp}}\right] \exp[-(\delta_{ijp} + \mu_{ijp})\cdot(t - t_{p-1})] \tag{9}$$

for $t_{p-1} \leq t < t_p$

Through recursive application of this equation, $q_{ij}(t)$ can be calculated for any time t in period p if we know $q_{ij}(0)$ and woman i's remarriage and dissolution rates in every period.

Let us first consider $q_{ij}(0)$, the probability that woman i is unmarried when IM program j is introduced. This is the same for all programs. If we wished to predict the time path of the proportion of unmarried women in the SIME/DIME sample, we could choose $q_{ij} = 0$ if the woman was married at enrollment and $q_{ij} = 1$ if she was unmarried at enrollment. Since we are not interested in the time path for the sample per se (which we know to have an over-representation of married Black couples and an under-representation of married White and Chicano couples), we instead choose $q_{ij}(0)$ to be the predicted probability that woman i would be unmarried in the control environment, based on her values of the other causal variables and our estimates of their effects in the time-independent rate models of remarriage and marital dissolution. This causes women in the sample to begin in what we estimate to be their "normal" nonexperimental (control) situation.

As equation (9) indicates, the probability of being unmarried at time t, $q_{ij}(t)$, also depends on the effect of an IM program on rates of marital status change in time period p and in previous periods, too.

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*See Cox and Miller, 1965, p. 272.

There is an infinite number of potential patterns of time-variation in the impacts of an IM program on rates of marital status change. We consider the consequences of three reasonable hypotheses. Hypothesis 1 (Immediate Adjustment) is that adjustment to an IM program is instantaneous so that the effects of IM on rates of marital status change do not depend on the length of time that IM has been in effect. This corresponds to the time-independent rate model [see Equation (1)] for which estimates are given in column 1 of Tables 1 and 2. The next possibility, suggested by our time-dependent analyses, is that adjustment to a new IM program takes six months, but after this period the effects of IM on dissolution and remarriage rates are time-independent. We refer to this as Hypothesis 2 (Six-Month Adjustment). We can obtain approximate estimates of effects under this hypothesis by assuming that the effect of a support level after six months is the average effect in the second through fourth six-month periods, as estimated in our time-dependent analyses (see columns 3 through 5 in Tables 1 and 2). A final possibility (Hypothesis 3, Eighteen-Month Adjustment) is that adjustment to an IM program takes 18 months, and that rates of marital status change are time-independent thereafter. Our time-dependent analyses also supply estimates for this hypothesis.

For each hypothesis we used equation (9) to calculate the probability that each woman in our sample would be unmarried in the steady-state condition (i.e., if $dq_{ij}(t)/dt = 0$) if she were on IM programs with $3800, $4800, and $5600 support levels, and if she remained in the control situation.* Table 3 gives the arithmetic mean of these probabilities. It approximates the expected proportion of unmarried women in a steady-state. The means are given by race-ethic group, hypothesis and support level.

Although our sample is not representative of the U.S. population, it is interesting to note that our predictions of the expected proportion of unmarried women in the steady-state control environment are extremely similar to those reported for low income families in the U.S. in 1974 by

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*Note that the equilibrium probability is just $\delta_{ijp}/(\mu_{ijp} + \delta_{ijp})$ where p is the last time period.

Table 3

EXPECTED PROPORTION OF UNMARRIED WOMEN IN A POPULATION LIKE THE SIME/DIME SAMPLE BY RACE-ETHNICITY, HYPOTHESIS,* AND IM SUPPORT

White

IM Support	Hypothesis 1	Hypothesis 2	Hypothesis 3
None (control)	.340	.351	.302
$3800	.473	.452	.247
$4800	.478	.461	.383
$5600	.491	.471	.477

Black

IM Support	Hypothesis 1	Hypothesis 2	Hypothesis 3
None (control)	.634	.657	.725
$3800	.687	.692	.645
$4800	.632	.665	.546
$5600	.641	.702	.664

Chicanas

IM Support	Hypothesis 1	Hypothesis 2	Hypothesis 3
None (control)	.381	.336	.297
$3800	.614	.650	.514
$4800	.539	.522	.443
$5600	.626	.635	.584

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*Hypothesis 1: Immediate Adjustment
Hypothesis 2: Six-Month Adjustment
Hypothesis 3: Eighteen-Month Adjustment
For complete definitions of the hypotheses, see text.

Ross and Sawhill (1975, p. 68). They report that this proportion is 0.33 for Whites and 0.63 for Nonwhites (most of whom would be classified as Blacks in our analyses). Our predictions range from 0.30 to 0.35 for Whites and from 0.63 to 0.73 for Blacks.*

Consider the first two hypotheses. Table 3 shows that for every race-ethnic group the expected proportion of unmarried women is higher for each support level than in the control environment. The support level-control differences are especially large for Chicanas, somewhat less large for Whites, and smallest for Blacks. For the most part, differences among support levels are relatively small and, except for Blacks, considerably smaller than the support level-control differences.

Hypotheses 1 and 2 give very similar results. The direction of the differences between the two agrees with what one would expect from the time-dependent patterns described in Section IV. For Whites, the expected proportion of unmarried women in the steady-state under IM is higher for the Immediate Adjustment Hypothesis than for the Six-Month Adjustment Hypothesis. For Blacks, the expected proportion of unmarried women under IM is lower for the Immediate Adjustment Hypothesis than for the Six-Month Adjustment Hypothesis.

Both Hypotheses 1 and 2 suggest that IM would raise the expected proportion of unmarried Chicana and White women in populations like our sample. The results for Hypothesis 3 (Eighteen-Month Adjustment) call this into question. According to the predictions in Table 3, if Hypothesis 3 is correct, then the expected proportion of unmarried women in the steady-state under IM would be lower than in the control environment for Blacks, higher than in the control environment for Chicanas (though not as high as if the first two hypotheses were correct), and possibly higher or lower than in the control environment for Whites, depending on the support level adopted.

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*The control predictions in Table 3 vary with the hypothesis since we permitted the constant term to vary with the time period in the time-dependent analyses.

It is important to note that the support level predictions vary considerably more under Hypothesis 3 (Eighteen-Month Adjustment) than under the other two hypotheses. This partly results from greater sampling error in the estimates generating the predictions under Hypothesis 3. Under Hypothesis 3, permanent IM effects are estimated by data on marital status changes in a single six-month period instead of twenty-four or eighteen months, as in Hypotheses 1 and 2, respectively. Because of this, the predictions under Hypothesis 3 must be interpreted cautiously.

The generation of accurate projections of the long-term impact of an IM program is not the purpose of these calculations. Rather, the message of Table 3 is that different patterns of time-variation in experimental effects on rates of marital status change have quite different implications for the long-term consequences of an IM program's effect on the proportion of unmarried women in a population.

Though the long-term expected proportion of unmarried women in a population is of particular policy relevance, it is also important for policy-makers to have some idea of the time path from the steady-state environment when IM is instituted until a new steady-state is reached. Figures 7 through 15 display these time paths by support level for Hypotheses 1 through 3 for Whites, Blacks, and Chicanas, respectively. These graphs have several implications worth noting.

First, for both hypotheses, every race-ethnic group and every support level found to affect rates of change in marital status, it takes many years to reach the steady-state. The steady-state is reached slowly because of the relatively long average time that a woman spends in a particular marital status. The slowness with which the steady-state is reached, even when there are no transitory effects of IM on rates of marital status change, should be considered if a national IM program is instituted and then evaluated a few years later. If this evaluation occurred five years after the program began and only took into consideration the current proportion of unmarried women in the population, then it might seriously underestimate the ultimate effects of the program.

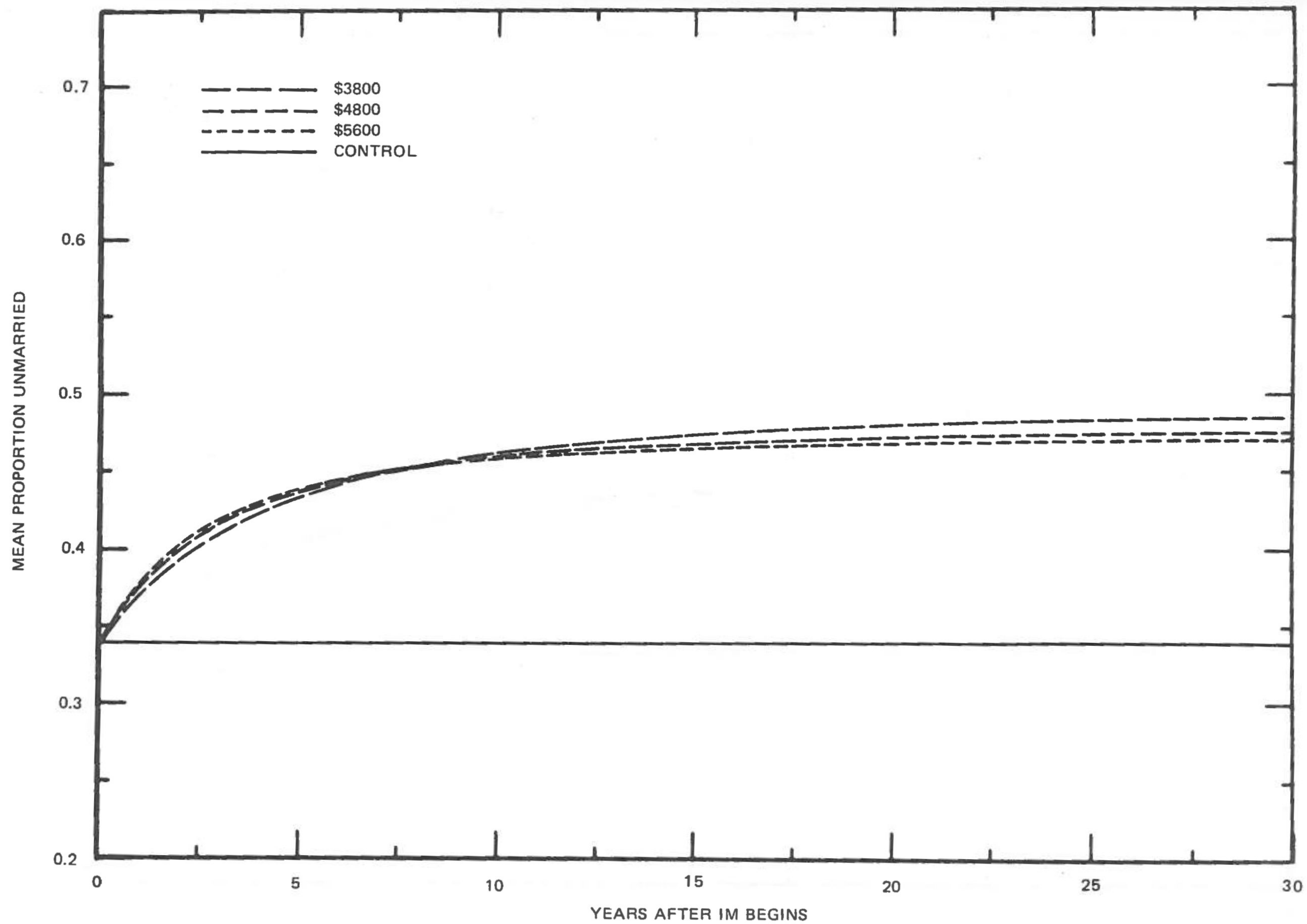

FIGURE 7 IMPLICATIONS OF HYPOTHESIS 1 (IMMEDIATE ADJUSTMENT) FOR THE MEAN PROPORTION OF UNMARRIED WHITE WOMEN OVER TIME

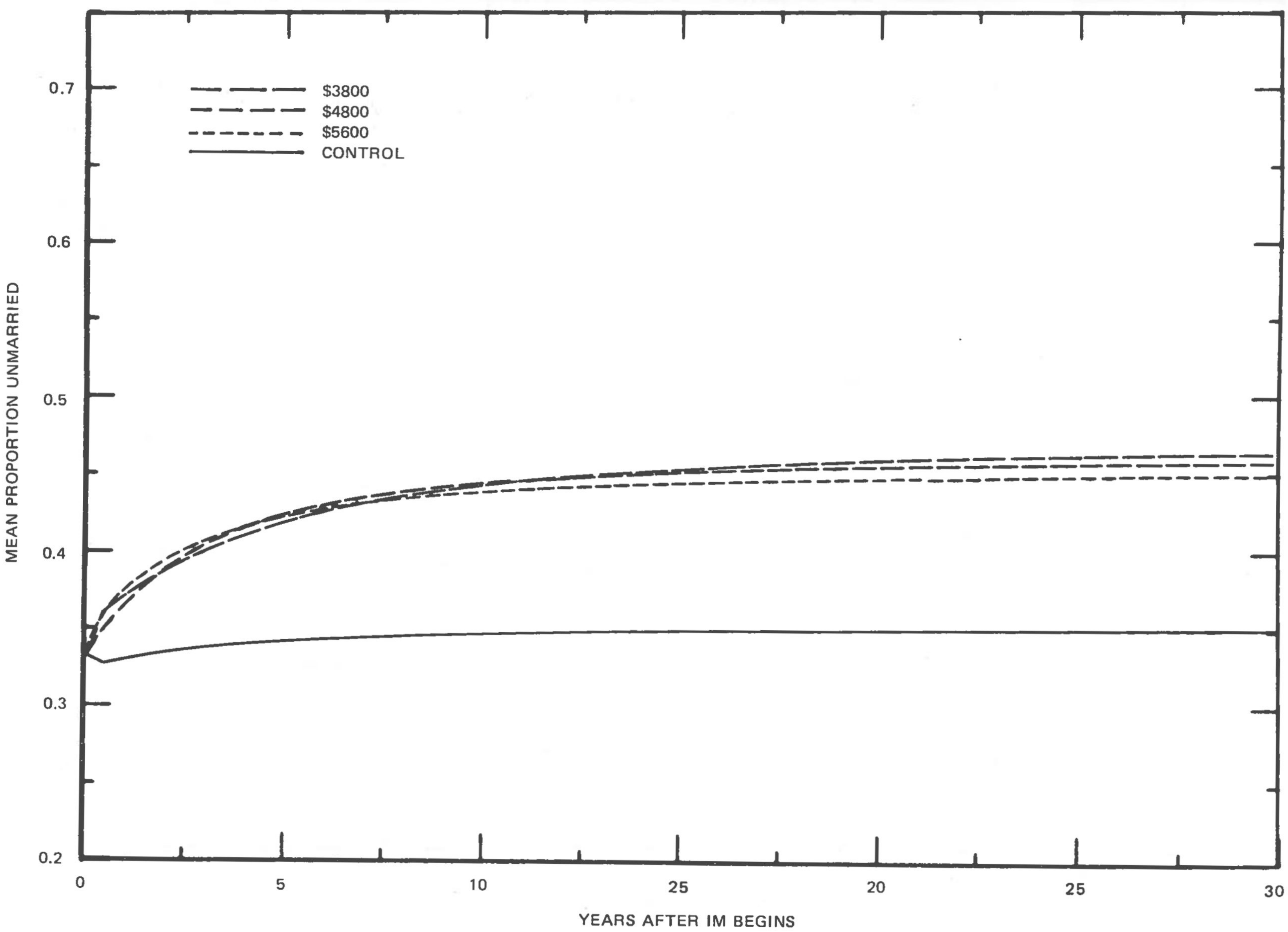

FIGURE 8 IMPLICATIONS OF HYPOTHESIS 2 (SIX-MONTH ADJUSTMENT) FOR THE MEAN PROPORTION OF UNMARRIED WHITE WOMEN OVER TIME

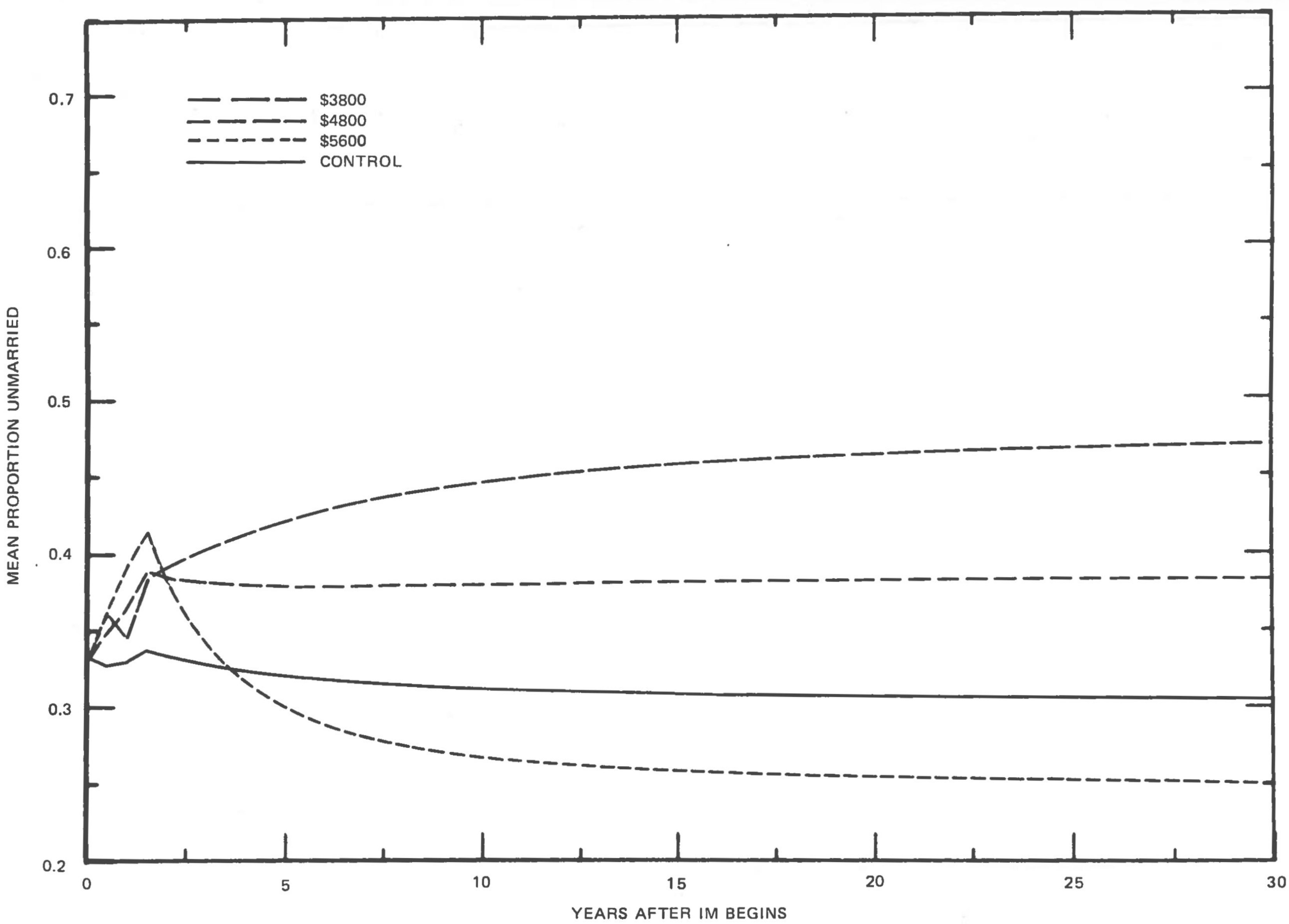

FIGURE 9 IMPLICATIONS OF HYPOTHESIS 3 (EIGHTEEN-MONTH ADJUSTMENT) FOR THE MEAN PROPORTION OF UNMARRIED WHITE WOMEN OVER TIME

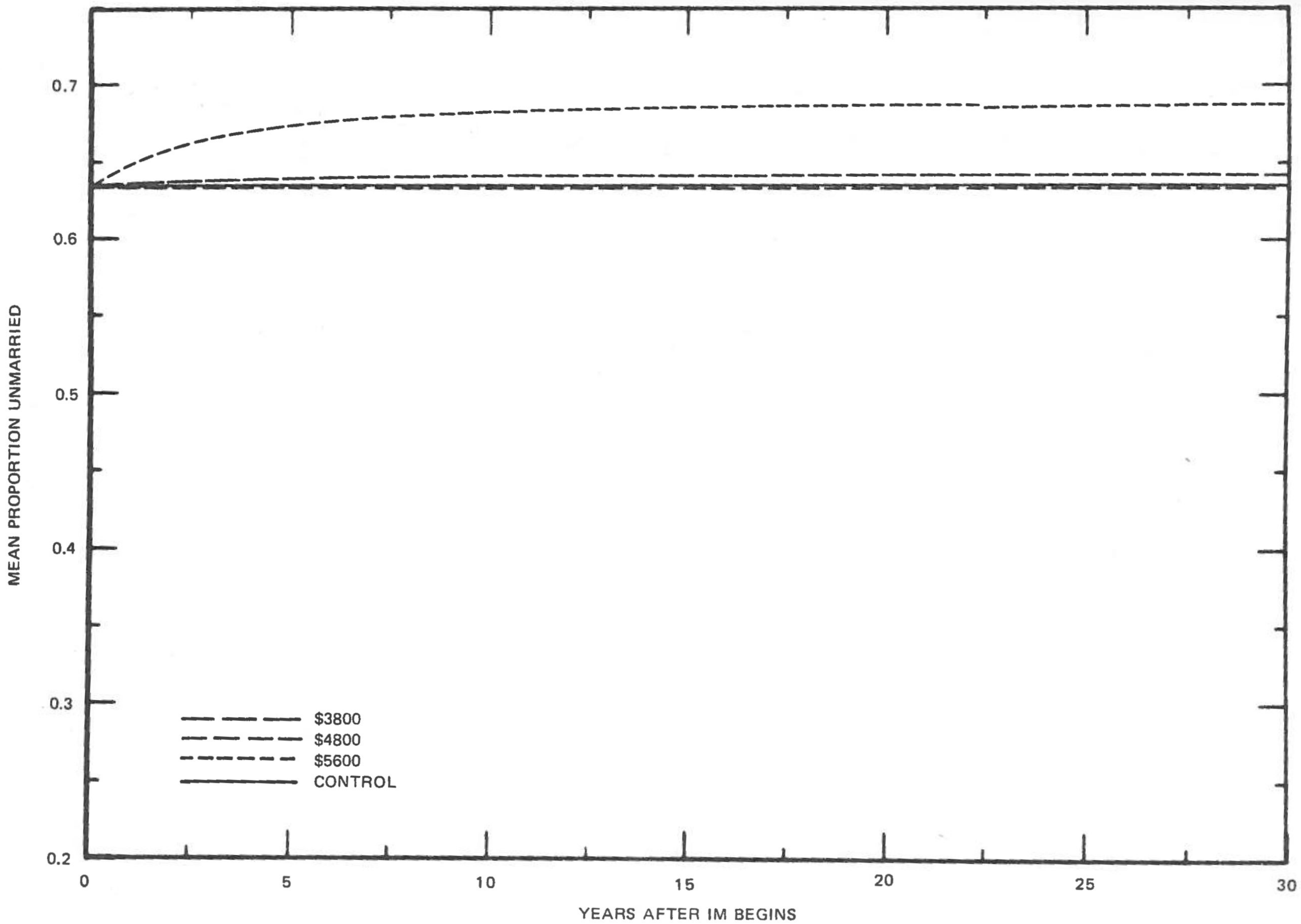

FIGURE 10 IMPLICATIONS OF HYPOTHESIS 1 (IMMEDIATE ADJUSTMENT) FOR THE MEAN PROPORTION OF UNMARRIED BLACK WOMEN OVER TIME

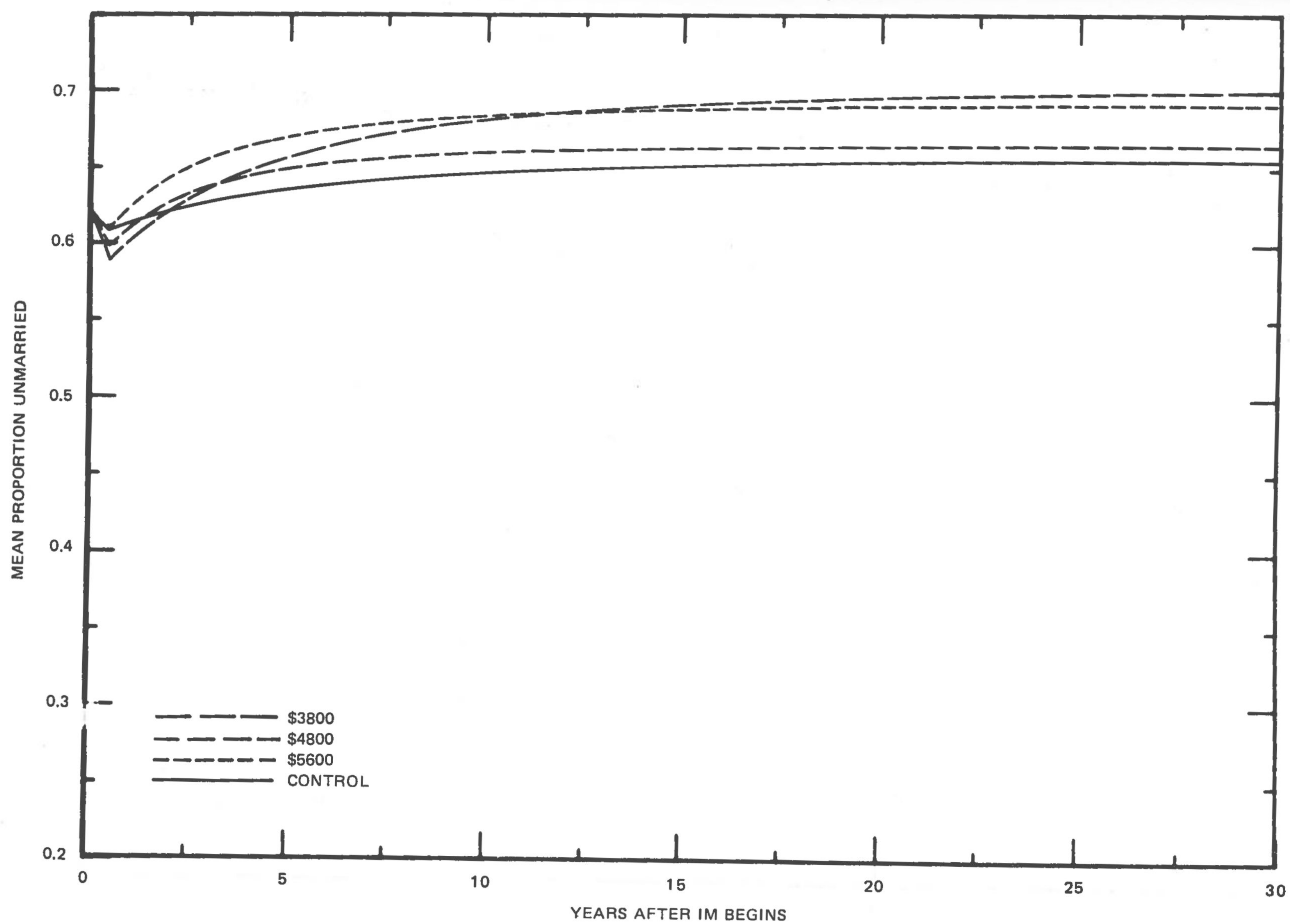

FIGURE 11 IMPLICATIONS OF HYPOTHESIS 2 (SIX-MONTH ADJUSTMENT) FOR THE MEAN PROPORTION OF UNMARRIED BLACK WOMEN OVER TIME

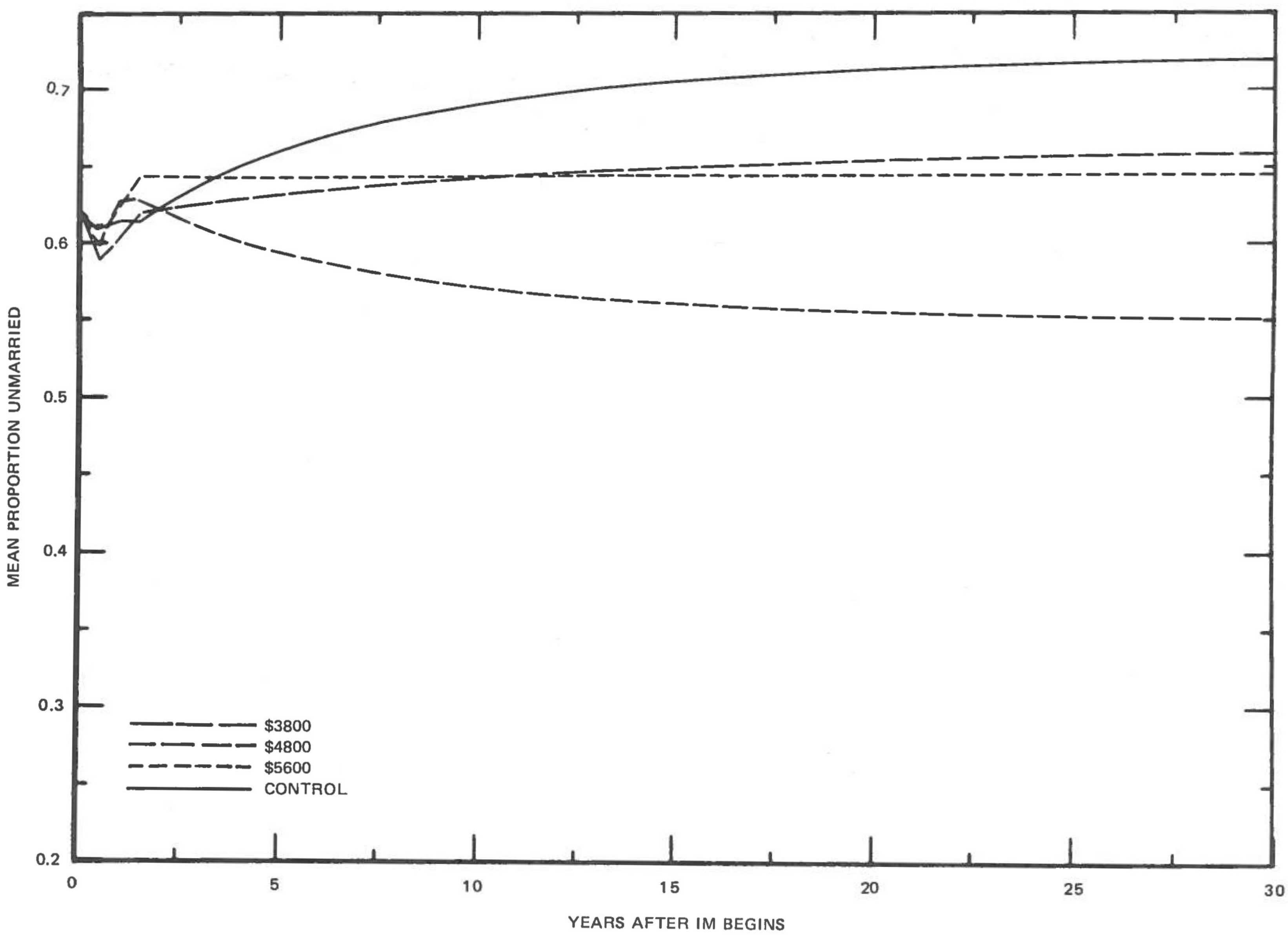

FIGURE 12 IMPLICATIONS OF HYPOTHESIS 3 (EIGHTEEN-MONTH ADJUSTMENT) FOR THE MEAN PROPORTION OF UNMARRIED BLACK WOMEN OVER TIME

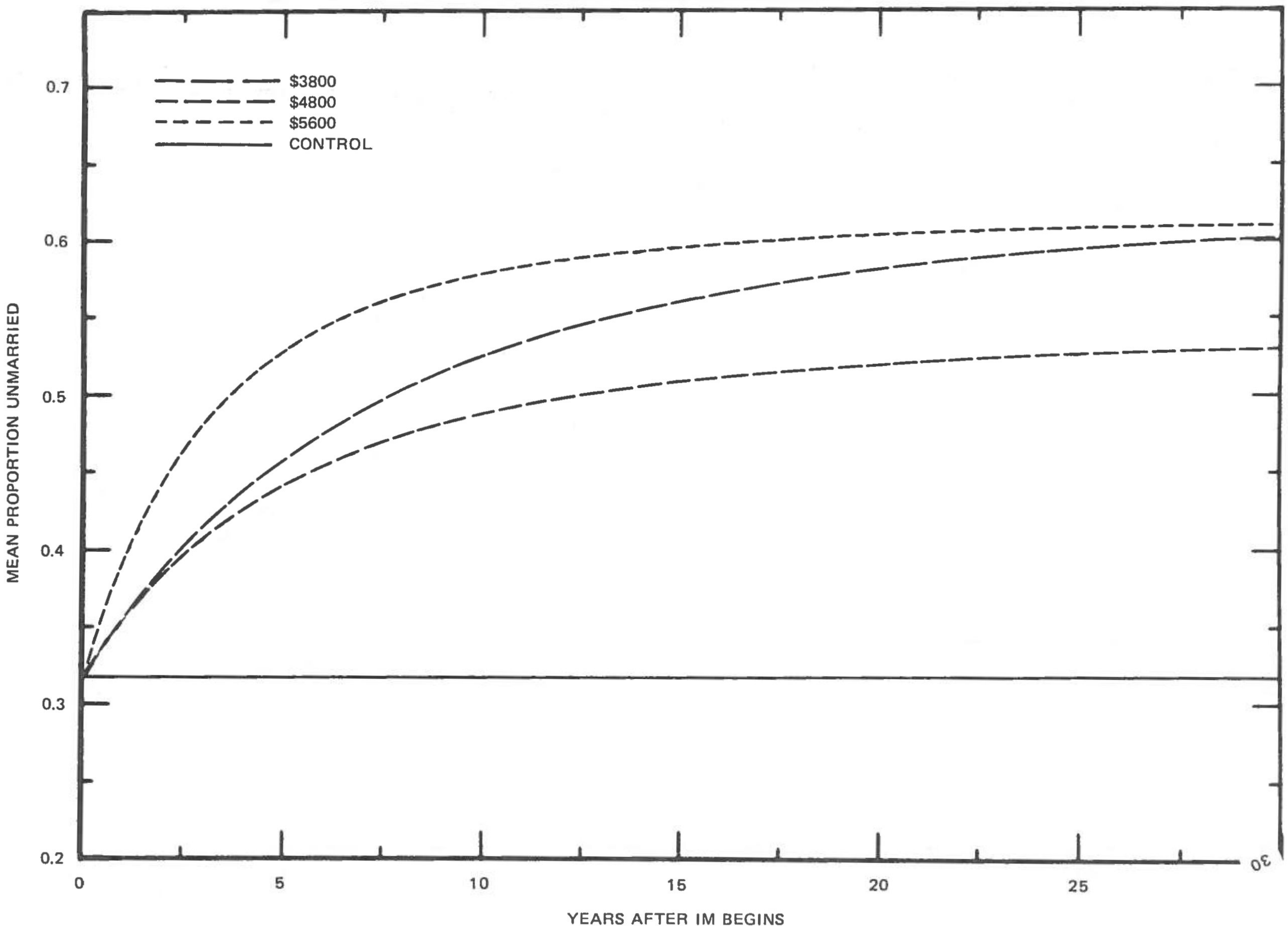

FIGURE 13 IMPLICATIONS OF HYPOTHESIS 1 (IMMEDIATE ADJUSTMENT) FOR THE MEAN PROPORTION OF UNMARRIED CHICANA WOMEN OVER TIME

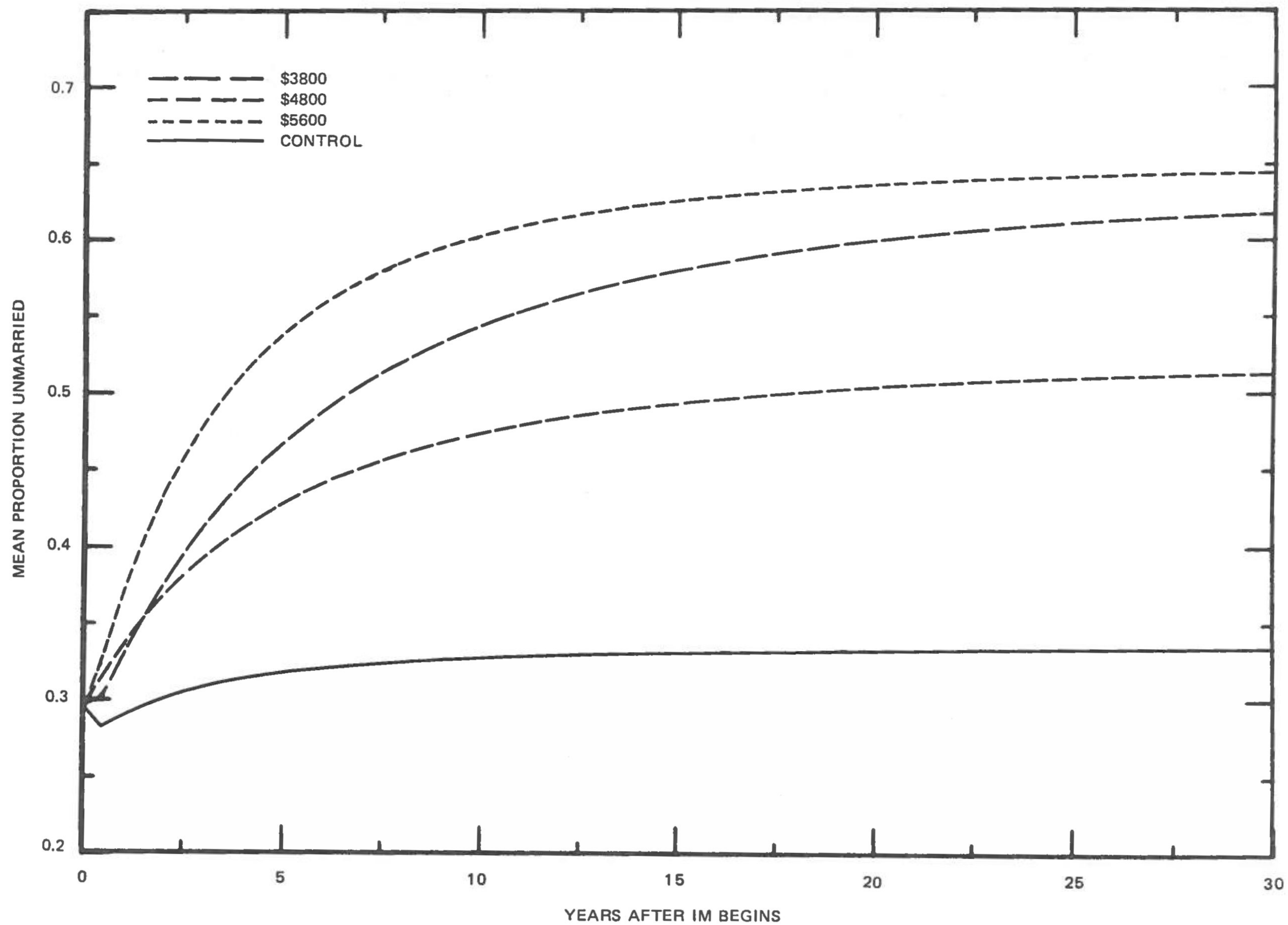

FIGURE 14 IMPLICATIONS OF HYPOTHESIS 2 (SIX-MONTH ADJUSTMENT) FOR THE MEAN PROPORTION OF UNMARRIED CHICANA WOMEN OVER TIME

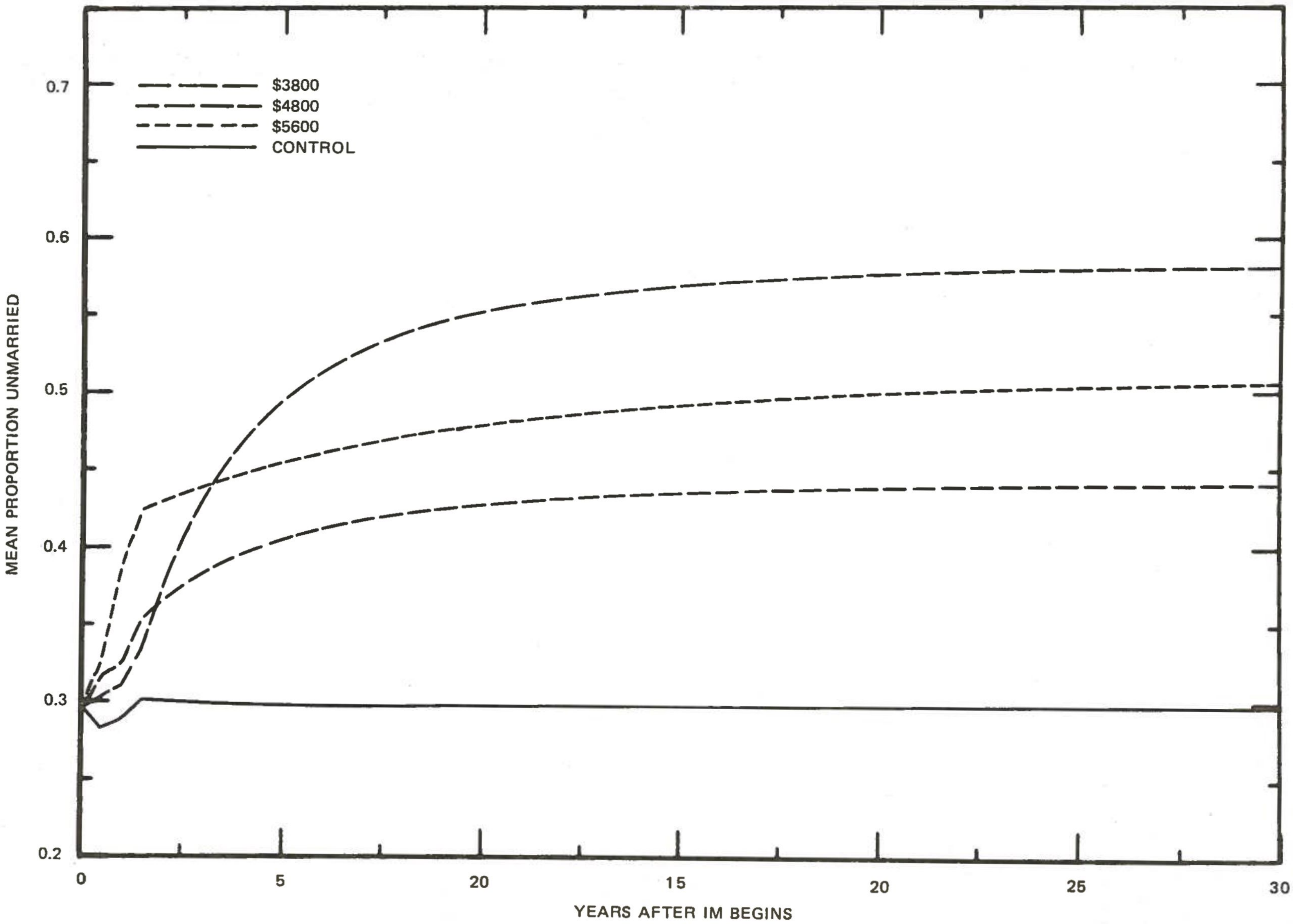

FIGURE 15 IMPLICATIONS OF HYPOTHESIS 3 (EIGHTEEN-MONTH ADJUSTMENT) FOR THE MEAN PROPORTION OF UNMARRIED CHICANA WOMEN OVER TIME

Second, these displays reveal interesting differences in the time paths, depending on whether or not there are transitory effects of IM on rates of marital status change. While the time path of the expected proportion of unmarried women steadily increases (though at a decelerating rate of increase) under Hypothesis 1 (Immediate Adjustment), this is not always the case under Hypothesis 3 (Eighteen-Month Adjustment). Though the expected proportion of unmarried White women rises more rapidly in the first year under Hypothesis 3 than under Hypothesis 1, thereafter it increases less rapidly. Moreover, for the $3800 and $4800 support levels, it actually begins to decline and eventually fall below the control proportion. Thus, if there are exaggerated but transitory effects of income maintenance on rates of marital status change, then its effects on the proportion of unmarried women in the population in the first year or so after the program begins may lead not only to exaggerated estimates of the long-term effects, but also to a misperception of the direction of the ultimate impact. Overall, these figures suggest that a national IM program should be accompanied by both careful monitoring of its effects on rates of marital status change (and not just on the current proportion of unmarried women in the population), and also considerable attention to variation in these effects with the length of time since the program began.

Hypotheses 1 through 3 assume that after some initial adjustment period the effects of IM on rates of marital status change stabilize and then continue indefinitely. However, the effects of IM could exist for a period of time and then terminate. This could occur because the IM program is terminated, or because the effects of IM on rates of marital status change are purely transitory.

Whatever the reason for their termination, it is useful to know how long it would take until the pre-IM level of the proportion of unmarried women would return. We will consider the implications of two hypotheses for this. The first, which we denote as Hypothesis 1A, is that the effects of IM do not vary within the first two years of the program and agree with our time-independent estimates, and thereafter are zero. The second, which we denote as Hypothesis 3A, is that the effects of IM vary from one six-month period to another during the first two years of the program and agree without time-dependent estimates, and thereafter are zero.

For both of these hypotheses, the predicted proportion of unmarried women during the first two years of the program will be the same as the time paths displayed in the corresponding graphs for Hypotheses 1 and 3 (see Figures 8, 10, 11, 13, 14, and 16). These graphs show that, under either hypothesis, the maximum difference between support level and control predictions after two years is 0.13 (for Chicanas on the $3800 support level). Calculations based on both Hypotheses 1A and 3A indicate that the maximum difference is reduced to 0.04 three years after termination of effects and to 0.01 eight years after the effects cease. Thus, the pre-IM levels of the proportion of unmarried women would be closely approximated within an additional three to eight years if effects of IM ended after two years. The slowness of adjustment to an altered equilibrium is again due to the comparatively long time that a woman can expect to spend in any marital status.

The slowness with which a new equilibrium level of the proportion of unmarried women in a population is attained has both desirable and undesirable features from the policy-maker's viewpoint. On the one hand, it allows decision-makers to anticipate potentially damaging changes in the steady-state proportion of unmarried women and to adjust policy before the current proportion reaches an unacceptable level. At the same time, it means that improvements will also occur slowly. In balance, the slowness of this adjustment should probably be viewed neutrally. The only real error would be to ignore it entirely and to mistake the situation within a few years after a program change as the permanent situation.

APPENDICES

Appendix A

MEANS AND STANDARD DEVIATIONS OF VARIABLES USED IN THE ANALYSIS
BY RACE AND MARITAL STATUS

(a) Married	Whites (N = 1367)		Blacks (N = 976)		Chicanas (N = 601)	
Variable	Mean	(S.D.)	Mean	(S.D.)	Mean	(S.D.)
$3800 Support	0.18	(0.38)	0.18	(0.39)	0.23	(0.42)
$4800 Support	0.23	(0.42)	0.24	(0.43)	0.23	(0.42)
$5600 Support	0.14	(0.34)	0.13	(0.33)	0.16	(0.37)
Three Year Treatment	0.38	(0.48)	0.37	(0.48)	0.44	(0.50)
Normal Income Level:						
$0-$999	0.03	(0.17)	0.04	(0.20)	0.02	(0.15)
$1,000-$2,999	0.08	(0.28)	0.06	(0.23)	0.11	(0.32)
$3,000-$4,999	0.18	(0.38)	0.16	(0.37)	0.23	(0.42)
$5,000-$6,999	0.27	(0.44)	0.24	(0.43)	0.29	(0.45)
$7,000-$8,999	0.26	(0.43)	0.27	(0.44)	0.19	(0.40)
Unclassified	0.02	(0.13)	0.02	(0.14)	0.01	(0.11)
One if on AFDC before enrollment	0.15	(0.36)	0.15	(0.36)	0.20	(0.40)
One if any children under 6 years	0.62	(0.49)	0.58	(0.49)	0.65	(0.48)
Number of children	2.22	(1.33)	2.45	(1.46)	2.66	(1.27)
Woman's age (yrs)	29.76	(9.11)	32.13	(10.08)	28.88	(8.87)
Woman's education (yrs)	11.56	(2.05)	11.25	(1.89)	9.81	(2.09)
Woman's wage ($/hr)	2.02	(0.57)	2.15	(0.57)	1.96	(0.28)
One if Denver	0.45	(0.50)	0.53	(0.50)	1.00	(0)

(b) Single	Whites (N = 902)		Blacks (N = 1046)		Chicanas (N = 407)	
Variable	Mean	(S.D.)	Mean	(S.D.)	Mean	(S.D.)
$3800 Support	0.27	(0.44)	0.27	(0.44)	0.31	(0.46)
$4800 Support	0.26	(0.44)	0.24	(0.43)	0.22	(0.42)
$5600 Support	0.10	(0.30)	0.10	(0.30)	0.12	(0.32)
Three Year Treatment	0.45	(0.50)	0.44	(0.50)	0.46	(0.50)
Normal Income Level:						
$0-$999	0.14	(0.35)	0.14	(0.35)	0.12	(0.33)
$1,000-$2,999	0.21	(0.41)	0.20	(0.40)	0.23	(0.42)
$3,000-$4,999	0.23	(0.42)	0.25	(0.44)	0.25	(0.43)
$5,000-$6,999	0.21	(0.41)	0.18	(0.38)	0.23	(0.42)
$7,000-$8,999	0.14	(0.34)	0.14	(0.34)	0.09	(0.28)
Unclassified	0.04	(0.18)	0.04	(0.20)	0.04	(0.20)
One if on AFDC before enrollment	0.36	(0.48)	0.40	(0.49)	0.43	(0.49)
One if any children under 6 years	0.42	(0.50)	0.54	(0.50)	0.57	(0.50)
Number of children	2.14	(1.20)	2.58	(1.39)	2.54	(1.29)
Woman's age (yrs)	33.17	(10.31)	32.55	(9.50)	31.37	(10.13)
Woman's education (yrs)	11.44	(1.95)	11.35	(1.85)	9.77	(2.35)
Woman's wage ($/hr)	2.12	(0.64)	2.24	(0.61)	2.00	(0.41)
One if Denver	0.44	(0.50)	0.54	(0.50)	1.00	(0)

Appendix B

EFFECTS OF OTHER CAUSAL VARIABLES
ON DISSOLUTION RATES BY RACE-ETHNICITY

Variable	Coefficient†		
	Whites	Blacks	Chicanas
Normal Income Level:			
$0-$999	4.04***	1.66	0.21
$1,000-$2,999	2.80***	1.28	0.72
$3,000-$4,999	2.16**	1.06	1.07
$5,000-$6,999	1.75*	0.69	1.06
$7,000-$8,999	1.27	0.66*	1.26
Unclassified	4.25***	1.22	1.90
One if on AFDC before enrollment	1.51**	1.40*	1.86***
One if any children under 6 yrs.	0.95	0.89	0.85
Number of children	0.91	1.01	1.00
Woman's age (yrs)	0.95***	0.95***	0.95***
Woman's education (yrs)	0.97	0.97	0.91
Woman's wage ($/hr)	1.33**	1.44***	1.42
One if Denver	0.88	1.31*	--

†Coefficients are the multipliers of the rate for a unit change in a variable.

*$0.10 \geq p > 0.05$

**$0.05 \geq p > 0.01$

***$0.01 \geq p$

Appendix C

EFFECTS OF OTHER CAUSAL VARIABLES
ON REMARRIAGE RATES BY RACE-ETHNICITY

Variable	Coefficient† Whites	Blacks	Chicanas
Normal Income Level:			
$0-$999	0.32**	0.31***	0.21**
$1,000-$2,999	0.49	0.22***	0.20**
$3,000-$4,999	0.54	0.24***	0.36*
$5,000-$6,999	0.56	0.33***	0.48
$7,000-$8,999	0.64	0.35***	0.45
Unclassified	0.66	0.35**	0.04***
One if on AFDC before enrollment	1.09	0.82	0.76
One if any children under 6 yrs.	0.84	0.81	0.55**
Number of children	1.20**	1.02	1.22*
Woman's age (yrs)	0.91***	0.95***	0.91***
Woman's education (yrs)	1.02	0.96	0.99
Woman's wage ($/hr)	1.05	0.82	0.91
One if Denver	0.93	1.15	--

†Coefficients are the multipliers of the rate for a unit change in a variable.

*0.10 ≥ p > 0.05

**0.05 ≥ p > 0.01

***0.01 ≥ p

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