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COHORT ANALYSIS OF RECENT TRENDS IN LABOR FORCE PARTICIPATION

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Abstract—The age-period-cohort accounting framework is used to describe labor force participation patterns for the sex-color groups over the interval 1969–1979, using data from the March Current Population Survey. A model with a special type of age-period interaction, in addition to main effects of the three indexing variables, is presented as a means of capturing the transitory period "shocks" which differentially influence participation odds for young and old age groups. Findings show that younger cohorts of nonblack men, nonblack women, and black women have greater "intrinsic" tendencies to participate than older cohorts, while younger cohorts of black men have lesser "intrinsic" tendencies to participate than older cohorts. The results are used to decompose across-time change into a part due to cohort effects and a part due to period effects.

INTRODUCTION

Women's labor force participation patterns have been considered by many researchers since Durand's (1948) discussion of the effects of modern occupational structure and the Second World War upon them. Examples include the treatises by Bancroft (1958), Bowen and Finegan (1969), Oppenheimer (1970), and Sweet (1973), works that are characterized by the emphasis given to women, by the focus on across-time change and trend, and by the inferences drawn about cohort differentiation with respect to a variety of indexes of labor force attachment. Robinson (1980) has recently presented labor force participation data for women covering the time interval from 1890 through the 1970s, summarizing information on completed or nearly completed cohort histories. Robinson's work effectively updates the descriptive empirical evidence now available, and with this data he assesses some of the long-term predictions made by Durand and others. Some researchers recently have attempted to describe labor force participation trends (for women) in terms of models for cohort analysis, thereby avoiding the logical difficulties associated with drawing inferences about cohort "effects" from period trends or cohort histories (Duncan, 1979; Farkas, 1977; Gustafsson, 1979).

A comparable literature on the labor force participation of men has failed to develop, either in terms of descriptive studies or in terms of more analytical studies attempting to separate age, period, and cohort effects. Researchers apparently have concluded that different cohorts of men are essentially similar in terms of their "intrinsic" degree of attachment to the labor force. Perhaps this conclusion has been based on the fact that period trends in men's labor force participation patterns do not show the same degree of secular change as those for women. But period change, or lack thereof, is insufficient evidence from which to make any definitive conclusion about *cohort* tendencies, implying that men's labor force participation patterns are also deserving of analytical appraisal. The objective of this paper is to present a comprehensive view of recent trends in labor force participation by considering men as well as women, and blacks as well as nonblacks. The reference period is the 10-year interval from 1969 through 1979. The primary concern is with estimating the "intrinsic" cohort tendencies to participate in the labor force, and then examining the way that cohort differentiation conditions the period change that can be directly observed.[1] Results will indicate that cohort factors, summarized as cohort "effects" in the models, were very influential in producing the trend observed. Results will also indicate that the conventional "discouraged worker" hypothesis, discussed by many researchers over the past three decades, cannot provide a convincing explanation of the way that labor force participation fluctuates during times of recession or growth.

There is much controversy about the methodology of cohort analysis used in this paper (see, e.g., Duncan and Winsborough, 1982), and so we begin with a rationale for the specific methods that will be used.

A RATIONALE FOR COHORT ANALYSIS OF LABOR FORCE PARTICIPATION

Cohort analysts have been careful to note that age, period, and cohort are merely indicators of other variables which actually "cause" the observed variation in the dependent variable under study. The age-period-cohort framework is properly interpreted as an accounting scheme, not a "causal model," but we hasten to add that the proper application of this framework enables a deeper analysis of proximate causal mechanisms. No claim is made here for unraveling causal mechanisms which have produced recent trends: the scope of this paper is entirely demographic in character, and the causal imagery employed is invoked only at the level of speaking of the "effects" of the age, period, and cohort indexing variables. To some readers our results will appear to traverse only a short distance from the empirical data to a convincing causal analysis, but we believe the limited objectives of our research are nevertheless justified in view of present knowledge. Duncan (1979), among others, does attempt to go beyond the accounting framework to a form of analysis that rests upon the specification of causal mechanisms which underlie the indexing variables. Our view is that the selection of the proper causal variables to be considered in a modeling procedure is a most difficult theoretical task, one that is at least as difficult as applying the age-period-cohort accounting framework.

The age variable is merely biological, but aging is closely related to experience and schooling, the two most important factors that determine "offered wage" rates for prospective workers, and hence participation tendencies, according to the human capital view (Mincer, 1962). Aging is also related to various status-role transitions that accompany a cohort's progression through an inherently age-graded life cycle (e.g., see Winsborough, 1978). Age differentials in participation (net of period and cohort) are probably not produced to any great extent by purely biological maturation. But once this much has been acknowledged, age differentials can be explained in causal terms only by properly specifying *and measuring* the life cycle events that account for age variability in a more precise causal sense. A method that incorporates the age variable explicitly into a model allows the determination of the combined influences of all the variables for which age is a proxy. But a method which incorporates the more proximate causal variables instead of the age variable must first be justified by carefully considering the proper specification of these variables, their functional relationship to the response, and their operational measurement.

The period variable is merely composed of observational units, but there are various meanings that are associated with it in the context of labor force participation. The "effects" of the period variable (net of age and cohort) could be due to the labor market variables (measured currently or in some specified lag scheme), which cause transitory "shocks" in labor force participation. But these period "effects" could be due as well to social movements not so closely associated with labor market conditions; the women's movement is certainly a period-specific phenomenon that could be relevant. The period "effects" could be due to many similar causal mechanisms as well, and these could all be considered jointly, if there were sufficient theoretical knowledge indicating which variables should be included, how these variables should be measured, and how these variables are functionally related to participation. Farkas (1977) and Duncan (1979) are led to conceive of period effects as those which are closely associated with certain (current) indicators of the condition of the labor market. But the decision regarding *which* indicators should be used poses a vexing problem, and we can never be certain that *all* of the period effect is captured by the set of indicators used. That these researchers conceive of period effects as *only* those which are related to the (current) economic condition of the labor market, excluding, say, period-specific social movements like the women's movement, is a serious problem. (Farkas and Duncan actually develop their analytic strategies in order to circumvent the identification problem in cohort analysis. But suppose that the true effect of the period variable P were to be included in an additive model via a term bP, and that this term were *exactly* equal to $\Sigma_i\, b_i\, x_i$, where the x_i are the relevant variables which cause period change. Since $C = A - P$ where A is age and C is cohort, we must have $C = A - (\Sigma_i\, b_i\, x_i)/b$. That is, the additive age-period-cohort model would be underidentified if causal variables which captured *all* of the period effect were used in place of the period variable. If only a subset of the x_i's were used in some model, then this model would be identifiable, but the specification error made in selecting only a subset of the relevant x_i's implies that there would be bias in the estimated effects.)

The cohort variable is likewise one whose effects on participation can be viewed as a composite of more proximate causal effects. Cohort size (Easterlin, 1978, 1980; Welch, 1979), cohort-specific socialization influences, peculiar cohort experiences (the Depression, the Vietnam War), and human capital endowments (Weiss and Lillard, 1978) are only a few of the cohort attributes that might be relevant (also see Glenn, 1977; Ryder, 1965, 1969). Heckman's (1974) analysis of the determinants of the "shadow price" of time (asking wages) for nonparticipants also reveals a possible role for cohort-specific causal agents. Of the several determinants of "shadow prices" that he considers, the possible hours that a nonparticipant might supply to the work force, the wages of those on whom the nonparticipant depends, and household asset income might all be conceived at least partly as cohort-specific causal agents. Once again there are the problems of which cohort characteristics are really the proper causal agents, how they should be measured, and how they are functionally related to participation.

Because of the difficulties in specifying what it is that the age, period, and cohort variables actually represent in terms of proximate causal agents, an analysis that makes as few assumptions as possible about the causal agents is certainly legitimate. The labor force participation variable is analyzed herein solely from the age-period-cohort accounting framework (Fienberg and Mason, 1978; Pullum, 1977). On the basis of the preceding remarks, we do not here abandon this framework in favor of the approach implicit in the work of Farkas (1977) or Duncan (1979). As is well known, there is an identification problem that must be squarely faced if unbiased parameter estimates are to be obtained. I believe that some of the methods discussed will be useful in resolving this issue in an acceptable way, at least in the context of labor force participation analysis.

After data sources and methods are discussed, age, period, and cohort effects will be estimated for the sex-color groups. One model is used which allows for a kind of age-period interaction, in addition to the main effects of age, period, and cohort. The parameters of this model are used to assess both the "intrinsic" cohort tendencies to participate and the period-by-period shocks on the participation of young and old workers. We conclude with a discussion of findings and some conjectures about strictly demographic approaches that might be used to explain the pattern of results obtained.

DATA

The data were obtained from the March Current Population Survey for the years 1969–1979. To ensure that the sampling weights were being used properly, statistics from the retrieved data were compared with published figures derived from the same data (U.S. Bureau of Labor Statistics, 1969–1979). The contingency tables subjected to analysis were obtained in a way that ensured that the actual sample sizes were approximated by the period marginals.

The black-nonblack dichotomy was used to code the color variable, a rationale for which is that nonblack nonwhites are probably more similar to whites than to blacks in terms of labor force participation patterns. Because nonblack nonwhites are only a small fraction of the nonblack group used here, the results for nonblacks should be nearly comparable to those that would be obtained for whites.[2]

The standard definition of the labor force was applied to dichotomize the population into participants and nonparticipants (U.S. Bureau of the Census, 1978). Other definitions could certainly be used (Clogg, 1979; Clogg and Sullivan, 1982; U.S. Commission on Employment and Unemployment Statistics, 1979), but the standard definition was used to make our results comparable to previous research. The age range chosen for analysis was 20–64 inclusive; the very young (16–19 years of age) and the very old (65 and over) were omitted because of the marked irregularities in labor force participation patterns for these groups.

The Current Population Survey is actually obtained by a very complicated form of sampling, but no attempt was made here to use some "sampling design" factor to adjust chi-square statistics measuring the agreement between observed and estimated expected frequencies. Fay (1979) has presented some jackknife procedures for analyzing individual CPS files which can be used to obtain appropriate statistical tests of log-linear models, but we believe that these methods would be very complicated to apply to an 11-year *series* of CPS files. Another complication arises from the fact that approximately one-half of the persons sampled at any given period were also sampled at the immediately preceding period. Fay's procedures apparently do not allow us to take this feature of the sampling scheme into account. For these reasons the chi-square statistics are to be taken merely as indexes of fit. However, most of the results are dramatic enough that sampling error is not a serious factor, and the samples are quite large. Each of the periods of observation contains over 10,000 blacks and over 50,000 nonblacks in the relevant age ranges. The sample sizes are certainly large enough to allow analysis of single-year age (and cohort) data (cf. Farkas, 1977, p. 34).

Other data sets have been used to analyze labor force participation in the recent U.S. experience, some of which are repeated cross-sectional in form (like the CPS) and some of which are longitudinal. It is now apparently feasible to use the year-to-year matching subfiles of the

CPS to obtain partially longitudinal data relevant for the analysis of "flows" to and from the labor force (Eck, 1980; U.S. Bureau of Labor Statistics, 1980). With the possible exception of these CPS matched data, the data used here appear to be the best currently available for analyzing participation trends in the recent U.S. experience.

THE MAIN EFFECTS MODEL

The models used to estimate the effects of age, period, and cohort on labor force participation are analogous to those used in a similar context by Clogg (1979). They are identical in most respects to those discussed by Pullum (1977, 1980) and Fienberg and Mason (1978), and these in turn are related to the diagonals parameter model of Goodman (1972). We let Ω_{ijk} denote the expected odds that a member of the ith age, jth period, and kth cohort will be participating in the labor force, for $i = 1, \ldots, I$; $j = 1, \ldots, J$; $k = 1, \ldots, K$ (with $K = I + J - 1$, the number of cohorts in a cross-classification with I age categories and J periods). For the situation here all indexes refer to single years, and the cohort indexes range from the youngest to the oldest cohort. There is a linear relationship between the indexes of the form $k = i - j + J$, and this accounts for the identification problem encountered in a model with main effects of the three indexing variables on the Ω_{ijk}.

The model that we consider first is

$$\Omega_{ijk} = \alpha\beta_i\gamma_j\delta_k, \tag{1}$$

where α is a constant and β_i, γ_j, δ_k denote age, period, and cohort parameters, respectively. There are $1 + I + J + K$ parameters in (1), and four (not three) restrictions must be imposed on them in order to achieve identifiability. For any given set of (unidentifiable) parameters in (1), define

$$f = (\delta_{k'}/\Delta\delta_K)^{(K-k')^{-1}}, \tag{2}$$

for some cohort category $k' \neq K$ and some positive constant Δ. The parameters in (1) are then transformed by taking

$$\tilde{\alpha} = \alpha\beta_I\gamma_J\delta_K f^{J-1}, \tag{3a}$$

$$\tilde{\beta}_i = (\beta_i/\beta_I)f^{I-i}, \tag{3b}$$

$$\tilde{\gamma}_j = (\gamma_j/\gamma_J)f^{j-J}, \tag{3c}$$

and

$$\tilde{\delta}_k = (\delta_k/\delta_K)f^{k-K}. \tag{3d}$$

It can be verified that

$$\Omega_{ijk} = \tilde{\alpha}\tilde{\beta}_i\tilde{\gamma}_j\tilde{\delta}_k, \tag{4}$$

using the relationship $k = i - j + J$, and that four independent restrictions have been imposed:

$$\tilde{\beta}_I = \tilde{\gamma}_J = \tilde{\delta}_K = 1, \tag{5a}$$

and

$$\tilde{\delta}_{k'} = \Delta. \tag{5b}$$

In addition to the three trivial restrictions of (5a) which serve to define the *effects* of age, period, and cohort category on the Ω_{ijk}, the k'th cohort effect has been set equal to Δ. Pullum (1980) actually considers situations where $\Delta = 1$, i.e., where the effect of the k'th cohort is assumed to be equal to that of the Kth cohort; Fienberg and Mason (1978) impose similar kinds of equality restrictions. But in view of these algebraic relationships, it is not necessary to assume equality of effects. Using algebraic identities like these, it is not difficult to derive the biases in the parameters which result from making an error of any specified amount in choosing Δ, a fact that will be exploited later. The reader is referred to Clogg (1982a) for algebraic details.

The necessary inputs to the model of (1) are thus two: the cohort category k' whose effect is to be restricted and the value Δ that is to be assigned to this effect. These specifications are very important. It can be shown that an error in choosing Δ produces systematic bias in all of the effects in the model, that this bias differs according to the cohort category k' chosen (with the bias being more severe when k' is not close to K, for a given error in specifying Δ). This bias depends as well on the number of age categories and the number of period categories used (a "design" effect). Therefore special care must be taken to choose k' and Δ. Here we choose $k' = K - 1$ (the next-to-oldest cohort), and we choose Δ in a manner that reflects a priori assumptions about intercohort trends in labor force participation. To assist in choosing Δ, note that if Ω_{ijK} is the value of the odds for the oldest cohort in any given age i and period j, then $\Omega_{ijk'}{}^* = \Delta\Omega_{ijK}$ is the "pseudo-odds" that describes the expected odds, under the model, if cohort k' could actually be observed for age i and period j. The ratio of the second quantity to the first is Δ, and so prior assumptions about the relative magnitude of these odds can be used to obtain a plausible value for Δ.

A MODEL WITH AGE-PERIOD INTERACTION

The model of (1) allows for main effects of age, period, and cohort, but it does not allow for any interactions among these variables. The model thus makes assumptions that are worth elaborating more carefully, at least in so far as it is to be applied in the labor force participation context.

1. The age effects, which indicate the age-graded character of labor force participation, or "life cycle" effects (see Winsborough, 1978), are assumed to be homogeneous across periods and cohorts. Thus no allowance is made for the possibility that period-specific shocks alter the participation behavior of different age groups in different ways. This assumption will appear quite implausible to labor force researchers for at least two reasons. First, this assumption is not consistent with notions about the changing volume of "discouraged workers" (see Flaim, 1972), who are measured as being *outside* the labor force. Discouraged workers are principally composed of the young and the old; their numbers fluctuate considerably with economic growth and recession. Using the standard definition of the labor force implies that *measured* participation rates will fluctuate in proportion to the changing numbers of discouraged workers. An assumption of null age-period interaction is thus inconsistent with the available knowledge about discouraged workers. Second, retirement patterns change with time, implying that a special allowance should be made for age-period interaction for older workers. Decisions to retire might be influenced by general labor market conditions, for example, and this implies an interaction of age and period for older workers.

2. A second assumption is that the period effects, which refer to the transitory labor market shocks as well as to other period-specific influences, are homogeneous across age and cohort. There is only one parameter that governs the entire influence of a given period's socioeconomic condition on participation, and this parameter exerts its influence independently of age or cohort. The implausibility of this assumption in regard to null age-period interaction has been commented on previously.

3. A third assumption is that the cohort effects of the model, referring to "intrinsic" cohort tendencies to participate ("separate" from age or period effects and determined exogenously to the periods of observation), are assumed to be homogeneous across age and period. These cohort tendencies do not change as the cohort ages; they do not adjust to the peculiar period experiences which cohorts encounter as they age. It can be noted that a period-cohort interaction appears to describe what Durand (1948, pp. 123–124) had in mind when discussing changing participation patterns of women subsequent to the Second World War. His conjecture was that the abnormally high work rates of young adult women during the war "transmitted" itself across the life cycle, leading to

increased rates of participation in later ages. If period-cohort interaction of this kind exists, then the model of (1) is clearly misspecified. But for the relatively short time interval considered herein, we can think of no peculiar period shocks as dramatic as the Second World War, and so no attempt is made to incorporate period-cohort interactions into the model. As regards the assumption of no cohort-age interaction, we can likewise find no compelling reasons to question the assumption. However, the possibility of age-period interactions cannot be ignored, because of fluctuations in the numbers of young and old discouraged workers and because of changing retirement patterns.

Fienberg and Mason (1978, p. 5) note that it is difficult to add interactions to the main-effects model in (1) because of estimability (or identifiability) problems. However, an age-period interaction is required here in order to take proper account of the way that participation is responsive to labor market conditions. Conditions of "full employment" should increase *measured* participation at both ends of the age range (but not so much at the middle ages), while recessions should have just the opposite effects. We can also surmise that retirement trends, certainly partially responsive to current economic conditions, will also require treating the older ages differently from the young and middle ages.

We assume that there exist ages i' and i'' such that for $i' < i < i''$ no age-period interaction exists. For our analysis we actually take $i' = 34$ and $i'' = 50$, the assumption being that no age-period interaction exists for the prime working ages 35–49. (We did consider another specification where $i' = 29$ and $i'' = 54$, and we found results that differed only slightly from this specification.) The choice of these ages is warranted, in our view, because the 35–49 age range is precisely that where unemployment is the lowest, where job security is the greatest, and where discouraged workers are virtually nonexistent. Workers in these ages are typically those with much experience, and their job skills are not yet outmoded. On the other hand, labor force participation is still in the process of being established for many persons under age 35, and it is in the process of being terminated for many persons over age 49. The assumption is then that the transitory period shocks on participation will exert their influences in different ways for each of the three age groups (20–34, 35–49, 50–64).

The model can be described by

$$\Omega_{ijk} = \begin{cases} \alpha\beta_i\gamma_j\delta_k(1/\theta_{1j})^{i-c}, & \text{for ages 20–34,} \\ \alpha\beta_i\gamma_j\delta_k, & \text{for ages 35–49,} \\ \alpha\beta_i\gamma_j\delta_k\theta_{2j}{}^{i-c}, & \text{for ages 50–64.} \end{cases} \tag{6}$$

where c is a constant (equal to 23, the age index that corresponds to age 42). The θ_{1j} and θ_{2j} refer to age-period interactions for period j, for young and old workers, respectively. The age index i appears as a power of the thetas, incorporating the added assumption that there is a *geometric* relationship between the period-specific shocks θ_{1j} and θ_{2j} and age, for the relevant age ranges. Note that $\theta_{1j} > 1$ implies that the period shock operates to increase the participation of younger persons relative to middle-aged persons, with a stronger relationship being exerted for the youngest persons; with $\theta_{1j} < 1$ the period shock operates to decrease the participation of younger persons relative to middle-aged persons. Similar comments apply to the θ_{2j} parameters governing the period shocks on older workers. The shocks are the greatest for the extreme ends of the ages 20–64, and they taper off to nothing in the center of the distribution.[3] By virtue of the fact that there is no age-period interaction for ages 35–49, the model of (6) is identifiable, once the necessary restrictions have been imposed on the main-effects parameters.

To estimate the parameters by the method of maximum likelihood, an algorithm presented by Goodman (1972) was applied. The algorithm is based on Newton's elementary (one-dimensional) numerical procedure (see Clogg, 1982b; Goodman, 1979), and it is related to the Newton-Raphson procedures discussed by Haberman (1978). A set of (unidentified) parameter estimates was first obtained, and then the main-effects parameters were transformed into identifiable parameters using (2) and (3a)–(3d). (Note that the J θ_{1j} and the J θ_{2j} parameters are identifiable by virtue of the restricted age ranges to which they apply.)

RESULTS

Indexes of Fit

Table 1 presents the degrees of freedom and indexes of fit for several models applied to the sex-color groups. One index is the chi-square statistic (L^2) based on the likelihood-ratio criterion, while the other is the index of dissimilarity (D) between observed and estimated expected frequencies. The symbols A, P, C, and AP^* denote parameters included in the models; the model of (1) is described by (A, P, C) while the model of (6), incorporating a kind of age-period interaction, is described by (A, P, C, AP^*). When the sample sizes are taken into account, it is clear that the more comprehensive models perform rather well, although the participation experience of blacks is not accounted for as well as is that of whites. (We have D = .66 percent, 1.17 percent for nonblacks, but D = 2.91 percent, 3.38 percent for blacks, for the model of (6).)

Because of the logical relationships among the indexing variables A, P, and C (with $C = A - P$), it is not appropriate to use the indexes of Table 1 in some ad hoc way to determine whether or not certain effects *exist*. Such a model-building recipe might be suited for cases where the independent variables are orthogonal, but it is potentially misleading when there are interdependencies among the independent variables of the kind encountered here. Consider the three models (A), (A, P) and (A, C) as an example. For each sex-color group, the (A, C) model does better than the (A, P) model. If some blind model-fitting strategy were used, the researcher might conclude that period effects are essentially null, and that the lion's share of what can be explained is explained by age and cohort. But this does not imply that if the period effects were added to the model (A, C), the coefficients would be nil as compared to, say, a criterion of standard error. Similarly, the difference in chi-square statistics $[L^2(A, C) - L^2(A, P, C)]$ might be moderate, as it is in fact, but this is not an appropriate assessment of

Table 1.—Indexes of Fit

Model	df	Nonblack Males L^2	Nonblack Males D	Nonblack Females L^2	Nonblack Females D	Black Males L^2	Black Males D	Black Females L^2	Black Females D
Null	494	38,766	6.73%	16,380	6.62%	3,793	7.76%	2,435	7.31%
A	450	1,909	1.40	4,861	3.63	1,032	3.83	835	4.47
P	484	38,328	6.70	13,738	5.73	3,576	7.51	2,321	7.05
A, P	440	1,489	1.20	2,106	2.40	811	3.40	730	4.09
C	440	13,967	3.85	5,477	3.77	1,680	5.12	1,208	5.22
A, C	396	764	.81	609	1.33	776	3.30	587	3.70
P, C	430	12,800	3.46	4,933	3.49	1,347	4.43	1,121	4.90
A, P, AP*	418	708	.79	921	1.65	769	3.31	653	3.85
A, P, C	387	580	.71	519	1.21	645	2.96	541	3.48
A, P, C, AP*	365	509	.66	471	1.17	620	2.91	517	3.38

the contribution of the P variable either, again because of the logical relationships among the three variables. (The standard result on partitioning hypotheses H and H^* states that $L^2(H|H^*) = L^2(H) - L^2(H^*)$ is a chi-square test of H *assuming* that H^* is true. Here, H and H^* are difficult to distinguish, and $H|H^*$ is difficult to distinguish from H^*, with $H = (A, C)$, $H^* = (A, P, C)$. Similar comments apply to the other conditional tests which might be formed from the L^2 values in Table 1.)

In sum, theoretical considerations must govern the choice of a model, and we believe that there is enough evidence to suggest that each of the variables A, P, and C exert main effects on participation, and the argument of the preceding section indicates that the AP^* term is required as well. In the next sections certain parameter estimates will be presented, and the magnitudes of these quantities will serve as another basis from which the model (A, P, C, AP^*) can be defended.

Table 2 presents some measures based on the indexes of Table 1 which are somewhat analogous to the usual kinds of R^2 statistics. Evidently, the (A, P, C, AP^*) model of equation (6) performs very well for nonblacks, but less well for blacks. It can be noted that the AP^* term contributes modestly to the (A, P) model, but it contributes negligibly to the (A, P, C) model. Some might take this as evidence that the AP^* terms do not need to be included, but we remind the reader that the AP^* term has some "shared variance" with the C term, since both are special kinds of A–P interaction. We believe that the AP^* terms are required for theoretical reasons, and as we shall later demonstrate, the magnitudes of the estimated age-period interactions are not trivial. In addition, it can be shown that the estimated *cohort effects* are quite different under the model (A, P, C) and the model (A, P, C, AP^*). Because the latter model "partials out" age-period interactions which should not be confounded with cohort effects, it is this model that is used to motivate all subsequent discussion.

Choosing Identifying Restrictions

The procedure used to identify the parameters is now outlined in some detail, bringing to light the prior information that we thought should be incorporated into the analysis. Actually, for the case of nonblack males it was concluded after the analysis that our prior assumptions were questionable, but we air them here nevertheless. Certain quantities will be considered later which do not depend so much on the exact truth of these prior assumptions, so that even if the reader takes exception to our reasoning here, the effort will not be altogether superfluous. For each sex-color group, the 54th cohort (aged 63 in 1969) was singled out

Table 2.—Some R^2-Type Measures Using the Null Model as a Baseline

Model	Nonblack Males		Nonblack Females		Black Males		Black Females	
	L^2	D	L^2	D	L^2	D	L^2	D
A, P	.96	.82	.87	.64	.79	.56	.70	.56
A, C	.98	.88	.96	.80	.80	.57	.76	.49
P, C	.67	.49	.70	.47	.64	.43	.54	.33
A, P, AP*	.98	.88	.94	.75	.80	.57	.73	.47
A, P, C	.99	.89	.97	.82	.83	.62	.78	.52
A, P, C, AP*	.99	.90	.97	.82	.84	.63	.79	.54

NOTE: Quantities in the "L^2" columns are based on the likelihood-ratio chi-square statistic. Quantities in the "D" columns are based on the index of dissimilarity.

for the purpose of imposing the identifying restriction; the 54th cohort effect is thus restricted in relationship to the 55th cohort (aged 64 in 1969).[4]

First consider nonblack males. Our prior assumption about them is that more recent cohorts will have lesser intrinsic tendencies to participate than older cohorts. Such a line of reasoning is certainly consistent with the observed period change for this group, although we hasten to add that the observed period change was not the criterion used to arrive at our assumption. Rather, the assumption made here seems to go hand in hand with the accepted notion that younger cohorts of nonblack females have greater intrinsic tendencies to participate; the labor force cannot expand without bounds; therefore an adjustment in nonblack male cohort tendencies of this kind seems worth entertaining. The value of Δ assigned to the 54th cohort effect is .98: the assumption is that the participation odds for the 54th cohort is 98 percent that of the 55th cohort. To see what this figure implies for labor force participation rates, suppose that $p = .90$ is the "intrinsic" labor force participation rate for the 55th cohort. (This value is approximately the average participation rate for nonblack males over the interval 1969–1979, and all that we require is that this value be somewhere near the "true" labor force participation rate for this cohort.) Now with $\Delta = .98$, the rate $\tilde{p} = p + \varepsilon$ for the 54th cohort is .18 percent smaller than that of the 55th cohort. For values of p in the neighborhood of .90, $\Delta = .98$ corresponds to about .2 of a percent decline in the participation rate. If intercohort change were constant, in 55 years there would be about a 10 percent drop in the participation rate (to about 80 percent).

For nonblack females, the necessary prior information is much more clearcut: newer cohorts of women must have a greater intrinsic tendency to participate than older cohorts. Such a prior assumption is consistent with earlier research and theory (see the references cited in our introduction). The value of Δ is taken at 1.02, and if the oldest cohort's rate of participation is $p = .55$, then this value of Δ implies that the rate $\tilde{p} = p + \varepsilon$ for the 54th cohort is .49 percent larger than p. If intercohort change were constant, then over 55 cohorts there would be a 27 percent increase in the participation rate (to an average of 82 percent, or approximately the figure that we earlier entertained as the ultimate rate for nonblack males). Suffice it to say that $\Delta = 1.02$ appears credible for nonblack females (and $\Delta = 1.00$ would not be credible).

For black males, we believe that the weight of the evidence supports a view that newer cohorts have lesser intrinsic tendencies to participate than older cohorts.[5] We choose $\Delta = .99$ to reflect these assumptions, and note that if $p = .85$ is the 55th cohort's intrinsic rate of participation, then the rate for the 54th cohort will be .13 percent less than this. If the intercohort change were constant over 55 cohorts, this would imply a 7 percent drop in labor force participation.

For black females, nearly the same prior information used for the case of nonblack females can be applied, although the pace of change for black women is probably not so great as for nonblack women. We choose $\Delta = 1.01$, and note that if $p = .60$ is the 55th cohort's intrinsic rate of participation, then the rate for the 54th cohort is .24 percent greater than that of the 55th. Assuming constant intercohort change, the rate of participation would rise 13 percent over 55 cohorts.

Estimated Cohort Effects

Table 3 presents the estimated cohort effects (on the logarithmic scale) for each of the sex-color groups. The values for $\log \hat{\delta}_{54}$ (−.02, +.02, −.01, +.01) are merely the restrictions that we attempted to defend earlier. The other quantities measure the effect of membership in a given cohort when contrasted with membership in the oldest cohort.

Table 3.—Cohort Effects (Logarithmic Scale) Obtained from the Model Incorporating Age-Period Interaction

Cohort	Nonblack Males	Nonblack Females	Black Males	Black Females
1	2.61	4.23	-10.19	12.60
2	2.66	4.16	-10.32	12.07
3	2.50	4.14	-10.41	11.79
4	2.39	4.19	- 9.92	11.74
5	2.35	4.08	- 9.82	11.55
6	2.16	4.08	- 9.86	11.43
7	2.10	4.03	- 9.45	11.12
8	2.01	3.92	- 9.39	11.03
9	1.78	3.89	- 9.17	10.93
10	1.74	3.83	- 8.96	10.63
11	1.70	3.71	- 8.92	10.35
12	1.64	3.59	- 8.64	10.11
13	1.56	3.49	- 8.31	9.83
14	1.46	3.39	- 8.26	9.54
15	1.56	3.23	- 8.05	9.39
16	1.59	3.14	- 7.76	8.99
17	1.58	3.03	- 7.62	8.79
18	1.61	2.95	- 7.45	8.58
19	1.24	2.82	- 6.72	8.37
20	1.43	2.79	- 6.61	8.08
21	1.47	2.69	- 7.00	7.68
22	1.39	2.50	- 6.72	7.64
23	1.33	2.44	- 6.20	7.26
24	1.26	2.36	- 6.20	7.18
25	1.16	2.27	- 6.03	6.91
26	1.08	2.10	- 6.16	6.56
27	0.98	2.03	- 5.54	6.44
28	1.11	1.97	- 5.22	6.22
29	0.90	1.84	- 5.35	6.07
30	0.95	1.72	- 4.96	5.69
31	0.86	1.67	- 4.75	5.38
32	0.96	1.68	- 4.37	5.12
33	0.77	1.62	- 4.35	5.01
34	0.72	1.44	- 4.58	4.91
35	0.56	1.48	- 3.98	4.46
36	0.61	1.34	- 3.91	4.31
37	0.62	1.33	- 3.87	4.09
38	0.59	1.25	- 3.69	3.89
39	0.47	1.18	- 3.05	3.58
40	0.41	1.11	- 3.10	3.28
41	0.44	1.06	- 3.04	3.35
42	0.38	0.87	- 2.55	3.16
43	0.40	0.81	- 2.22	2.66
44	0.28	0.76	- 2.26	2.51
45	0.29	0.71	- 2.00	2.37
46	0.21	0.59	- 2.02	2.20
47	0.32	0.56	- 1.68	1.99
48	0.10	0.41	- 1.37	1.63
49	0.15	0.36	- 1.35	1.45
50	0.04	0.30	- 1.12	1.24
51	0.05	0.11	- 0.68	0.71
52	0.09	0.19	- 0.39	0.76
53	-0.09	0.12	- 0.64	0.36
54[a]	-0.02	0.02	- 0.01	0.01
55	0.00	0.00	0.00	0.00

[a]The value for cohort 54 is restricted. See text for details.

For nonblack males the tendency is for younger cohorts to have a *greater* intrinsic tendency to participate than older cohorts. This is admittedly contrary to our prior line of reasoning. One immediately suspects that the identifying restriction is faulty, and so some attention must be given to the possible bias in the estimates. Let us suppose that Δ should have been taken as 1.00 instead of .98; results of Clogg (1982a) indicate that the effects would then be in error by a factor of $-(55 - k) \log (.98/1.00) \doteq (55 - k)(.02)$. Thus, $\log (\hat{\delta}_1) = 2.61$ would be too low by $54 \times .02 = 1.08$, and the "corrected" estimate for $\log \delta_1$ would be 3.69. Similar kinds of upward adjustments would be required for each of the other estimated effects.

If interest were to focus on intercohort changes in effects, $\log (\delta_k) - \log (\delta_{k+1})$, then it can also be shown that these can be off no more than the error made in imposing the identifying restriction (Clogg, 1982a). That is, the estimated difference $\log (\hat{\delta}_k) - \log (\hat{\delta}_{k+1})$ is equal to the unbiased estimate of the difference plus $\varepsilon[(55 - k) - (55 - k - 1)] = \varepsilon$, where ε is the log of the error made in imposing the identifying restriction. Similarly, the average intercohort change in effects (i.e., $[\Sigma_{k=1}{}^{54} (\log (\hat{\delta}_k) - \log (\hat{\delta}_{k+1}))/54]$) can be off by no more than ε. The average intercohort change in effects calculated from the first column of Table 3 is +.048. The true value of Δ would have to be .953 ($= e^{-.048}$) before our inference about intrinsic cohort tendencies would change to a conclusion of no average intercohort change. We believe that such a value for Δ is incredible (recall that we began by assuming Δ = .98), and so we conclude that the *sign* of the average intercohort change cannot be questioned.

For nonblack females the expected pattern of results appeared, and so there is little need to reconsider the value of the identifying restriction used. The average intercohort change in effects is +.078. We do not see how any plausible identifying restriction could be used that would alter the substantive conclusion.[6] Younger cohorts of nonblack females are thus much more likely to participate in the labor force than older cohorts.

The cohort effects for blacks are not as reliable as those for nonblacks, owing partly to the smaller samples and partly to the less satisfactory performance of the models. However, the pattern of results obtained for both groups is consistent with prior expectations. Younger cohorts of black males have lesser intrinsic tendencies to participate and younger cohorts of black females have greater intrinsic tendencies to participate. The average intercohort change in effects was found to be −.189 for black males and +.233 for black females. The magnitudes of these numbers appear somewhat suspect, but it is inconceivable to us how different values of Δ—chosen within the bounds of plausibility—could change the signs of these quantities.

To summarize, the average intercohort change in effects was estimated at +.048, +.078, −.189, and +.233 for nonblack males, nonblack females, black males, and black females, respectively. The quantity for nonblack males (+.048) might be somewhat too low, but none of the substantive inferences would change if any reasonable specifications of the identifying restriction were used. In view of the magnitudes of these effects, which are "separate" from age, period, and age-period effects, the conclusion must be that cohort differentiation exerts a strong influence on what is observed as across-time change.

How Cohort Effects "Translate" the Observed Period Change

Ryder (1964) referred to the "demographic translation" of cohort rates into period rates. In this section a simple method is used to show how the cohort effects "translate" the period rates, or equivalently, how the cohort effects confound the interpretation that might be given to the "crude" period rates actually observed.

Consider the odds of participation defined by (6). These expressions contain cohort effects, and so we define "purged odds" that have the cohort effects removed but which still depend on all of the other effects in the model (see Clogg, 1978). That is,

$$\Omega_{ijk}^* = \Omega_{ijk}/\delta_k. \qquad (7)$$

These "purged odds" are next converted into "purged" age-period specific rates, $r_{ij}^* = \Omega_{ijk}^*/(1 + \Omega_{ijk}^*)$, and the claim is that—under the model—these r_{ij}^* are adjusted for the influence of the cohort effects. The r_{ij}^* are then applied to the observed age distributions, period-by-period, and overall period rates (say, $r_{.j}^*$) are derived. These rates depend on all of the effects in the model, except for the cohort effects, and so in this sense they indicate the "true" period-specific influences on participation. The observed period-specific participation rates (the "crude" rates, $r_{.j}$) can be compared with the adjusted rates $r_{.j}^*$, thereby inferring the components of period change that are due to period effects and to cohort effects. This procedure is merely a means of assessing the effects of period and cohort in terms different from the odds and logits used hitherto, and in terms consistent with "components" methods of demography (see Wunsch and Termote, 1978).

The successive differences of both the observed period rates and the adjusted rates are compared in Table 4. The estimated cohort effects used to calculate the purged odds of (7) were those presented in Table 3; no attempt was made to "correct" these estimates for possible bias.

For nonblack males, the observed participation rate declined in 8 of the 10 intervals; the average interperiod change was −.2 percent. When the cohort effects are removed, the interperiod change is usually much more negative, the average being −1.7 percent. Thus the cohort effects *translate* the period rates by *offsetting* the effects of the period-specific influences; the combined influence of both kinds of effects operated to modestly reduce the participation of nonblack males. The average interperiod change observed minus the average "purged" interperiod change yields an estimate of the general influence of the cohort effects; this quantity is −.2 percent − (−1.7 percent) = 1.5 percent. Thus, if there were no period effects at all, nonblack males would increase their

Table 4.—Observed Period Change in Participation and Period Change Purged of the Cohort Effect

	Period Interval										Mean Change
	1	2	3	4	5	6	7	8	9	10	
Nonblack males											
Observed	- .1%	- .5%	- .3%	- .3%	- .2%	- .5%	- .5%	.1%	- .3%	.4%	- .2%
Purged	- .4	-2.7	-3.6	-1.8	-2.8	-2.9	- .2	-2.1	-2.8	2.3	-1.7
Nonblack females											
Observed	1.2	.1	.8	.6	1.6	1.1	.9	1.6	1.4	1.9	1.1
Purged	- .3	- .7	- .5	-1.8	- .4	- .7	- .7	- .1	- .4	- .1	- .6
Black males											
Observed	.3	-1.4	-1.5	.5	- .5	-3.2	-2.5	3.9	- .1	- .8	- .5
Purged	1.7	- .2	- .0	- .1	.9	- .4	.2	1.0	+ .0	- .2	.3
Black females											
Observed	1.3	-1.0	- .6	1.4	-1.3	.6	2.2	+ .0	2.5	.2	.5
Purged	- .6	- .4	- .6	- .3	- .4	- .2	- .1	- .2	- .2	- .1	- .3

participation on the average by 1.5 percent per annum. Alternatively, if there were no cohort effects, they would decrease their participation on the average by −1.7 percent per annum.

For nonblack females, the observed interperiod change is everywhere positive, with the average interperiod change being 1.1 percent. However, when the cohort effects are removed the interperiod change is everywhere negative, with the average interperiod change being −.6 percent. The average period-specific influences (−.6 percent per annum) are thus offset by cohort influences which average 1.1 percent − (−.6 percent) = 1.7 percent per annum.

Table 4 also presents results for black males and females. Over the entire interval of time covered by our data, the observed participation rate of black males declined by an average of .5 percent per annum, while the observed participation rate of black females rose on average by the same amount. However, the data purged of the cohort effects is quite different; the actual influence of period-specific effects averages +.3 percent for black males and −.3 percent for black females. Thus the observed period change is conditioned to a great extent by cohort effects for these groups as well. The results of Table 4 demonstrate the importance of a sociodemographic perspective in the analysis of recent trends in labor force participation. For each sex-color group, but especially for the nonblack groups, the effects of cohort differentiation with respect to "intrinsic" labor force attachment must be reckoned with in order to give a proper accounting of across-time change. The usual economic view that attributes the bulk of across-time change in participation to labor market factors of supply and demand—period effects crudely conceived—is simply not supported by our analysis. Cohort effects, which by definition are not attributed to aging or period effects, must be considered in order to understand the very character of recent trends. According to these results, an analysis of recent trends which incorporates only period-specific causal agents is bound to be misleading.

To further examine how period-specific influences and cohort-specific influences combine to produce across-time change, Table 5 is presented. There the average of the differences, log $(\hat{\beta}_{j+1})$ − log $(\hat{\beta}_j)$, are compared with the average differences in cohort effects, log $(\hat{\delta}_k)$ − log $(\hat{\delta}_{k+1})$. Both kinds of differences are relevant for describing why a given period of time is different from the preceding period, although here only the "general" period effects (log $(\hat{\beta})_j$)) are considered.[7] The two sets of quantities are seen in every case to be of the opposite sign, and this fact can serve in part to explain the results in Table 4.

Period Shocks on Participation for Young and Old Age Groups

A distinctive feature of the model of (6) is that it incorporates a kind of age-period interaction, in addition to the main effects of age, period, and cohort. Such a model appeared to be consistent with notions about discouraged workers (see Flaim, 1972; Levitan and Taggart, 1974; Sullivan, 1978), since these "marginal" workers are principally composed of young and old "nonparticipants," measured as being outside the labor force given the conventional definition. It also appeared to be consistent with trends in the age at retirement, at least in so far as withdrawal from the labor force in ages 50–64 is indicative of retirement.

The "period shocks" on the participation of young and old age groups were denoted as θ_{1j} and θ_{2j} in (6). To summarize the effect of these parameters on participation, the following contrasts will be utilized. Under the model expressed in logarithmic form, 14 log θ_{1j} describes the effect of the jth period on participation for a 20 year old in comparison with a 34 year old, net of age effects. If this quantity is negative, the period shock is unfavorable to the participation of young

Table 5.—Average Intercohort Change Compared with Average Interperiod Change, Effects on Logarithmic Scales

Group	Average Change in Cohort Effects	Average Change in General Period Effects
Nonblack males	.048	-.081
Nonblack females	.078	-.032
Black males	-.189	.133
Black females	.233	-.186

persons; if it is positive, the period shock is favorable to the participation of younger persons. Similarly, 14 log θ_{2j} describes the relative favorableness of period j for the participation of 64 year olds (in comparison to 50 year olds). The estimates of these quantities appear in Table 6.

First consider the quantities 14 log $\hat{\theta}_{2j}$ describing the period shocks on the participation of older persons. The pattern of results obtained appears remarkably consistent with prior expectations. For nonblack males, nonblack females, and black females, after 1974 or 1975 the period shocks operate to *reduce* the participation of older workers. That is, from about the middle of the decade onward the socioeconomic conditions of each period encouraged earlier withdrawal from the labor force for older workers. The magnitudes of these quantities range from around −.20 to +.15, and exploiting the relationship between log-odds

Table 6.—Transitory Period Shocks on Participation for Young and Old Age Groups

Period	Nonblack Males		Nonblack Females		Black Males		Black Females	
	Young	Old	Young	Old	Young	Old	Young	Old
1969	-.26	-.01	-.10	.13	-.12	-.13	.02	.16
1970	-.21	.01	-.14	.10	-.05	-.01	-.08	.10
1971	-.15	.10	-.18	.10	.01	.21	-.11	-.07
1972	-.04	.07	-.16	.07	-.13	-.09	-.09	.04
1973	.03	.06	-.08	.01	-.09	-.12	-.04	-.06
1974	.03	.03	-.03	-.01	.12	.08	-.04	.05
1975	.08	-.05	-.02	-.06	.09	.05	-.11	-.04
1976	.06	-.03	-.02	-.12	.08	.10	-.01	-.03
1977	.10	-.04	.02	-.18	.01	.00	-.01	-.17
1978	.02	-.11	.03	-.22	-.08	.01	.08	-.24
1979	-.04	-.11	.10	-.20	-.02	-.22	.07	-.14

NOTE: The quantities in the "Young" columns are equal to 14 log $\hat{\theta}_{1j}$ and are interpreted as the period effect on a 20-year-old's log-odds of participation in comparison to a 34-year-old. Thus a negative number signifies that the period shock was unfavorable to the participation of young persons, while a positive number signifies that the period shock was favorable to their participation. The quantities in the "Old" columns are equal to 14 log $\hat{\theta}_{2j}$. A positive number thus indicates that the period shock was favorable to participation of older persons, while a negative number indicates that the period shock was unfavorable to participation of older persons. See text for further explanation.

and rates indicates that each increment of .05 accounts for about 1 percent increase in participation. Thus a value of $-.10$ in Table 6 corresponds to a period shock that decreases participation for 64 year olds by about 2 percent in comparison with 50 year olds. Taking the quantities in Table 6 as a whole, it is clear that the period shocks on the older age group can alone account for a variability in age-period-specific participation rates from about $+4$ percent to -4 percent, which indicates that the effect of age-period interaction for older persons is nontrivial. Even though the interaction terms did not appear to contribute "significantly" to fit as judged by chi-square comparisons (see Table 1), nevertheless the coefficient estimates are of a magnitude that should not be ignored.

It should also be noted that the period shocks on the participation of older black males do not exhibit the same patterns as those for the other groups. Indeed, it is difficult to formulate any general statement about the trend in the period shocks for older black males.

Next consider the quantities 14 log $\hat{\theta}_{1j}$ describing the period shocks on the participation of younger persons. Our initial hypothesis was that these should fluctuate with economic growth and recession, being positive for periods with much employment opportunity and negative for periods of recession. For the time series considered here, 1969 best represents a "full employment" economy (unemployment was around 3.5 percent for the year), and 1974–1975 best represents recession experience (unemployment was around 8–9 percent for this interval). Inspection of results indicates that this initial hypothesis is false: the employment opportunity of the several periods simply does not correlate with the estimated period shocks. For example, 1969—a "full employment" year—was actually unfavorable to the participation of 20 year olds in comparison to 34 year olds for three of the groups, and it was only negligibly favorable to the participation of younger persons for the remaining group (black females). During the 1974–1975 recession, the period shocks were actually *favorable* to the participation of younger nonblack male and black male groups, although they were unfavorable to the participation of younger females of both color groups.

At least three different explanations of these findings can be offered. First, while it is true that younger persons are more susceptible to being discouraged workers, probably the majority of discouraged workers are actually located in the 16–19 age category, a range of ages that was deliberately excluded here (see Clogg, 1979, ch. 2). Second, the data pertain only to *civilian* persons, and the inferences drawn here would only be strictly appropriate if the excluded military persons were representative of the included civilian persons. In the initial years of our series, there was probably a *net* movement of military persons into the work force, and this would confound the inference in complicated ways. (This cannot suffice to completely explain the findings, however, in part because it does not explain the unexpected finding for women, and in part because it does not explain the unexpected finding for the years after mid-decade when the military-civilian flows could be regarded as being inconsequential.) Third, the "additional worker" hypothesis (cf. Bancroft, 1958; Long, 1958, p. 181) might actually be a better description of how young persons as a group react to economic growth and recession. This hypothesis states that a certain segment of the population exists as a "ready reserve," forced into participating when recessions erode the economic well-being of those on whom they depend, and easily separated from the work force when conditions of "full employment" strengthen the economic base of those on whom they depend. When a recession occurs, there are surely those who withdraw from the labor force because of discouragement in job seeking, but there are also those who are forced to enter the work force to "make ends meet" in their households.[8] In considering the recession of 1974–1975, it appears that there were more "additional workers" than "discouraged workers" when considering males, and that there were more "discouraged workers" than "additional workers" when considering females. Nearly the largest effects occurred during the relatively prosperous years 1969–1970. By virtue of the *negative* values of these effects for the first three groups (−.26, −.10, −.12, for 1969), the inference would be that the conditions of full employment actually led to less favorable participation changes for young persons relative to middle-aged persons. The period shocks for young workers in Table 6, certainly not trivial in magnitude, are clearly parameters that must be explained further: the discouraged worker idea simply does not appear to capture the salient properties of these estimates.

It would be desirable to take explicit account of "discouraged workers" and "additional workers" to explicate the findings reported above for the young age group. While the March CPS does provide some information about "discouraged workers" (see Levitan and Taggart, 1974), it does not provide suitable information about "additional workers." While it is possible to locate "new entrants" to the labor force with the March CPS, it is not possible to ascertain whether these "new entrants" attached themselves to the labor force for other than normal reasons (e.g., for reasons like economic hardship of the households of which they are a part). At the very least, our results suggest the need to reconsider the discouraged worker vs. additional worker controversy: recent research on the labor force apparently has concentrated too much on the discouraged worker phenomenon.

DISCUSSION

This paper has applied the age-period-cohort accounting framework to analyze recent trends in labor force participation. The substantive conclusions were that (a) cohort differentiation is of a nontrivial magnitude, (b) cohort effects play a major role in producing the time trends observed, and (c) a convincing "causal" analysis of recent trends cannot possibly be conducted with reference to only period-specific "causal" variables, ostensibly economic in form. In this concluding section certain criticisms of the methods used will be addressed, some implications of our findings will be drawn out, and suggestions will be made concerning how strictly demographic concepts might be used to explain the results.

Throughout we have relied on the age-period-cohort accounting framework as discussed by Pullum (1977) or Fienberg and Mason (1978). These methods have been criticized chiefly because they do not incorporate interactions of the indexing variables (e.g., age-period interactions) and because there is a nontrivial identification problem in estimating the parameters of interest. To make the model more realistic, age-period interaction was included in our modifications of these methods, exploiting certain relationships that appeared suitable in the given context. The identification problem was also circumvented by using prior information to restrict the value of one cohort effect in the model and qualifying the inferences drawn by considering possible biases that might accompany a given restriction. Our results are certainly dependent on the identifying restrictions used, but for each sex-color group it was demonstrated that the substantive inference about the general *trend* in cohort effects could not change given *plausible* alternative specifications.

A principal finding of this research was that cohort effects and period effects (including age-period interactions) had just the *opposite* influences on time trend for *each* sex-color group (see Tables 4 and 5). For nonblack males, e.g., period effects operated to *decrease* participation by an average of 1.7 percent per annum, while cohort effects operated to largely offset the period effects, producing an observed time trend that showed a slight reduction in participation over the period. For nonblack females, period effects also operated to decrease participation (by an average of .6 percent per annum), but cohort effects more than compensated for the period effects, producing the increase in participation that was actually observed. For nonblack males the period effects were actually much more negative than the observed time trend would indicate, while for nonblack females the period effects were actually the *opposite* of that indicated by the observed time trend (see Table 4). At the very least, these results can be used to justify the use of a sociodemographic perspective (incorporating ideas of intrinsic cohort differentiation) in an analysis of recent trends. A narrow economic view, geared toward explaining trend in terms of supply-demand conditions in the labor market, is impossible to justify if cohort factors are considered.

Results also pointed to the fact that period-specific shocks began to be favorable to earlier labor force withdrawal among older workers commencing at mid-decade (an exception to this was found for black males). Whether or not these period shocks were largely caused by labor market conditions or by social factors is a question that our data and methods simply cannot address. The parameters used to make this inference "partialed out" the influence of general age effects, cohort effects, and general period effects, implying that the conclusions drawn from them are in certain respects more rigorous than corresponding inferences made from period-specific age-participation profiles (see Reimers, 1976). It was demonstrated that these period shocks had nontrivial effects on participation, which is also consistent with findings reported by Duncan (1979).

Unanticipated results were obtained when estimating the period shocks on the participation of the young age group (20–34). It was demonstrated that a "discouraged worker" hypothesis was not consistent with the estimated pattern of effects, and it was suggested that the "additional worker" hypothesis might better explain the findings for some groups. Clearly, more research needs to be done on the way that labor market conditions affect the participation behavior of young persons. Long's (1958, ch. 10) dismissal of the additional worker hypothesis appears to be incorrect for recent labor force behavior.

In conclusion, some strictly demographic approaches to explaining the pattern of results can be offered, particularly regarding the cohort effects estimated from the data (see Table 3). Perhaps the first factor that comes to mind is how absolute or relative cohort size might play a role in explaining results (see Easterlin, 1980). It requires no deep understanding of past fertility history to conclude that cohort size, or relative size, cannot by itself explain the cohort effects. Indeed, the pattern of cohort effects appears to be rather regular throughout the range of cohorts considered, whereas if cohort size could serve as the "explanatory factor," a curvature in the pattern of effects would be in evidence. The relative sizes of cohorts born from 1910 to 1929, 1930 to 1940, or 1945 to 1959, which are known to be quite different, are simply not reflected at all in the pattern of estimated cohort effects. (Our youngest cohort corresponded to birth cohort 1959, while our oldest cohort corresponded to birth cohort 1904). Thus, to explain trends in participation, the cohort size variable can probably be ruled out as being a substantial "cause" of cohort differentiation.

NOTES

[1] The term "intrinsic cohort tendency" is used repeatedly in this paper. It refers to the tendency to participate in the labor force (expressed in terms of odds or log-odds) that can be described solely as a *cohort attribute*. It cannot be attributed to age or period influences *of the age groups or periods*

considered. (Intrinsic cohort tendencies could, however, be partly due to the age, period, or age-period influences of age groups or periods prior to the age groups or periods under study. Thus cohort tendencies inferred in an analysis of 1969–1979 data might actually be due to lagged period effects, i.e., influences of period-specific causal variables operating prior to 1969.)

[2] Nonblack nonwhites constitute only about 10 percent of the nonwhite figure. No attempt was made to isolate Hispanics for separate treatment.

[3] These assumptions imply, e.g., that a 20 year old is more susceptible to a period shock than a 34 year old, that a 64 year old is more susceptible than a 50 year old. Such a specification is similar to Clogg's (1979, ch. 9) model of "constant age elasticity," and it merely states that participation is more transitory at the ends of the age distribution. A reviewer correctly notes that a variety of functional forms might be plausible for the age-period interaction parameters, even when the interaction is restricted to the age ranges 20–34, 50–64. For example, it is possible to obtain a set of parameters coding single-year-of-age age-period parameters, using methods like those used to define the main effect parameters (i.e., "dummy variable" coded effects). Many other functional forms could be considered as well. We suggest the present one because it is simple, because it accords well with the notion that there is a regular relationship of the period shocks with age, and because it is analogous to many of the recent strategies for dealing with ordered categorical data (Clogg, 1982b; Goodman, 1979).

[4] In restricting the relationship between the two oldest cohorts, we are assuming in effect that these cohorts probably do not differ to a very large extent in their intrinsic tendencies to participate in the labor force. Surely these cohorts, socialized during a time when the tempo of social change was generally less than in recent times, could be expected to differ only negligibly from each other. This accounts in part for the values of the restrictions actually used. For related comments, see Clogg (1979, p. 139).

[5] The literature on the black "underclass" (see Wilson, 1980), coupled with the observation that black male rates of participation are declining over time, quite possibly indicates that younger cohorts of black males have a lesser intrinsic tendency to participate. A reviewer notes that it would be difficult to find evidence for any *numerical* magnitude to be assigned to this effect, but the literature does seem to indicate that the sign of the effect would be in the indicated direction. Later it will be demonstrated that reasonable values of this identifying restriction would all lead to the same general conclusion.

[6] The true value of Δ would have to be .92 ($= e^{-.078}$) before we would conclude that there was no average intercohort change in the effects. When translated into participation rates, this value can be seen to be incredible (recall that the assumed value for Δ was 1.02). Also see note 5.

[7] A reviewer correctly notes that the comparison of period effects here can be somewhat misleading because the model contains an age-period interaction term, as well as main effects of period. However, the general period effects are much larger in magnitude than the age-period interaction terms (see Table 6), and so we can safely state that the actual influence of period effects is the opposite of that of cohort effects. It should be noted here too that the average trend in the estimated period and cohort effects (Table 5) is consistent with the summary rates calculated in Table 4. In each case, the period effects and the cohort effects are in the *opposite direction*; the summary rates calculated in Table 4 are thus merely different quantities that describe the implications of the parameter estimates developed for the odds or logit models.

[8] Long (1958, ch. 10) surveyed evidence on the "additional worker"–"discouraged worker" controversy and concluded that the "discouraged worker" theory was more consistent with his data. His study focused on several national labor forces, but relied on data from the 1935–1950 period (largely census data). Also, he did not disaggregate into sex-color groups as we have done here (in his analysis of the additional worker theory).

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