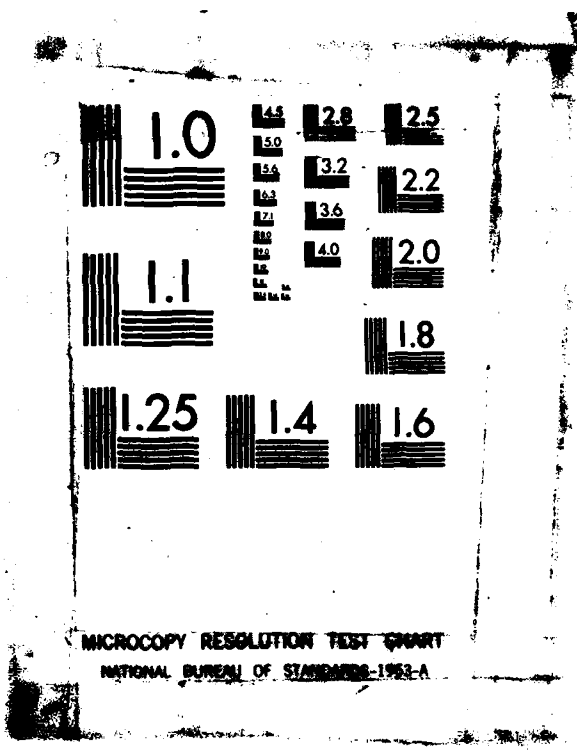
AD A155 420 FORECASTING THE WAGES OF YOUNG MEN: THE EFFECTS OF COHORT SIZE(U) RAND CORP SANTA MONICA CA H TAN ET AL. MAY 85 RAND/R-3115-ARMY MDA903-80-C-0652 1/

UNCLASSIFIED F/G 5/3 NL

END
DATE
FILMED
8-85


1.0
1.1
1.25
1.4
1.6
1.8
2.0
2.2
2.5
2.8
3.2
3.6
4.0
MICROCOPY RESOLUTION TEST CHART
NATIONAL BUREAU OF STANDARDS-1963-A

AD-A155 420

Forecasting the Wages of Young Men

The Effects of Cohort Size

Hong W. Tan, Michael P. Ward

Rand

The research described in this report was sponsored by the Department of the Army under Contract MDA903-80-C-0652.

Library of Congress Cataloging in Publication Data

Tan, Hong W., 1949-
Forecasting the wages of young men.

"Prepared for the Department of the Army."
"R-3115-Army."
"May 1985."
1. Wages—Young men—United States—Forecasting—Mathematical models. 2. Wages—United States—Forecasting—Mathematical models. I. Ward, Michael Paul, 1947- . II. Rand Corporation. III. Title.
HD4975.T36 1985 331.3'42973 84-9873
ISBN 0-8330-0574-X

Published by The Rand Corporation

2

R-3115-ARMY

Forecasting the Wages of Young Men

The Effects of Cohort Size

Hong W. Tan, Michael P. Ward

May 1985

Prepared for the
Department of the Army

Rand
1700 MAIN STREET
P.O. BOX 2138
SANTA MONICA, CA 90406-2138

PREFACE

This report was prepared as part of the study program of Rand's Defense Manpower Research Center. The Center has as its purpose the development of both broad strategies and specific solutions for dealing with present and future military manpower problems. This research was conducted for the Department of the Army under Contract MDA903-80-C-0652.

In this study, the authors develop forecasts of civilian wage structure over the next two decades for a variety of different scenarios. They focus on how the wage structure will change as the demographic trend reverses itself, i.e., as the smaller post-baby-boom birth cohorts enter the labor market in the 1980s and 1990s. These wage forecasts, and their implications for compensation policy in the coming decades, should be of interest to decisionmakers and personnel managers in the U.S. Department of Defense and in the private sector.

SUMMARY

The post-war baby boom and the baby bust that followed have produced dramatic changes both in the size and in the age composition of the U.S. workforce. These demographic changes have been accompanied by changes in the age structure of civilian wages. The evidence suggests that as members of the baby-boom cohorts entered the labor market in the 1960s and 1970s, their wages fell in comparison with the wages of prime age workers. The question that we address in this report is how the wage structure will change as the demographic trend reverses itself, i.e., as smaller birth cohorts enter the labor market in the 1980s and 1990s.

This issue is of interest to military manpower planners because civilian pay opportunities are believed to be important in the decisions of individuals both to join the Armed Forces and to reenlist. Thus, policy makers would like to be able to predict and anticipate changes in the structure and levels of wages of relevant demographic groups so that appropriate compensatory measures in military pay can be taken. The objective of this research is to provide the civilian wage estimates needed to begin formulation of such policies.

We approached this question by first estimating the magnitude of cohort size effects on the earnings of white males. We did this using regression models estimated on *Current Population Survey* data covering the period from 1967 to 1980. In each schooling group, mean wages were related to cohort size, work experience, and a variety of macroeconomic variables such as unemployment, GNP growth, and the size of the civilian labor force. The estimated cohort-size wage elasticities were then used, in a second step, to forecast wages over the next two decades for four alternative economic scenarios.

Our results indicate that the recent trend toward smaller entry cohorts will produce a flattening of wage profiles in the 1980s and 1990s. For example, we project that by 1990 the relative wages of new high school graduates will rise 5 percentage points in comparison with mature workers, and 8 percentage points by the year 1995. In the most likely scenario, real (inflation-adjusted) weekly wages of young high school graduates are projected to rise 17 percent by 1990 as compared to 1980, and over 30 percent by 1995. In contrast, weekly wages of peak earners are projected to rise only 7 percent by 1990, and 14 percent by 1995. These results are generally robust to alternative assumptions about the future course of the economy or schooling continuation

rates. In other words, while our wage forecasts will vary depending upon scenario, the differences should be fairly small.

We also explored two competing explanations for the observed decline in relative youth wages over the 1967–1980 period—schooling inflation and rising female labor force participation. As schooling levels rise over time, the average unobserved market "ability" within a schooling group would be expected to fall, and their wages would decline compared with older cohorts of similar education. A second explanation attributes declining relative youth wages to rising labor force participation of women who, by increasing the competition for entry-level jobs, bid down the relative wages of male youth. Although we find some empirical support for these alternative hypotheses, and especially for the wage effects of women, they do not materially change our projections of relative youth wages over the next two decades. However, the projected rise in overall wage levels is slightly lower when we account for the effects of rising female labor force participation.

The implications of these results for the military are that declines in cohort size will raise the future cost of attracting new recruits. And these costs will come on top of projected declines in the number of high-quality male accessions. Employers, both civilian and military, will face recruiting problems over the next decade and a half. But, because the services rely almost entirely on youth for recruits, they will be disproportionately hard hit by the rising cost and shrinking supply of youth.

The projected flattening of the age structure of civilian wages has other implications for civilian and military "pay comparability." At present, pay comparability is maintained by linking across-the-board military pay increases to changes in an index of *mean civilian wages*, such as the PATC (Professional, Administrative, Technical and Clerical) index or the proposed Employment Cost Index (ECI). Being indices of mean wages over all age groups, they would neither detect changes in the age structure of civilian wages, nor offer any guidance on how military compensation (both pay and bonuses) should be structured in response to the flattening trend in civilian age-wage profiles. Over the next decade, continuation of the current pay policy should maintain rough pay comparability for careerists but make starting military pay increasingly less competitive with civilian wages.

How the services respond—whether through changes in pay, bonuses, enlistment or retention policies, or some combination of the above—will depend in large part upon the desired force structure. If the services continue to rely on a large first-term enlisted force, targeting of military pay increases (or bonuses) at the enlistment point may be necessary to meet future accession requirements. On the other hand, if the force should become more career and less first-term oriented, other compensation structures may be needed to enhance retention of career personnel. In either case, knowledge of the civilian wage structure, and how this wage structure will change over time, is crucial. This study provides the estimates necessary to begin formulation of compensation and personnel policies appropriate to achieving the desired force structure.

ACKNOWLEDGMENTS

This research has benefited from the contributions of a number of people. Useful suggestions were made by James Hosek, Glenn Gotz, Dale M. Landi, and James Smith at Rand, Finis Welch and other members of the Labor Workshop at UCLA. The first draft of this study was reviewed by Joyce Shields, Curtis Gilroy, and the staff of the Army Research Institute, and by Paul Hogan at OASD (MIL). We are grateful to them for their thoughtful comments on that draft. Rand colleagues Richard Fernandez and James Dertouzos served as formal reviewers for this report. We are indebted to them for their insightful criticisms and comments, many of which found their way into the final draft. Frank Berger provided swift and expert programming assistance in this project.

CONTENTS

FIGURES

TABLES

I. INTRODUCTION

The post-war baby boom and the baby bust that followed have produced dramatic changes both in the size and in the age composition of the U.S. workforce. These demographic changes have been accompanied by changes in the age structure of civilian wages. The evidence suggests that as members of the baby-boom cohorts entered the labor market in the 1960s and 1970s, their wages fell in comparison with the wages of prime age workers.[1] The question that we address in this study is how the wage structure will change as the demographic trend reverses itself, and as smaller birth cohorts enter the labor market in the 1980s and 1990s.

Some figures from the *Current Population Survey* (CPS) will serve to highlight the magnitude of these changes. Consider, for example, the working population of white males with a high school education.[2] In 1967, 17.1 percent of this population was made up of youth with 1 to 5 years of work experience. As large baby-boom cohorts made the transition from high school to work, this fraction rose to a peak of 22.6 percent in 1978, or an *increase of 32 percent* in just over ten years. Over this period, the wages of new entrants fell 11 percent relative to the wages of prime age workers (those with 23–27 years of experience). The question then is how youth wages will fare as the proportion of new entrants falls to a projected 17.4 percent by 1990, and to 15.1 percent by the year 1995?

This issue is of interest to military manpower planners because civilian pay opportunities are believed to be important in the decisions of individuals both to join the Armed Forces and to reenlist. Thus, policy makers would like to be able to predict and anticipate changes in the structure and levels of wages of relevant demographic groups so that appropriate compensatory measures in military pay can be taken. The objective of this research is to provide the civilian wage estimates needed to begin formulation of such compensation policies.

We approach this question by first determining the magnitude of cohort size effects on the earnings of white males. We do this by

---

[1]For example, see J. P. Smith and F. Welch, "No Time to be Young: The Economic Prospects for Large Cohorts in the United States," 1981; and F. Welch, "Effects of Cohort Size on Earnings: The Baby Boom Babies' Financial Bust," 1979.

[2]Throughout this work, our empirical results pertain to white males only. This was dictated by data limitations in analyzing nonwhite populations. The thrust of our conclusions applies, however, to all race groups.

estimating wage models using *Current Population Survey* data covering the period from 1967 to 1980. In each schooling group, we relate mean wages to a measure of cohort size, work experience, and controls for a variety of macroeconomic variables such as unemployment, GNP growth, and the size of the civilian labor force. The cohort-size wage elasticities that we estimate are then used, in a second step, to forecast wages over the next two decades.

A great deal of uncertainty is inherent in any kind of forecasting. Our efforts to project wages out to the year 2000 are no exception. To minimize this uncertainty, we develop different sets of wage projections for alternative scenarios in which we vary assumptions about schooling continuation rates, GNP growth, and unemployment rates. This allows us to isolate (and compare) the effects of higher schooling continuation rates or faster economic growth on forecasted wages. More importantly, we can investigate the sensitivity of projected wage changes induced by cohort size to reasonable variations in these other mitigating variables.

We also explore two other competing explanations for the observed decline in relative youth wages that might challenge the presumption that the reversal of demographic trends will reverse wage trends. The first explanation attributes this decline to the rising schooling attainment of recent birth cohorts. As schooling levels rise, the average unobserved market "ability" within a schooling group falls, and this may explain why wages of recent birth cohorts have declined relative to older cohorts. The second explanation attributes this decline to the rising labor force participation of women. By increasing the competition for entry-level jobs, females may bid down the relative wages of youth for whom they are perhaps most substitutable. If both (or any) of these effects are important, the cohort-size wage elasticities may overestimate the depressant effects of large baby-boom cohorts in the 1970s and, in consequence, also overstate the rise in relative youth wages forecasted for the 1980s and 1990s. This overstatement could spill over into our wage forecasts if female labor force participation rates continue to rise in future. How much of an impact this will have depends on the extent to which rising participation rates of women offset the shrinking size of *both* male and female cohorts. We extend the cohort size wage model to address both of these issues.

The study is organized into several sections. In Section II, we describe the survey data used to create a working file for the analysis. Based on this file, we paint a broad overview of how cohort size and relative wages have changed over the 1967–1980 period. In Section III, we discuss the wage model used and highlight the main empirical results. The assumptions and approach used to forecast wages are detailed in Section IV. Here, we also present wage projections for four alternative scenarios which combine different assumptions about the future course of GNP growth, unemployment, and schooling attainment. In Section V, we extend the wage model to investigate two alternative explanations for the observed decline in youth wages. In the last section, we conclude with a summary of the main findings and their implications for military compensation policy.

II. DATA AND OVERVIEW

Our data consist of fourteen years of information from the March *Current Population Survey* (CPS), covering the period from 1968 to 1981. Each year, the CPS surveys a representative sample of the civilian non-institutional population—between 130,000 and 150,000 persons. This broad-based coverage, and the detailed information on personal attributes, employment status, hours of work, and income which it contains, makes the CPS well suited for our purpose. In this section, we first describe the sample selection criteria and how the main variables were created. The data are then used to trace the major movements in wages and cohort size that have taken place over this period.

THE DATA AND VARIABLES

Our analysis of cohort size effects is based on a selected sample of white males, aged 15 through 65, who were in the civilian labor force during the survey week. We chose to look only at the white sample for several reasons. Aside from small sample size, many different forces—such as affirmative action, migration from the south, schooling quality and the like—were operating on the earnings of blacks over this period. The results we get for whites should broadly be applicable to blacks as well, but it would have be difficult empirically to sort out the effects of these forces from cohort size effects. Our decision to look only at the civilian labor force was dictated, in large part, by the data. Military personnel were not part of the sampling design of the CPS, and thus military bases were not surveyed. Nonetheless, a number of servicemen living off-base were questioned; these individuals were dropped from our sample.[1]

We used this sample to produce a time series aggregate data set organized by level of schooling completed, single years of labor market experience, and year. We define five schooling categories—1–7, 8–11, 12, 13–15, and 16 or more years of schooling—but focus only on the

[1]We recognize, but do not address, the potential problem of focusing only on the civilian labor force. The civilian population, on which our cohort size variable is based, is determined not only by demographic changes but also by military personnel requirements. And this has changed over the observation period, notably with the winding down of the Vietnam War and with the changeover from a draft system to the All Volunteer Force (AVF).

last four.[2] Like most census-style surveys, actual labor market experience is not reported and must be inferred from other information on the individual's age, schooling and year of birth. We did this following the methodology suggested by Welch and Gould (1976).[3] For each individual in a schooling group j, we calculated the full density of probabilities, P_{ij} (where $i = 1$ to 45), that he was in the i-th year in the labor market. In each schooling group, these probabilities were conditioned on age and year of birth, the latter to account for the trend toward earlier completion of schooling in more recent birth cohorts. These data were then aggregated into single year of experience cells so that in a given survey year, the number of people with i years of experience and schooling j is calculated as the sum over all individual probabilities. We restricted the sample to the first 44 cells because the last observation contains the open-ended interval of 45 or more years. Pooling across experience cells yielded a dataset with 616 observations over 14 survey years for each schooling group.

We calculated a cohort size variable measuring the experience distribution in each schooling group. It is likely that the wages of a particular cohort are affected both by its own size and by the size of other surrounding cohorts. To allow for this possibility, we calculated cohort size as a relative measure: first normalizing cell counts by the size of the workforce in that schooling group, and then smoothing these fractions by computing a moving average with inverted-V weights.[4] Each fraction thus measures the size of that (experience) cohort relative to the whole workforce. Changes in cohort size are assumed to change the structure of relative wages across experience levels by changing the experience distribution within each schooling group. Large entry cohorts skew the distribution toward inexperienced workers whereas small entry cohorts reduce their relative numbers.

Our interest is in explaining changes in the wages of those who are full time in the civilian labor force. We chose to exclude those with low labor force attachment, such as students or the retired, whose

---

[2]The fifth group—those with less than a junior high school education—are included in subsequent analysis in Section V and in projections of future schooling distributions.

[3]Welch and Gould (1976) use macro panel data to estimate the probability distribution of schooling completion by age and year of birth. While complex, this specification of work experience improves upon the conventional use of potential experience proxies defined as age minus schooling minus 6. We gratefully acknowledge the use of their estimates in this study.

[4]Except for new entrants, the size of each experience cohort j is calculated using the weights 0.33 (0.33, 0.66, 1.0, 0.66, 0.33) centered on cohort j and including the experience cohorts $j-2$ to $j+2$. For new entrants, the weights of succeeding cohorts are not defined, and we scale the remaining weights accordingly to sum to one. The construction of this variable follows the methodology described in Welch (1979).

wages may be misleadingly low because of their part-time work status. Thus, in calculating mean wages for each cell, we restricted the sample to persons not in school as the primary activity in the survey week, who worked 50 to 52 weeks last year, or if they worked fewer weeks did so involuntarily, i.e., for reasons other than schooling and retirement. We also excluded those who worked but failed to report earnings. The CPS imputed to these individuals the mean earnings of those with ostensibly similar personal attributes, but a severe problem with the imputation procedure advised against their use.[5] We know little about how wages are determined in the informal sector, and therefore dropped all individuals who were self-employed, agricultural workers, or working without pay. Finally, we excluded individuals with such low or high earnings (less than $10 or greater than $2000 a week) that we presume their incomes were miscoded.

For each cell, we calculated the (geometric) mean weekly wage and annual earnings for those observations with income. The *Current Population Survey* reports annual earnings and weeks worked last year, from which we can calculate weekly wages. The weekly wage measure more closely approximates a "wage rate" whereas annual earnings include the effects of variations in weeks worked. Because changes in cohort size can affect both "wage rates" and labor supplied, we report the results for both wage measures. These variables are computed using CPS sampling weights and expressed in constant 1970 dollars using the GNP deflator. To account for possible nonrandom exclusion of persons without usable income data, we also constructed two variables to represent the fraction in each cell dropped because of income imputation or nonwork. These variables are included in subsequent regression analyses as controls for potential selectivity bias.

AN OVERVIEW OF CHANGES IN WAGES AND COHORT SIZE

Table 1 introduces the major changes that have taken place in the experience composition of the labor force over the last decade and a half. For each schooling group, the table reports the percent of the workforce with 1 to 5 years of experience (we call them "new labor market entrants"). In all schooling groups, the proportion of new labor market entrants increased dramatically over the first half of this

[5]This is shown in recent research by Lillard, Smith, and Welch (1982).

Table 1

PERCENT OF THE LABOR FORCE WITH 1-5 YEARS OF EXPERIENCE BY SCHOOLING CATEGORY (1967-1980)

Schooling Level Completed	1967	1968	1969	1970	1971	1972	1973	1974	1975	1976	1977	1978	1979	1980
Grade school	9.9	11.5	12.4	14.7	15.6	17.1	17.6	17.5	18.8	19.7	20.4	21.8	22.3	21.5
High school	17.1	17.0	17.7	18.5	20.3	20.6	21.0	21.4	21.8	22.3	22.3	22.6	22.3	21.9
Some college	22.3	22.4	23.7	26.0	27.7	29.1	29.7	27.6	25.9	25.4	24.5	23.0	22.6	21.1
College graduate	18.8	20.1	20.4	21.6	23.2	22.8	23.5	23.9	23.3	22.7	20.8	20.4	18.8	17.14

NOTE: Workforce restricted to those with 1-44 years of experience.

period, reflecting primarily the influx of baby-boom cohorts into the labor market in the the late 1960s and 1970s.[6]

The turnabout in entry cohort size is just apparent by the late 1970s, at least for those with a high school degree or less. For those with 9–11 years of schooling, the proportion of new entrants increased from 9.9 percent in 1967 to a peak of 22.3 percent in 1979, and fell thereafter. Likewise, for high school graduates, the fraction of new entrants increased from 17 percent to a peak in 1978 of 22.6 percent, and subsequently dipped back to 21.9 percent by 1980. Since we know their age—between 18 and 23—in any survey year we can also determine their year of birth. New entrants in the peak years are easily identified as the baby-boom cohorts born in the years immediately preceding the 1961 baby bust.

The peak years occur much earlier for those with higher education. The largest new entrant cohort is in 1973 for those with some college (29.7), and in 1974 for college graduates (23.9). By 1980, the proportion of new entrants in both schooling groups had shrunk back to levels prevailing in 1967. These earlier peaks may be attributed, in part, to declines in progression rates to college, declines that were large enough to offset the increased numbers of baby-boom cohorts that could have enrolled in college. This interpretation is consistent with evidence that continuation rates to college peaked in 1969. It is perhaps not coincidental, then, that the proportion of new labor market entrants with some college education also peaked some 4 or 5 years later.

Tables 2 and 3 show how relative weekly wages and annual earnings of new labor market entrants changed over this period. Their wages are measured relative to those of peak earners, defined as those with 23–27 years of work experience. A comparison of these figures with those in Table 1 is suggestive. Relative wages and annual earnings both exhibit declining trends over the first half of this period, mirroring increases in the size of entry cohorts noted earlier. Further, declines in relative wages reach their low points in years that coincide (roughly) with peak years when the new entrant cohort size is largest. In Table 2, for example, note that relative weekly wages are lowest in 1977 for high school graduates. For college graduates, relative wages are lowest in 1974 (0.49), thus tracking rather well the early peak noted previously.

Nonetheless, the relationship between relative wages and cohort size is less than perfect. Although relative weekly wages and relative

---

[6] Part of the increase in cohort size over this period may also reflect the return of Vietnam era veterans to the labor force.

Table 2

WEEKLY WAGES OF NEW ENTRANTS RELATIVE TO PEAK EARNERS BY SCHOOLING CATEGORY (1967-1980)

Schooling Level Completed	1967	1968	1969	1970	1971	1972	1973	1974	1975	1976	1977	1978	1979	1980
Grade school	0.62	0.56	0.56	0.53	0.46	0.50	0.50	0.51	0.52	0.47	0.48	0.47	0.51	0.48
High school	0.68	0.69	0.66	0.61	0.60	0.59	0.57	0.58	0.56	0.58	0.56	0.57	0.59	0.57
Some college	0.60	0.59	0.57	0.55	0.55	0.52	0.51	0.52	0.54	0.51	0.53	0.55	0.55	0.52
College graduate	0.60	0.61	0.60	0.59	0.55	0.53	0.53	0.49	0.52	0.51	0.49	0.51	0.51	0.51

NOTE: New entrants (1-5 years experience). Peak earners (23-27 years experience).

Table 3

ANNUAL EARNINGS OF NEW ENTRANTS RELATIVE TO PEAK EARNERS BY SCHOOLING CATEGORY (1967-1980)

Schooling Level Completed	1967	1968	1969	1970	1971	1972	1973	1974	1975	1976	1977	1978	1979	1980
Grade school	0.61	0.59	0.57	0.53	0.52	0.53	0.52	0.53	0.51	0.52	0.50	0.53	0.55	0.52
High school	0.51	0.49	0.48	0.45	0.38	0.41	0.43	0.42	0.41	0.39	0.40	0.40	0.43	0.39
Some college	0.57	0.57	0.52	0.50	0.50	0.49	0.48	0.49	0.50	0.48	0.50	0.53	0.53	0.49
College graduate	0.59	0.59	0.58	0.56	0.53	0.51	0.51	0.49	0.50	0.49	0.50	0.49	0.51	0.50

annual earnings decline smoothly in the first half of the period, they bounce around a great deal in the second half of the 1970s, possibly because of the economic dislocation that prevailed over this period. Thus, although there is a slight rising trend in relative youth wages in the late 1970s (as predicted by the turnabout in cohort size), this relationship between cohort size and relative wages is confounded by cyclical instability. In our characterization of cohort size effects, then, we want to control for the wage effects of macroeconomic factors such as unemployment or GNP growth.

While not a substitute for formal analysis, these tables do highlight the central relationships that we want to model. They show that increases in the size of entry cohorts have been accompanied by declines in the relative wages of youth. They also make the point that comparisons of this sort are confounded by changes in schooling progression rates and by cyclical instability. What these tables do not show are other secular changes that have also taken place and which may well have contributed to the observed decline in relative youth wages over this period. These include the potentially depressing wage effects of rising schooling attainment among more recent birth cohorts and increased competition from women reentering the labor force. These concerns are addressed in the following sections.

III. MODEL SPECIFICATION AND EMPIRICAL RESULTS

In this section, we discuss the main features of the model used to estimate the effects of cohort size on male wages.[1] The idea is to associate the wage movements noted earlier with changes in own cohort size, controlling for the wage effects of macroeconomic conditions. We then present and discuss the main results of our regression analyses.

THE CONCEPTUAL MODEL

The hypothesis that relative wages are affected by cohort size rests on the premise that workers with different amounts of work experience substitute imperfectly for each other.[2] Otherwise, a large cohort entering the labor market would depress both the wages of its members and that of older cohorts, but leave the age structure of wages unchanged. If, on the other hand, younger and older workers are somehow different inputs in production, then a large entry cohort would tend to have larger wage effects on its own members than on other cohorts.

Human capital theory provides one justification for this lack of substitution among workers with differing amounts of experience. If most skills are acquired in the early work career, as is plausible, then workers will be very different depending upon where they are in their life cycle. First, we would expect new entrants and prime age workers to be poor substitutes for each other, either because they do different jobs (training versus working) or because they possess different amounts of job skills. Furthermore, we expect these differences to be more pronounced among youth than among older workers. Five years of work experience, for example, would add more to a youth's skills than it would to a more experienced worker whose job training is behind him. These results suggest a specific pattern of cohort size wage effects

[1]The specification of this model closely follows that of Welch's pioneering study of the effects of cohort size on wage structure (1979). However, we extend the sample period to 1981 and include controls for GNP and the size of the civilian labor force.

[2]Another explanation, attributable to Easterlin (1978), is that a large cohort creates "environmental" changes for its members which reduce their market earnings ability. Their diminished ability stems from relative deprivation, both at home (e.g., increased competition for limited family resources and time) and in school (overcrowding and limited seats in universities). There is little solid evidence that diminished ability was responsible for the observed declines in the wages of large cohorts (Smith and Welch, 1981).

varying with years of work experience. That is, while a large entry cohort depresses the wages of its members relative to older cohorts, these cohort-size wage effects are ameliorated as large cohorts acquire more experience and become more substitutable for (i.e., similar to) smaller cohorts around them.

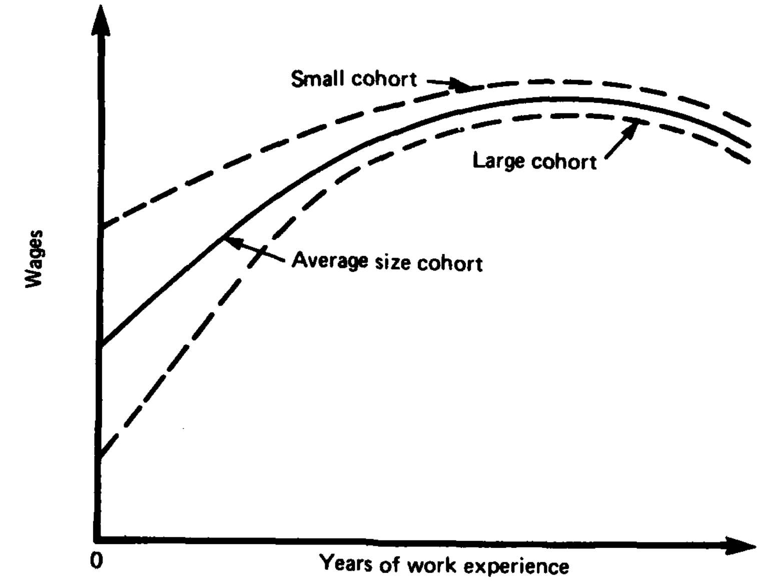
Figure 1a is a schematic of this model. The solid line shows the wage profile of an average size cohort, with the typical pattern of wage growth over the life cycle, i.e., first rising and then flattening out in later years. The wage profile of a large cohort is represented by the bottom (dashed) line. Here, starting wages are depressed below those of the average cohort, but the differential shrinks as work experience is acquired. This case is, we argue, the situation that prevailed over the late 1960s and 1970s as baby-boom cohorts entered the labor market. For small cohorts, such as those predicted for the 1980s and 1990s, it is likely that they will face wage profiles like the one depicted in the top graph. There, the effect of smaller entry cohorts is to raise youth wages relative to wages of more experienced workers.

THE EMPIRICAL MODEL

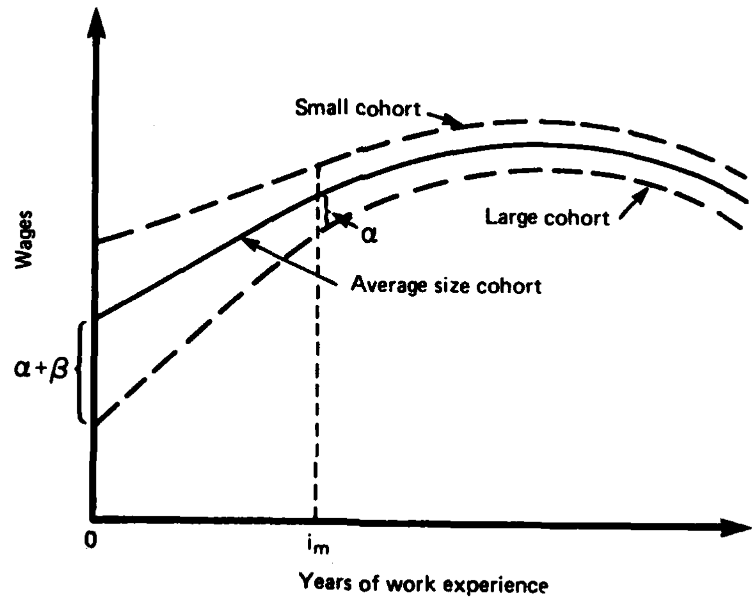
The wage model that we use relates mean wages in each schooling-experience category to several sets of explanatory variables. These variables and their definitions are listed in Table 4. The dependent variable is the natural logarithm of the (geometric) mean weekly wage or annual earnings, expressed in constant 1970 dollars.[3] The explanatory variables include a measure of cohort size, a set of experience variables, controls for different sample exclusions, and several macroeconomic variables. We now discuss these variables in turn and relate them to Fig. 1b.

Cohort size, as noted earlier, is defined relative to the total male workforce with similar amounts of completed schooling. We assume that the wages of a particular cohort depend not only on its own size but on the size of all other competing experience cohorts. Other things equal, we should expect to find a negative relationship between mean wages and cohort size. We note that this focus on (within schooling group) changes ignores the possibility of substitution across schooling groups and between males and females. In the regression analysis, however, we do allow the wages of one group to be affected by other schooling groups and by females via the civilian labor force variable (more on this later).

[3]The GNP deflator is used to express wages in 1970 prices.



(1a) Schematic of the model



(1b) Model with early career spline

Fig. 1—The effects of cohort size on wage profiles

Table 4

VARIABLE DEFINITIONS

Variable Name	Variable Definition
Dependent variables	
Log real wage	Logarithm of weekly wage (1970 dollars)
Log real earnings	Logarithm of annual earnings (1970 dollars)
Explanatory variables	
Log cohort size (ij)	Logarithm of fraction of workforce with schooling i and experience j
Exp, exp-squared	Quadratic specification of years of work experience
Early career spline (i)	Value of spline = 1 - (exp/n) for exp < n, and spline = 0 otherwise; n varies with schooling level i
Cohort × spline	Interaction of log(cohort) and early career spline
Unemployment rate	Aggregate unemployment rate for males
Log GNP	Logarithm of real gross national product in 1970 billion dollars
Log labor force	Logarithm of total civilian labor force
Control variables	
Nonwork	Fraction of cell count excluded because of low labor force attachment
Allocated incomes	Fraction of cell count excluded because of missing or miscoded incomes

We use a quadratic specification of years of experience to capture the average life cycle pattern of wage growth. To this, we add an early career spline variable S, whose value equals 1 on entry into the labor market, declines linearly to 0 after n years, and then remains at 0 thereafter. Following Welch (1979), we define different values of n for each schooling group: 6 years for those with 9–11 years of grade school, 7 for high school graduates, 8 for 1–3 years of college, and 9 years for college graduates.[4] This early career spline serves two purposes. First, it adds flexibility to the specification of experience that several researchers have found to inadequately describe early career wage growth. The spline variable allows for more rapid wage growth between 0 and n years of experience, after which the wage profile follows a quadratic path. This wage profile is shown as the middle graph in Fig. 1b. Second, when interacted with the cohort size variable, the spline also allows the estimated wage effects of cohort size to vary with years of experience.[5] We would expect this interaction term to be negative if younger and older workers are imperfect substitutes for each other, as argued earlier. Such a result will imply a large cohort wage profile corresponding to the bottom line in Fig. 1b and a small cohort wage profile like the top line.

We also include variables to capture the wage effects of business cycles and secular growth in labor productivity. It is important to distinguish between these wage effects and those of cohort size since we want to forecast future wages for a variety of alternative macroeconomic scenarios. The aggregate male unemployment rate is used to control for cyclical fluctuations in economic activity.[6] Economic recession puts downward pressure on wage levels while a boom raises them, so we expect a negative relationship between wages and unemployment rates. Aggregate labor productivity is measured by two separate variables: Gross National Product expressed in billions of 1970 dollars, and the total civilian labor force. The latter variable measures the number of males and females, age 16 and over, who are in the non-institutional civilian labor force. Given the size of the labor force, a larger GNP implies higher per capita labor productivity. An expansion of the economy should therefore also have an expansionary effect on wage levels. Conversely, holding GNP constant, an increase in the size of the labor force should reduce wages. We note that as specified these macreconomic variables have a uniform effect on wages at all levels of work experience but not on wage structure itself.

The civilian labor force variable serves a second purpose, that of allowing for substitution across schooling and sex groups. As defined,

---

[4]Welch used an iterative procedure to determine n for each schooling group. Essentially, this involved repeated estimation of the wage model using alternative values for n until one was found that minimized mean squared error, i.e., that maximized R-squared.

[5]In the regression analysis, both wages W and cohort size C are expressed in natural logarithms, so we can write the wage elasticity of cohort size as $dW/dC = \alpha + \beta S$, where α and β are estimated parameters. Cohort size has an "initial wage effect" of $\alpha + \beta$ on entry into the labor market. As experience accumulates (and as S approaches 0), this initial effect declines to α, the "persistent wage effect."

[6]The reason for not including unemployment rates that are age-education specific is that they are *outcomes* of cohort size effects.

the cohort size variable focuses only on the effects of (within schooling group) changes in experience composition. The idea behind including an all-inclusive measure of the labor force—one that measures the total number of males and females in all schooling groups—is to allow different groups to affect each other's wage levels although the effects are presumed equal for all competing groups.

Finally, we include a set of variables to control for the wage effects of various exclusions. These variables measure the proportion of observations in an experience cell that was excluded in calculating mean wages because of income imputation, because the individuals did not work or, if they did, reported income data that were not usable (see Section II). The inclusion of these variables in the wage model represents an attempt to control for the well-known problems of selectivity bias in the reporting of income and in work decisions.

REGRESSION RESULTS

The wage model was estimated separately for each schooling group, using (log) weekly wages or, alternatively, annual earnings as the dependent variable. All regressions were weighted by the number reporting usable income data in each experience cell.[7] Four schooling groups were considered: grade school (8–11 years of school), high school graduates, some college (13–15 years of school), and college graduates (16 or more years of school). Our discussion focuses primarily on the weekly wage results but references are made to the annual earnings results when important differences are noted.

Table 5 presents the estimated cohort size wage elasticities. We distinguish between two wage elasticities: the entry level effect and the effect that persists over the life cycle. To get the initial effect, we sum the coefficients of cohort size (the main effect) and its interaction with the early career spline variable. The persistent effect is simply the main effect of the cohort size variable. In general, we find that cohort size has the expected effects on wages. For example, Table 5 suggests that a 1 percent increase in cohort size reduces the starting weekly wages of high school graduates by 0.357 percent. As they acquire work experience, this penalty diminishes to a lower level of about 0.114 per-

[7]In estimating these wage models by ordinary least squares, we recognize (but do not address) the possibility of both serial and contemporaneous (cross-equation) correlations in the residuals, in large part because of the lack of appropriate software. Since computed standard errors are probably biased, we do not present confidence intervals for our wage forecasts in Section IV. The coefficient estimates, however, are unbiased and are not a problem for these forecasts.

Table 5

EFFECTS OF COHORT SIZE ON WEEKLY WAGES AND ANNUAL EARNINGS

Model Specification/ Cohort Size Effects	Grade School	High School	Some College	College Graduate
	Schooling Level Completed			
Weekly wages				
Initial effects	-0.258	-0.357	-0.376	-0.257
	(11.08)[a]	(10.29)	(13.98)	(6.24)
Persistent effects	-0.099	-0.114	-0.103	-0.159
	(7.05)	(9.47)	(11.38)	(10.92)
Annual earnings				
Initial effects	-0.273	-0.201	-0.441	-0.378
	(8.89)	(4.51)	(14.72)	(8.64)
Persistent effects	-0.186	-0.153	-0.156	-0.168
	(10.03)	(9.98)	(15.54)	(10.84)

NOTE: *The wage effects of cohort size in the two panels are taken* from the results reported in Tables A.1 and A.2 in Appendix A.

[a]Absolute value of t-statistics is in parentheses.

cent that persists over their lifetime. Each of the other schooling groups also shows initial cohort size effects exceeding persistent ones.[8]

The permanent losses in lifetime weekly wages are most pronounced for college graduates (–0.159), whereas those who have not graduated from high school experience the smallest decline (–0.099). However, these differences across schooling groups disappear when we look at the wage elasticities for annual earnings. The persistent effect is twice as large for those with grade school (–0.186) and about unchanged for college graduates (–0.168). This result suggests that large cohorts have an impact not only on wage rates but also on weeks worked, and that the reduction in weeks worked is greatest for those with the least amount of schooling.

[8]Berger (1981), on the other hand, finds that the wage effects of cohort size increases (rather than decreases) with years of experience. We speculate that differences in model specification and variable definition (work experience, in particular) may be responsible for these diametrically opposite results.

The estimated wage elasticities track well the observed decline in wages from 1967 to 1977. For instance, we noted in the Introduction that the proportion of high school graduates with 1 to 5 years of experience rose about 32 percent over this period. The 10 percent decline in their wages predicted by these wage elasticities compares favorably with the 11 percent decline actually observed.[9] The difference arises in part because we are not accounting here for the mitigating wage effects of other macroeconomic changes that were controlled for in the regression analysis.

Table 6 presents the estimated wage effects of three macroeconomic variables: Gross National Product, the total civilian labor force, and aggregate unemployment. The results generally accord with our expectations. Controlling for the size of the total labor force, GNP growth has a wage elasticity of approximately one, suggesting that aggregate labor productivity growth is associated with an equal percentage increase in both weekly wages and annual earnings. The only exception appears to be those with some college education. Further, an increase in the size of the labor force holding GNP constant reduces wages and earnings. The unemployment rate variable was intended as a control for cyclical fluctuations in business activity, but the estimated effects are small and, in several cases, of the wrong sign. We suspect that collinearity between this and other time-trended variables may be responsible for this anomalous result.

---

[9]Since the proportion of new entrants increases from 17.1 percent in 1967 to 22.6 percent in 1978, the estimated wage elasticity of 0.357 implies a $[\ln(17.1) - \ln(22.6)] \times 0.357$ or 9.96 percent decline in relative wages.

Table 6

EFFECTS OF MACROECONOMIC VARIABLES ON WEEKLY WAGES AND ANNUAL EARNINGS

	Schooling Level Completed			
Model Specification/ Macroeconomic Effects	Grade School	High School	Some College	College Graduate
Weekly wages				
Log(GNP)	0.821	0.993	0.340	0.887
	(6.01)[a]	(9.37)	(2.92)	(6.38)
Unemployment rate	0.006	0.003	-0.002	0.007
	(2.33)	(1.83)	(1.14)	(2.66)
Log total civilian labor force	-1.024	-1.291	-0.536	-1.365
	(5.29)	(8.78)	(3.28)	(6.99)
Annual earnings				
Log(GNP)	1.021	1.233	0.355	0.782
	(5.68)	(9.08)	(2.71)	(5.28)
Unemployment rate	-0.015	0.009	-0.013	0.000
	(4.64)	(3.77)	(5.50)	(0.01)
Log total civilian labor force	-1.444	-1.615	-0.574	-1.205
	(5.67)	(8.57)	(3.15)	(5.81)

NOTE: The wage effects of macroeconomic variables in the two panels are taken from the results reported in Tables A.1 and A.2 in Appendix A.

[a]Absolute value of t-statistics is in parentheses.

IV. FORECASTING THE WAGES OF YOUNG MEN

The estimates of the cohort size model presented in the preceding section are now used to forecast the wages of white males for 1981 through to the year 2000. For each schooling category, we forecast weekly wages and annual earnings under a variety of assumptions about school progression rates and the future path of macroeconomic variables. In this section, we first discuss the assumptions underlying these forecasts and describe how the main explanatory variables were created. Next, we highlight the principal results using the forecasts for high school and college graduates. We also investigate the sensitivity of our wage forecasts for 1990 to changes in scenario.

ASSUMPTIONS AND FORECASTING METHODOLOGY

The wage forecasts are made under several assumptions about the future path of the explanatory variables. For some variables, such as the controls for different exclusions, we have no way of predicting their future behavior and therefore assume that they remain unchanged at 1980 levels. For the main variables of interest, however, we have population projections from the Census Bureau and macroeconomic forecasts from Data Resources, Incorporated (DRI). To use this information, certain assumptions have to be made and we now discuss them in turn.

Cohort Size Projections

Our projections of future cohort size are based on three pieces of information: population projections by the Census Bureau[1] and two assumptions about the size of the Armed Forces and future school progression rates. We assume that:

- The size and age distribution of the Armed Forces remains constant at 1979 levels up to the year 2000.[2]

[1]These *Projections of the Population of the United States: 1980 to 2050* are the most recent estimates available, and they incorporate census counts from the 1980 Population Census. Except for minor changes in future mortality rates and immigration, there should be little forecast error in their age, race, and sex projections.

[2]These figures are weighted counts taken from the *1979 DoD Survey of Officers and Enlisted Personnel*, 1981. The survey's detail on age, race, and sex characteristics recommended its use over alternative published data sources.

- Schooling attainment follows one of two possible time paths: (1) remaining at the average levels prevailing in 1980, or (2) trending over time but leveling off by 1990.

To project future cohort size, we first subtract military personnel from the Census Bureau's population projections.[3] For simplicity, we assume that the age, race, and sex composition of the Armed Forces remains constant; although alternative assumptions about the future force structure are readily incorporated into these projections, they are unlikely (we suspect) to have an appreciable effect on our wage forecasts. In a second step, we develop two sets of projections of the civilian workforce by level of schooling completed, corresponding to the alternative assumptions about future school progression rates.[4] In the stationary case, we use the mean distribution of completed schooling prevailing between 1978 and 1980, averaging over three years to minimize random year-to-year variations due to sampling error. In the second case, we project schooling attainment using the trends prevailing over the last 10 years of the CPS data. We assume that past trends (whether increasing or decreasing) will continue into the future, following a parabolic path over the next 10 years and then plateauing.[5] Thus, for example, if high school completion rates were increasing over time, we projected that they would rise further (although at a slower pace) until 1990, leveling off after 1990.

Table 7 provides counts of the projected number of young men in selected years. The counts are aggregates over five age groups, and ages range from 15–19 years for those with 8–11 years of grade school to 21–25 years for college graduates. Panel A represents the counts for the case where schooling distribution is fixed at 1980 levels, Panel B for the case where schooling continuation rates are trended. These figures highlight the sensitivity of cohort size projections to alternative assumptions about future school progression rates.

---

[3]Let N_{jt} be the projected number of white males in the population of age j in year t, and M_j be the number of white males age j in the military, assumed to be constant over time. Subtracting M from N, we get the civilian population age j in any given year. Further, let P_{ijt} be the fraction of the total civilian population age j, not in school in the survey week and with i years of completed schooling. If we know the future time path of P_{ijt}, then the number of white civilian male workers by age and education in year t, C_{ijt}, is $(N_{jt} - M_j) \times P_{ijt}$.

[4]Developing a model of school enrollment to forecast the future educational attainment of the workforce was beyond the scope of this study. One possible approach has been explored by Wachter and Washer (1982). They investigate how school enrollment rates in the United States respond to peaks and troughs in the demographic cycle.

[5]For each schooling group i of age j, let this trend be represented by the difference between the average P's of 1968–1970 and 1978–1980, and call it D_{ij}. P_{ij} is assumed to change by 30 percent of D_{ij} over the 1981–1985 period, by another 20 percent from 1986 to 1990, and remain constant thereafter until the year 2000.

Table 7

NUMBER OF YOUNG MEN PROJECTED UNDER DIFFERENT ASSUMPTIONS ABOUT SCHOOLING CONTINUATION RATES

(Thousands)

Schooling Group	1985	1990	1995	2000
A. Constant 1978-1980 Schooling Progression Rates				
Grade school	1196	1106	1075	1197
High school graduate	2247	2073	1789	2001
Some college	899	785	679	731
College graduate	992	841	794	704
B. Trended Schooling Progression Rates				
Grade school	1162	1138	1159	1291
High school graduate	2443	2399	2005	2315
Some college	896	812	703	756
College graduate	956	846	798	708

NOTE: Young men are defined by level of schooling completed: Grade school = ages 15-19, high school graduate = ages 17-21, some college = ages 18-22, and college graduate = ages 21-25.

To illustrate, consider how the number of high school graduates age 17–21 change over the next two decades. In Panel A, the number of young men are projected to decline from 2,247,000 in 1985 to 1,789,000 in 1995, and to increase to 2,001,000 by the year 2000. We have fixed schooling distribution at 1980 levels, so changes in projected numbers are driven only by demographic changes in which cohort size first declines and then increases again, but not until the late 1990s. This reversal in entry cohort size at the end of this century is attributable to the recent increase in birth rates. The projections in Panel B are larger, reflecting the rising trend in high school completion rates over the last decade. For college graduates, differences between Panels A and B are less apparent because trends in college enrollment have been flat, and even slightly declining, over the last decade after they peaked in 1969.

Total Civilian Labor Force

Projections of the total civilian labor force, which includes both males and females, were made using the same data base provided by the Bureau of the Census. We assumed, again, that the Armed Forces would remain at 1979 force levels over the next two decades and subtracted them from the projected population. Since we are interested in the size of the labor force, and not the civilian population, we assumed that past trends in male and female labor force participation rates (LFPR) would continue through to 1990 (though at a slower pace), and then level off from then on. These LFPR figures were used to project the total civilian labor force variable out to the year 2000.

Projections of Macroeconomic Variables

The macroeconomic projections that we use are, at best, no more than informed guesses about the future. Changes in fiscal and monetary policy or developments overseas could have a large (and unknown) impact on the future course of GNP growth and unemployment. To mitigate the uncertainty inherent in such forecasts, we used two sets of macroeconomic projections. The first set of projections was taken from DRI's *U.S. Long-Term Review.* Their forecasts of the U.S. economy, at least those up to 1982, have turned out to be too optimistic. In that year, DRI projected a less severe recession, and an unemployment rate (9.4 percent) that was lower than the 10.3 percent rate that was actually recorded. The second set of projections assumes a much stronger recovery to 1985, after which real GNP is assumed to grow at a rate 0.5 percentage points higher, with unemployment rates one percentage point lower, than the first set of DRI figures. These macroeconomic projections are presented in Table 8.[6]

The four assumptions—two about future macroeconomic activity and two about schooling continuation rates—are combined to form four possible scenarios. These scenarios are listed in Table 9. The first scenario, which we believe to be the most plausible of the four, combines the DRI forecasts with our assumption of a trended change in schooling continuation rates. The second scenario envisages a more robust economy over the next two decades. The third and fourth scenarios combine a stationary schooling distribution with alternative assumptions about the future course of the economy.

WAGE FORECASTS

In the discussion that follows, we first consider how wages change over time under scenario 1, and then explore the consequences of making different assumptions about the future course of macroeconomic

[6]Presumably, one could have picked different (larger) GNP growth rates. However, note that even a 0.5 percentage point increase translates into a 4 percent increase in *real* GNP by 1990, 6 percent by 1995, and 9 percent by the year 2000.

Table 8

GNP AND UNEMPLOYMENT RATE ASSUMPTIONS: 1981-2000

	Assumption 1[a]			Assumption 2		
Year	UNEM[b] Rate(%)	GNP Growth	Real GNP[c]	UNEM Rate(%)	GNP Growth	Real GNP
1981	7.6	1.9	1379	7.6	1.9	1376
1982	9.4	-1.5	1359	9.4	-1.5	1359
1983	9.1	3.1	1401	9.1	3.5	1406
1984	8.4	4.1	1459	8.2	4.6	1471
1985	8.0	3.6	1510	7.5	4.1	1532
1986	7.6	3.4	1562	6.6	3.9	1591
1987	7.4	2.9	1607	6.4	3.4	1645
1988	7.1	3.5	1664	6.1	4.0	1711
1989	6.8	3.2	1718	5.8	3.7	1774
1990	6.7	2.5	1761	5.7	3.0	1828
1991	6.6	2.8	1811	5.6	3.3	1888
1992	6.5	2.7	1859	5.5	3.2	1948
1993	6.5	2.3	1903	5.5	2.8	2003
1994	6.5	2.3	1947	5.5	2.8	2059
1995	6.4	2.4	1997	5.4	2.9	2199
1996	6.4	2.4	2045	5.4	2.9	2180
1997	6.4	2.4	2094	5.4	2.9	2243
1998	6.4	2.4	2144	5.4	2.9	2309
1999	6.4	2.3	2193	5.4	2.8	2373
2000	6.5	2.3	2244	5.5	2.8	2440

[a] Assumption 1 unemployment and real GNP growth rate forecasts come from Data Resources, Incorporated, *U.S. Long-Term Review*, Fall 1982, Table 1.

[b] UNEM is the aggregate unemployment rate.

[c] Real GNP estimates are in 1970 billions of dollars.

activity and schooling attainment. For expositional simplicity, we focus only on the projected weekly wages for high school and college graduates.

Figures 2 and 3 introduce the results of our wage forecasts. They show how weekly wages *at each level of work experience* are projected to increase from 1980 to several selected years in the future. The horizontal axis in each graph marks out single years of experience in the labor force. The vertical axis measures the extent of wage increase relative to wage levels prevailing in 1980. Thus, for example, a value of 120 indicates a 20 percent increase over 1980. The graph for each year therefore traces out the percentage increase in wages for every level of

Table 9

ALTERNATIVE SCENARIOS FOR WAGE FORECASTS

Scenario 1:	A.	DRI forecasts of unemployment and GNP growth rates
	B.	Trended schooling distribution
Scenario 2:	A.	DRI forecasts
	B.	Stationary schooling distribution at 1978-1980 levels
Scenario 3:	A.	"Optimistic" forecasts of unemployment and GNP growth rates
	B.	Trended schooling distribution
Scenario 4:	A.	"Optimistic" forecasts
	B.	Stationary schooling distribution

work experience. Note that these are real wage increases which net out the effects of nominal price increases due to inflation. Further, since the effects of macroeconomic variables are assumed to be uniform for all individuals, differential wage increases across experience levels are attributable only to changes in the cohort size variable over time.

Figure 2 summarizes our forecasts of the weekly wages of high school graduates over the next two decades. To illustrate, consider first the group of peak earners with 23–25 years of work experience. Reading up the figure to the 1985 graph, and across to the vertical axis, we see that peak earners' wages increase 1 percent over 1980 levels by 1985. We project larger increases subsequently: 7.5 percent by 1990 and 14 percent by 1995. In contrast, weekly wages of new labor market entrants—those with 1–3 years of experience—are projected to increase much more rapidly, rising 5 percent by 1985, 18 percent by 1990, and 33.5 percent by 1995. While less pronounced, declines in cohort size have similar effects on the weekly wages of college graduates. As shown in Fig. 3, the wages of peak earners rise about 1 percent by 1985, 5 percent by 1990, and 8 percent by 1995. These increases may be compared to increases of 2, 15, and 27 percent, respectively, for new entrants.

Wage increases do not decline monotonically with years of experience. In both Figs. 2 and 3, there is a second "hump" corresponding to more rapid wage growth by those with over 30 years of experience. This unexpected humping effect is due to the small cohorts born

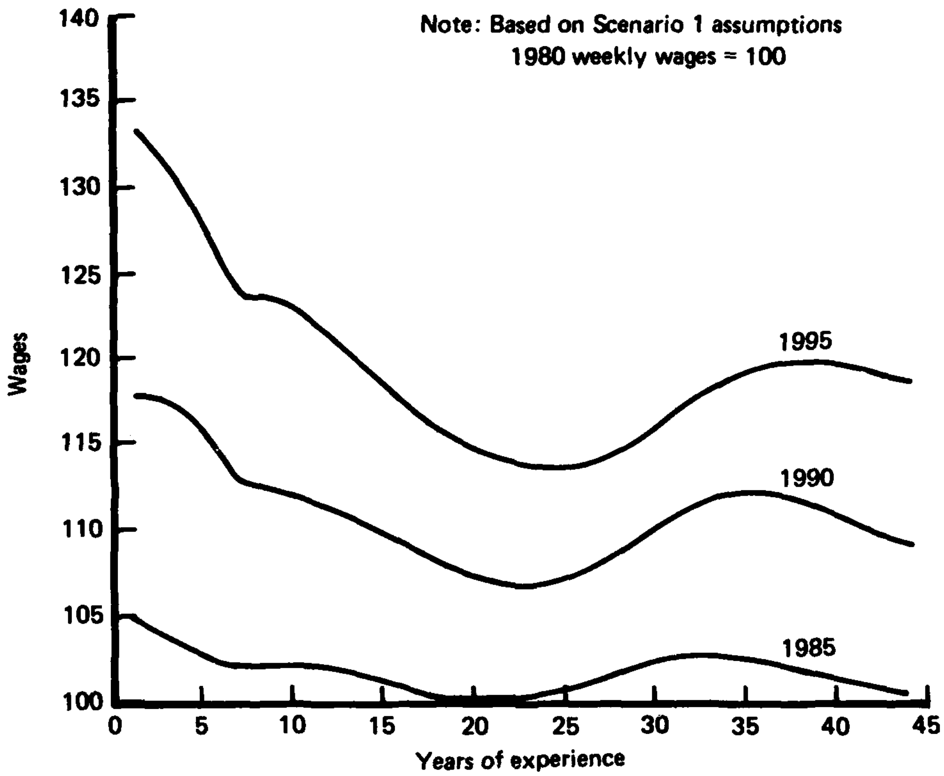


Fig. 2—Wage increase over time by years of experience: high school graduates

during the Great Depression of the 1930s when birth rates fell to historically low levels. This is readily verified in Fig. 2. The 1985 hump peaks at 32 years of experience, meaning that this individual entered the labor market in 1953. If he graduated from high school at 18 or 19, we would place his year of birth at 1934 or 1935, in the middle of the depression. Similar results obtain in Fig. 3 for college graduates. Since they have four extra years of schooling, the hump peaks earlier at 28 years of experience in 1985.

Figures 4 and 5 present forecasts of wage increases from 1980 to 1990 under each of the different scenarios. First, compare scenarios 1 and 2. They differ only in their assumption about school progression rates: trended in scenario 1 but frozen at 1980 levels in scenario 2. As noted earlier, schooling continuation rates have risen over the 1970s for high school graduates but have slightly declined for college

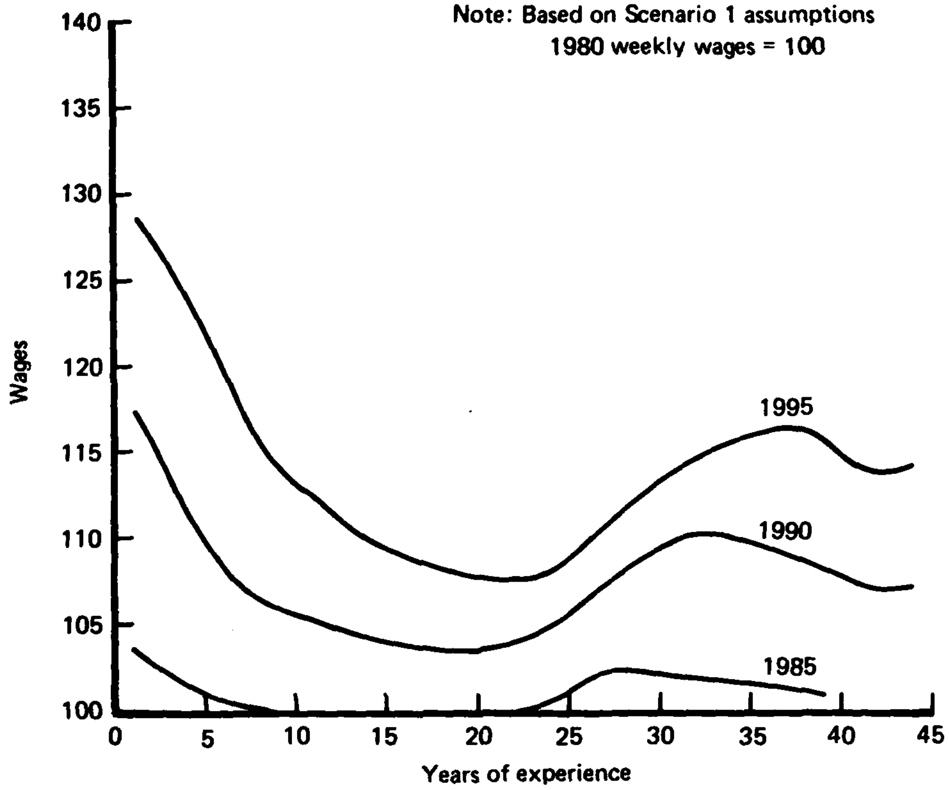


Fig. 3—Wage increase over time by years of experience: college graduates

graduates. Scenario 2 therefore projects smaller entry cohorts of high school graduates and larger entry cohorts of college graduates than scenario 1. It follows, then, that scenario 2 starting wages are higher for high school graduates (Fig. 4) and lower for college graduates (Fig. 5) as compared to scenario 1. Comparisons of scenarios 1 and 3 (or 2 and 4) show how wage forecasts change when we vary the macroeconomic assumptions. Scenarios 3 and 4 embody a more optimistic set of assumptions about GNP growth and unemployment, so larger wage increases are projected for these scenarios compared with 1 and 2. Since these variables are constrained to have a uniform percent wage effect at all experience levels, varying the macroeconomic assumption shifts the wage profiles up or down by the same percentage amount.

The magnitude of these changes are small compared with the wage effects of demographic changes in cohort size. Freezing school

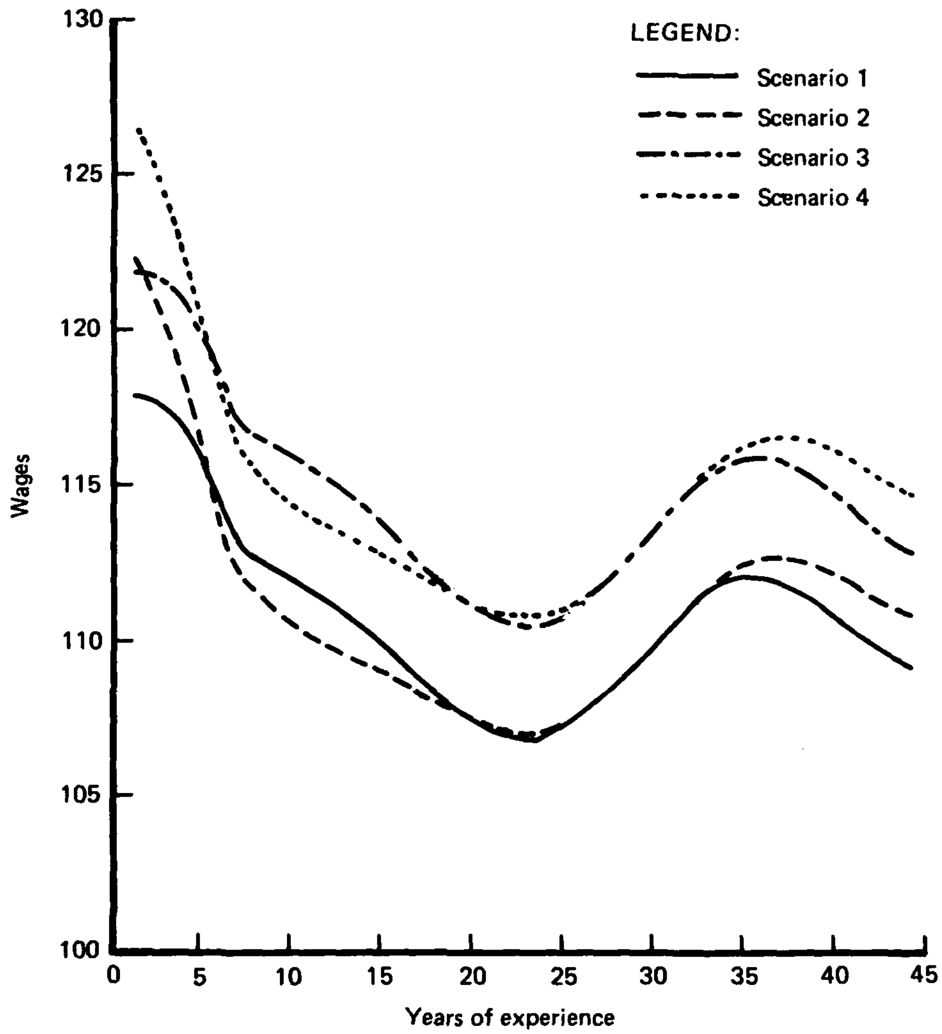


Fig. 4—Wage increase from 1980-1990 by scenario: high school graduates

progression rates increases starting wages of high school graduates by 4 percentage points (from 118) and reduces starting wages of college graduates by no more than 1 percentage point (from 117). In scenarios 3 and 4, the wage effects of more rapid economic growth and lower unemployment rates are small—no more than 4 percent—relative to the effects of demographic factors. These projections show that the effects of cohort size dominate any reasonable variations in other mitigating variables.

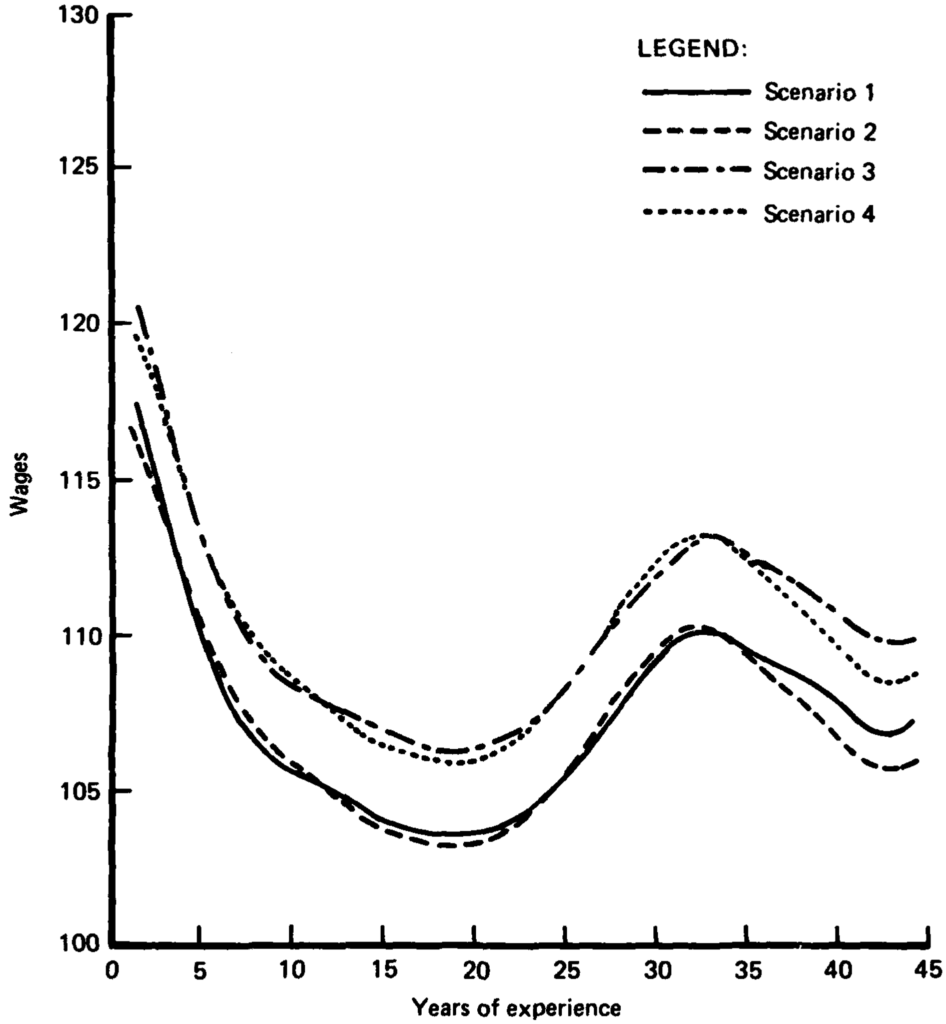


Fig. 5—Wage increase from 1980-1990 by scenario: college graduates

These wage forecasts, while showing a flattening trend over time, nonetheless should be viewed with some caution. Not only are the projections based upon just 14 years of historical data, but they do not incorporate potentially important changes in the structure of the economy, in schooling attainment, and in female labor force participation. In the next section we explore the effects of some of these secular changes.

V. OTHER COMPETING EXPLANATIONS

So far, we have attributed the fall in relative male youth wages primarily to the labor market entry of large male baby-boom cohorts. There are, however, at least two other competing explanations for this phenomenon: one associated with the secular increase in schooling continuation rates, and the other with the rise in female labor force participation rates.[1] Both hypotheses are predicted to have a disproportionately larger impact on the wages of youth relative to more experienced workers. If their wage effects are important, the wage elasticities reported in Section III may overestimate the initial depressant effects of large cohorts in the 1970s and, in consequence, also overstate the rise in relative youth wages forecasted for the 1980s and 1990s.

In this section, we extend the wage model to address both of these issues. In the first case, we include measures of mean cohort schooling attainment to control for the postulated fall in the average ability of more recent cohorts that may have accompanied the rise in schooling continuation rates. In the second case, we expand the cohort size variable to include both the number of males and females with given levels of work experience. This specification incorporates the effects of rising female labor force participation since we use trends in this variable to estimate the experience distribution of female workers. The idea here is to compare the cohort-size wage elasticities estimated with and without controls for the two competing hypotheses. If differences are found, we then ask how our wage projections in Section IV are changed by using the new elasticity estimates.

EFFECTS OF RISING SCHOOLING ATTAINMENT

The first alternative hypothesis attributes the decline in relative youth wages to rising trends in schooling attainment by recent birth

[1]We note, but do not address, a third hypothesis—that structural change may also affect relative wages in the future. The growth of high technology industries would probably create more employment opportunities for recent labor market entrants. On the other hand, declines in manufacturing might depress the wages of more experienced workers who have obsolete skills. In this view, such structural change would tend to accentuate the flattening trend in the wage structure predicted by the cohort size model. In another view, this trend could be offset (at least in part) by recent shifts toward low-wage service-sector employment by youth.

cohorts.[2] Although schooling attainment has risen over time, there is little reason to believe that the same sort of change has occurred in the underlying distribution of "ability." As schooling levels rise, the average ability within a schooling group will fall, as long as unobserved labor market ability and education are positively associated with one another. This is another way of noting that the average ability of (say) high school graduates has declined over the past 30 years because fewer and fewer people halt their education at high school. Those who do so in 1980 are likely to be of lower ability than the high school graduate of 1950.[3] In short, the decline in within-schooling-group relative youth wages may simply reflect the lower average ability of more recent birth cohorts compared with older cohorts.

To address this hypothesis, we include measures of mean cohort schooling in the wage model as controls for unobserved ability. In the literature, there is evidence that the returns to schooling estimated by wage studies are biased upward by the omission of ability (see Griliches and Mason, 1972).[4] Although we do not have measures of ability in our data, we know (or can estimate) the mean schooling attainment of each birth cohort, YOBSMEAN. This variable is, by assumption, uncorrelated with ability. We can also calculate another variable, YOBSDEV—the difference between actual schooling completed and the cohort mean—which will be positively correlated with ability. With data pooled across schooling groups and across different birth cohorts, there is enough variation in these two variables to both identify the "true" returns to schooling and to control for omitted ability bias. More importantly, we will be able to determine if the estimated cohort size wage elasticities are altered when we account for unobserved ability.

We estimate two specifications of a fully interacted wage model pooled across five schooling groups. The first specification includes, in addition to four schooling group intercept terms, a full set of interactions between all variables in the wage model and schooling group dummy variables. Since the coefficient estimates for each group vary freely, this specification replicates the results reported in Section III of

---

[2]The mean educational attainment of birth cohorts has risen dramatically since the turn of the century. In the CPS data, the oldest birth cohort of 1903 had an average of 9.2 years of schooling. Just 27 years later (for the 1930 birth cohort), this figure had risen to 11.9 years. We estimate mean schooling attainment of the 1960 birth cohort at about 14 years. Thus, over a 60 year period, mean cohort schooling attainment will have increased nearly five years.

[3]The assumptions and predictions of this model are discussed in greater detail in Appendix B.

[4]In essence, this ability bias arises because the omitted variable, ability, is thought to be positively correlated with both schooling and with wages.

estimating separate wage models for each schooling group. In the second specification, we replace the schooling group intercepts with YOBSMEAN and YOBSDEV. Table 10 summarizes the cohort-size wage elasticities estimated in these two wage models (the results are reported fully in Table B.2 of Appendix B).

Two major findings are suggested by this table. First, the returns to schooling are substantially overstated by failure to control adequately for ability. Since mean ability is assumed to be invariant across birth cohorts, the coefficient of YOBSMEAN (3.35 percent) is an estimate of the "true" average return to schooling for all year-of-birth cohorts.

Table 10

COHORT SIZE WAGE EFFECTS WITH CONTROLS FOR MEAN COHORT SCHOOLING[a]

	Equation 1		Equation 2	
Variable Description	Coefficient	T-ratio[b]	Coefficient	T-ratio
Cohort size				
Grade school	0.0919	2.89	0.1256	4.71
Non-high school	-0.0994	8.15	-0.1097	8.30
High school graduate	-0.1136	11.03	-0.0908	8.35
Some college	-0.1026	8.81	-0.1068	8.86
College graduate	-0.1592	10.91	-0.1497	10.17
Interaction with spline				
Grade school	0.0714	0.40	0.0249	0.14
Non-high school	-0.1591	7.03	-0.1497	6.42
High school graduate	-0.2439	8.13	-0.2557	8.36
Some college	-0.2730	7.49	-0.2636	7.20
College graduate	-0.0975	2.15	-0.0867	1.90
Schooling intercepts				
Grade school	--	--		
Non-high school	-0.5492	1.57		
High school graduate	-0.2758	0.82		
Some college	1.1200	3.09		
College graduate	0.8781	2.47		
Birth cohort schooling				
Mean cohort education (YOBSMEAN)			0.0335	2.13
Deviation from mean cohort education (YOBSDEV)			0.0743	6.45

[a]The full results are reported in Table B.2 in Appendix B.

[b]T-ratios reported are absolute values.

The coefficient of YOBSDEV (7.4 percent) provides an alternative measure of the return to schooling, but one that is confounded with the effects of ability. The extent of ability bias, as measured by the difference between the two estimates (4.1 percent), is well over 100 percent. A second result more relevant to the major interest of this study is that the coefficients of the cohort size variables are virtually unaffected by controls for mean schooling differences across birth cohorts, i.e., average within-group ability. The one exception—those with 0–7 years of grade school—is not statistically different. In short, we find no evidence that the wage forecasts are biased by our failure to account for unobserved ability.

EFFECTS OF RISING FEMALE LABOR FORCE PARTICIPATION

The second hypothesis attributes the decline in observed relative youth wages to the secularly rising labor force participation of women. By increasing the competition for entry-level jobs, women with little labor force experience may bid down the relative wages of similar males. The wage model used in Section III did not address this point: females were allowed to affect the average level of male wages via the civilian labor force variable but not relative wages. The issue, then, is whether the initial wage effects of cohort size are overestimated by our failure to control for the rising numbers of women in the labor force.

We model the effects of rising female labor force participation rates (LFPR) on male wages through changes in the cohort size variable, which now includes both males and females. For all females in our CPS data, we first develop a distribution of work experience by age, cohort, and schooling group.[5] This imputation procedure is discussed more fully in Appendix C. We then construct a new cohort size variable where each male and female experience cohort is defined relative to the combined total workforce of both males and females. In the regression analysis that follows, both the male and female components of this new cohort size variable (and their interactions with the early

---

[5]Our focus on work experience differs from that of other studies which estimate substitution relationships for worker groups segregated on the basis of age and sex (see Hamermesh and Grant, 1981). Using age as a criterion clearly yields very heterogeneous competing groups, particularly among females who exhibit greater diversity in labor force patterns over their life cycle as compared with men. For example, mature women could include both recent reentrants and those with continuous work histories. And if patterns of female labor force participation change, as they have over time, the experience mix of this group of mature women could also vary from year to year (see Smith and Ward, 1984). The expected work experience variable that we develop in Appendix C overcomes this important shortcoming.

career spline) are entered separately to allow for different impacts on male wages.

Table 11 summarizes the wage elasticities of the expanded cohort size variable (Panel B) and compares them to the estimates based on the male-only cohort size variable (Panel A).[6] The first point to note is

Table 11

THE WAGE EFFECTS OF DIFFERENT COHORT SIZE VARIABLES: A DECOMPOSITION[a]

	Schooling Level Completed			
Model Specification/ Cohort Size Effects	Grade School	High School	Some College	College Graduate
A. Male-only cohort size				
Initial effects	-0.189	-0.305	-0.318	-0.269
	(11.31)b	(11.17)	(13.59)	(6.49)
Persistent effects	-0.128	-0.122	-0.096	-0.152
	(9.12)	(10.72)	(11.40)	(11.53)
B. Male and female cohort size				
Initial effects	-0.172	-0.258	-0.269	-0.244
	(9.56)	(9.55)	(12.01)	(6.63)
Male component	-0.198	-0.176	-0.178	-0.056
	(5.73)	(5.03)	(5.99)	(1.36)
Female component	0.026	-0.083	-0.091	-0.188
	(0.75)	(4.00)	(5.59)	(7.00)
Persistent effects	-0.085	-0.102	-0.111	-0.159
	(4.41)	(8.31)	(5.41)	(6.54)
Male component	-0.167	-0.241	-0.075	-0.126
	(8.42)	(10.72)	(4.64)	(7.18)
Female component	0.081	0.139	-0.036	-0.034
	(3.12)	(5.81)	(2.00)	(3.78)

NOTE: The wage elasticities in Panel A differ slightly from those reported in Table 5 because the cohort size variable was not smoothed.

[a]Based on Table C.1 in Appendix C.

[b]Absolute value of t-statistics is in parentheses.

[6]The wage elasticities in Panel A are not directly comparable with those reported in Section III since moving averages were not used to smooth the different cohort size vari-

that the total initial wage effects are lower in the expanded specification of cohort size, but these differences are not large. For example, the initial effects for high school graduates are estimated to be −0.259 when we account for females (Panel B) versus −0.305 when we ignore them (Panel A). The corresponding figures are −0.244 and −0.269 for college graduates. The results, however, are mixed for the estimated persistent effects: lower for those with high school and less, and higher for those with some college or more.

A pattern of own and cross-group effects also emerges from a comparison of the separate male and female cohort size components. They suggest, quite plausibly, that the size of male cohorts has a greater effect on the level of male wages than female cohort size. With one exception, initial and persistent effects of own cohort size are always larger. For high school graduates, own (male) initial and persistent effects are −0.176 and −0.241, as compared with female effects of −0.083 and 0.139, respectively. For the one exception—college graduates—own initial effects are small and measured very imprecisely (standard error of 0.04). The pattern of female wage effects across schooling groups is intriguing, being positive or insignificant for lower schooling groups, but negative and significant for those with higher schooling attainment. They suggest that female and male workers with a high school education or less are complementary inputs in production whereas those with more schooling are substitutes.

It is not obvious, from these findings, how large an impact not accounting for rising female labor force participation rates (LFPR) will have on our wage forecasts. Assuming that past trends in LFPR continue, the numbers of working females should rise, but this increase will be offset, in part, by the shrinking size of female entry cohorts. The impact would also vary across schooling groups depending upon future changes in the schooling completion rates of females. To estimate these effects, we use the new wage elasticities from the expanded (male and female) cohort size model to forecast weekly male wages for 1990 and 1995. We base this exercise on scenario 1 assumptions about the future path of the major variables: trended schooling attainment for both males and females, and the DRI forecasts of GNP and unemployment rates. In addition, we assume that female LFPR continues to rise to 1990 (though at a slower rate), after which they level off.

ables. This decision was motivated by the algorithm used to calculate female experience. Because we impute low work experience to a large number of females, smoothing the cohort size variable would tend to dampen precisely the effect (of rising female LFPR) that we are trying to model, i.e., the early "spike" in the female experience distribution. Thus, in this exercise, both the expanded and male-only cohort size variables are left unsmoothed.

Specifically, we allow female LFPR to rise by half as much over the next 10 years as they did in the past decade (1970–1980).[7]

Table 12 summarizes our wage forecasts for two groups: new entrants with 1–5 years of work experience and prime age males with 23–27 years of experience. Panel A reports the percentage increase in real weekly wages of these groups for the periods 1981–1990 and 1981–1995. Panel B shows the wages of new entrants compared with wages of prime age males in 1990 and 1995. To facilitate comparisons with these figures, the results using the male-only cohort size model are also shown in this table.

Two important results emerge from these comparisons. First, the male-only cohort size model is seen to overstate the rise in weekly wages of *both* new entrants and prime age males. For example, for the 1981–1990 period, the real wages of high school graduates are projected to rise 16.1 percent for new entrants and 6.9 for prime age males when we use the male cohort size model. When we account for the rising number of females, the corresponding wage increases are estimated to be 12.5 percent and 4.6 percent, respectively. A similar pattern also holds for the longer period from 1981 to 1995. Second, because we overstate the wage increases of both groups, the net result is that projected relative youth wages are only marginally lower in the expanded cohort size model. Using the case of high school graduates again, Panel B indicates that we overpredict the rise in relative youth wages by only 0.6 percent (0.576–0.570) by 1990, and by 1 percent (0.605–0.595) by 1995. This "prediction error" is small compared to the increases in relative youth wages projected under the male cohort size model—4.6 percent by 1990 and 7.5 percent by 1995. In other words, over the next two decades rising female labor force participation will not substantially mitigate the projected flattening of civilian wage profiles, although it may affect the rate at which average civilian wages rise.

To summarize, the wage models that we have estimated in Section III appear to capture reasonably well the movements in wage structure that have taken place over the sample period. In this section, we tested two competing hypotheses and found no cause to prefer them, at least for forecasting purposes, over the simpler model based on male cohort size. We are confident that this model yields robust projections about future changes in relative youth wages and, in particular, the flattening trend in civilian wage profiles over the next two decades. And while it tends to overstate real increases in the average level of civilian wages, as compared with the expanded cohort size model, the differences are not large.

---

[7]Recent evidence for the 1980s suggests that this is an accurate assumption.

Table 12

SENSITIVITY OF WAGE FORECASTS TO RISING FEMALE LABOR FORCE PARTICIPATION

	Schooling Level Completed			
Model Specification	Grade School	High School	Some College	College Graduate
A. Forecasted wage increase (%)				
1981-1990				
Male model				
New entrants	8.1	16.1	10.4	13.7
Prime age males	5.5	6.9	1.1	5.9
Male and female model				
New entrants	5.6	12.5	7.4	9.1
Prime age males	5.2	4.6	0.6	4.6
1981-1995				
Male model				
New entrants	16.2	29.2	16.2	25.7
Prime age males	11.0	13.2	1.6	9.4
Male and female model				
New entrants	13.4	24.6	11.7	18.2
Prime age males	11.0	11.1	0.4	7.0
B. Relative youth wages				
1990				
Male model	0.401	0.576	0.569	0.598
Male and female model	0.393	0.570	0.557	0.581
1995				
Male model	0.410	0.605	0.596	0.639
Male and female model	0.400	0.595	0.580	0.615

NOTES: Forecasted wages are based on scenario 1 assumptions. Female labor force participation rates are assumed to rise by half as much from 1981 to 1990 as they did in the previous 10-year period, and flatten out thereafter.

VI. SUMMARY AND IMPLICATIONS

SUMMARY

As members of baby-boom cohorts entered the labor market in the 1960s and 1970s, their wages fell in comparison with the wages of prime age workers. This observation motivated us to ask the question: How will the age structure of civilian wages change as the demographic trend reverses itself in the 1980s and 1990s? We approached this question by estimating wage models for white males, using *Current Population Survey* data for 1967 to 1980. In these models, estimated separately for several schooling groups, we distinguished between the wage effects of cohort size and changes in macroeconomic variables, such as business cycles and secular growth in aggregate labor productivity. Having developed these estimates, we were able to project wage changes over the next two decades for a variety of alternative economic scenarios.

Our regression analysis indicated that cohort size has much larger wage effects for new labor market entrants than for more experienced workers, a finding common to all schooling groups studied. On average, a 1 percent increase in cohort size *reduces* the weekly wages of new entrants by between 0.25 and 0.35 percent. As work experience is acquired, however, the cohort size effect falls to about the 0.10 to 0.15 range. We also found evidence that large cohorts affect not only wages but also weeks worked per annum, and that the reduction in labor supply is greatest for those with the least schooling.

The estimated cohort-size wage elasticities accorded well with the observed decline in relative youth wages on from 1967 to 1977. Over this period, for example, the proportion of high school graduates with 1 to 5 years of experience rose over 30 percent. Our prediction of a 10 percentage point decline in their wages, relative to wages of prime age males, compares favorably with the 11 percent fall actually observed. These, however, are the partial wage effects of cohort size controlling for the effects of other macroeconomic changes. Of these, secular increases in aggregate labor productivity over this period may have been the most important. Controlling for the size of the labor force, a 1 percent increase in GNP was found to elicit about an equal percentage increase in weekly wages. Changes in the unemployment rate (a control for business cycles) had mixed effects on wages, but these effects were not measured precisely.

The main conclusion to be drawn from our forecasts is that wage profiles will flatten as the demographic trend toward smaller entry cohorts reverses itself. For example, the proportion of new entrants in the workforce with a high school education peaked in 1978 at over 22 percent, and is projected to decline to 17 percent by 1990, and to 15 percent by the year 1995. We forecast that these changes will be accompanied by a rise in relative youth wages of about 5 percentage points by 1990, and 8 percentage points by the year 1995. Further, these shifts in wage structure will translate into substantially larger real wage increases for young men than for more experienced workers. In the most likely scenario, we see weekly wages of high school new entrants increasing 17 percent from 1981 to 1990, and over 30 percent by 1995. In contrast, weekly wages of peak earners are projected to rise only 7 percent by 1990, and 14 percent by 1995.

How sensitive are these wage forecasts to possible changes in the future economic environment? We attempted to answer this question by developing wage projections for four alternative scenarios which combined different assumptions about the future course of economic growth, unemployment, and schooling continuation rates. The differences in wage forecasts across these four scenarios were found to be small relative to the effects of demographic factors. From this experiment, we concluded that the effects of cohort size are likely to dominate any reasonable assumptions about other mitigating variables.

These findings were not changed materially when we accounted for two competing explanations for the 1967–1978 decline in relative youth wages. One explanation, associated with rising trends in schooling attainment, attributes this decline to a fall in the average market "ability" of more recent birth cohorts. When we included birth cohort-specific controls for unobserved ability, we found some evidence of ability bias in the estimated returns to schooling, but no indication of bias in the estimated cohort-size wage elasticities. A second explanation links declines in youth wages to the rise in female labor force participation, arguing that females may have bid down the wages of younger males. We found some support for this hypothesis. Including females in the model reduced the initial wage effects of cohort size, but did not change persistent wage effects in a systematic fashion. Using these new wage elasticities, we developed a new set of wage projections which assumed a continued (though lower) rise in female labor force participation rates. Compared with the previous estimates, we projected somewhat lower wage increases for *both* new entrants and prime age males, but only marginally lower relative youth wages. In other words, rising female labor participation over the next two decades will not substantially mitigate the projected flattening in civilian wage profiles, though it may affect the rate at which average civilian wages rise.

IMPLICATIONS FOR THE ARMED FORCES

The implications of these findings for the military are that declines in cohort size will raise the future cost of attracting new recruits. And these costs will come on top of future declines in the number of "high quality" male accessions projected by studies of enlisted supply such as Fernandez (1979) and Cotterman (1984).[1] Employers, both civilian and military, will face recruiting problems over the next decade and a half. But because the services rely almost entirely on youth for recruits, they will be particularly hard-hit by the rising cost and shrinking supply of youth.

The projected changes in civilian wage structure will also affect the comparability of civilian and military pay structures. Under the present system, overall "comparability" is maintained (at least in principle) by linking across-the-board military pay increases to changes in an index of mean civilian wages, such as the Employment Cost Index (ECI) or the Professional, Administrative, Technical and Clerical (PATC) index which it is to replace. Being indices of *mean wages*, they offer no information on the age structure of civilian wages and, therefore, would not have picked up the flattening trend in wage structure that has now begun. Their continued use in the face of these trends implies that military pay by, say, length of service will become increasingly steeper relative to civilian wage profiles. And while this would tend to keep military pay of careerists roughly comparable to the wages of their civilian counterparts, it would make starting military pay significantly less competitive with civilian wages.

The following two figures illustrate these points. Figure 6 shows the future military pay increases required to maintain accessions of high school graduates at current levels. This assumes comparability of starting military and civilian pay in 1983, the base year. The bottom bar represents military pay adjustments required to keep up with *real* increases in average civilian wages. Under scenario 1 assumptions, this projected increase is about 13.5 percent from 1983–1990, and 22.4 percent from 1983–1995. Two additional entry-level pay increases will be required because of future declines in entry cohort size. The middle

[1]"High quality" accessions are typically defined as those with a high school diploma who test in the Category 1 through 3a range in the Armed Forces Qualifying Test (AFQT).

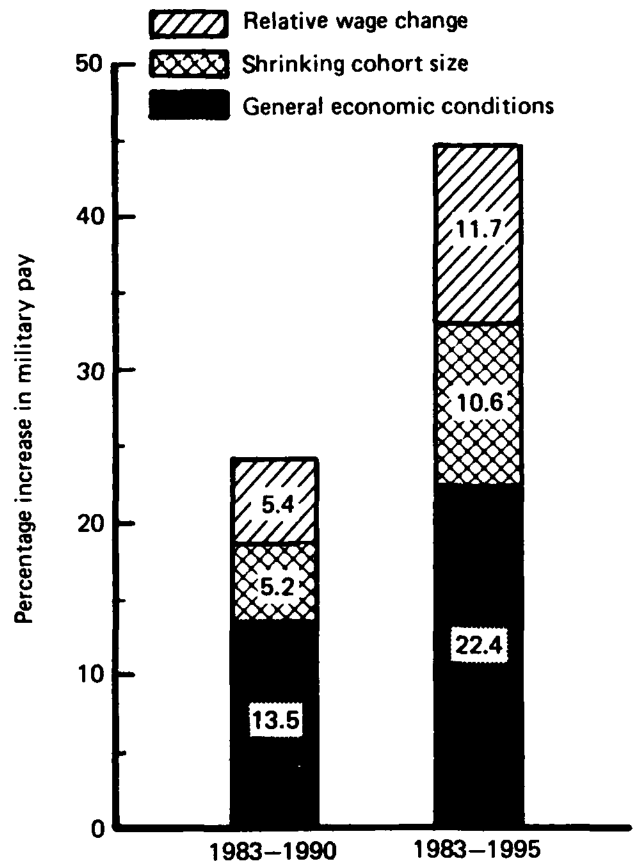


Fig. 6—Military pay increases required to maintain current accessions

bar represents the pay increase required just to offset the projected decline in the accession pool and recruit the same number of high school graduates. Implicit in this calculation are an assumed unitary accession pay elasticity and a population elasticity of one half.[2] The top bar is the pay adjustment required to compensate for the projected rise in youth wages relative to the civilian average. Taken together, these cohort-size-related adjustments should add 10.6 percentage points to military pay increases by 1990, and 22.3 percentage points by 1995. These pay adjustments will be over and above those required to keep

[2]Given the projected 10.5 percent decline in the high school accession pool by 1990, the population elasticity of 0.5 implies that a pay increase of over 5.2 percent (0.5 × 10.5) will be needed to attract the same number of high school graduates.

up with average civilian wage increases due to general economic conditions.

Figure 7 indicates that different pay increases will also be required to maintain first and second term pay comparability. This is because the entire civilian wage structure will be affected by future changes in cohort size. As before, the bottom bar represents the pay increase due to general economic conditions whereas the top bar measures the additional pay adjustment required to compensate for relative wage changes. These numbers are based on the assumption that the personnel force structure remains unchanged over time. The figure shows that military pay increases tied to an index of mean civilian wages produces *rough* pay comparability in the second term. This result is not surprising since the length of service of second termers is reasonably close to the average experience of the labor force as a whole. *Without*

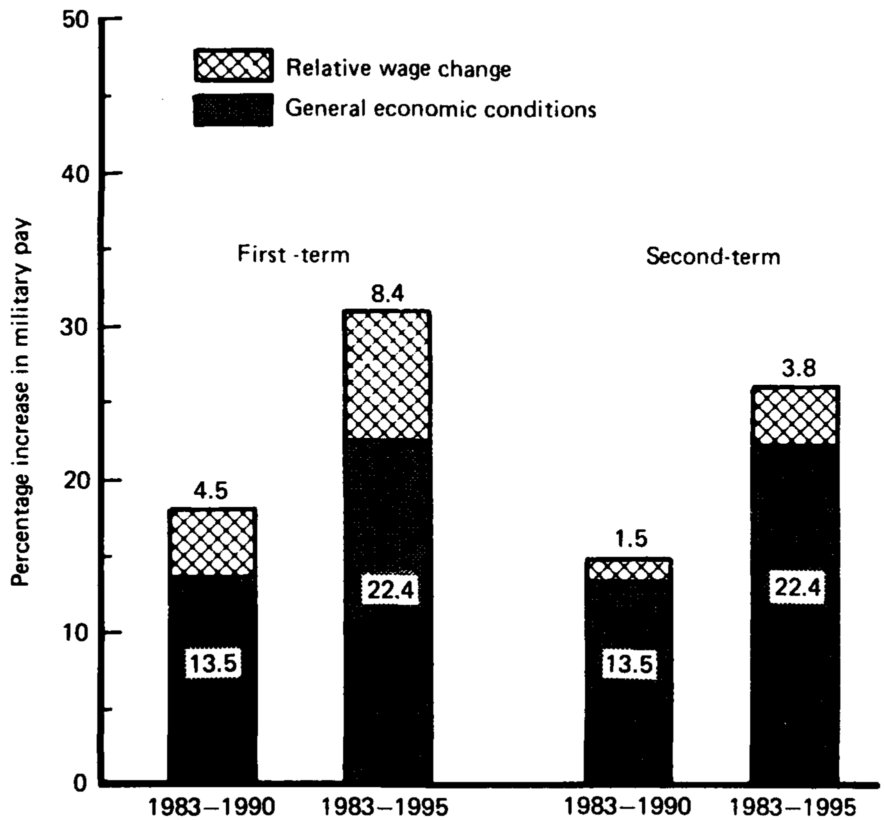


Fig. 7—Military pay increases required to maintain pay comparability

compensating pay increases, however, comparability will be eroded significantly for first termers, and still more for accessions (as shown in the previous figure).

The tentative nature of this exercise should be emphasized. Being illustrative, these figures depend importantly on our assumptions about the future economy (scenario 1), specific pay and population elasticities, and an unchanged personnel force structure. Outcomes could well change (perhaps greatly) with a different set of assumptions.

How the services respond will depend, in large part, on the desired personnel force structure. If the services continue to rely on a large first-term enlisted force, targeting of military pay increases and bonuses at the enlistment point may be necessary to meet future accession requirements. Alternatively, if the force should become more career and less first-term oriented, other compensation structures may be needed to enhance the retention of career personnel. In either case, knowledge of the age structure of civilian pay, and how this wage structure will change over time, is crucial. This study provides the estimates needed to begin formulation of compensation policies appropriate for the desired force structure.[3]

[3]The work by Hosek, Fernandez, and Grissmer (1984) on the prospects for future enlisted supply is a promising start in this direction. Using a force structure model with enlistment and retention modules, they examine the implications of alternative military pay scenarios (based on the wage forecasts provided by this study) for the enlisted personnel force structure through the end of the current decade.

Appendix A

REGRESSION RESULTS OF THE WAGE MODEL

Table A.1

DETERMINANTS OF THE WEEKLY WAGES OF WHITE MALES: 1967-1980

Independent Variable	Schooling Level Completed			
	Grade School	High School	Some College	College Graduate
Log cohort size:				
Main effect	-0.0994 (7.05)	-0.1136 (9.47)	-0.1027 (11.38)	-0.1592 (10.92)
Interaction with spline	-0.1591 (6.08)	-0.2439 (6.98)	-0.2730 (9.67)	-0.0975 (2.15)
Years of experience:				
Early career spline	-1.0574 (11.55)	-1.0772 (9.94)	-1.1893 (14.27)	-0.4398 (3.04)
Experience	0.0411 (50.23)	0.0373 (46.15)	0.0390 (36.36)	0.0579 (39.28)
Experience squared	-0.0006 (34.99)	-0.0008 (44.88)	-0.0008 (38.00)	-0.0014 (37.71)
Fraction excluded:				
Nonwork	-0.0818 (0.93)	0.7410 (6.17)	1.2900 (7.51)	2.5356 (9.35)
Allocated income	-0.2529 (3.18)	-0.1832 (3.01)	-0.2938 (4.19)	-0.2768 (3.59)
Log real GNP	0.8207 (6.01)	0.9929 (9.37)	0.3402 (2.92)	0.8873 (6.38)
Unemployment rate	0.0059 (2.33)	0.0034 (1.83)	-0.0023 (1.14)	0.0068 (2.66)
Log civilian labor force	-1.0242 (5.29)	-1.2906 (8.78)	-0.5360 (3.28)	-1.3648 (6.99)
Intercept	2.9264	3.1997	4.5956	4.3537
R-square	0.985	0.986	0.988	0.979

NOTE: Dependent variable = log real weekly wage in 1970 dollars; absolute values of t-statistics are in parentheses.

Table A.2

DETERMINANTS OF THE ANNUAL EARNINGS OF WHITE MALES: 1967-1980

Independent Variable	Schooling Level Completed: Grade School	High School	Some College	College Graduate
Log cohort size:				
Main effect	-0.1859	-0.1534	-0.1564	-0.1680
	(10.03)	(9.98)	(15.54)	(10.84)
Interaction with spline	-0.0870	-0.0475	-0.2849	-0.2098
	(2.53)	(1.06)	(9.04)	(4.36)
Years of experience:				
Early career spline	-0.9752	-0.5525	-1.2814	-0.8535
	(8.10)	(3.98)	(13.77)	(5.55)
Experience	0.0564	0.0431	0.0390	0.0592
	(52.39	(41.56)	(32.58)	(37.75)
Experience squared	-0.0009	-0.0009	-0.0009	-0.0014
	(37.13	(40.16)	(35.37)	(36.61)
Fraction excluded:				
Nonwork	0.2217	0.4534	0.6322	2.4222
	(1.92)	(2.95)	(3.30)	(8.40)
Allocated income	-0.2057	-0.1954	-0.0791	-0.2462
	(1.97)	(2.50)	(1.01)	(3.01)
Log real GNP	1.0209	1.2332	0.3522	0.7816
	(5.68)	(9.08)	(2.71)	(5.28)
Unemployment rate	-0.0154	0.0091	-0.0125	0.0000
	(4.64)	(3.77)	(5.50)	(0.01)
Log civilian labor force	-1.4444	-1.6153	-0.5741	-1.2054
	(5.67)	(8.57)	(3.15)	(5.81)
Intercept	6.7738	6.7018	8.4133	8.2875
R-square	0.986	0.985	0.988	0.979

NOTE: Dependent variable = log real annual earnings in 1970 dollars; absolute values of t-statistics are in parentheses.

Appendix B

SCHOOLING INFLATION AND UNOBSERVED ABILITY

In this appendix, we consider a competing hypothesis that attributes declines in within-schooling-group relative youth wages over the 1967–1980 period not to cohort-size changes but to the secular increases in schooling attainment of more recent birth cohorts. This "schooling inflation" hypothesis argues that increases in schooling continuation rates tend to draw into any given schooling group individuals that are, on average, of lower innate ability. Thus, declines in relative youth wages observed *within a schooling group* may simply reflect a lowering of the mean ability of more recent birth cohorts. The interest here is in determining the extent to which our estimates of cohort-size wage elasticities are biased by omitting controls for unobserved market ability.

Figure B.1 illustrates the main features of this model. We make two assumptions: (1) the distribution of unobserved ability is invariant across year-of-birth cohorts and (2) in a given year-of-birth cohort, individual ability and schooling attainment are positively correlated. If this correlation remains unchanged over time, then (1) and (2) imply that the mean schooling attainment of each birth cohort is uncorrelated with ability. These assumptions may be depicted in the figure as two ellipsoids, one for an older cohort (I) and another for a more recent birth cohort (II). The second ellipse is shifted over to the right to represent the rise in schooling attainment of more recent birth cohorts, but in such a way as to keep mean ability equal in both ellipsoids. From this figure, we can show that a rise in school attainment rates draws down the mean ability (A) in *each and every schooling group*. For any arbitrary schooling interval, such as S1 and S2, average ability in the more recent birth cohort (A-II) is always lower than that in an older cohort (A-I).

Our empirical approach in addressing this competing hypothesis is to extend the wage model to include measures of schooling attainment for each year-of-birth cohort.[1] Several researchers (e.g., Griliches and Mason, 1972) have noted that the estimated returns to schooling are upward biased by the omission of ability, which is thought to be

[1]The formal model is available on request from the authors.

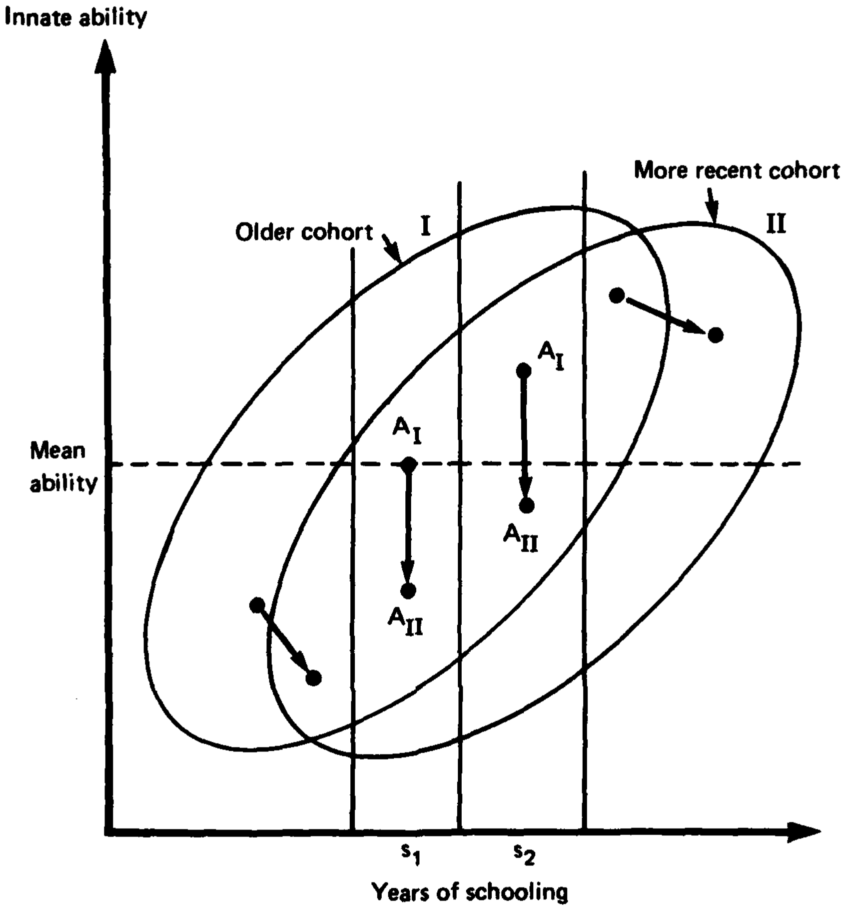


Fig. B.1—Schooling and ability in two year-of-birth cohorts

positively correlated with schooling attainment. We do not have measures of ability. However, by assumption, we know that the mean schooling attainment of each birth cohort (YOBSMEAN) is uncorrelated with ability. In each birth cohort, we can also calculate the difference between actual schooling completed and the cohort mean, a variable YOBSDEV that will be positively correlated with ability. With time-series cross-section data pooled across schooling groups and several year-of-birth cohorts, these two variables should allow us to both identify the "true" returns to schooling and control for omitted ability bias.

Our estimates of mean schooling completed for different birth cohorts are reported in Table B.1. These figures are based on CPS data. Whenever possible, mean cohort schooling (YOBSMEAN) is calculated at age 30 when, presumably, most individuals would have completed schooling. For cohorts older than 30 in 1968 (when our data begin), YOBSMEAN is computed from reported schooling in the 1968 survey. For the more recent cohorts (those aged 29 or less in 1981), we predict mean schooling completed from a regression of YOBSMEAN on a quadratic specification of year of birth (YOB). The table shows that mean schooling attainment rose steadily in the first half of the century, and accelerated in the latter half. The dramatic rise in YOBSMEAN, particularly for cohorts born after 1940, is consistent with the well publicized increase in post-war school continuation rates.

We estimate two specifications of a fully interacted wage model which pooled data from all five schooling groups. In one, the baseline specification, we replicate the results (reported in Section III) of estimating separate wage models for each schooling group. This is because the fully interacted model allows different coefficient estimates and constant terms for each schooling group. In the second specification, we replace the schooling group constant terms with the two

Table B.1

MEAN YEARS OF SCHOOLING COMPLETED BY SELECTED YEAR-OF-BIRTH COHORTS (WHITE MALES)

Year of Birth (YOB)	Mean Years of Schooling (YOBSMEAN)
1903	9.25
1905	9.71
1910	10.09
1915	10.44
1920	11.11
1925	11.52
1930	11.96
1935	12.23
1940	12.28
1945	12.71
1950	13.54
1955a	13.62
1960a	13.94
1965a	14.24

[a]A predicted value for mean cohort schooling attainment.

cohort schooling variables, YOBSMEAN and YOBSDEV. The regression results for the two specifications are reported in Table B.2 and discussed in the text (see Section V).

Table B.2

POOLED WAGE REGRESSIONS WITH MEAN SCHOOLING BY YEAR-OF-BIRTH COHORT

	Equation 1		Equation 2	
Independent Variable	Coefficient	T-ratio	Coefficient	T-ratio
Intercept	3.474	11.01	2.714	17.49
Log cohort size				
Grade school	0.091	2.89	0.125	4.71
Non-high school	-0.099	8.15	-0.110	8.30
High school graduate	-0.114	11.03	-0.091	8.35
Some college	-0.103	8.82	-0.107	8.86
College graduate	-0.159	10.91	-0.149	10.17
Cohort size × spline				
Grade school	0.071	0.40	0.025	0.14
Non-high school	-0.159	7.03	-0.150	6.42
High school graduate	-0.243	8.13	-0.256	8.36
Some college	-0.273	7.49	-0.263	7.20
College graduate	-0.097	2.15	-0.087	1.90
Early career spline				
Grade school	-0.320	0.41	-0.591	0.74
Non-high school	-1.057	13.35	-1.035	12.35
High school graduate	-1.077	11.57	-1.139	11.89
Some college	-1.189	11.04	-1.124	10.41
College graduate	-0.439	3.04	-0.362	2.50
Years of experience				
Grade school	0.038	18.30	0.035	16.47
Non-high school	0.041	58.04	0.039	38.25
High school graduate	0.037	53.73	0.034	35.28
Some college	0.039	28.16	0.037	24.12
College graduate	0.058	39.23	0.056	33.02
Experience squared (× 100)				
Grade school	-0.067	21.08	-0.069	21.59
Non-high school	-0.065	40.43	-0.655	40.53
High school graduate	-0.077	52.26	-0.077	51.86
Some college	-0.083	29.42	-0.083	29.37
College graduate	-0.136	37.66	-0.135	37.67
Allocated income				
Grade school	-0.431	4.48	-0.403	4.37
Non-high school	-0.253	3.68	-0.167	2.61
High school graduate	-0.183	3.50	-0.114	2.37
Some college	-0.294	3.24	-0.505	6.03
College graduate	-0.277	3.59	-0.406	5.56

Table B.2—continued

	Equation 1		Equation 2	
Independent Variable	Coefficient	T-ratio	Coefficient	T-ratio
Nonwork				
Grade school	0.387	4.07	0.418	4.87
Non-high school	0.082	1.08	-0.000	0.00
High school graduate	0.741	7.19	0.905	8.74
Some college	1.290	5.86	1.191	5.37
College graduate	2.536	9.34	2.320	8.42
Log real GNP				
Grade school	1.644	7.32	1.817	9.61
Non-high school	0.821	6.95	0.640	6.26
High school graduate	0.993	10.91	0.819	10.65
Some college	0.340	2.26	0.873	7.48
College graduate	0.887	6.37	1.203	10.73
Unemployment rate				
Grade school	0.023	5.55	0.028	6.75
Non-high school	0.006	2.69	0.003	1.69
High school graduate	0.003	2.13	0.002	1.57
Some college	-0.002	0.88	0.004	1.73
College graduate	0.007	2.66	0.011	4.81
Log civilian labor force				
Grade school	-2.348	7.39	-2.385	8.37
Non-high school	-1.024	6.11	-0.741	4.79
High school graduate	-1.290	10.22	-0.968	8.54
Some college	-0.536	2.54	-1.061	5.99
College graduate	-1.365	6.99	-1.647	9.82
Schooling group dummy				
Non-high school	-0.549	1.57		
High school graduate	-0.276	0.82		
Some college	1.120	3.09		
College graduate	0.878	2.47		
Birth cohort schooling				
Mean cohort education (YOBSMEAN)			0.033	2.13
Deviation from mean cohort schooling (YOBSDEV)			0.074	6.45
R-square	0.989		0.989	
Mean square error	0.108		0.109	
F-ratio	5236		5402	

NOTES: Dependent variable: logarithm of real weekly wages. Regressions are estimated by weighted least squares. T-statistics are absolute values.

Appendix C

WAGE EFFECTS OF MALE AND FEMALE COHORT SIZE

In this appendix, we discuss the methodology that we used to model the effects of rising female labor force participation on male wages. In the text (see Section V), we noted a shortcoming inherent in many studies which estimate substitution relationships for worker groups segregated on the basis of age and sex (e.g., Hamermesh and Grant, 1981). If work experience is what matters in how well one group substitutes for another, then an age criterion would clearly yield very heterogeneous competing groups, particularly among females who exhibit greater diversity in labor force patterns over their life cycle as compared with men. Thus, the group of mature women could include both recent reentrants and those with continuous work experience. The experience mix of mature women could also vary from year to year if patterns of female labor force participation change, as they have over time.

We avoid this shortcoming by developing an expected work experience variable for females in our data, one that accounts explicitly for both birth cohort and age and schooling-specific changes in their labor force participation rates over time. In this way, the effects of rising female labor force participation are incorporated into a variable which measures the changing experience composition of the female workforce. This allows us to broaden the definition of cohort size to include both males and females and ask: To what extent are the relative wages of males affected by variations in the stock of experience capital in the labor market, both male and female?

Our approach builds on earlier work by Heckman and Willis (1977) on the labor force participation of women. From panel data, they estimate the probability that a woman works j out of n years by a beta distribution which can be characterized by just two parameters, α and β. We use their results to write these parameters as

$$\alpha = 0.232 \text{ and } \beta = [(1 - P)/P] \times 0.232$$

where P is the mean annual labor force participation rate (LFPR). Given P, the two parameters can be used to generate the full density of probabilities needed to calculate expected work experience for women in our CPS data.

For this calculation, we developed a dataset on annual LFPR by schooling group and age interval for the 1940 to 1980 period. Weekly LFPR data by these attributes are available from 1959 to 1979 in the *Current Population Survey* special reports on educational attainment; for the earlier years, similar kinds of data can be found in the *Population Census* of 1940 and 1950. To fill in the gaps in this series, we used aggregate data on weekly LFPR by age which is available for the entire period from 1940 to 1980. These data were constructed using information from the *Historical Statistics of the United States: Colonial Times to 1957* and from the *Handbook of Labor Statistics*. We regress age and education-specific LFPR on aggregate weekly LFPR by age, and use the coefficient estimates to infer weekly LFPR by age and education for the intervening years from 1941 to 1958. An expansion factor—based on the fraction of women who worked zero hours in 1970—is then used to convert weekly data to annual LFPR.

The next step was to calculate the full density of probabilities for females in our CPS dataset. In each of the 14 survey years, we know the number of women, their age, and their schooling attainment. From this information, we also know their year of birth and can infer the maximum number of years n that they could have worked from graduation to the survey year. In each survey year, we first calculate P (the average LFPR) for each schooling-birth cohort using annual LFPR data prevailing over these n years. In these calculations, we assume that women who enter the labor force prior to 1940 had LFPR that were similar to those prevailing in 1940. Next, we use P and the parameters α and β to estimate the matrix of probabilities that they worked j out of n years. We then repeat these calculations for the next survey year, updating P as women gain another calendar year of potential work experience n. In the final step, we use this matrix to create a new dataset arrayed by schooling group, single full years of work experience, and survey year. Like the male dataset, the number of females in the i-th year of experience is the sum across individual i-th year probabilities.[1]

We use these data to create a new cohort size variable that includes both males and females. This time, the numbers of males and females in each experience cohort are measured relative to the combined total

[1]Implicitly, these calculations assume that work experience is not depreciated by time out of the labor force. Thus, two five-year interrupted work segments are treated like 10 years of continuous work experience. In future research, these calculations could be generalized to incorporate alternative assumptions about depreciation from time out of the labor force.

workforce of both males and females. In a new set of regressions, we include the male and female components of the new cohort size variable (and their interactions with the early career spline) separately to see if they have a differential impact on male wages. These results are reported in Table C.1 together with those using the male-only cohort size variable.

One difference between this exercise and earlier analyses should be noted. The new cohort size variables are not smoothed using the procedure described in Section II. This decision was motivated by the algorithm used to calculate female experience. Because we impute low work experience to a large number of females, smoothing the cohort size variable would tend to dampen precisely the effect we are trying to

Table C.1

THE WAGE EFFECTS OF DIFFERENT COHORT SIZE MEASURES

Model Specification/ Cohort Size Effects	Schooling Level Completed: Grade School	High School	Some College	College Graduate
A. Male cohort size model				
Cohort size	-0.1279	-0.1219	-0.0963	-0.1529
	(9.12)	(10.72)	(11.40)	(11.53)
Interaction with spline	-0.0611	-0.1834	-0.2220	-0.1164
	(2.91)	(6.58)	(8.88)	(2.62)
B. Male and female cohort size model				
Cohort size				
Male component	-0.1668	-0.2414	-0.0752	-0.1259
	(8.42)	(10.72)	(4.64)	(7.18)
Female component	0.0812	0.1394	-0.0359	-0.0337
	(3.12)	(5.81)	(2.00)	(3.78)
Interaction with spline				
Male component	-0.0310	0.0656	-0.1030	0.0696
	(0.69)	(1.42)	(2.90)	(3.78)
Female component	-0.0551	-0.2222	-0.0551	-0.1541
	(1.23)	(6.66)	(2.52)	(5.65)

NOTES: Absolute values of t-statistics are in parentheses. Cohort size variables in the two models have not been smoothed by moving averages. As such, the wage elasticities in Panel A will not correspond with those reported in the text and in Table A.1.

model, i.e., the early "spike" in the female experience distribution. Thus, in this exercise, both the expanded and male-only cohort size variables are left unsmoothed. Consequently, the wage elasticities in Panel A do not correspond with those reported in Section III.

BIBLIOGRAPHY

Berger, Mark C., "The Effect of the Baby Boom on the Earnings of Young Males," NBER Conference Paper, No. 98, February 1981.

Cotterman, Robert F., *A Time Series of Cross Sections Model for Forecasting Enlisted Supply,* The Rand Corporation, R-3162-MIL, forthcoming.

Data Resources, Incorporated, *United States Long-Term Review,* Cambridge, Massachusetts, Fall 1982.

Doering, Z. D., D. W. Grissmer, J. A. Hawes, and W. P. Hutzler, *1978 DoD Survey of Officers and Enlisted Personnel: User's Manual and Codebook,* The Rand Corporation, N-1604-MRAL, January 1981.

Easterlin, Richard, "What Will 1984 Be Like? Socio-economic Implications of Recent Twists in Age Structure," *Demography,* 4, November 1978, 397–432.

Fernandez, Richard L., *Forecasting Enlisted Supply: Projections for 1979–1990,* The Rand Corporation, N-1297-MRAL, September 1979.

Freeman, Richard D., "The Effect of Demographic Factors on Age-Earnings Profiles," *The Journal of Human Resources,* 14, Summer 1979, 289–318.

Griliches, Zvi, and William Mason, "Education, Income and Ability," *Journal of Political Economy,* 3, May/June 1972, S74–103.

Hamermesh, Daniel S., and James Grant, "Labor Market Competition among Youths, White Women, and Others," *Review of Economics and Statistics,* 68, August 1981, 354–360.

Heckman, James J., and Robert J. Willis, "A Beta-logistic Model for the Analysis of Sequential Labor Force Participation by Married Women," *Journal of Political Economy,* 1, February 1977, 27–58.

Hosek, James R., Richard L. Fernandez, and David W. Grissmer, *Active Enlisted Supply: Prospects and Policy Options,* The Rand Corporation, P-6967, March 1984.

Lillard, Lee, James Smith, and Finis Welch, "What Do We Really Know about Wages: The Importance of Non-reporting and Income Imputation," The Rand Corporation, unpublished draft.

Smith, James P., and Finis Welch, "No Time to be Young: The Economic Prospects for Large Cohorts in the United States," *Population and Development Review,* 7, March 1981, 71–83.

Smith, James P., and Michael P. Ward, "Time Series Changes in the Female Labor Force," *Journal of Labor Economics*, forthcoming.

U.S. Bureau of the Census, *Historical Statistics of the United States Colonial Times to 1957*, Washington, D.C., 1960.

U.S. Bureau of the Census, *Projections of the Population of the United States: 1982 to 2050 (Advance Report)*, Washington, D.C.

U.S. Department of Labor, Bureau of Labor Statistics, *Handbook of Labor Statistics*, Washington, D.C., 1980.

Wachter, Michael L., and William L. Washer, "Leveling the Peaks and Troughs in the Demographic Cycle: An Application to School Enrollment Rates," unpublished notes.

Welch, Finis, "Effects of Cohort Size on Earnings: The Baby Boom Babies' Financial Bust," *Journal of Political Economy*, 87, October 1979, S65-S97.

Welch, Finis, and W. Gould, "An Experience Imputation or an Imputation Experience?" The Rand Corporation, unpublished draft.

SECURITY CLASSIFICATION OF THIS PAGE (When Data Entered)

REPORT DOCUMENTATION PAGE		READ INSTRUCTIONS BEFORE COMPLETING FORM
1. REPORT NUMBER R-3115-ARMY	2. GOVT ACCESSION NO.	3. RECIPIENT'S CATALOG NUMBER
4. TITLE (and Subtitle) Forecasting the Wages of Young Men The Effects of Cohort Size		5. TYPE OF REPORT & PERIOD COVERED Interim
		6. PERFORMING ORG. REPORT NUMBER
7. AUTHOR(s) Hong Tan, Michael Ward		8. CONTRACT OR GRANT NUMBER(s) MDA903-80-C-0652
9. PERFORMING ORGANIZATION NAME AND ADDRESS The Rand Corporation 1700 Main Street Santa Monica, CA 90406		10. PROGRAM ELEMENT, PROJECT, TASK AREA & WORK UNIT NUMBERS
11. CONTROLLING OFFICE NAME AND ADDRESS U.S. Army Research Inst. for the Behavioral & Social Sciences 5001 Eisenhower Ave., Alexandria, VA 22333-5600		12. REPORT DATE May 1985
		13. NUMBER OF PAGES 58
14. MONITORING AGENCY NAME & ADDRESS (if different from Controlling Office)		15. SECURITY CLASS. (of this report) Unclassified
		15a. DECLASSIFICATION/DOWNGRADING SCHEDULE

16. DISTRIBUTION STATEMENT (of this Report)

Approved for Public Release; Distribution Unlimited

17. DISTRIBUTION STATEMENT (of the abstract entered in Block 20, if different from Report)

No Restrictions

18. SUPPLEMENTARY NOTES

19. KEY WORDS (Continue on reverse side if necessary and identify by block number)

Salaries
Income
Military Personnel

20. ABSTRACT (Continue on reverse side if necessary and identify by block number)

see reverse side

DD FORM 1 JAN 73 1473 EDITION OF 1 NOV 65 IS OBSOLETE

SECURITY CLASSIFICATION OF THIS PAGE (When Data Entered)

SECURITY CLASSIFICATION OF THIS PAGE(When Data Entered)

In this study, the authors develop

forecasts of the civilian wage structure over the next two decades for a variety of different scenarios. They focus on how the wage structure will change as the demographic trend reverses itself, i.e., as the smaller post-baby-boom birth cohorts enter the labor market in the 1980s and 1990s. Section II of the report describes the survey data used to create a working file for the analysis. Based on this file, the authors paint a broad overview of how cohort size and relative wages have changed over the 1967-1980 period. Section III discusses the wage model used and highlights the main empirical results. The assumptions and approach used to forecast wages are detailed in Section IV. Section V extends the wage model to investigate two alternative explanations for the observed decline in youth wages. The last section concludes with a summary of the main findings and their implications for military compensation policy. Additional Keywords: Salaries, Income, and Military personnel.

SECURITY CLASSIFICATION OF THIS PAGE(When Data Entered)

RAND/R-3115-ARMY

ATE

LMED

-8

DD FORM 1 JAN 73 1473 EDITION