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Econometric Models of Armed Forces Enlistment Levels

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Econometric Models of Armed Forces Enlistment Levels

by

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This study has had two primary objectives. First, to conduct a detailed review of the major, past, time-series analyses which covered the 1958-1965 time frame, attempting to discern differences among them in methodology, time intervals, and in the data used. This review was followed by an attempt to reconcile some of these differences and to reproduce these analyses using a common data base and a more nearly standardized set of variable specifications. (over)

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PRICES SUBJECT TO CHANGE

The second objective was to attempt a pooled, time series cross-sectional analysis of enlistments. This was accomplished for certain preferred enlistment groups for the Army and Navy using nine census regions and five yearly time intervals (1970-1974). The results of this model were to be compared to prior modeling efforts.

Finally, the significance of these studies was reviewed in the context of the enlistment environment of the mid-1970's and forward into the 1980's.

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EXECUTIVE SUMMARY

by

Daniel F. Huck and Kenneth D. Midlam

BACKGROUND

In recent years several studies have been undertaken to estimate the relationships between certain key economic and policy variables and the supply of enlistees to the armed forces. The impetus for most of the early studies came from the need to determine the feasibility of an all-volunteer force and the subsequent need to know cost-effective ways to obtain the enlistment levels necessary to sustain such an all-volunteer force.

Of primary interest in these studies were the effects on the voluntary enlistment supply of the following variables:

- Compensation - expressed either in absolute terms for civilian pay and military pay (including in kind as well as subsistence and quarters allowances), or as the ratio of military to civilian pay for the age groups which constitute the apparent preferred supply groups.
- Unemployment Rate - usually estimated specifically for the 17-21 year old age group either by conversion from overall unemployment rates or by direct measurement. Occasionally the overall unemployment rate is used.
- Draft Pressure - expressed as the fraction of enlistees motivated to enlist by an apparent likelihood of being drafted. This had been estimated or scaled in various ways by the several studies.

- Recruiting Resources - usually measured in terms of the number of active duty recruiters.

Though these definitions appear reasonably consistent here, a major potential source of variation in the results of past studies has been the different ways these variables have actually been measured or estimated.

The kinds of analyses that have been attempted fall into two groups — time-series and cross-sectional. Typically, the time-series analyses have measured variables at the gross, national level over time. The time intervals have usually been quarterly, though some recent studies have used monthly data. The cross-sectional studies have measured the variables over a single, fixed time interval for various geographic breakdowns (usually, census regions; sometimes, states).

OBJECTIVES OF THIS STUDY

This study has had two primary objectives. First, to conduct a detailed review of the major, past, time-series analyses which covered the 1958-1965 time frame, attempting to discern differences among them in methodology, time intervals, and in the data used. This review was followed by an attempt to reconcile some of these differences and to reproduce these analyses using a common data base and a more nearly standardized set of variable specifications.

The second objective was to attempt a pooled, time series cross-sectional analysis of enlistments. This was accomplished for certain preferred enlistment groups for the Army and Navy using nine census regions and five yearly time intervals (1970-1974). The results of this model were to be compared to prior modeling efforts.

Finally, the significance of these studies was reviewed in the context of the enlistment environment of the mid-1970's and forward into the 1980's.

REVIEW OF PREVIOUS TIME-SERIES STUDIES

Many of the studies reviewed were done as part of feasibility studies on moving to an all-volunteer force. Particular attention in this review was given to the analyses of A. C. Fisher,[1] A. E. Fechter,[2] B. J. Klotz,[3] and D. W. Grissmer, et al.[4] The first three studies attempted time series analyses on quarterly data covering time intervals from 1958 to 1965 (basically prior to the major Vietnam build-up). In general, each used enlistments as a fraction of the Qualified Military Availables (QMAs) as a dependent variable and used various functional formulations of compensation, unemployment and draft pressure as independent variables. Fisher and Klotz dealt with total DOD enlistments; Fechter, with Army enlistments. All used Mental Group I-III enlistees without further disaggregation by such factors as race and education. The fourth study analyzed all four services and broke enlistments down further by education. That study covered 1971-1973.

The results of the various studies have yielded a wide variety of estimates of the relationships between these factors and the supply of enlistees. Generally speaking, they have found a statistically significant relationship between levels of relative compensation and enlistment levels and a large but statistically doubtful relationship between unemployment and the enlistment supply. Those studies which included recruiting variables found large, significant relationships between recruiting levels and enlistment supply. Estimates of the fraction of enlistees draft-induced ranged from 15-40% in the early 1960's.

Because of the gross level of aggregation of enlistments over mental groups, education levels and services, the first three studies, though important to the development of econometric analyses of enlistment behavior, are not nearly as relevant under the current (1976) enlistment environment. Better, more disaggregated data on more recent enlistments has led to the clear conclusion (Grissmer and other recent studies) that both pay and unemployment effects are inversely related to the quality level of the enlistee, i.e., the preferred enlistment group (Mental Group I-III, high school graduates) is the least responsive to changes in relative military pay and to civilian employment opportunities. Also, this preferred group appears more responsive to the activities of recruiters in the field and to the availability of program options.

Based on the Grissmer analyses, the relative pay (ratio of military to civilian compensation) elasticity for Army Mental Group I-III high school graduates is about .6, and for the Navy, about .45. The elasticity is defined as the rate of change in a dependent variable in response to a change in an independent variable. As an example, an elasticity of .45 for relative pay with respect to Navy enlistments means that a 10 percent increase in relative pay should result in a 4.5 percent increase in Navy Mental Group I-III diploma enlistments. The employment rate elasticity for the Army is estimated to be about 3.7 (about .5 for unemployment). No significant unemployment relationship could be established for Navy enlistments.

REVIEW OF PREVIOUS CROSS-SECTIONAL STUDIES

Three major cross-sectional studies are reviewed here.[4,5,6] These studies were all based on limited data from the early 1970's. As with the time-series studies, serious inadequacies are apparent in those studies which did not disaggregate enlistments on education and mental quality. In general, they report elasticities for pay and employment which, for the gross aggregations of enlistments, are larger than found in Ref 4 using only Mental Group I-III, high school graduates. This latter cross-sectional study, using 1973 data, estimates elasticities for relative pay and unemployment of .68 and .70, respectively, for the Army's preferred enlistment group. As with all the time-series studies reviewed, statistical significance was almost always stronger for relative pay than for the unemployment variable.

RESULTS OF POOLED TIME-SERIES/CROSS-SECTIONAL ANALYSIS

A pooled time-series, cross-sectional analysis was conducted by GRC to attempt to capitalize on the best features of the separate analyses described earlier. By combining the time-series and the cross-sectional data, a significant increase in observations was obtained. This frequently yielded an increase in variability in some of the independent variables which may have been quite "flat" in a time-series or single cross-sectional set of observations. In this study, data were obtained on an annual basis for 1970-1974 enlistments, QMAs, civilian pay, youth unemployment, recruiting, paid advertising and the black proportion of the MA for each of the nine census regions for Army and Navy enlistees in the preferred quality groups.

The results of these analyses differ strikingly from those discussed earlier. In no case was a statistically significant relative pay or employment effect obtained. In the Army analysis only two factors are statistically significant — regional differences and recruiters per QMA. In other words, regional variations in either propensity to enlist or in disqualification rates are much more significant than variation in pay or unemployment rates in measuring enlistment rates for these quality groups. Estimates of the recruiter elasticity range from .1 to .9 for Mental Group I-III high school graduates. Pay elasticities, though never statistically significant, range from .10 to .18. Unemployment elasticities were usually in the wrong direction, but not statistically significant.

In the analysis of Navy enlistments, the only statistically significant factor was recruiters per QMA, with elasticities in the range of .60 - 1.0 for the preferred enlistment groups. Relative pay and unemployment effects were consistently in the wrong direction and not significant. Regional differences were relatively minor.

SUMMARY OF FINDINGS

- Based on the previous studies reviewed, increases in the level of military pay should cause some positive response in the Army and Navy's supply of preferred enlistees. This response most probably has an elasticity less than .5. The pooled time-series/cross-sectional model did not find any statistically supportable relationship between pay and enlistment rates as defined in this study for either Army or Navy. The appearance of the recruiter variables in supply models has the effect of significantly lessening the impact of relative pay measurements from those obtained in other models which excluded the recruiting and advertising variables. It seems probable that the pay and recruiting programs are supportive rather than simply additive; increases in the recruiting program have effects which are amplified by an improved military compensation system.

- The increases in recruiter levels and distribution in recent years appear to have had a significant impact on the level of enlistees. The observed measurements do not, however, preclude the possibility of a direct tracking of the enlistment supply by the services. That is, it remains to be determined how much this is a reflection of continuing improvements in the geographical assignment of recruiters, or how much of this is due to the ability of recruiters to foster a more favorable general propensity to enlist.

- Few of the studies discussed in this report were able to demonstrate any statistically significant relationship between unemployment levels and enlistment rates. The most reliable estimates show the higher quality enlistee groups being responsive to changes in employment levels; but these estimates are not always obtained from stable parameters. Since it is generally assumed that there must be some relationship, it must be concluded that the specification of the unemployment variables has been inadequate. Several difficulties with measuring unemployment effects have been discussed by the various authors. A basic problem is simply that observed unemployment rates may be poor estimators of an individual's perceived attractiveness of his short-term, future, civilian employment opportunities. A second problem arises in measuring the effects of unemployment on volunteer enlistment rates when there is a sizeable variation in accession levels over time. High levels of public sector employment are bound to affect employment conditions in the private sector. This was particularly true during the Vietnam era. Tracking the effects of unemployment on volunteer enlistments during this period can lead to spurious results; it is better to obtain measurements of unemployment effects during periods of stable accession rates with varying unemployment rates. The 1972-76 time period is probably best for this kind of measurement since accession levels were moving into a steady-state period and unemployment was rising with the onset of the economic recession. Possibly studies that incorporate more recent enlistment data will provide more reliable results.

- Sociodemographic characteristics of potential enlistees are significant determinants of enlistment levels and response to economic and policy variables for the Army and Navy. Econometric models used in forecasting supply levels should disaggregate the supply by (at least) the racial and educational characteristics of the population.

- A pooled time-series, cross-sectional model can be useful in providing estimates of supply parameters across census regions. The analyses of Army accessions show clearly that regional differences in accession levels are more clearly defined and persistent than the policy variables included in the analysis.

- These regional variations can represent several factors affecting the supply enlistees, including differences in higher education enrollment rates (which to some extent may also reflect differences in racial distributions, as evidenced by the fact the pooled time-series, cross-sectional analysis found the black percentage of the QMA a significant variable only when basic regional variations were not included explicitly as independent variables).

RELEVANCE OF FINDINGS TO CURRENT POLICY-MAKING ENVIRONMENT

The early econometric analyses of accessions were motivated by the need to assess the feasibility of maintaining an all-volunteer force. Primary attention was given to the question of the effects of military compensation and civilian employment levels on accessions and the estimation of the proportion of recent enlistees who had been draft-motivated.

These feasibility studies gave little attention to questions of the quality of the enlistees and seldom attempted to separate supply effects for the individual services. They did, in general, demonstrate that the all-volunteer force was at least numerically feasible and that achieving a more attractive balance between military and civilian compensation could significantly enhance the likelihood of success of an all-volunteer force.

Generally speaking, the all-volunteer force has now been achieved with increasingly favorable distributions of enlistees by education and mental quality. Since it has been clearly shown that enlistments necessary to sustain a force of the current size can be achieved, current attention is focused more on raising the overall quality of accessions and on determining the most nearly optimal way of obtaining the desired higher quality enlistees. Recent studies (e.g., Refs 7 and 8) have demonstrated that even with an assumption of a large relative pay elasticity, the cost-effectiveness of a military pay increase as a tool for increasing accessions is very unfavorable when compared to alternative ways of obtaining enlistees such as paying quality and/or skill specific enlistment bonuses or increasing resource allocations to the recruiting and advertising programs. These recent studies have clearly shown the attractiveness of selective adjustments in military compensation when compared to the very costly across-the-board increase. Table S1 compares the cost-effectiveness of several programs for increasing high school graduate, Mental Group I-III enlistments based on results in Refs 4 and 8.

The major reason for the impracticality of increasing basic pay to attract more volunteers is the compensation policies that are currently in effect. The very large marginal cost for basic pay is in large part

Table S1

MARGINAL COST PER ENLISTMENT

MENTAL GROUP I-III, HIGH SCHOOL GRADUATE

Program	Cost of One Additional Enlistee	
	Army	Navy
Increase relative pay (elasticity — Army .6 Navy .45	$104,700 [a]	$146,600[a]
Increase recruiters	4,000	2,300
Increase paid advertising	7,000	5,200
Offer $1,500 combat arms bonus for four-year enlistment	3,900	---

[a] Assumes the increase is in enlisted pay only.

a reflection of the need to inflate the entire pay structure to maintain the integrity of the compensation system. Thus, if a pay raise were given to new entrants, all remaining enlisted members would be entitled to a similar percentage increase if differential salaries among pay grades were to be preserved.

Because the pay elasticities differ for each service, a free market approach to compensation would yield radically different pay scales. Under this approach, basic pay would be adjusted to bring supply and demand for new accessions into rough balance for all services. Obviously this represents a significant change in military compensation policy by DOD and we are not suggesting that such a practice be adopted. We raise the point only to demonstrate that manpower policies that are optimal in an economic sense are often unacceptable by other criteria — social equity being one.

Whatever the time frame considered, the effect of variations in the unemployment rate continues to be an important issue and investigations to refine estimates of the relationship between unemployment rates and enlistments are clearly indicated. Recent investigations [7,9] are shedding new light on this relationship and have produced results which have been more statistically significant than the results of analyses reviewed in this study. These recent studies also show more clearly the substitution of higher quality enlistees for lower quality in the presence of high unemployment rates.

Inasmuch as the size of the QMA will be declining in future years, a significant decline in the unemployment rate could generate a serious shortfall of high quality enlistees in those services with the least basic attractiveness to the preferred quality group.

RECOMMENDATIONS FOR FUTURE ANALYSIS

Continuing studies of enlistment levels in the Armed Forces should concentrate on the issues which currently occupy the attentions of policy makers and program planners. Such studies should concentrate on the following:

- Understanding more thoroughly the effects of changes in unemployment rates on enlistment levels of the various quality groups.
- Projecting future enlistment levels with primary attention given to current policy options.
- Evaluating the extent to which the significant effects of recruiting on accessions are simply reflections of improved positioning of recruiting resources by the services. This can probably not be done using the methodologies which are

the subject of this study. It will require a more micro-analytic cross-sectional study of the kind used to assess the effectiveness of paid advertising campaigns. It may, in fact, require planned experimental variation in regional recruiter assignments rather than simply an analysis of available historical data.

- Refining the QMA and other key policy variables by explicitly considering race, education and mental quality.
- Determining the effect renewed emphasis on recruiting for the Reserves is going to have on active duty accessions. Given the appropriate data, this issue could be studied using many of the basic techniques discussed in this study.
- Making a more detailed assessment of inter-service competition for quality enlistees.
- Continuing to compare, in a more nearly uniform way, the cost-effectiveness of enlistment policy options.
- Continuing to recognize that responsiveness to economic as well as policy variables differs widely among various quality and socioeconomic subgroups of the total potential supply of enlistees.

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Chapter I
INTRODUCTION
by
Dorothy Amey

The several econometric analyses of the supply of volunteer enlistments to the Armed Forces which have surfaced over the past eight years have reported a broad range of estimates for supply parameters. Most notable, the estimates of military and civilian pay, and unemployment elasticities have been as diverse as the methodologies employed by the different studies. Consequently, the decision maker analyst has been presented with a confusing choice of options in deciding the best model specification to use for policy making. The main accomplishment of this study has been to review, explain and reconcile in many ways the breadth of findings resulting from the major studies or enlistment supply that were performed in recent years.

This report consists of an extensive review of most of the published studies on enlistment supply. A brief description of the types of econometric models and methods employed to estimate supply parameters is presented in Chapter II. The basic assumptions that are made in studies on the supply of enlistees to the services are explained and examined in Chapter II. In Chapter III, the different analyses reviewed are listed and presented with the major findings from each. Some comparisons of variable specifications, model structure and functional form, and methodology are made in Chapter III. In addition, the detailing of the differences found among the findings of these studies is set forth in Chapter III, thereby providing the reader with the major discrepancies which have so concerned decision makers.

In an effort to reconcile all of the major differences among the findings of the studies reviewed, a reworking of the analyses performed by Fisher and Klotz is reported in Appendix C. The findings from these analyses based on a uniform data base are compared with the results of other studies which are reviewed but not duplicated here.

Perhaps the most disputed difference in the methodology of the various econometric analyses of enlistment supply has been the choice of time series models versus cross sectional models. In Chapter IV, results of an analysis of the supply of volunteer enlistments using a pooled time series cross sectional model are reported. The analysis was made in an effort to determine measurements of variable impact which might bridge the gap between disputed findings of studies which differed mainly in the structural form of the model.

Interpretations of the findings of this study and the conclusions made from our survey of econometric analyses of enlistment supply are included in the chapters and in the summary of this report.

Chapter II

REVIEW OF LITERATURE: SOME METHODOLOGICAL ISSUES

by

Alan Fechter

Many of the enlistment supply studies reviewed here were generated as part of the two major governmental studies of the feasibility of moving to an all-volunteer armed force.[1] These supply studies concentrated their efforts on the derivation of estimates of two important factors required to calculate the incremental cost to the government of moving to an all-volunteer military; the supply of enlistments in the absence of a draft and the responsiveness of enlistments to changes in military pay.[2] The move to an all-volunteer armed force in January 1973 has made moot the issue of draft-motivated enlistments. In response to new policy issues, more recent studies, using estimates of voluntary enlistments, have concentrated their attention on deriving estimates of the impact of other enlistment determinants, such as recruiting effort and advertising, and on the socio-demographic distribution of enlistments.[3]

THE THEORY OF ENLISTMENT SUPPLY

The studies of enlistment supply are based on an economic theory of occupational choice in which individuals are assumed to pick their occupations so as to maximize their utility, subject to the constraint that they can only engage in one occupation at a time and that they can only devote some fraction of their time to working in that occupation. One of the earliest expositions of this theory applied to enlistment behavior is the work of Fisher.[4] Since most of the literature reviewed is based more or less on a model like the one developed by Fisher, general aspects of his model are summarized below.

Fisher assumes that a potential enlistee is faced with two options: to enlist or not to enlist. He defines the returns associated with these two options as follows: Let V_M be the present value of the pecuniary and nonpecuniary returns to enlisting; let V_C be the present value of the pecuniary and nonpecuniary returns to not enlisting. V_M and V_C can be written as follows:

$$V_M = \sum_{t=1}^{n} \frac{W_{Mt}}{(1 + i)^t}$$

$$V_C = \sum_{t=1}^{n} \frac{W_{Ct}}{(1 + i)^t}$$

where W_{Mt} is the expected pecuniary and nonpecuniary return to enlisting for any given year t, W_{Ct} is the expected pecuniary and nonpecuniary return to not enlisting in any given year t, and i is the subjective rate of discount. He further refines his definition of V_M to account for the finite duration of the first enlistment decision, m.

$$V_M = \sum_{t=1}^{m} \frac{W_{Mt}}{(1 + i)^t} + \sum_{t=m+1}^{n} \frac{W_{MCt}}{(1 + i)^t}$$

where W_{MCt} is the expected post-military returns in year t to an enlistee after completing his first term of enlistment. He assumes that W_{MC} is roughly equal to W_C based on a comparison of earnings of veterans and nonveterans made by Gilman.[5] This allows him to conclude that the relevant comparison for potential enlistees is between

$$\sum_{t=1}^{m} \frac{W_{Mt}}{(1 + i)^t} \quad \text{and} \quad \sum_{t=1}^{m} \frac{W_{Ct}}{(1 + i)^t}$$

He further assumes that the time profiles of these two streams of return are roughly similar. This allows him to ignore discounting and to conclude that the appropriate comparison for potential enlistees is between

$$\sum_{t=1}^{m} W_{Mt} = W_M \quad \text{and} \quad \sum_{t=1}^{m} W_{Ct} = W_C .$$ [6]

Given Fisher's assumptions, a potential enlistee will prefer enlisting over not enlisting if $W_M > W_C$.

In order to drive empirically testable implications from this model, Fisher shifts his attention to the pecuniary returns to enlisting and not enlisting, W_{MP} and W_{CP}. He introduces a variable d, which represents the net nonpecuniary disadvantage to enlisting. Thus, if $W_{MP} > W_{CP}$ (1+d) a potential enlistee will prefer to enlist. He assumes that the distribution of d among potential enlistees is lognormal and that it is stable over time. This allows him to derive an enlistment function of the following form:[7]

$$(1) \quad \frac{E}{P} = f(\ln W_{MP}, \ln W_{CP})$$

He further assumes that equation (1) is linear and that the enlistment response to a one percent change in W_{MP} is equal in magnitude, but opposite in sign compared to the enlistment response to a one percent change in W_{CP}. This allows him to derive the following explicit enlistment equation:

$$(2) \quad \frac{E}{P} = f\left(\ln \frac{W_{MP}}{W_{CP}}\right) = \alpha' + \beta_1' \ln \frac{W_{MP}}{W_{CP}} + \varepsilon$$

Fisher further assumes that W_{CP} is full-time earnings and makes the following modification for unemployment. He adjusts W_{CP} for unemployment using the following identity:

$$(3) \qquad W_{CP*} = p_e W_{CP} + p_u W_{CPU}$$

where W_{CP*} is full-time earnings, adjusted for the probability of not being employed, p_e is the probability of being employed, p_u is the probability of not being employed, and W_{CPU} is the earnings received while unemployed.

If $W_{CPU} = 0$, then $W_{CP*} = p_e W_{CP}$. And equation (2) can be rewritten:

$$(4) \qquad \frac{E}{P} = \alpha' + \beta_1' \ \ln \left(\frac{W_{MP}}{W_{CP*}} \right) + \varepsilon$$

Equation (4) describes the enlistment function in a no-draft environment. To account for the presence of the draft, Fisher modifies his definition of the pecuniary returns to not enlisting as follows:

$$(5) \qquad W_{CP'} = p_c \ W_{CP*} + p_d \ W_{MP}$$

where $W_{CP'}$ is the expected pecuniary return to not enlisting, p_c is the probability of remaining a civilian, and p_d is the probability of not remaining a civilian.[8] From equation (3) (assuming W_{CPU} is zero) and equation (5), equation (2) can be rewritten to account for unemployment and the draft as follows:

$$(6) \qquad \frac{E}{P} = \alpha' + \beta_1' \ \ln \left(\frac{W_{MP}}{W_{CP'}} \right) + \varepsilon$$

$$= \alpha' + \beta' \ \ln \left(\frac{W_{MP}}{p_e p_c W_{CP*} + p_d W_{MP}} \right) + \varepsilon$$

Further manipulation of (6) produces:[9]

$$(7) \quad \frac{E}{P} = \alpha' + \beta' \left(\ln p_c + \ln p_e + \ln \frac{W_{CP}}{W_{MP}} \right) + \varepsilon$$

Fisher further assumes the following:

$$(8) \quad p_c = \delta_1 \, (1-A/P)^{\delta_2}$$

$$(9) \quad p_e = \gamma_1 \, (1 - U)^{\gamma_2}$$

where A/P is the observed military accession rate, defined to include enlistments and inductions, and U is the observed unemployment rate. This allows him to rewrite (7) as follows:

$$(10) \quad \frac{E}{P} = \alpha + \beta_1 \ln \frac{W_{CP}}{W_{MP}} + \beta_2 \ln (1 - U) + \beta_3 \ln \left(1 - \frac{A}{P}\right) + \varepsilon$$

where:

$$\alpha = \alpha' + \delta_1 + \gamma_1$$

$$\beta_1 = \beta_1'$$

$$\beta_2 = \beta_1' \gamma_2$$

$$\beta_3 = \beta_1' \delta_2$$

The relative pay elasticity, η_p, can be estimated from equation (10):

$$\eta_p = \beta_1 \cdot \frac{1}{\frac{E}{P}}$$

The elasticity of expectations with respect to changes in probabilities of being employed (γ_2) and remaining a civilian (δ_2) can also be computed from equation (10):

$$\gamma_2 = \beta_2 \div \beta_1$$

$$\delta_2 = \beta_3 \div \beta_1$$

The effect on enlistments of eliminating the draft can be derived from

$$\frac{\beta_3}{1 + \beta_3}.^{10}$$

Practically all enlistment studies have included as arguments of the enlistment function variables indexing military and civilian pay, employment conditions in civilian labor markets, and the likelihood of being drafted. Moreover, many of these studies have also included arguments that index factors affecting (1 + d). Recall that Fisher assumed d was stable over time. Thus, the general enlistment function estimated by econometric models can be described as:

$$(11) \quad E = e(M, C, U, D, P, X)$$

where:

E = enlistments

M = first term military pay

C = alternative first-term civilian pay

U = unemployment rates

D = draft probability

P = population of eligible enlistees

X = other enlistment determinants.

The testable hypotheses are:

$$\frac{\delta E}{\delta M} > 0, \qquad \frac{\delta E}{\delta D} > 0$$

$$\frac{\delta E}{\delta C} < 0, \qquad \frac{\delta E}{\delta P} > 0$$

$$\frac{\delta E}{\delta U} > 0, \qquad \frac{\delta E}{\delta X} \gtrless 0$$

SPECIFICATION OF AN ESTIMATING EQUATION

A number of decisions must be made in moving from (11) to an estimating equation. First, an explicit functional form must be chosen; then methods for estimation of the variables must be decided on; finally, an appropriate estimation technique must be employed.

Choosing a Functional Form

The theory of enlistment supply developed by Fisher suggests that the enlistment have a nonlinear form. Fisher selects a semilogarithmic function based on an assumed lognormal distribution of W_{CP} and $(1 + d)$. Other studies have not constrained themselves to this particular functional form. By selecting alternative functional forms, these studies imply alternative, but unspecified, assumptions about these distributions.

Specifying the Dependent Variable

Most other studies, following Fisher's lead, assume that the enlistment function is homogeneous of degree one in P and use an enlistment rate as their dependent variable.

Specifying the Effect of Pay

Fisher assumes that equiproportionate changes in M and C will generate equal, but opposite in sign, responses in enlistments. This allows him to specify his enlistment function in terms of the ratio, M/C. This particular specification of enlistment behavior was convenient for many of the studies reviewed here because they lacked sufficient variation in M to derive a statistically reliable estimate of the enlistment response to changes in M. The problem was particularly acute in the cross section studies, in which M was assumed to be a constant that was independent of region. The relative pay specification was also convenient to the early time series studies because of the limited amount of variation in military pay observed during the period of their analyses.[11] This data limitation necessitated reliance on the variation in C for estimating the effect of military pay on enlistments. However, there are reasons for questioning the validity of this assumption.

Fechter argues that the pecuniary return to not enlisting is not fully captured by C. He suggests that the expected pecuniary return to remaining in school is a particularly relevant return to not enlisting and that failure to account for it could conceivably bias estimates of military pay elasticity derived only from variations in C.[12]

These arguments can be used to defend an alternative specification of the effect on enlistments of M and C that does not rest on the assumption of symmetrical enlistment response to equiproportionate changes in M and C. An alternative specification would describe the enlistment function in terms of the levels of both M and C. Two studies, those of Fechter and Withers, experiment with enlistment functions specified in terms of absolute pay.

An additional argument in opposition to the assumption can be based on an assumption that nonpecuniary factors, such as working conditions, are superior goods.[13] The military can be characterized as comparatively risky and regimented. There is evidence that, on the average, workers want to be compensated for such risks as the likelihood of injury or death.[14] On this account alone, the expected nonpecuniary return to enlisting might be lower than the expected nonpecuniary return to not enlisting. Thus, if non-pecuniary returns are assumed to be superior goods and if the military can be characterized as offering relatively small amounts of nonpecuniary returns, one could expect enlistment rates to decline with equiproportionate changes in M and C even when C is correctly specified. This would imply that the enlistment response to changes in C would be greater in absolute terms than the enlistment response to equiproportionate changes in M.

METHODS OF ESTIMATING VARIABLES

The studies of enlistment supply differ most dramatically in their methods of estimating the arguments of equation (11). Since these differences could be a major reason for variations in findings summarized above, they are discussed in some detail.

Enlistments

Enlistments were assumed to be demand determined during the period of the draft. In general, most of the studies followed Fisher in assuming that enlistments from the lowest two of the five mental categories, category 4 and category 5, were demand determined.[15] Some studies limited their estimate of enlistments to mental categories 1-3;[16] others used enlistments classified by mental category and level of school completed (high school graduate vs. non high school graduate);[17] still others used all enlistees and controlled for quotas by including a screening variable in their enlistment function.[18]

Eligible Population

The population of eligibles (P) can, in principle, be identified as the product of the total population (TP) and the fraction of the total population who are able to meet the mental and physical standards for enlistment (q). The time series studies, following Fisher, assume a stable q and estimate P from the civilian noninstitutional male population age 17-20. (A notable exception is the time series study by Kim, et al., in which P is estimated as the number of Selective Service registrants between the ages of 19 and 26 who are classified as 1-A.) The cross section studies generally adjust their estimates of TP by an estimate of q derived from the results of preinduction examinations or examinations of enlistment applicants. A problem with this procedure is that potential inductees or enlistment applicants are not a representative sample of TP because they exclude those who have been prescreened before their exams and those who are able to obtain deferments. One cross section study, by Altman, further restricts its estimate of P to full-time members of the civilian labor force.

Time Horizons, Discount Rates, and Expectations Functions for Pay Variables

Estimates of M and C require assumptions about the time horizon of the potential enlistee. All studies follow Fisher in assuming that the time horizon is the term of the initial enlistment contract. With the exception of Cook, who uses a 4 year time horizon, the studies uniformly assume a 3 year time horizon. Recall that Fisher is able to limit the time horizon to this period by assuming that the present value of post-military returns to enlisting is equal to the present value of post-military returns to not enlisting.

His evidence in support of this assumption is an unpublished study by Gilman which compares earnings of veterans and nonveterans. This comparison was made for comparable males standardized by age and education. It is possible that there are systematic differences between enlistees and nonenlistees in the amount of school they complete. Some nonenlistees may choose not to enlist in order to remain in school because of the perceived higher present value of post-school returns relative to the value of comparable returns to enlisting. Other nonenlistees may remain in school as a means of avoiding the draft. Both sets of behavior on the part of nonenlistees would be consistent with a higher amount of school completed by nonenlistees -- and a resultant greater post-military return to not enlisting.

Offsetting this tendency however, many enlistees take advantage of the provisions of the G.I. Bill and return to school upon completion of their military service.

It is not clear, a priori, which of these tendencies would be dominant; however, both tendencies operate on Fisher's assumption of equality in post-military returns. Therefore, factors affecting these tendencies ought to be considered in evaluating this formulation of M and C. In particular, policy changes with respect to the G.I. Bill and evidence related to the returns to additional schooling would be relevant, particularly to the time-series studies.

An additional factor to consider in estimating M and C is the appropriate discount rate to use. Studies have varied in their choice of discount rates to apply to first term pay. Some studies have eschewed discount rates for simple averages;[19] others have used rates of 20 and 30 per cent.[20] A final factor relevant to estimation of both M and C is the nature of the expectations function used to form the variables. Most studies-particularly the cross-section studies-use a static formulation in which expected future values of M and C are based solely on the current value. Some of the time-series studies use linear interpolation of annual values of M and C to derive quarterly values.[21] Still other time-series studies utilize a more dynamic procedure in which current trends are extrapolated to produce estimates of M and C beyond the initial year of enlistment.[22]

Military Pay

Methods of estimating M have differed mainly in the components of pay included. A unique feature of military pay is the substantial amount of pay that is provided in-kind in the form of quarters, subsistance, and medical services. Roughly half of military pay was provided in the form of these in-kind benefits in the early 1960's. This fraction has fallen as a result of the pay increases instituted to move to an all-volunteer force; but it is still considerably larger than it is in the civilian sector. Studies have varied considerably in their assessment of these in-kind benefits. Most studies have valued these benefits according to the allowances awarded for them by the military for enlistees who must purchase them on the open market.[23] Other studies, notably the Cook study, completely excluded them from their estimates of M.[24] The cross-section studies form a unique sub-set by assuming that their estimates of M, however defined, are constant over regions. This means their estimates of military pay elasticity must be derived from variations in estimates of C. The analytic problems and potential biases associated with the method of estimation are discussed below.

Estimating Civilian Pay: The Risk of Unemployment

Estimation of C involves a slightly different set of analytic issues. First, while enlisting assures employment in the military, not enlisting involves the risk of enduring some periods of unemployment. One must therefore determine a method for accounting for the effects of these periods of unemployment on C. Recall that Fisher discounted an estimate of full-time earnings by the complement of the unemployment rate (1-U) to account for unemployment. An alternative method, used by Hause, adjusts full-time earnings according to the duration of unemployment (DU).[25]

These methods of accounting for unemployment assume that the risk of unemployment operates on enlistments primarily through its effects on C. Under these conditions, C can be measured as the product of full-time earnings, C*, and the duration of employment, proxied by (1-U) or (1-DU). Some have argued that the risk of unemployment also operates on enlistments through its effects on d. Assuming risk-aversion on the part of potential enlistees, variations in the risk of unemployment can be expected to be inversely related to d; i.e., the higher the risk of unemployment, the smaller the net nonpecuniary disadvantage to enlisting. Under these conditions, the enlistment response to changes in the risk of unemployment that are equivalent to a given change in C will be greater than the enlistment response to an equivalent change in C*.

The argument that the risk of unemployment is a determinant of the net nonpecuniary disadvantage to enlisting can be used to justify specifying an enlistment function in which C* and (1-U) are included as separate arguments. Recall that Fisher's estimating equation is so specified--only Fisher's justification is based on differences between measured and expected values of his variable, p_e. which he proxies by (1-U). Both assumptions, risk aversion and differences in expectations, lead to formulation of an enlistment function like (11) in which U appears as an explicit argument. However, the risk aversion assumption leads to an expectation that:

$$\frac{\delta E}{\delta(1-U)} < \frac{\delta E}{\delta C^*}$$

Most enlistment studies include U or 1-U as an explicit independent argument in their enlistment functions. Most studies use global unemployment rates for teenagers to index U; some studies have used unemployment rates of youth whose major activity was other than school.[26]

Civilian Pay: Full-Time Earnings

Most studies have followed Fisher in estimating C* from the incomes of age-specific year round full-time workers. Some studies base their estimates on C* on manufacturing payroll data. Some use manufacturing payrolls bench-marked to age-specific statistics. All estimates are biased in that they include the earnings experience of potential enlistees in mental category 5 and in other excluded mental categories. Consequently, the estimated enlistment response to a given change in measured C* may be a biased estimate of the enlistment response to a given change in the mental-group-specific C*. In addition, the studies which base their estimate of C* on manufacturing payrolls are further biased by their exclusion of potential enlistees in non-manufacturing industries. These biases may be important in evaluating the validity of estimates of military pay elasticities generated from variations in civilian earnings only. They will be particularly relevant for the cross section studies; they will also be important in both time series and cross section to understanding differences in pay elasticities generated for different enlistment groups from estimates of relative military pay (M/C*).

Draft Probability

There are a number of methodological issues associated with estimating this probability. First, the rules by which Selective Service is administered would lead one to expect little or no regional variation in D. Thus, some cross section studies were constrained to estimates of the enlistment effect of the draft that are derived from survey data. Enlistees who indicated on the survey that they would not have enlisted in the absence of the draft are considered "draft-motivated" enlistees. These draft-motivated enlistees, expressed as a proportion of all enlistees, can be considered an estimate of the enlistment elasticity with respect to D; i.e., the proportional change in enlistments that can be expected from a 100 per cent decline in draft probability. These studies relied on a 1964 survey of first-term enlistee for their estimates of draft-motivated enlistments. Thus, their findings reflect the factors affecting D that were prevalent during the 1961-64 period. Estimates of the enlistment elasticity with respect to D derived from surveys taken at other periods might differ from those used in these cross-section studies. A fortiori, estimates generated from time-series analyses based on observed variations in indexes of D can also be expected to differ from those generated from the 1964 survey.

The time series studies relied mainly on an induction rate as an index of draft-probability. As noted earlier, some of the time series studies, in order to minimize the possibility of simultaneous equations bias, used an overall accession rate as their index of D. While these rates are undoubtedly important elements in an index of D, they are not necessarily the only elements. The effect of a given draft-probability on the enlistment behavior of a potential draft-motivated enlistee will depend on the options he has available to avoid the draft. Historically, these options included staying in school, getting married and having children, and, for a brief period, just getting married. Considering draft-avoidance as an option to enlisting can result in a more complex formulation of D. In principle, the effect of any draft-avoidance option may depend on the level of D; i.e., the effort one makes to take advantage of these options may depend on the perceived likelihood of being drafted. Moreover, the effect may differ between peacetime and wartime environments. This would suggest some sort of interactive model of D. Some studies have tried to include draft-avoidance options in their analysis in relatively simple ways[27]; most time-series studies do not address the issue.

Other Enlistment Determinants

A number of studies have assumed that d varies with race, region, risk of injury, level of school completed, public opinion, or resources devoted to advertising and recruiting. Cross section studies focussed on race, region and recruiting;[28] time series studies examined the effects of risk, public opinion, and level of school completed.[29] The race effect was estimated by including an estimate of the racial composition of a region as an additional argument in the enlistment function. The regional effect was estimated by Gray as a South-nonSouth effect. He used a dummy variable to indicate Southern states. The recruiting effect was estimated through a recruiter variable, the ratio of recruiters to P, the number of qualified military available. The effect of risk was estimated through dummy variables indexing the Berlin crisis, the Cuban missile crisis and the involvement of U.S. forces in Viet Nam. Viet Nam involvement was also proxied by the total number of U.S. military casualties in Southeast Asia.

Seasonality in Enlistments. Time series enlistments display strong seasonal patterns reflecting two kinds of seasonal events; 1) the surge in enlistments in June, when most enlistment eligibles complete school, and 2) the seasonal dip in enlistments in December, when the Armed Forces Enlistment and Examination Stations (AFEES) are closed for the Christmas and New Years holidays, followed by a surge in January, when the AFEES catches up on the backlog created by the December holidays.

With one notable exception, the Klotz study, all the time series studies control for this seasonal pattern in enlistments by including seasonal dummy variables. As we shall see, the method of dealing with seasonality has a strong effect on estimates of several key parameters, the estimated coefficient of U and the estimated coefficient of D. Moreover, the sensitivity of these parameters also affects the estimates of the voluntary enlistment supply elasticity of M and C.

The reason for this sensitivity lies in the strong seasonal components found in the series used to estimate U and D. This seasonality makes them highly collinear with the seasonal dummies included to control for the seasonal effects on enlistments described above. Including the seasonal dummies assumes that there is enough independent nonseasonal variations in them. Excluding the seasonal dummies results in attribution of the entire seasonal effect on enlistments to these seasonally sensitive variables. One could speculate that including the seasonal dummies should produce a lower-bound estimate of the effects of D and U, whereas excluding the seasonal dummies should produce an upper-bound estimate of their effects.

SIMULTANEOUS EQUATIONS BIAS

The issue of potential simultaneous equations bias arises from the possibility that some of the variables on the right hand side of (11) may be affected by enlistments, in which case ordinary least squares techniques of estimating the parameters (11) may be biased. The possibility of such bias is strongest in the case of the variables C, U, and D.

The potentiality of an enlistment effect on C and U arises from the role of enlistments as a shift variable in the civilian labor supply function of potential enlistees. Other things equal, this supply can be expected to be inversely related to the number of enlistments; i.e., an increase in enlistments is expected to shift the civilian labor supply schedule to the left. This shift is expected to reduce unemployment rates and, under the appropriate demand conditions, to increase C. Given the first-order partials hypothesized for (11), the result of this reverse causality would be to bias the estimated effect of U and C toward zero. The bias could be expected to be more pronounced for the variable U because of short-run wage rigidities and fairly elastic long-run demand functions for potential enlistees in the civilian sector. Some of the time-series studies (notably, Fisher, Klotz, and Kim, et al.) have attempted to minimize the potential simultaneous equations bias by lagging their estimates of C, U, and M one quarter. Otherwise, the studies have not attempted to deal with the problem of simultaneous equation bias in C and U. From the evidence, we shall see that such neglect is probably justifiable in the case of C, but is riskier in the case of U.

The issue of simultaneous determination of E and D would arise from an assumed fixed demand for military accessions. Given this demand, one would expect to find an inverse relationship between enlistments and inductions, i.e., given the demand for accessions, higher enlistments are expected to result in a smaller number of inductions. The result of this reverse causality would be to bias the estimated impact of inductions on enlistments toward zero. Steady-state accession demand arises from the need to replace individuals who are expected to separate from service. This steady-state replacement demand is modified by an appropriate increment (or decrement) when the desired stock of military manpower changes. Thus, steady-state accessions are augmented for force build-ups and are reduced for force reductions. Replacement demand is determined largely by first-term accessions lagged an appropriate number of years. The length of the lag would be the length of obligated service of the average first-term accession. Force build-ups and reductions can be assumed to be exogenously determined. Thus, the issue of simultaneity between E and D does not appear to be one that requires special treatment. Some studies attempt to minimize the problem by using accessions as an index of D.[30]

FOOTNOTES, CHAPTER II

1. The first, initiated by President Johnson, was undertaken by the Department of Defense in 1964. The second, initiated by President Nixon, was undertaken by the President's Commission on an All-Volunteer Armed Force (hereafter referred to as the Gates Commission) in 1969. The findings of these analyses are summarized in: US Congress, House Committee on Armed Service, Hearings, Review of the Administration and Operation of the Selective Service System (89th Congress, 2d Session, 1966), pp. 9923-35 (hereafter abbreviated House hearings); US President's Commission on an All-Volunteer Armed Force, The Report of the President's Commission on an All-Volunteer Armed Force (Washington: US Government Printing Office, 1970), (hereafter abbreviated Report).

2. The literature pertaining to these studies includes: A.C. Fisher, "The Cost of the Draft and the Cost of Ending the Draft," American Economic Review, LIX, No. 3 (June 1969), pp. 239-255; S.H. Altman, "Earnings, Unemployment, and the Supply of Enlisted Volunteers," Journal of Human Resourses, IV, No. 1 (Winter 1969), pp. 38-59; B.J. Klotz, "The Cost of Ending the Draft: Comment," American Economic Review, LX, No. 5 (December 1970), pp. 970-979; A.C. Fisher, "The Cost of Ending the Draft: Reply," loc. cit., pp. 979-983; B.C. Gray, "Supply of First-Term Military Enlistees," Studies Prepared For the President's Commission on an All-Volunteer Armed Force, Vol. 1 (Washington: US Government Printing Office, 1970), (hereafter abbreviated Studies); A.E. Fechter, "Impact of Pay and Draft Policy on Army Enlistment Behavior," loc. cit.; and Army Enlistments and the All-Volunteers Force: The Application of an Econometric Model, Institute for Defense Analysis Paper P-845 (Arlington, Va.: Institute for Defense Analysis, February 1972); A.A. Cook, Jr., "Supply of Air Force Volunteers," Studies;and "Quality Adjustments and the Excess Supply of Air Force Volunteers," Review of Economics and Statistics, LIV, (May, 1972), No. 2, pp. 166-171; A.A. Cook, Jr., and J.P. White, "Estimating the Quality of Airman Recruits," Studies, K.H. Kim, S. Farrell, E. Clague, The All-Volunteer Army: An Analysis of Demand and Supply (New York: Praeger Publishers, 1971), pp. 79-120.

2. See, for example, J.T. Bennett, S.E. Haber and P.J. Kinn, "The Supply of Volunteers to the Armed Forces Revisited," unpub. ms., George Washington University, School of Engineering and Applied Science Institute for Management Science and Engineering, March, 1972; D.W. Grissmer, D.M. Amey, R.L. Arms, D.F. Huck, J.F. Imperial, L.D. Koenig, W.S. Moore, G.P. Sica, R. Szymanski, An Econometric Analysis of Volunteer Enlistments by Service and Cost Effectiveness Comparison of Service Incentive Programs, General Research Corporation Report OAD-CR-66, October, 1974. Glenn A. Withers, "International Comparisons in Manpower Supply," Unpub. ms., Institure of Advanced Studies, Australian National University, December, 1975.

4. Fisher, "The Cost of the Draft...," loc. cit., esp. pp. 240-246.

5. Ibid., p. 241.

6. Ibid.

7. The lognormal distribution also allows him to define W_{CP} as median rather than mean earnings. Ibid., pp. 243-244.

8. Fisher further assumes that the draftee and the volunteer receive the same compensation, W_{MP}.

9. Full details of the derivation are presented in A.C. Fisher, The Supply of Enlisted Volunteers for Military Service, unpublished Ph.D. dissertation, Columbia University, 1968.

10. If $\frac{A}{P}$ is close to zero, in $\ln(1-\frac{A}{P})$ can be approximated by $-\frac{A}{P}$. Thus, equation (10) can be written:

$$(10')\quad \frac{E}{P} = \alpha + \beta_1 \ln \frac{W_{CP}}{W_{MP}} + \beta_2 \ln (1-U) - \beta_3 \frac{A}{P}$$

since $\frac{A}{P}$ is the sum of the enlistment rate, $\frac{E}{P}$, and the induction rate, $\frac{I}{P}$, we can rewrite (10) in terms of the induction rate as follows:

$$(10'')\quad \frac{E}{P} = \frac{\alpha}{1+\beta_3} + \frac{\beta_1}{1+\beta_3} \ln \frac{W_{CP}}{W_{MP}} + \frac{\beta_2}{1+\beta_3} \ln (1-U) - \frac{\beta_3}{1+\beta_3} \frac{I}{P}$$

Thus, if $\frac{I}{P}$ becomes zero, $\frac{E}{P}$ will fall by an amount equal to $\frac{\beta_3}{1+\beta_3}$ times the current level of $\frac{I}{P}$.

11. First term military pay, indexed by the basic pay of an E-1, was unchanged from 1952 to 1965. It rose by about ten percent in 1965 and at above five percent per year in 1966-67. Major pay increases were not experienced until 1969, when basic pay was raised by about ten percent. The most dramatic pay increase occured in 1971 when basic pay was doubled to move to an all-volunteer force. See U.S. Congress, House Armed Services Committee, Pay and Allowances of the Uniformed Services, (Washington: U S Government Printing Office, 1973, pp. 91-93.

12. Fechter, Army Enlistments ..., pp. 61-62, and Appendix C. The exact nature of the bias will depend on the first order partial differential between Rs and C, holding constant the other independent variables in the enlistment function.

13. Richard B. Freeman, The Market for College Trained Manpower: A Study in the Economics of Career Choice (Cambridge: Harvard University Press, 1969), pp. 14-15.

14. Richard Thaler and Sherwin Rosen, "Estimating the Value of a Life," unpub. ms., October, 1973; Robert S. Smith, "Compensating Wage Differentials and Hazardous Work," Technical Analysis Paper No. 5, Office of Evaluation, Office of the Assistant Secretary for Policy and Research, August, 1973.

15. Mental categories are percentile distributions based on performance of enlistment applicants and potential inductees on tests designed to assess the ability of new recruits to absorb training within a specified period. The five mental categories used by the military in classifying potential recruits are

Mental Category	Percentile
1.	93-100
2.	66-92
3.	31-65
4.	11-30
5.	0-10

A discussion of the tests used to generate these mental categories can be found in B. Karpinos, "Mental Test Failures," in S. Tax, ed., The Draft: A Handbook of Facts and Alternatives, (Chicago: University of Chicago Press, 1967), pp. 35. 35-49. Applicants in mental category 5 were not qualified for either induction or enlistment. For several years--between January, 1959 and May, 1963--the Army accepted no regular enlistments from mental category 4. The other services accepted applicants from the upper half of mental category 4, but only if they achieved minimum scores on aptitude test batteries or if they had high school diplomas. The Army adopted a similar policy in 1962. A detailed chronology of mental standards since 1948 can be found in H. Wool, The Military Specialist: Skilled Manpower for the Armed Forces, (Baltimore: Johns Hopkins University Press: 1968), pp. 63-69; See also Hearings before the House Committee on Armed Services, Review of the Administration and Operation of the Selective Service System, June, 1966 (Washington: Government Printing Office, 1966), pp. 10019-10020.

16. Fisher; op. cit.; Klotz, op. cit.; Kim et al., op. cit.; Fechter, op. cit.; Altman, op. cit.; Gray, op. cit.

17. Grissmer, et al., op. cit.; Grissmer, op. cit.

18. Cook, op. cit.

19. Fisher, op. cit.; Altman, op. cit.; Klotz, op. cit.; Kim, et al., op. cit.

20. Fechter, op. cit.; Cook, op. cit.; Gray, op. cit.; Grissmer, op. cit.

21. Fechter, op. cit.

22. Cook, op. cit.

23. Fechter, op. cit.; also includes the tax advantage on these allowances.

24. Cook, op. cit.

25. John C. Hause, "Enlistment Rates for Military Service and Unemployment," Journal of Human Resources, VIII, (Winter, 1973), No. 1; pp. 98-109.

26. Altman, op. cit.; Gray, op. cit.

27. Fechter, op. cit.; Cook, op. cit.

28. Altman, op. cit.; Gray, op. cit.; Bennett, et al., op. cit.; Grissmer, et al., op. cit.

29. Fechter, op. cit.; Cook, op. cit.; Withers, op. cit.; Grissmer, op. cit.

30. Fisher, op. cit.; Klotz, op. cit.; Kim, et al., op. cit.

Chapter III

REVIEW OF LITERATURE: SUMMARY OF FINDINGS

by

Alan Fechter and Dorothy Amey

The findings of 17 American enlistment studies[1] are summarized in this chapter. Studies selected for review were those in which there was some econometric model either explicitly or implicitly stated from which the parameters of the variables summarized in (11) were estimated. Table 1 catalogs those studies and provides information about the service for which the analysis was performed, the time period of the analysis, the units of observation, and the particular enlistees studied. For a number of reasons, it is desirable to stratify the studies into those based on time series data and those based on cross section data.

First, the cross section studies faced problems in deriving estimates of military pay elasticities because they could not observe variation in military pay at a moment in time. They were therefore constrained to deriving their estimate of military pay elasticity from the enlistment effect of the civilian pay variable. To some extent, the earlier time series analyses were also so constrained. As noted earlier, derivation of military pay elasticities from civilian pay parameters creates the possibility of bias arising from the assumption of symmetric enlistment responses to equi-proportionate pay changes and from potential systematic measurement error in the civilian pay variable. The time series analyses, on the other hand, were able to observe at least some variability in the military pay variable. Thus, they were able to base their estimates on some of the observed variation.

Table 1

SUMMARY OF ENLISTMENT STUDIES

Author	Service[a]	Period of Analysis	Units of Observation	Units of Analysis
Time series				
Fisher	D	3:57-4:65	Quarters	Enlistees, Mental Groups I-III[b]
Klotz	D	3:57-4:65	Quarters	Enlistees, Metnal Groups I-III
Kim, et al.	D, A	3:57-4:65	Quarters	Enlistees, Mental Groups I-III
Fechter	A	1:58-3:68	Quarters	White enlistees in Mental Groups I-III
Cook	AF	1:59-2:67	Quarters	Enlistees in Mental Groups I-II, I-III, I-IV
Hause	D	1:57-4:63	Quarters	Enlistees in Mental Groups II.
Withers	D, A	2:66-4:73	Quarters	Enlistees, Mental Groups I-IV
Grissmer, et al.	A, N, MC, AF	1:71-12:73	Months	Voluntary enlistees by Mental Group and level of school completed
Grissmer	A, N, MC, AF, D	1:70-12:75	Months	Voluntary enlistees by mental group and level of school completed
Cross Section				
Altman, Oi	D, A	1963	Census regions	Enlistees in Mental Groups I-II, I-III
Kim, et al.	D, A	1963		Enlistees in Mental Groups I-III
Gray	D, A, N, MC, AF	1964	State-groups	White enlistees in Mental Groups I-III
	N	1967	State-groups	Enlistees by level of school completed
Bennett, et al.	D, A, N, MC, AF	1970		
Grissmer,	A, N, MC,	1972, 1973	State-groups	Enlistees by age (17-18, 19-21) Enlistees by Mental Groups (I-II, I-III) Enlistees by level of school completed (high school grad.- Black enlistees by level of school completed
Lockman, et al.	N	1973	Navy recrui- Districts	All enlistees, School-eligible high-school graduates.

Notes: a. D=all Services; A=Army; N=Navy; MC=Marine Corps; AF=Air Force.
b. See Fechter, op. cit., p. 48, fn 8, for a description of Mental Groups.

Second, the cross section and the time series studies differed in methods used to estimate the impact of D on enlistments. Most of the time series studies derived their estimates from the estimated regression coefficient of some index (or indexes) of draft probability. (A few even tried to control for other changes in draft policy, such as shifts in priority of enlistment). With the advent of the lottery draft, direct estimation of voluntary enlistments became possible, eliminating the need to estimate the impact of D. Thus, the later time series analyses (notably those of Grissmer et al. and Grissmer) finesse the problem by estimating voluntary enlistments from the fraction of total enlistees with low-priority lottery numbers. As noted in the preceding chapters, most of the early cross section studies, notably those of Fisher, Kim et al., and Gray, used responses of first-term enlistees to a survey question about their enlistment motivation to derive their estimates of voluntary enlistments.[2]

Third, the time series studies had to come to grips with the problem of seasonality in the enlistment series and in the series used to estimate D and U. In contrast, the cross section studies were able to work with observed variability that was by and large nonseasonal in nature. Thus, the time series results reflects in one way or another, depending on how seasonality is handled in the study, this seasonal factor, whereas the cross section results can be considered relatively free of seasonal influences.

In addition, there is reason to expect the all-service results to differ from the service-specific results because of what may be called the "interservice" effect. This effect can be attributed in large part to changes in the return to enlisting in service i relative to the return to enlisting in the other services. This variation in relative returns should result in a change in the share of total enlistment supply going to service i. The presence of an inter-service effect implies that total service elasticities should be no larger than service-specific elasticities and frequently, they should be smaller.

Moreover, results can be expected to differ from service to service according to the amount of excess supply faced by any individual service. The conventional wisdom of military manpower experts is that the Air Force and Navy, because of their relatively good conditions of service (and the relatively high nonpecuniary returns they imply), are most likely to enjoy circumstances of excess supply.

For these reasons, it would also be desirable to analyze the results by service.

TIME SERIES STUDIES

There are really four distinct sets of data on which time series analysis were performed; the Fisher, Klotz, and Kim et al. studies were performed on total enlistment data for the period 1958-1965; the Fechter and the Cook studies were performed on total enlistment data for the period 1958-1968 and 1959 and 1967, respectively; the Withers study was performed on total enlistment data for the period 1966-1973; and the Grissmer et al. and the Grissmer study was of voluntary enlistments for the periods 1971-1973 and 1970-1975, respectively and used monthly rather than quarterly data. Moreover, these monthly studies were able to stratify by level of school completed. The results are reported for high school graduates in Mental Groups 1-3. In addition, the Fechter and the Cook studies were of individual services; the Fisher and Klotz studies were for all services. And Withers and Kim et al. studied both Army and DOD enlistments. Finally, because of the peculiar nature of the Air Force enlistment market, Cook had to deal with the possibility of an excess supply of enlistment applicants.

The results of these studies are summarized in Table 2. In order to standardize for methodological differences among models, the findings related to enlistments in Mental Groups 1-3 are reported, a group that has generally been assumed to be supply-determined. Moreover, the summary is restricted to relative pay models that assume instantaneous adjustment. Finally, only the results of models in which male civilian populations (or their equivalents) are used to estimate P are discussed. Other findings are reported only when they differ notably from the findings summarized in Table 2.

Several general aspects of these findings are notable at the outset. First, there is a consistent tendency to find a significantly positive relative pay elasticity, a negative but statistically insignificant employment rate elasticity, and a significant enlistment elasticity with respect to draft pressure. Klotz is the only exception to the relative pay and employment rate findings. Withers is the only exception to the draft pressure finding. Only Fisher and Withers find a pay elasticity consistently less than one.[3]

Third, although generally not statistically significant, the employment rate elasticities exceed the relative pay elasticities in eight of the twelve cases displayed in Table 2, and these employment rate elasticities exceed one in ten of these twelve cases.[4] The tendency for the employment rate elasticity to exceed the relative pay elasticity is consistent with the risk averting behavior on the part of potential enlistees with respect to employment discussed in the preceeding chapter.[5]

The statistical insignificance of the coefficient of (1-U) may be the result of measurement error; i.e., (1-U) may be a very poor proxy for p_e.

Table 2

SUMMARY OF TIME-SERIES FINDINGS: ELASTICITIES WITH RESPECT TO EXPECTED PECUNIARY RETURNS TO ENLISTING (W_{MP}/W_{CP}), THE EXPECTED PROBABILITY OF BEING EMPLOYED (1-U) AND THE PROBABILITY OF BEING INDUCTED (A/P OR I/P)

			Elasticities W_{MP}/W_{CP}		(1-U)		
Study	Equation estimated	Type of Service[a]	With draft	No draft	With draft	No draft	Proportion Draft-motivated
Fisher	Semi-log	D	.47	.62	-.59[b]	-.78[b]	.24
Klotz	Semi-log	D	.87[b]	1.47[b]	-2.08	-3.53	.41
Kim, et al.	Semi-log	D	1.68	2.78	-1.92[b]	-3.15[b]	.38
		A	1.94	2.40	-2.88	-3.56	.20
Fechter	Linear	A	1.39	1.74	c	c	.20
Cook	Log-linear	AF	2.23	2.23	-1.36[b]	-1.36[b]	n.a.
Withers	Linear	D	.52	.55	c	c	.05
		A	.28[b]	.35[b]	c	c	.18
Grissmer, et al.	Non-linear[d]	A	n.a.	.62	n.a.	3.67	n.a.
		N	n.a.	.44	n.a.	n.a.	n.a.
		MC	n.a.	.15	n.a.	2.09	n.a.
		AF	n.a.	.53	n.a.	2.16	n.a.
Grissmer	Log-linear	A	n.a.	1.33	n.a.	1.47[b,c]	n.a.
		N	n.a.	1.45	n.a.	n.a.	n.a.
		MC	n.a.	.38	n.a.	c	n.a.
		AF	n.a.	.94	n.a.	3.40[e]	n.a.

Sources: Fisher, "The Cost of the Draft...," p. 248; Klotz, "The Cost of Ending the Draft...," pp. 972-973; Kim, et al., The All-Volunteer Army..., pp. 94-95, 200-201; Fechter, Army Enlistments..., pp. 93-94, 97; "Impact of Pay...," pp. 24, 27, 29, 34, 37; Cook, "Supply of Air Force Volunteers...," p. 17; Withers, "International Comparisons...," p. 12; Grissmer, et al., "An Econometric Analysis...," pp. 108, 111, 128, 140, 153, 161.

Notes: a. D=all Services (DOD); A=Army; AF=Air Force
b. Based on a regression coefficient that was not statistically significant from zero at the .05 level.
c. Regression coefficient had "wrong" sign.
d. Grissmer, et al., "An Econometric Analysis...," p. 169.

It seems reasonable to assume that p_e, which is a long-run concept, would be better approximated by a function that included lagged values of (1-U) (such as a moving average or a distributed lag function).

These general tendencies with respect to sign and significance hide a considerable amount of variation among these studies in the magnitude of the relationships. Much of the variation arises from differences in units of analysis. To standardize for this, comparisons are made of the findings from the all-service enlistment equations estimated by Fisher, Klotz, Kim et al., and Withers; then findings from the Army equations estimated by Kim et al., Fechter, and Withers are compared. Finally differences between the all-service and the specific-service enlistment equations are discussed.

Within the all-service enlistment equations there were several notable differences in research methods that could generate differences in findings. Klotz, for example, excludes seasonal dummies from this estimating equations and derives his estimate of draft-motivated enlistments from the coefficient of the induction rate (rather than from the coefficient of the military accession rate). Kim, et al. use an estimate of P that encompasses a broader age group, but excludes potential inductees who were found to be ineligible for induction. Withers uses a different method of estimating pay, includes variables to account for quality and taste variation, and estimates his parameters from a later sample used. Both Fisher and Withers derive relative pay elasticities that range between 0.47 and 0.62. The effect of Klotz's modification of Fisher's analysis is summarized in Table 3. His use of an induction rate to estimate draft-motivated enlistees raises the pay and the employment elasticities.[6] However, it lowers the estimate of the fraction of enlistees who were draft-motivated. The latter finding can be attributed to biases inherent in the alternative specifications of draft pressure. In Fisher's case, A/P is equal to the sum of E/P and I/P. Thus, regresssing (1-A/P) on E/P produces a certain amount of spurious built-in negative correlation that biases Fisher's estimated proportion of draft-motivated enlistees upward. In Klotz's case, given a level of required accessions, one would expect to find a negative correlation between E/P and I/P. Thus, deriving draft-motivated enlistments from I/P could bias Klotz's estimated proportion of draft-motivated enlistees downward.

Table 3

SUMMARY OF EFFECTS OF KLOTZ MODIFICATIONS

	Elasticities W_{MP}/W_{CP}		(1-U)		
	With draft	No draft	With draft	No draft	Proportion draft-motivated
Fisher	.47	.62	-.59	-.78	.24[a]
Klotz - using I/P	.67	.83	-.84	-1.05	.20[b]
Klotz - excluding seasonal	.87	1.47	-2.08	-3.53	.41[b]

Source: Fisher, loc. cit.; Klotz, loc. cit.

Notes: a. Estimate based on equation using ln (1-A/P).
b. Estimate based on using as the appropriate argument of the original enlistment function.

Klotz's exclusion of the seasonal dummies increases all elasticities and also increases his estimate of draft-motivated enlistments. In addition, it makes his pay coefficient statistically insignificant and his unemployment coefficient statistically significant--exactly the reverse of Fisher's findings (see Table 2). The effects are most dramatic for the probability of being employed (1-U) and for the induction rate (1-I/P), suggesting that seasonal variations in unemployment and induction activity have been important determinants of the seasonal pattern of enlistments and that Fisher's use of seasonal dummies tended to obscure this relationship.

The pay elasticities derived from the Kim, et al. study are considerably larger than those derived by Fisher and Klotz. In part, this is the result of differences between the two studies in their method of estimating the eligible population. Fisher uses a male civilian population age 17-20, whereas Kim, et al. use an estimate of qualified and available Selective Service registrants age 19-26. A comparison of the two estimates shows that the Kim, et al. estimate is consistently larger than the Fisher estimate and that the difference increases from slightly under 4 percent in 1958 to slightly over 40 percent in 1965, averaging around 10 percent.

Since estimates of elasticities are inversely related to estimates of E/P in semi-logarithmic functions, one would expect part of the difference between these two studies to reflect the smaller E/P of the Kim, et al. study. Moreover, the secular widening of the differential could have caused differences in some of the estimated parameters of variables displaying strong trends, such as the relative pay variable. Since the average enlistment rates of the two studies differ by only ten percent, it is apparent that only a small fraction of the observed difference in pay elasticities can be attributed to differences in estimated population. The remainder must be due to factors that produce differences in estimated parameters of the enlistment function.

Turning to the Army studies, we find that the Kim et al. pay elasticities are roughly 35 percent higher than Fechter's, whereas the Withers elasticities are 80 percent lower than Fechter's. Again, part of the difference can be attributed to differences between Kim et al. and the others in methods of estimating P. The employment rate elasticities of Kim, et al. are over 150 percent higher then those of Fechter and Withers and are statistically significant. In contrast, their induction rate elasticities are exactly the same as Fechter's, and both elasticities are substantially below those of all service enlistment studies. The Fechter and Withers study differ from the Kim, et al., study in their use of the civilian male population, age 17-20, to estimate P. And the Fechter study also differs from the Kim, et al. study in its restriction to white enlistees, and in its coverage of a longer time period. Both Fechter and Withers also include additional independent variables in their analyses. Reconciliation of these findings can be accomplished by reestimating these Army enlistment equations in a manner that standardizes for the differences discussed above.

Fechter also experiments with absolute pay models and distributed lag models and finds that his estimates of pay elasticity are sensitive to model specification. In particular, he finds that military pay elasticities estimated from absolute pay models are higher than military pay elasticities estimated from equivalent relative pay models. He also finds that, in the absolute pay models, the point estimate of the military pay elasticity exceeds the point estimate of the civilian pay elasticity in three of the four cases presented. This finding is inconsistent with the hypothesis that the net nonpecuniary returns to enlisting would result in a declining enlistment rates with equiproportionate increased in W_{MP} and a correctly specified W_{CP}. However, since W_{CP} may not be correctly specified, this finding can not be considered strong evidence refuting this hypothesis. The distributed lag model increases Fechter's estimate of the proportion of enlistees who are draft-motivated.

The most direct comparison between DOD enlistments and specific service enlistments can be made by examining the findings of Kim, et al. They find that point estimates of the fraction of DOD enlistments that are draft-motivated is almost double that of the Army. This finding is strengthened when one broadens the comparison to include the less comparable specific-service study of Fechter and DOD studies of Fisher and Klotz.

Their finding with respect to pay elasticity is less clear. The point estimate of the no-draft DOD pay elasticity exceeds the point estimate of the no-draft Army pay elasticity by 16 percent, but the point estimate of the DOD pay elasticity of enlistment in a draft environment is 13 percent less than the point estimate of the equivalent Army pay elasticity. Extending the comparison to the less comparable studies noted earlier, we find that the specific-service studies of Fechter and Cook produce point estimates of pay elasticities that exceed those of the DOD studies of Fisher and Klotz, but fall short of the DOD pay elasticity estimated by Kim, et al. Recall that a major difference between the Kim, et al. study and the other studies in their estimates of P. Also note that the Fechter and Cook studies extend their samples beyond 1965, whereas the samples of the other studies stop with 1965. Finally, consider the fact that both Fechter and Cook include variables to control for the influence of the cold war, and the hot war in Southeast Asia, and that Cook also includes variables to control for a possible excess supply of enlistment applicants and for the effects of a change in the draft-priority of married men. Thus, while the comparisons between Fechter-Cook and Fisher-Klotz are standardized so that there are no differences in their estimates of P, one cannot standardize them for differences in periods of analysis and model specification without reestimating the enlistment equations.

A hypothesis explaining the observed difference in pay elasticities is the longer period of analysis in the studies by Cook and Fechter. Military pay remained virtually unchanged during the years 1958-1963 and began to advance rapidly thereafter. Since military pay is not estimated precisely it is possible that its error component was a larger fraction of its total variation from 1958-1965 than it was from 1958-1968. Table 4 summarizies some of the salient aspects of military pay variation for the two periods discussed above. By any criterion, the studies by Cook and Fechter had more pay variation to work with than did the earlier studies. Measurement errors in variables can, under the appropriate circumstances, bias estimated regression coefficients toward zero. And this bias will be stronger, the larger is the proportion of the observed variation that can be attributed to the error component. Thus, it is conceivable that the observed higher pay elasticities estimated by Cook and Fechter can be attributed in part to larger amounts of military pay variation in their data.

Unfortunately, there is no way to test this hypothesis directly. One must either reestimate the parameters of the Cook and Fechter equations on a subset of their data or one must reestimate the parameters of the earlier studies using the more recent data. In addition, comparisions between Kim et al., and Fechter and Cook will require further adjustments for differences in their estimates of P.

The Cook study also departs from earlier studies by accounting for the possibility of excess supplies of enlistment applicants to the Air Force. He assumes that recruiters "cream" enlistment applicants by taking applicants from the top mental group first, and attempts to control for this effect by using an estimate of average score on the mental test as an additional independent variable in his analysis. Unfortunately, it is not possible to determine whether inclusion of this variable significantly effects his estimate of the military pay elasticity since he does not present results in which the test score variable is excluded from his analysis.

Table 4

COMPARISON OF MILITARY PAY VARIATION, 1958-65 AND 1958-68

	1958-65	1958-68
Average	4688	4987
Variance (in thousands)	20	301
Standard deviation	143	548
Coefficient of variation	.03	.11
Range	659	1872

Source: A. Fechter, Army Enlistments..., p. 93.

CROSS SECTION STUDIES

The major cross section studies are those of Altman, Oi, Kim et al., Gray, Bennett et al., and Grissmer et al. Altman, Oi, and Kim et al. each fit enlistment functions to nine regional observations for the year 1964. Gray fit enlistment functions to the same nine regional observations and to 34 "state" observations for the same year. Bennett, et al. and Grissmer, et al. are the first studies to use direct estimates of voluntary enlistments in their analysis. Bennett, et al. fit enlistment functions to 29-33 state groups using data for the year 1970. Grissmer, et al. fit enlistment functions to 47 state observations for the years 1972 and 1973.

In addition to being able to fit enlistment functions to voluntary enlistment data, these latter two studies also included measures of recruiting resources spent by the services in each of these areas. However, offsetting their strength in the superior enlistment data they use, they are weak in the methods used to estimating their pay variables.

Table 5 summarizes the cross section findings. We confine ourselves to the findings for all services and for the Army since the findings for the other services were either statistically insignificant or had the wrong sign. Like the time series studies, the pay elasticities are positive and, by and large, significant whereas the employment rate elasticities are generally positive, but based on coefficients that are not statistically different from zero. Eighteen of the twenty-two pay coefficients reported were significantly positive, whereas only five of the sixteen employment rate coefficients are significantly positive. The cross section pay elasticities are considerably lower than the time series elasticities, exceeding one in only six of the cases reported. However, as in the time series findings, the pay elasticity for the Army consistently exceeds the pay elasticity for all services and the pay elasticities for volunteers exceeds the pay elasticity for total enlistees. Only Gray reports reasonably consistent pay elasticities in excess of one. Like the time series findings, the point estimates of the employment rate coefficient, although generally not statistically different from zero, usually exceed the point estimates of the relative pay coefficient, falling short of one in only four of the cases reported. These employment rate findings are also consistent with the speculations concerning risk and measurement error discussed above in the interpretation of time series findings. The lower relative pay coefficients

Table 5

SUMMARY OF CROSS-SECTION FINDINGS: ELASTICITIES WITH RESPECT TO EXPECTED PECUNIARY RETURNS TO ENLISTING (W_{MP}/W_{CP}), THE EXPECTED PROBABILITY OF BEING EMPLOYED (1-U), AND THE PROPORTION OF TOTAL ENLISTEES WHO WERE DRAFT-MOTIVATED (DM)

			Elasticities W_{MP}/W_{CP}		(1-U)		
Study	Equation estimated	Type of Service[a]	With draft	No draft	With draft	No draft	DM
Altman	Log-linear	D	.38	.80[b]	-1.51	-2.73[b]	.39
		A	.54[b]	1.10[b]	-2.02[b]	-3.23	.45
	Log-complement	D	.44	.88[b]	-1.13[b]	-2.08[b]	.39
		A	.62	1.18	1.35[b]	-2.59[b]	.45
Kim, et al.	Semi-log	D	.12[b]	.82	n.a.	-2.44[b]	.39
		A	.46	1.19	n.a.	-2.44[b]	.45
Gray	Linear	D	.84	1.05	c	c	.39
		A	1.54	1.77	c	c	.45
Bennett, et al.	Log-linear	A[d]	.35	.65	.72[b]	-.48	n.a.
		A[e]	.45	.75	.96[b]	-.48	n.a.
Grissmer, et al.	Log-linear	A[f]		.20[b]		-2.30	n.a.
		A[g]		.68		-.70[b]	n.a.

Sources: Altman, "Earnings, Unemployment and the Supply of Volunteers," p. 56; Kim, et al., The All Volunteer Army..., p. 105; Gray, "Supply of First-Term Military Enlistees," pp. 15-16; Bennett, et al., "The Supply of Volunteers to the Armed Forces Revisited," p. 9; Grissmer, et al., An Econometric Analysis..., pp. 21-22.

Notes:
a. D=all services (DOD); A=Army.
b. Based on a regression coefficient that was not significantly different from zero at the .10 level.
c. Not included in this estimating equation.
d. Recruiting effort is included as a regressor.
e. Recruiting effort is excluded as a regressor.
f. Estimates are for fiscal year 1972; recruiting effort is excluded as a regressor.
g. Estimates are for fiscal year 1973.
n.a. Not applicable.

seem to be the result of the tendency in the cross section studies to adjust P for potential enlistees who may not be qualified to enlist.

As in the time series findings, these general tendencies with respect to sign and significance hide a considerable amount of variation among studies in their findings with respect to magnitude of the relationships. (e.g., Army vs. DOD) and some of these differences reflect methodological differences or differences in periods or units of observations among studies. To standardize for these difference, like services are compared within similar time periods. First the findings of the three early studies (Altman, Kim et al., and Gray) are compared and then the two later studies (Bennett et al. and Grissmer et al.) are compared.

The findings of the Altman study and the Kim, et al., study differ only because of the difference between them in the assumed explicit functional form of the enlistment function. Gray, however, differs from these studies in several ways. He estimates a different functional form; he confines his analysis to white enlistees only; he uses state rather than regional data; and he estimates W_{CP*} directly rather than estimating separate components of W_{CP*} (i.e., W_{CP} and p_e). He finds much higher pay elasticities than the other two early cross section studies. He reports that his findings are not sensitive to alternative specifications of functional form. This makes intuitive sense. Unless there is a large variance in the independent variables, elasticities estimated at mean values will not differ by much even if one functional form might better describe the entire range of the function than do the others. Moreover, the relatively small difference between Altman and Kim, et al. in their findings reinforces the likelihood that Gray's higher elasticity is not merely a reflection of the functional form he used.

The effect on Gray's findings using state data is summarized in Table 6. This table also summarizes the combined effect on Gray's findings of his use of white enlistees only, his inclusion of enrolled persons in his estimate of eligible enlistees, and his use of W_{CP*} as his independent variable. The effects of these methodological differences between Altman and Gray (other than their choice of units of observation) can be observed by comparing the first two rows of the table. The combined effect of the three major differences between these studies is to raise the estimated pay elasticity in three of the five cases presented--and to raise their elasticities by a substantial amount for the DOD enlistment functions. The effect of using state data can be determined by comparing the second and third row of the table. His use of state data substantially increases his estimated pay elasticities for Army enlistees, but reduces his estimates for DOD enlistees.

Table 6

COMPARISON OF ALTMAN AND GRAY ESTIMATES OF RELATIVE PAY ELASTICITY AND COMPARISON OF GRAY ESTIMATES OF RELATIVE PAY ELASTICITIES DERIVED FROM REGIONAL AND STATE-GROUP DATA

	All Services		Army	
	With draft	No draft	With draft	No draft
Altman	.38	.80	.54	1.10
Gray - 9 Census regions	.91	1.78	.43	1.31
Gray - 34 state-groups	.84	1.05	1.54	1.77

Source: Table 5 and Gray, Table 4.

Based on evidence presented above, we would expect Gray's inclusion of school enrollees in his estimate of eligibles and his use of W_{CP*} to increases his relative pay elasticity. The inclusion of school enrollees raises his estimate of P, lowering his estimate of E/P. If enlistment elasticities are inversely related to E/P, this should raise the pay elasticity. The use of W_{CP*} should increase the estimated pay elasticities because the employment rate elasticity, although statistically insignificant, has been shown to be substantially larger than the elasticity of W_{MP}/W_{CP} in the cross section studies (Table 5).

On a priori grounds, we would not be able to predict the effect of race on enlistment elasticities. We would expect enlistment elasticities for blacks to be smaller than enlistment elasticities for whites if black enlistment rates exceed white enlistment rates. One might expect black enlistment rates to exceed white enlistment rates because of discrimination against blacks in civilian labor markets. However, offsetting this tendency, one might expect black enlistment rates to be lower than white enlistment rates because of higher disqualification rates for blacks (resulting in a smaller number of eligibles, other things equal.)[7] There is evidence from the Grissmer, et al. study that enlistment elasticities are larger for blacks,[8] suggesting that the disqualification rate effect overwhelms the discrimination effect. However, methodological shortcomings in this particular study make this evidence very weak.

In contrast to the earlier cross section studies, the studies of Army enlistments by Bennet, et al. and Grissmer, et al. find substantially lower pay and unemployment rate elasticities. Major difference between the earlier and later cross section studies include differences in the time period of the analysis and the inclusion in the latter studies of variables to control for recruiting effort. Comparisons of findings from the Bennett, et al. study (Table 6) suggest that inclusion of the recruiter variable lowers the estimate pay elasticities slightly. However, their estimates of pay elasticity from the equation that excludes the recruiter variable suggests that the other differences are also important. To standardize for differences in estimates of eligibles, compare the Bennett, et al. findings with those of Gray. Gray's estimated pay elasticities are three times as large as the elasticities of Bennett, et al. Gray's study is restricted to white enlistees only and Gray estimates W_{CP*}. In addition, Bennett, et al. use a much cruder proxy for W_{CP}, the year-round, full-time equivalent of average hourly wages in manufacturing. Gray's restriction of his analysis to white enlistees only should have resulted in a lower pay elasticity, other things equal, if one is willing to accept the validity of the findings of Grissmer, et al. noted earlier. Thus, the lower pay elasticities of Bennett, et al. are probably attributable to differences in methods of estimating W_{CP}. The lower pay elasticity estimated by Bennett et al. appears to reflect measurement error in their estimate of W_{CP}.

Similar conclusions apply to the study of Grissmer, et al., except that their proxy for W_{CP} is even cruder than the one used by Bennett, et al. In brief, they create estimates of W_{CP} by adjusting national benchmarks of incomes of 17-21 year olds by the ratio of state-specific to national wages in manufacturing.[9] They estimated the national benchmarks on the basis of incomes of all persons 14-19 and 20-24 rather than on the basis of incomes of year-round, full-time workers. This biases their estimate downward since many persons in these age groups are voluntarily working part-time or part-year while they attend school.

In addition, Grissmer, et al. include variables proxying for the effect of enrollment in college and the presence of military facilities in their analysis. They find that both variables have significant effects on Army enlistments; college enrollment tends to inhibit enlistment and a military presence tends to encourage enlistments. Both of these variables can be justified as appropriate enlistments determinants. The college enrollment variable reflects an aspect of the return to not enlisting that is not captured by W_{CP}; the return to investment in higher education. The military presence variable can be interpreted as reflecting differences in perceived nonpecuniary returns to enlisting (i.e., taste differences) arising from exposure to the military environment. Since both are found by Grissmer, et al. to be statistically significant, they could have affected the coefficient of the relative pay variable-particularly if they are statistically related to relative pay.

FOOTNOTES, CHAPTER III

1. There have been several studies of foreign enlistment behavior which have been excluded from this review on the grounds that differences across countries enlistment standards and practices would make it difficult to reconcile the findings of the foreign enlistment experience with the findings of the studies of U.S. enlistment experience. For further information on the foreign enlistment studies, see Withers, "International Comparisons ...," E.S. Lightman, "Economics of Supply of Canada's Military Manpower," <u>Industrial Relations</u>, May, 197 , pp. 209-219.

2. See <u>supra</u>.

3. Fisher only reports the relative pay elasticity with a draft. I derive the no-draft pay elasticity as follows:

$$\eta_P^{ND} = \frac{\beta_1}{(1-d)\cdot\frac{E}{P}}$$

where: η_P^{ND} = no-draft enlistment rate elasticity with respect to Fisher's estimate of W_{CP}/W_{MP}

4. To derive the employment rate elasticities from the unemployment rate elasticities. the following transformation was used:

$$\frac{\partial\frac{E}{P}}{\partial(1-U)} \cdot \frac{(1-U)}{\frac{E}{P}} = \frac{\partial\frac{E}{P}}{\partial U} \cdot \frac{\partial U}{\partial 1-U} \cdot \frac{(1-U)}{U} = \frac{\partial\frac{E}{P}}{\partial U} \cdot \frac{U}{\frac{E}{P}} \cdot \frac{(1-U)}{U}$$

5. See <u>supra</u>.

6. I should note that my estimates of Klotz's no-draft elasticities differ from his reported estimates. This is because his estimates are based on an estimated 62 percent true volunteer enlistment, derived from survey data (Klotz, <u>op. cit.</u>, p. 972, whereas my estimates are based on an estimated 20 and 41 percent true volunteers, derived from the estimated parameters of the equation reported by Klotz for the Army and for all services, respectively.

7. Disqualification rates by race are summarized below for 1972.

	Mental Only	Physical Only	Mental & Physical	Administrative	Total
Nonblack	4.6	37.5	2.4	1.2	45.7
Black	31.8	17.6	11.8	1.4	62.6

Source: Office of the Surgeon General, Department of The Army, "Results of the Examination of Youths for Military Service, 1972"; Supplement to <u>Health of the Army</u>, September 1973, p. 3.

8. Grissmer, et al., _op. cit._, pp. 85-90.

9. _Ibid._, pp. 184-195. Unfortunately, details of how they generated the income estimate for 17-21 years old and of exactly which manufacturing wage statistic was used are not presented.

Chapter IV

RECONCILIATION OF THE FINDINGS OF THE FISHER AND KLOTZ STUDIES

by

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In this chapter, the Fisher and Klotz studies are reviewed in great detail. An attempt is made to reconcile the differences among their findings with respect to the impact of military and civilian pay, employment conditions and the draft, and the results of other studies discussed in the literature review of Chapter III. Among the differences investigated are differences in functional form, differences in sample period, differences in units of analysis, and differences in the model and variable specifications.

In Chapter III, a wide range in published estimates of the impact of pay, employment conditions and other variables was reported. The process of reconciliation of the findings consists of attempts to standardize for the methodological differences among the studies by means of a uniform data base compiled for this purpose. First, the original data for each of the studies are replicated whenever possible. The findings are then benchmarked to the uniform data base. The uniform data base is then used to investigate the sensitivity of the findings to a number of methodological differences among the studies in deriving parameter estimates.

THE FISHER STUDY

The estimating equation from the original Fisher model had the form:

$$(1)\quad \frac{E}{P} = \beta_0 + \beta_1 \ln\frac{W_C}{W_M} + \beta_2 \ln(1 - U) + \beta_3 \ln(1 - \frac{A}{P}) + u$$

where:

E = DOD enlistments, all races, Mental Groups 1-3

P = Civilian male population, age 18-19

W_C = Average earnings of civilian male year-round, full-time workers, age 18-20[1]

W_M = Average earnings of first-term enlistees[2]

U = Unemployment rate, male civilian labor force, age 18-19

A = Total DOD accessions[3]

ν = Error term[4]

The equation was fit to quarterly data in which the variables W_C/W_M and (1-U) were lagged one quarter in order to reduce possible simultaneous equations bias. Dummy variables were also included to control for seasonal effects. Fisher reports he fit this equation to quarterly data for the period starting with the third quarter of calendar year 1957 (:57) and ending with the fourth quarter of 1965 (4:65).

In analyzing the Fisher data base, it was discovered that Fisher deflated his estimate of W_C by the Consumer Price Index (CPI) so that civilian earnings are expressed in real terms. (Details of this analysis are described in Appendix A.) In contrast, Fisher does not deflate his estimate of W_M by the CPI; rather, he leaves his estimate of military pay in current dollars. There is no published explanation of why he opted for this method of estimating relative military pay. However, Fisher has indicated that he chose this method of account for income in kind. An alternative procedure would express both W_C and W_M in either real or nominal dollars. Consequently, his estimating equation was fit using alternative estimates of W_C/W_M. Moreover, the size of Fisher's sample was uncertain. While he reports having fit the equation to a sample beginning in 3:57, experiments with that sample did not conform very well to the results he reported. Consequently, experiments were made with two sample frames: one beginning in 3:57, and one beginning in 4:57.

Results with Original Fisher Data Base

Table 7 compares the findings reported by Fisher with the findings generated by the GRC experiments. FISH A summarizes the results using Fisher's estimate of relative pay on a sample beginning in 3:57; FISH B summarizes the results of the experiment using Fisher's estimate of relative pay on a sample beginning in 4:57; FISH C summarizes the results using GRC's estimate of relative pay on a sample beginning in 4:57.

FISH B produces a better approximation of Fisher's findings than FISH A. Also, while the relative pay coefficient is sensitive to the method used to estimate relative pay, the other results are relatively insensitive to this issue. The estimated relative pay coefficient falls from -0.00706 in FISH B to -0.00502 in FISH C, a drop of 29 percent. This reduces the relative pay elasticity estimated using the Fisher model on the Fisher data base from .67 to .48 with a draft, and from .82 to .59 without a draft.

Table 8 summarizes estimates of the elasticities of W_C/W_M and (1-U) both with and without a draft. The pay elasticities are less than unity. The elasticities with respect to (1-U), although derived from regression coefficients that were not statistically significant, always exceed, in absolute terms, the relative pay elasticities. The voluntary enlistment elasticities with respect to (1-U) exceed unity.

Fisher's model attributes the differences between the variables W_C/W_M and (1-U) in estimated elasticities to differences in the way expectations are formulated about the two variables. Recall that Fisher assumes the coefficient of (1-U) is composed of two components: a pecuniary component, reflected by the coefficient of W_C/W_M, and an "expectations"

Table 7

COMPARISON OF ESTIMATED ENLISTMENT SUPPLY PARAMETERS
DERIVED USING ORIGINAL FISHER DATA BASE

Estimated coefficients (and t-statistics)	Fisher (AER)[a]	FISH A[b]	FISH B[c]	FISH C[d]
$\ell n(W_C/W_M)_{t-1}$	-.00709[e]	-.00619[e]	-.00706[e]	-.00502[e]
	(-2.19)	(-2.02)	(-2.17)	(-2.27)
$\ell n(1-A/P)_t$	-.312[f]	-.311[f]	-.313[f]	-.311[f]
	(-7.61)	(-7.56)	(-7.54)	(-7.57)
$\ell n(1-U)_{t-1}$	-.00891	-.01160	-.00901	-.00881
	(-.87)	(-1.20)	(-.89)	(-.87)
R^2	.90	.91	.90	.91
Durbin-Watson Statistic	.31	.28	.28	.29

[a]Reported in Klotz, op. cit., p. 971.

[b]Computer from the sample, 3:57-4:65, using Fisher's estimate of W_C/W_M.

[c]Same as FISH A, only sample is 4:57-4:65.

[d]Same as FISH B, only W_C/W_M is estimated using nominal values of both W_C and W_M.

[e]t-statistic > |2.0|

[f]t-statistic > |3.0|

Table 8

COMPARISON OF ELASTICITIES WITH RESPECT TO W_C/W_M AND (1-U) WITH AND WITHOUT DRAFT AND OF THE PROPORTION OF ENLISTEES IN MENTAL GROUPS 1-3 WHO WERE DRAFT MOTIVATED (FROM FISHER MODEL USING ORIGINAL DATA BASE AND ALTERNATIVE SAMPLE PERIODS)

	Fisher (AER)	FISH A[a]	FISH B[a]	FISH C[a]
Elasticities (with draft)[b]				
W_C/W_M	-.68	-.58	-.67	-.48
(1-U)	-.85	-1.10	-.86	-.84
Proportion draft motivated[c]	.19	.19	.19	.19
Elasticities (no draft)[d]				
W_C/W_M	-.83	-.71	-.82	-.59
(1-U)	-1.04	-1.35	-1.25	-1.03

[a]See Table 1 notes b, c, and d for explanation of these models.

[b]Derived by dividing regression coefficient by $(1+\hat{\beta}_3) \cdot \bar{e}$, where β_3 is the regression coefficient of $\ln(1-A/P$ and (I/P) and $\bar{e}$ is the average enlistment rate for the sample period.

[c]Derived from the expression $\frac{\hat{\beta}_3}{1+\hat{\beta}_3} \cdot (\bar{I}/\bar{e})(I/P)$.

[d]Derived by dividing the elasticities (with draft) by 1-D, where D is the proportion of total enlistees in Mental Groups 1-3 who are draft motivated.

coefficient which relates the conceptual variable, the probability of employment (p_e) to the observable variable, (1-U).

Alternative explanations of the difference include the possible effects of risk-aversion or timing. As noted earlier, the variable (1-U) may be indexing the relative variability of expected returns to not enlisting as well as being a component of the expected value of these returns. This should serve to strengthen the effect of (1-U) on enlistments if potential enlistees tend to be risk-averse. In addition, it can be argued that employment conditions affect the timing of the enlistment decision as well as the decision whether or not to enlist. (That is, given a decision to enlist based on long-run expectations, it is best to enlist when opportunities are not good.) If the variable (1-U) reflects a timing decision as well as an enlistment choice decision, then its coefficient can be expected to be greater than that of other pecuniary variables reflecting only the enlistment choice decision.

Fisher derives his estimate of draft-motivated enlistments for the coefficient of $\ln(1-A/P)$. He makes the following assumptions:

(a) $\ln(1 - \frac{A}{P}) = -\frac{A}{P}$ for $\frac{A}{P}$ close to zero.

(b) $\frac{A}{P} = \frac{E}{P} + \frac{I}{P}$

Given (a) and (b), equation (1) can be rewritten:

$$(1') \quad \frac{E}{P} = \frac{\beta_0}{1+\beta_3} + \frac{\beta_1}{1+\beta_3} \ln \frac{W_C}{W_M} + \frac{\beta_2}{1+\beta_3} \ln(1-U) - \frac{\beta_3}{1+\beta_3} \frac{I}{P} + \frac{\nu}{1+\beta_3}$$

and the voluntary enlistment rate can be derived by subtracting $(\beta_3/1+\beta_3)(I/P)$ from the observed E/P. Fisher reports a voluntary enlistment rate of 0.0117. Given a total enlistment rate of 0.0153, it can be estimated, as Fisher does, that 24 percent of enlistees in Mental Categories 1-3 are draft motivated. However, an estimate of 19 percent based on average values of enlistment and induction rates is derived at GRC. The differences between GRC's estimates and Fisher's appears to be due to Fisher's use of the third and fourth quarter of 1965 for his estimates of I/P. These were quarters in which inductions were very high.

When reestimating the parameters of Fisher's enlistment supply function using his data base, the following assumptions are modified: (1) the assumption of symmetry of enlistment response to equiproportionate changes in W_C and W_M; (2) the assumption of stability in the mean of the taste distribution; and (3) the assumption of instantaneous supply adjustment.

The assumption of symmetry is tested by including as separate arguments in the estimating equation the real values of W_M and W_C (in place of the ratio, W_C/W_M). This estimating equation is not strictly comparable to the original Fisher estimating equation since Fisher uses the real value of W_C and the nominal value of W_M.

A major factor that could have caused a change in the taste distribution is the perceived risk of death or injury associated with enlisting. An event that occurred during Fisher's sample period that could have changes this perception was the U.S.-Russian confrontation in late 1961-early 1962 over the erection of the Berlin Wall. A dummy variable for this period of tension, during which draft calls were increased dramatically and 75,000 reservists were called to active duty, was added to the estimating equation to determine whether the conflict had any independent effect on enlistment behavior. Recall that both Fechter and Cook utilize this variable in their enlistment analyses.

The assumption of instantaneous adjustment is tested by adopting a distributed lag model of supply adjustment. This adjustment model was used with some success by Fechter in his analysis of Army enlistments. It assumes that the observed quarterly enlistment response to exogenous changes in enlistment determinants only represents some fraction of the long-run equilibrium response and that this fraction can be estimated by adding the enlistment rate (E/P) lagged one quarter as an additional independent argument to the enlistment function. Its estimated regression coefficient can be used to derive estimates of the fraction of the observed enlistment response that represents the long-run equilibrating response and can also be used, together with the other estimated regression coefficients, to estimate long-run response coefficients for the other independent variables included in the estimating equation.[5]

The parameters of the eight estimating equations representing the various combinations and permutations of these assumptions are displayed in Table 9. The results are disappointing. Relaxation of the symmetry assumption consistently produces military pay coefficients with the "wrong" sign (i.e., coefficients that differ in sign from what the theoretical model predicts). While these models produce the theoretically expected negative coefficients for civilian pay, these coefficients are only statistically significant in the models which assume instantaneous adjustment. Pay elasticities estimated from the significant coefficients are only slightly larger than the pay elasticities estimated from comparable relative pay models. Recall, however, that the estimating equations for the absolute pay models (equations (5)-(8)) are not strictly comparable to the original Fisher equation represented by equations (1)-(4). Equations (5)-(8) are expressed in terms of the real values of W_C and W_M, whereas equations

Table 9

SUMMARY OF RESULTS OF ALTERNATIVE MODEL ASSUMPTIONS ON PARAMETERS OF FISHER'S DOD ENLISTMENT SUPPLY MODEL, 4:57–4:65

Model	Regression Coefficient (t-statistic in parentheses) $\ln(W_C/W_M)_{t-1}$	$\ln(W_M)_{t-1}$	$\ln(W_C)_{t-1}$	$\ln(1-U)_{t-1}$	$\ln(1-A/P)_t$	Berlin	$(E/P)_{t-1}$	R^2	D-Statistic	Proportion Draft Motivated[a]
Relative pay, no Berlin										
(1) No lag	-7.057[b]			-9.014	-312.7[c]			.90	0.28	.19
	(-2.17)			(-0.89)	(-7.54)					
(2) Lag	-1.313			0.992	-255.6[c]		0.542[c]	.96	1.22	.31
	(-0.54)			(0.14)	(-4.93)		(5.63)			
Relative pay, Berlin										
(3) No lag	-7.110[b]			-8.556	-298.2[c]	1.036		.91	0.38	.17
	(-2.20)			(-0.85)	(-6.95)	(1.18)				
(4) Lag	-0.219			2.302	-261.0[c]	-1.118	0.640	.96	1.40	.40
	(-0.09)			(0.33)	(-8.97)	(-1.62)				
Absolute pay, no Berlin										
(5) No lag		-20.52[b]	-8.565[c]	2.671	-360.2[c]			.92	0.67	.23
		(-2.19)	(-3.49)	(0.26)	(-8.66)					
(6) Lag		-6.054	-2.324	3.548	-274.3[c]		0.495[c]	.96	1.19	.31
		(-0.78)	(-1.01)	(0.46)	(-7.47)		(4.49)			
Absolute pay, Berlin										
(7) No lag		-20.14	-8.505[c]	2.493	-358.5[c]	0.072		.92	0.67	.23
		(-1.87)	(-3.24)	(0.23)	(-7.45)	(0.08)				
(8) Lag		-10.93	-2.323	7.175	-292.6[c]	-1.398[b]	0.585[c]	.96	1.40	.41
		(-1.41)	(-1.07)	(0.95)	(-8.17)	(-2.0)	(5.16)			

[a] Equals $(e - \frac{\beta_3}{1+\beta_3} i) \div e$, where e is the mean total enlistment rate (volunteer and draft motivated), 15.258, i is the induction rate, 6.26, and β_3 is the estimated regression coefficient of $\ln(1-A/P)$.

[b] t-statistic > |2.01|

[c] t-statistic > |3.01|

(1)-(4) are expressed in terms of the nominal value of W_M.

The consistently negative coefficient for W_M may be due in part to the lack of statistical variation in W_M over Fisher's sample period. It is possible that the results with respect to W_M will be sensitive to a longer sampling period over which there will be enough statistical variation in W_M to produce coefficients with the theoretically expected signs.

The Berlin dummy is statistically significant in only one of the four estimating equations in which it appears, suggesting that it was not an important independent enlistment factor during this period.[6] Not surprisingly, therefore, it had little impact on the parameters of the other enlistment variables.

The coefficient of E/P lagged one quarter was statistically significant with the theoretically expected positive value between zero and one in all four estimating equations in which it appears. The magnitude of the coefficients suggests an adjustment lag that ranges from four to seven quarters. The coefficient of E/P lagged on quarter is particularly sensitive to the presence of serial correlation and can be a biased estimator of the adjustment process on this account. Since the estimating equations that assume instantaneous adjustment display significant positive serial correlation, these results should be treated with some caution. Introduction of the lagged enlistment rate also reduces the estimated effects of the pay variables to the point where they are no longer statistically significant and raises the estimated long-run effect of inductions on enlistments and the proportion of enlistees who were draft-motivated that is estimated from this long-run effect from a range that varies between 0.17 and 0.23 to a range that varies between 0.31 and 0.41.

The most satisfying results from a theoretical point of view are produced by the model using the original Fisher assumptions. This model produces statistically significant coefficients with the theoretically expected signs for the relative pay and the accession rate variables. The assumption of symmetry is not supported by the findings of equations (5)-(8); moreover, the unsatisfactory findings with respect to the military pay variable constrain us to deriving implications about the effects of military pay variable from the coefficient of the civilian pay variables -- which is equivalent to the assumption of symmetry that we are trying to relax in this experiment -- in which case, the equations embodying the original Fisher assumption would seem more appealing.

The evidence from the Berlin dummy does not strongly support any shift in the distribution of tastes for military service during the Berlin crisis. And the addition of the variable E/P, lagged one quarter, while increasing the explanatory power of the estimating equations and reducing the degree of positive serial correlation produced by these equations, also reduces the estimated coefficients of the pay variables to the point where they are no longer statistically significant.

The most satisfying results from the statistical point of view are those produced by the assumption of a lagged supply adjustment. These models produce uniformly higher R-squares and uniformly lower degrees of serial correlation.

None of the results described above alter two of the fundamental findings of the original Fisher study: pay elasticities that are less than one and statistically insignificant employment rate effects. However, the assumption of a lagged supply adjustment raises the estimated proportion of enlistees who are draft-motivated from around .20 to around .35. The latter estimates accord more closely to the findings based on survey data.

Next, Fisher's findings are benchmarked to those of the uniform data base. Table 10 compares the regression coefficients derived using the Fisher estimating equation and the Fisher data base over two sample periods: the Fisher sample period (3:57-4:65) and the sample period of the common base (2:58-4:65). Dropping 1957 from the sample period raises the relative pay elasticity from -0.68 to -0.99, an increase of almost 50 percent. Otherwise, dropping these observations has little effect on the findings.

Table 4 also summarizes the results of experiments with the original Fisher data base in which the original Fisher enlistment function is estimated using alternative nonlinear functional forms. The relative pay elasticity is also slightly sensitive to the choice of functional form. Fisher's semi-log function produces the highest elasticity; estimates of relative pay elasticity derived from the log linear and logit functions are -0.75 and -0.70, respectively, roughly 15-20 percent below the -0.99 estimate generated from the semi-log function. The estimated proportion of enlistees who are draft-motivated is also sensitive to choice of functional form. Estimates generated by the log-linear and logit functions range around 0.14, roughly 25 percent lower than the 0.19 estimated from the semi-log function.

The semi-log function produces a slightly larger R^2 and a slightly lower Durbin-Watson statistic, indicating more positive serial correlation than the log-linear and logit functions.

The findings from the Fisher data base may be summarized as follows:

1. Relatively inelastic pay elasticities which are somewhat sensitive to sample period and functional forms.

2. Relatively low estimates of draft-motivated enlistments which are quite sensitive to sample period and functional form.

Table 10

EFFECTS OF DROPPING 1957 AND USING ALTERNATIVE FUNCTIONAL FORM ON THE ESTIMATED PARAMETERS OF THE FISHER MODEL, 1:58-4:65

Equation	Regression Coefficient (t-statistic in parentheses)			Elasticity of $\ln(W_M/W_C)$	Proportion Draft Motivated[a]	R^2	D-Statistic
	$\ln(W_C/W_M)_{t-1}$	$\ln(1-U)_{t-1}$	$\ln(1-A/P)_t$				
(1) Original Fisher results	-7.057[b] (-2.17)	-9.014 (-0.89)	-312.7[c] (-7.54)	-.68	.19	.90	0.28
(2)-(1) less 4:57, 1:58	-10.437[b] (-2.83	2.698 (0.23)	-314.9[c] (-7.43)	-.99	.19	.91	0.35
(3)-(2), log-linear	-.750[b] (-2.81)	.179 (0.21)	-21.55[c] (-7.03)	-.75	.14	.88	0.51
(4)-(2), logit	-.639[b] (-2.50)	-.213 (-0.26)	-21.0[c] (-6.86)	-.70[d]	.33[e]	.88	0.50

[a]See Table 3, note a.

[b]t-statistic > |2.0|

[c]t-statistic > |3.0|

[d]Elasticity = $\frac{\hat{\beta}_K}{(1-e)}$ 3, where $\hat{\beta}_K$ is the estimated regression coefficient of $\ln(W_M/W_C)$ and e is the enlistment rate, 15.39.

[e]Proportion draft motivated = $\hat{\beta}_D \cdot \frac{1-\hat{e}}{1-\hat{i}} \cdot \frac{\hat{i}}{1-\hat{e}-\hat{i}}$, where $\hat{\beta}_D$ is the regression coefficient of $\ln(1-A/P)$, $\hat{e}$ is the mean enlistment rate, 15.39, and $\hat{i}$ is the mean induction rate, 6.26.

3. No evidence of a statistically significant employment rate effect.

Results with Uniform Data Base

The Fisher model parameters are reestimated now using the quarterly time series data base compiled at GRC for this study. The sensitivity of these parameters is investigated for variation in the size of the sample period, functional form, model assumptions, and estimates of variables. The particular sample periods that are examined include the Kim et al. sample (3:58-4:65), the Fechter sample (2:58-3:68), and the Cook sample (1:59-2:67). There are three additional time series studies: one by Withers, a quarterly time series for the years 1966-1973, one by Grissmer and others using monthly time series data for the period 1970-1973, and one by Grissmer using monthly time series data for the period 1970-1975. The latter two time series studies are unique in that they estimate enlistment functions for voluntary enlistees only. Separate enlistment functions are estimated for DOD, Army and Air Force enlistments in order to compare findings from Fisher's model with those of Fechter and Cook.

The theory of occupational choice used to generate our enlistment supply functions suggests a non-linear functional form. A number of functional forms (including a linear function) have been used to estimate enlistment supply equations. As part of this investigation, analyses are made with the following types of functional form: log linear, semi-log, and logit. We have seen that the choice of functional form has some effect on the estimated proportion of enlistees who are draft-motivated. Different functional forms may also produce different results when it comes to forecasting accuracy -- particularly in cases where the data in the forecasting period go beyond the range of the data used to estimate the original enlistment equation.

The most important factor to consider in extending the Fisher sample period is probably the effects of the Viet Nam war. Earlier studies produced inconclusive evidence about the effect of this war on enlistment behavior. Fechter found no effect on Army enlistments when he measured the war in terms of casualties; Cook found a significant negative effect on Air Force enlistments when he measured the war as a dummy variable.

The impact of the war on the Fisher results is evaluated in two ways: (1) by including a Viet Nam war dummy variable in equations estimated over the larger sample period; and (2) by testing for instability in the estimated enlistment supply parameters between the period encompassing the war and the remainder of the sample. (Parameter instability will be evaluated on the basis of the Chow test.)

Although the models that do not assume symmetric pay response and instantaneous supply adjustment produce disappointing results when fit to the original Fisher data base, better results may be obtained when a longer sample period is used. Therefore, these assumptions are tested again using the new data base. In addition to the three assumptions tested earlier, the assumption of no excess enlistment supply is also examined by following Cook's model and including an enlistment quality variable as an additional argument to capture the potential effect of "creaming."

Finally, the effects of alternative methods of estimating draft pressure, employment rates, and military pay on the estimated enlistment supply parameters was investigated. Recall that Fisher estimated draft pressure in terms of the accession rate (1-A/P), a method that tends to produce an upward bias in the estimated draft pressure coefficient because of the inclusion of E/P as a sizable fraction of A/P.

An alternative specification would cast the draft pressure variable in terms of the induction rate. (Klotz makes this argument in his comment on Fisher's study.) Unfortunately, this formulation may bias the magnitude of the draft pressure variable downward because of the possible simultaneous determination of E/P and I/P.

Fisher's peculiar treatment of military and civilian pay raises interesting questions about model assumptions with respect to these variables. The military is a unique occupation in that it pays a considerable amount of its wages in-kind and in the form of medical services, food and lodging. Fisher values these in-kind benefits on the basis of allowances awarded for the purchase of food and quarters by those not living on military posts and as a flat $253 per year for medical services. By refusing to deflate W_M he is implicitly (and correctly) arguing that, on the one hand, in-kind benefits ought to be inflated to account for their rising value to the enlistee while, on the other hand, cash benefits ought to be deflated to account for the erosion of their purchasing power. He estimates that roughly half of his estimated W_M is cash and half is in-kind. Therefore, the need to inflate the value of in-kind benefits is just about offset by the need to deflate the cash benefits, and the use of nominal (rather than real) military dollars is justified. His estimate of W_C, on the other hand, is in real dollars.

The issue is how should one treat military pay. Fisher has implicitly argued that, while one should include both cash and in-kind pay, one should use different methods of adjusting them for changes in purchasing power. Cook implicitly argued that only cash pay mattered and that potential enlistees evaluated in-kind benefits at a constant value over his sample period. (This value could, in principle, be zero.) Fechter implicitly argued that both cash and in-kind pay were important, and that both forms of pay should be treated similarly. The results of experiments with Fisher's original model (discussed earlier) indicate that the pay elasticities are sensitive to how one treats military pay. Recall that the pay elasticities estimated from the Fisher estimating equation in which both military and civilian pay are treated similarly were roughly one-third lower than the pay elasticity estimated using nominal values of W_M and real value of W_C.

Estimates of enlistment supply parameters derived from a given estimating equation over a common sample period using Fisher's original data base are first compared with those obtained using the data base prepared at GRC for this study. Fisher's data base consists of six variables -- enlistments, accessions, unemployment rates, population, military pay, and civilian pay. Detailed comparisons between Fisher's estimates and estimates generated for this study indicate that they approximate each other quite well. The relative difference between them is rarely more than 3 percent. (For details of this comparison, see Appendix B.) This suggests that there should be little difference between the estimated enlistment supply parameters generated by the original Fisher data base and parameters generated by the new data base. Table 11 compares these estimates. Qualitatively, the results are similar. The relative pay and the accession rate parameters are both negative and statistically significant, while the unemployment rate parameter is positive, but statistically insignificant. However, the point estimate of the relative pay elasticity derived from the new data base is -0.85, or roughly 15 percent lower in absolute terms than the -0.99 generated by the original Fisher data base.

Table 11

COMPARISON OF RESULTS FROM FISHER ESTIMATING EQUATION, USING ALTERNATIVE DATA BASES, ORIGINAL FISHER DATA, VS. NEW DATA FROM COMMON DATA BASE

Equation	Regression Coefficient (t-statistic in parentheses)			Elasticity of $\ln(W_M/W_C)$	Proportion Draft Motivated[a]	R^2	D-Statistic
	$\ln(W_C/W_M)_{t-1}$	$\ln(1-U)_{t-1}$	$\ln(1-A/P)_t$				
(1) Original Fisher data, 2:58–4:65	-10.437[b]	2.698	-314.9[c]	-.99	.19	.91	0.35
	(-2.82)	(0.22)	(-7.42)				
(2) New data, 2:58–4:65	-8.768[b]	8.742	-331.1[c]	-.85	.20	.89	0.38
	(-2.26)	(0.65)	(-6.91)				

[a]See Table 3, note a.

[b]t-statistic > |2.0|

[c]t-statistic > |3.0|

Standardizing for Differences in Samples and Differences in Military Service

An investigation of the effects of differences in sample period and differences in the branch of service used as the unit of analysis on the performance of the Fisher estimating equation is made next. Standardizing for differences in performance due to differences in functional form used as the estimating equation is accomplished by using a log-linear estimating equation. Recall the experiments with functional form using the original Fisher data base produced slightly lower absolute values of the total enlistment elasticity of W_C/W_M and slightly smaller estimates of the proportion of Mental Group 3 enlistees who were draft-motivated. The sensitivity of the findings to our military involvement in Viet Nam is also investigated.

Effects of Viet Nam

The analysis of the impact of Viet Nam on enlistment behavior consisted of including (1) a dummy variable for the period of our military involvement as an independent argument in the Fisher estimating equation; and (2) testing the parameters of the original Fisher estimating equation (excluding the Viet Nam dummy) for structural differences between the Viet Nam and the non-Viet Nam periods of the sample.[7]

Table 12 summarizes the findings. Viet Nam did not appear to be an important factor in enlistment behavior for any of the services analysed and for any of the samples. The estimated parameter of the Viet Nam dummy is generally negative, implying that, other things equal, enlistments were lower during the Viet Nam era. However, the coefficient is statistically significant in only one of the 12 cases examined.[8] The Chow test for differences in enlistment parameters between the pre-Viet Nam sample and the Viet Nam sample rejects the null hypothesis of no difference in only one

Table 12

SUMMARY OF ESTIMATES OF IMPACT OF VIET NAM WAR ON ENLISTMENT BEHAVIOR, ARMY, AIR FORCE, AND DOD

Sample Period	Army		Air Force (Dummy Variable)		DOD	
	Coefficient	t-value	Coefficient	t-value	Coefficient	t-value
2:58-4:65 (Fisher)	-.016	-.1	-.174	-.7	.065	.6
3:58-4:65 (Kim, et al.)	-.015	-.1	-.171	-.7	.066	.6
2:58-3:68 (Fechter	-.046	-.7	-.415[a]	-2.4	-.137	-1.9
1:59-2:67 (Cook)	-.034	-.4	-.243	-1.5	-.100	-1.4
			Chow Test (F-Statistic)			
2:58-3:68 (7, 28)	.454		1.331		2.500[b]	
1:59-2:67 (7, 20)	.345		1.548		2.475	

[a] t-statistic > |2.0|

[b] Exceeds critical value at .05 level.

instance — for DOD enlistments in the sample period 2:58-3:68. Since the war grew increasingly unpopular with the passage of time, a larger Viet Nam sample period is likely to produce stronger findings of structural differences in enlistment parameters.

Effects on Pay Elasticities

Table 13 summarizes the pay elasticities estimated for each of the services using alternative sample periods. The estimating equation for the Army gives the best results. The relative pay parameter is consistently negative, as predicted by the model, and statistically significant, with t-values in excess of 4.0. Point estimates of the relative pay elasticity range between 1.18 and 1.35. In contrast, the estimating equations for DOD are more disappointing. While the DOD estimating equation gives consistently negative relative pay parameters, they are not as stable as the Army elasticities. They are statistically significant in three of the four sample periods, but with t-values ranging around 2.0. Point estimates of the relative pay elasticity are considerably lower than those estimated for the Army, ranging between 0.19 and 0.71. The Air Force results are most disappointing. The relative pay parameter is consistently positive, contrary to theoretical expectations, and significantly positive in one of the samples.

Based on the results described above, we conclude that the relative pay parameters of the enlistment supply function are very sensitive to the particular service for which they are estimated. Moreover, while there are other methodological differences between Fechter and Fisher, it would appear that choice of service (i.e., DOD vs. Army) explains a large fraction of the difference in their relative pay findings.[9] On the other

Table 13

ESTIMATES OF RELATIVE PAY ELASTICITIES DERIVED FROM
FISHER ENLISTMENT MODEL, ARMY, AIR FORCE, DOD

Sample Period	Army		Air Force		DOD	
	Coefficient	t-value	Coefficient	t-value	Coefficient	t-value
2:58–4:65	-1.32[a]	-4.3	.51	.8	-.67[b]	-2.4
3:58–4:65	-1.35[a]	-4.5	.42	.7	-.71[b]	-2.6
2:58–3:68	-1.18[a]	-4.6	1.29	1.9	-.19	-.7
1:59–2:67	-1.29[a]	-4.5	.28	.5	-.52	-1.9

[a] t-statistic > 3.0

[b] t-statistic > 2.0

hand, adjustment for type of service (i.e., DOD vs. Air Force) widens the difference between Fisher and Cook in their relative pay findings.[10] This suggests that the difference between Fisher and Cook in their relative pay findings arises from other methodological differences between the two studies.

A likely reason for the unsatisfactory performance of the Fisher model on Air Force data is the assumption in Fisher's enlistment model of no excess supply of enlistees. This assumption is probably least tenable for the Air Force. Recall that Cook's enlistment model differs from Fisher's in that it makes explicit provision for the effect of rationing on enlistments.

Another possible reason for the difference between the Fisher and the Cook findings is the difference in their methods of estimating military pay. Cook confines himself to base pay only and completely ignores military pay received in kind, whereas Fisher includes in-kind pay in his estimate of military pay. Cook also forms his pay variables from a four-year stream of income, whereas Fisher uses a three-year stream. Since Air Force enlistment contracts are usually four-year contracts, Cook's method would be more appropriate for Air Force enlistments. In addition, Cook's formulation of civilian pay accounts for changes in the age composition of enlistments. Finally, his estimates of both military and civilian pay account for expectations about future levels of pay. Fisher's formulation ignores both age and expectations.

The causes of the difference among services in estimated relative pay elasticities are further illuminated by examining the results of absolute pay models estimated for these services over the same time periods.

Military and civilian pay elasticities estimated from the Fisher model are summarized in Table 14. The military pay elasticity is consistently negative, contrary to theoretical expectations, for the Air Force and DOD; it is also negative in the Army sample for the pre-Viet Nam period. The civilian pay elasticity is consistently negative, as theoretically expected, for the Army; but it is consistently and significantly positive, contrary to theoretical expectations, for the Air Force. It is negative, never statistically significant, in two of the three sample periods for DOD.

These results reveal that the major determinants of the relative pay results are a relatively well-behaved civilian pay variable and a poorly performing military pay variable. These findings support the speculation that the method used by Fisher to formulate his pay variables may have been inadequate. This speculation is reinforced by the fact (discussed in more detail in later sections of this chapter) that both Fechter and Cook were able to derive well-behaved and statistically significant military and civilian pay elasticities from absolute pay models which they estimated for the Army and the Air Force.[11]

<u>Effects on Employment Rate Elasticities</u>

Table 15, Panel A, summarizes estimates of employment rate elasticities by service for the sample periods of Fisher, Kim et al., Fechter, and Cook. The results stand in marked contrast to the pay results. Unlike the pay elasticities, the employment rate elasticities are about the same across services, given the sample period and vary considerably across sample periods for a given service.

The Fechter and the Cook samples generally produce satisfactory results. The elasticities have the correct sign in all the equations, with

Table 14

ESTIMATES OF ABSOLUTE PAY ELASTICITIES DERIVED FROM FISHER ENLISTMENT MODEL, ARMY, AIR FORCE, DOD

Sample Period	Army				Air Force				DOD			
	Military Pay		Civilian Pay		Military Pay		Civilian Pay		Military Pay		Civilian Pay	
	Coef.	t-value	Coef.	t-value	Coef.	t-value	Coef.	t-value	Coef.	t-value	Coef.	t-value
2:58-4:65	-2.24	-1.7	-.55	-1.4	-10.36[a]	-4.0	2.65[a]	3.5	-2.64[a]	-2.2	.05	.1
2:58-3:68	.61	1.3	-1.20[a]	-4.8	-5.34[a]	-5.1	1.15[a]	2.1	-1.45[a]	-3.5	-.25	-1.1
1:59-2:67	.46	.7	-1.24[a]	-4.3	-3.42[a]	-2.8	.49	.9	-1.29[a]	-2.5	-.40	-1.8

[a] t-statistic > |2.0|

absolute t-values that exceed 2.0 in four of the six equations. Moreover, the absolute value of the elasticity is greater than 2.0 in four of the six equations and is greater than 3.0 in one of these equations.

The Fisher and the Kim et al. samples consistently produce unsatisfactory results. The elasticity has the "wrong" sign in three of the six equations and the absolute t-value never exceeds 0.6.

The results reflect the pattern of variation in unemployment rates during the period 2:58-3:68. These rates ranged between 15 and 20 percent from 2:58-2:65. They fell abruptly thereafter, ranging between 9 and 12 percent from 3:65-3:68. While much of this decline can be attributed to a surge in economic activity in the late 1960's that reduced overall unemployment rates to rates below 4 percent, some of the decline could be attributed to our military involvement in Viet Nam. Quarterly military accessions, while never exceeding an annual rate of 500,000 prior to 1965, jumped to an annual rate that ranged between 750,000 and 1 million from 3:65 to 4:66. The annual rate ranged between 500,000 and 1 million in 1967 and 1968. The effect on enlistments of this increase in demand would be most pronounced in the Fechter and the Cook sample periods.

This dramatic increase in the demand for military manpower pulled substantial numbers of young men out of the civilian labor force and could have had an impact on unemployment rates that was independent of the impact of the expansion in business activity during this period. It is likely, however, that the effect of the increase in demand for military manpower was most pronounced on the unemployment rates of males in the 20-24 year old age group, the group that would be most affected by the draft.

In view of this, it is possible that the negative employment rate elasticity in the Fechter and the Cook samples is merely reflecting the negative enlistment effect of the war. The sensitivity of the estimated employment rate elasticity to the presence of a Viet Nam dummy variable in the estimating equation on the grounds that the employment rate variable might be picking up some of the enlistment effects of the war is examined next. The results are summarized in Panel B of Table 15. The effect of the Viet Nam dummy is to reduce the employment rate elasticities, especially for the Fechter and the Cook samples. In these samples, the elasticities continue to have the "correct" signs in five of the six equations, but the absolute t-values no longer exceed 2.0 in any of the equations and exceed 1.8 in only one equation.

Effects on Estimated Proportion Draft-Motivated

Estimates of the proportion of enlistees in Mental Categories 1-3 who are draft-motivated for the Army, Air Force, and DOD in each of the sample periods are summarized in Table 16. The estimates are much lower than comparable estimates reported in Table 11. The reason for the discrepancy appears to be that Fisher used an induction rate that was substantially larger than the average to compute his estimate. Fisher does not report the value of the induction rate he used, but his estimate of the voluntary enlistment rate is consistent with an induction rate of approximately 0.0147. The mean induction rate for these quarters is 0.0137, calculated from the common data base compiled for this study. The mean induction rate for the period 2:58-4:65 calculated from the common data base is 0.00626.

The estimates do not vary much across services for a given sample period. DOD generally has the largest proportion of draft-motivated enlistees and the Air Force usually has the lowest proportion. A notable

Table 15

ESTIMATES OF EMPLOYMENT RATE ELASTICITIES DERIVED FROM FISHER ENLISTMENT MODEL FOR ARMY, AIR FORCE AND DOD ENLISTMENTS

Sample Period	Army		Air Force		DOD	
	Coefficient	t-value	Coefficient	t-value	Coefficient	t-value
Panel A (No Vietnam)						
2:58-4:65	-.26	-.2	.70	.3	.58	.6
3:58-4:65	-.64	-.6	-.42	-.2	.15	.2
2:58-3:68	-.89	-1.6	-2.13	-1.4	-1.34[a]	-2.2
1:59-2:67	-2.01[a]	-2.5	-3.4[a]	-2.0	-2.05[a]	-2.7
Panel B (Vietnam)						
2:58-4:65	-.28	-.3	.48	.2	.66	.7
3:58-4:65	-.66	-.6	-.63	-.3	.23	.2
2:58-3:68	-.55	-.7	.96	.5	-.31	-.4
1:59-2:67	-1.76	-1.8	-1.44	-.7	-1.32	-1.5

[a] t-statistic > |2.0|

Table 16

ESTIMATES OF THE PROPORTION OF ENLISTEES WHO ARE DRAFT MOTIVATED[a] DERIVED FROM THE FISHER ENLISTMENT MODEL FOR ARMY, AIR FORCE AND DOD ENLISTMENTS

Sample Period	Army			Air Force			DOD		
	Coef.	t-value	Proportion	Coef.	t-value	Proportion	Coef.	t-value	Proportion
2:58-4:65	-19.9[b]	-5.3	.13	-14.7	-1.8	.09	-22.7[b]	-6.6	.15
3:58-4:65	-20.9[b]	-5.5	.13	-17.7[c]	-2.2	.11	-23.9[b]	-7.0	.15
2:58-3:68	-18.2[b]	-7.3	.15	-14.8[c]	-2.2	.12	-19.8[b]	-7.4	.16
1:59-2:67	-21.0[b]	-7.0	.16	-27.4[b]	-4.5	.20	-24.5[b]	-8.6	.18

[a]Estimates of proportion of enlistees who are draft motivated (DM) are derived from the log-linear estimating equation using the following formula:

$$DM = \hat{\beta}_{(1-a)} \cdot \frac{\hat{i}}{1-\hat{a}},$$

where: $\hat{\beta}_{(1-a)}$ = regression coefficient of $\ell n(1-A/P)$; $\hat{i}$ = inductions per male civilian noninstitutional population age 17-20; $\hat{a}$ = DOD accessions per male civilian noninstitutional population, age 17-20.

Estimates of the mean values of $\hat{i}$ and $1-\hat{a}$ for the different sample periods are given below:

	$\hat{i}$	$(1-\hat{a})$
2:58-4:65	.00626	.9771
3:58-4:65	.00615	.9771
2:58-3:68	.00804	.9742
1:59-2:67	.00729	.9754

[b]t-statistic > |3.0|

[c]t-statistic > |2.0|

exception is the Cook sample period, during which the Air Force had the largest prooportion. A possible reason for the pattern observed in Table 16 is spurious correlation. Enlistments in Mental Categories 1-3 constitute a part of total accessions, the variable used by Fisher to index draft pressure. Consequently, there should be spurious negative correlation between the enlistment rate and the complement of the accession rate tending to bias the regression coefficient toward -1. Moreover, the larger enlistments are as a fraction of accessions, the more serious this spurious correlation will be. Thus, one would expect to find more negative spurious correlation for DOD enlistments than for specific service enlistments. The findings in Table 16 are consistent with the presence of such spurious correlation.

There also does not appear to be much variability in the estimates over sample periods for a given service. The results from the Fechter and the Cook sample periods consistently produce larger estimates that the Fisher or the Kim et al. samples. This is especially notable for the Air Force. There are a number of possible reasons for this finding. First, as indicated in the footnote to Table 16, induction rates were higher during the Fechter and the Cook sample periods. Indeed, if one controlled for differences in induction rates, one would find practically no differences among sample periods for the Army and for DOD. Second, there was relatively less variability in the accession rate statistic during the Fisher and the Kim et al. sample periods. Lack of variability could have had the effect of biasing the estimated enlistment effect toward zero.

The results are not generally sensitive to inclusion of a Viet Nam variable. Including a Viet Nam variable raises the estimated proportion slightly for the Army and for DOD; it has a more substantial effect on the estimated proportion for the Air Force. The results for the Air Force are summarized below:

Sample Period	No Viet Nam	Viet Nam
2:58-4:65	.09	.12
3:58-4:65	.11	.14
2:58-3:68	.12	.22
1:59-2:67	.20	.25

The results in Table 10 suggest that the draft had a relatively small impact on enlistments much smaller than the impact suggested by the surveys. A number of reasons can be offered to explain this finding. Among them is the possibility that the parameters of the estimating equation are capturing short-run effects. The existence of lags in enlistment response would mean that long-run effects would be larger. The findings reported above in Table 3 for estimating equations which assume a distributed lag partial adjustment are consistent with this explanation. Estimates of the long-run enlistment impact of the draft are roughly twice the estimates of the short-run impact. An additional reason for expecting low estimates of the effects of the draft on enlistments is the possibility of simultaneous equations bias. This possibility was discussed earlier in Chapter II and was considered to be a highly unlikely prospect. A final reason, suggested by Klotz, is that the seasonal dummy variable may be picking up some of the effect of the accession rate variable (which has a strong seasonal component).

Effects of Differences in Estimates of Draft Probability

Recall that there are differences between Fisher and others in the method used to estimate the effects of the draft. Fisher assumes that the appropriate theoretical variable is the probability of remaining a civilian, p_c. His model has p_c operate through the expected pecuniary return to not enlisting:

$$W'_C = p_c W_C + (1 - p_c) W_M$$

where W_C is the expected pecuniary return to enlisting in the absence of a draft and W_M is expected first-term military pay. Fisher approximates p_c by using the complement of the military accession rate, (1-A/P). As we have noted, this estimate should tend to bias the regression coefficient of (1-A/P) toward -1 because the enlistment rate is part of the accession rate. Moreover, this bias should be more pronounced in circumstances where E/P represents a large fraction of A/P, such as in the case of DOD accessions.

Other studies have approximated p_c by using the induction rate I/P directly. As we have noted, this could bias the regression coefficient toward zero if inductions are not truly exogenous. Toikka has demonstrated that these two methods of measuring draft pressure represent two alternative assumptions about the length of the military planning horizon.[12] Using (1-A/P) assumes that military planners target their inductions to meet accession requirements over a one period planning horizon; using (1-I/P) assumes that military planners target their inductions to meet accession requirements over a very long planning horizon.

GRC investigates the effects of alternative methods of approximating p_c by substituting (1-I/P) for (1-A/P) in Fisher's estimating equation. Table 17 summarizes these findings. Panel A describes the regression coefficients and t-values derived from the modified estimating equation;

Table 17

EFFECTS OF USING (1-A/P) AS A MEASURE OF DRAFT PRESSURE ON ESTIMATES OF PROPORTION OF ENLISTEES WHO ARE DRAFT MOTIVATED[a] AND ON RELATIVE PAY ELASTICITIES FOR ARMY, AIR FORCE, AND DOD ENLISTMENTS

Sample Period	Army			Air Force			DOD		
	Coefficient		t-value	Coefficient		t-value	Coefficient		t-value
Panel A									
2:58-4:65	-19.7[b]		-2.7	-2.1		-.2	-23.5[c]		-3.2
2:58-3:68	-22.7[c]		-4.8	-6.1		-.6	-21.5[c]		-3.9
1:59-2:67	-28.5[c]		-4.5	-26.7[b]		-2.1	-30.2[c]		-4.2
Panel B	Proportion Draft Motivated								
		(1-I/P)			(1-I/P)			(1-I/P)	
	(1-A/P)	No VN	VN	(1-A/P)	No VN	VN	(1-A/P)	No VN	VN
2:58-4:65	.13	.12	.09	.09	.01	d	.15	.15	.09
2:59-3:68	.15	.18	.17	.12	.05	.11	.16	.16	.17
1:59-2:67	.16	.16	.19	.20	.20	.20	.18	.18	.21
Panel C	Relative Pay Elasticities								
2:58-4:65	-1.32[c]	-1.39[c]	-1.71[c]	.51	.63	.36	-.67[b]	-.76	-1.26
2:58-3:68	-1.18[c]	-1.32[c]	-1.34[b]	1.29	1.17	1.22	-.19	-.34	-.35
1:59-2:67	-1.29[c]	-1.47[c]	-1.50[c]	.28	.16	.16	-.52	-.69	-.72

[a] Estimates of the proportion of enlistees who are draft motivated (DM) are derived from the two log-linear estimating equations using alternative methods of estimating p_c using the formula described below:

For (1-A/P): $DM_A = \hat{\beta}_{(1-a)} \cdot \frac{\hat{i}}{1-\hat{a}}$ for (1-I/P): $DM_I = \hat{\beta}_{(1-i)} \cdot \frac{\hat{i}}{1-\hat{i}}$,

where $\hat{\beta}_{(1-a)}$, $\hat{i}$ and $\hat{a}$ are defined in Table 10, and $\hat{\beta}_{(1-i)}$ is the estimated regression coefficient of $\ell n(1-I/P)$.

[b] t-statistic > |2.0|

[c] t-statistic > |3.0|

[d] Regression coefficient had "wrong" sign.

Panel B compares estimates of the proportion of enlistees who are draft-motivated derived from the Fisher model to comparable estimates generated by the coefficients described in Panel A. Panel C compares estimates of relative pay elasticities derived from the estimating equations summarized in Panel B.

The t-values in Panel A are substantially lower in absolute value than the t-values reported in Table 16. In part, this reflects the removal of E/P from the estimating draft pressure variable in Table 11 and the consequent reduction in built-in spurious correlation.

Estimates of the proportion of enlistees who are draft-motivated are fairly similar to those found in Table 16 except for the Air Force, where use of (1-I/P) reduces this proportion to levels where they are no longer statistically significant for the Fisher and Fechter samples. The Air Force results may be reflecting the existence of quotas which are biasing the estimated regression coefficients toward zero. Panel B also indicates that the results are not particularly sensitive to inclusion of a Viet Nam dummy variable.

Panel C describes the sensitivity of the relative pay elasticities to methods of estimating draft pressure and to Viet Nam. Use of (1-I/P) raises the absolute value of the pay elasticities slightly for the Army and for DOD; it makes the elasticity less positive for the Air Force, which is moving it in the "right" direction. Adding the Viet Nam dummy variable has little impact on these elasticities except for the Fisher sample, where it raises the absolute values of the Army and DOD pay elasticities by 66 percent for DOD and by 23 percent for the Army. It is interesting to note that the combined effect of using I/P to estimate draft pressure and accounting for Viet Nam by means of a dummy variable is to raise Fisher's estimate of the DOD relative pay elasticity by almost 100 percent.

The Effects of Alternative Methods of Measuring Military Pay

In replicating Fisher's original findings, it was observed that Fisher's relative pay variable was estimated as the ratio of the real value of W_C to the nominal value of W_M. Further, the relative pay elasticity estimated using Fisher's relative pay variable was found to be roughly 25 percent lower in absolute terms than the relative pay elasticity estimated from a relative pay measure which treats W_C and W_M consistently. Since the time series studies that do not use the Fisher data base (i.e., Fechter, Cook, Withers, Grissmer et al., and Grissmer) use the latter measure of relative pay, we now report on findings derived from the Fisher estimating equation modified so that the latter relative pay measure is substituted for the Fisher relative pay measure.

Table 18 summarizes the effects of this substitution on estimated relative pay elasticities by service over differing sample periods for alternative measures of draft pressure and alternative assumptions about a Viet Nam effect.[13] The results are by and large similar to those reported in Table 13; the Army consistently shows the "best" results (i.e., coefficients that have the "correct" sign and t-values in excess of 3.0). The Air Force consistently displays the "worst" results (i.e., coefficients that have the "wrong" sign). And the DOD results, which reflect both the Army and the Air Force results, tend to fall between the Army results and the Air Force results.

Table 18

EFFECTS OF USING REAL W_M ON RELATIVE PAY ELASTICITIES FOR ALTERNATIVE MEASURES OF DRAFT PRESSURE AND ALTERNATIVE ASSUMPTIONS ABOUT A VIET NAM EFFECT FOR ARMY, AIR FORCE, AND DOD ENLISTMENTS

Sample Period	Army		Air Force		DOD	
	Nominal W_M	Real W_M	Nominal W_M	Real W_M	Nominal W_M	Real W_M
Panel A, A/P						
			No Viet Nam Effect			
2:58-4:65	-1.32[a]	-.96[a]	.51	.31	-.67[b]	-.50[b]
2:58-3:68	-1.18[a]	-.94[a]	1.29	.34	-.19	-.40
1:59-2:67	-1.29[a]	-1.03[a]	.28	-.02	-.52	-.56[b]
			Viet Nam Effect			
2:58-4:65	-1.29[a]	-.92[a]	.89	.51	-.82[b]	-.59[b]
2:58-3:68	-1.16[a]	-.96[a]	.79	.81	-.11	-.56
1:59-2:67	-1.26[a]	-1.06[a]	.51	.27	-.42	-.49[b]
Panel B, I/P						
			No Viet Nam Effect			
2:58-4:65	-1.39[a]	-.97[a]	.63	.36	-.76	-.60[b]
2:58-3:68	-1.32	-1.07[a]	1.17	.31	-.34	-.52
1:59-2:67	-1.47[a]	-1.22[a]	.16	-.14	-.69	-.74[b]

[a] t-statistic > |3.0|

[b] t-statistic > |2.0|

The models which measure draft pressure in terms of I/P and which assume a Viet Nam effect consistently produce the largest estimates of the relative pay elasticity (in absolute terms). Neither factor, i.e., either using I/P or accounting for Viet Nam, has an independent effect; it is their joint effect that increases the relative pay elasticities.

Again, the original Fechter finding of an Army relative pay elasticity of 0.87 may be compared to the Army relative pay elasticity of 1.16 reported in Table 18. Similarly, the original Cook finding of an Air Force relative pay elasticity of 2.23 may be compared to the -0.16 reported in Table 18. These comparisons reinforce the earlier conclusion that the Fisher model is probably not appropriate for explaining Air Force enlistment behavior because it fails to account for the existence of excess supply of applicants to the Air Force and because it is deficient in its method of estimating the military and civilian pay variables.

THE KLOTZ STUDY

Klotz modified Fisher's estimating equation (equation (1)) by dropping the seasonal dummy variables and reestimating the enlistment supply parameters. He argues that seasonal dummies do not belong in the estimating equation because they obscure the effects of the draft pressure surrogate, (1-A/P), which he claims has a strong seasonal component.[14] Although Klotz does not make it, a similar argument could have been made for the effects of the employment rate variable which, it can be argued, also has a strong seasonal component.[15]

Klotz claims to fit his modification of Fisher's estimating equation to Fisher's estimates of variables using Fisher's data base and reports that excluding the seasonal dummies raises the estimated impact of the draft pressure and the employment rate variables (Table 1). He further reports that dropping the seasonal dummies has no effect on the estimated relative pay coefficient. Finally, he reports that omitting the seasonal dummies reduces the explanatory power of the estimating equation (from an R^2 of 0.9 to an R^2 of 0.7) and eliminates serial correlation from the residuals. Table 19 summarizes these differences.

Klotz reports estimates of relative pay elasticities with respect to voluntary enlistments of -1.11 with seasonal dummies, and -1.44 excluding seasonal dummies.[16] The former estimate is based on the parameters of Fisher's estimating equation (see Table 7), but is considerably higher than the estimate of -0.74 reported by Fisher.[17] This is because Klotz estimates the adjustment factor for the draft differently than Fisher. Recall that Fisher derives his adjustment factor as $(\frac{\beta_3}{1+\beta_3})$, where β_3 is the coefficient of the draft surrogate, $\ln(1-A/P)$. This adjustment factor can be interpreted as the marginal effect of inductions on enlistments.[18] Using the parameters reported in Table 7, Klotz reports a value of .4394.[19] To derive voluntary enlistments, one subtracts $.4394 \cdot I/P$, or .0065, from the total enlistment rate of 0.015 reported by Klotz, giving a voluntary enlistment rate of .0084.[20] Dividing this voluntary rate into the structural estimate of the relative pay variable reported by Klotz, -0.01, produces an estimated elasticity of 1.13; Klotz reports an estimated elasticity of 1.11. The difference between GRC estimates and Klotz's appears to be due to rounding. As noted earlier, the difference between Fisher and Klotz in their reported estimates of relative pay elasticities is due to Fisher's erroneous method of deriving his elasticity for voluntary enlistments. Although he does not state his method explicitly, we

Table 19

COMPARISON OF ESTIMATED ENLISTMENT SUPPLY PARAMETERS
REPORTED BY FISHER AND KLOTZ

	Estimated Coefficient (t-statistic in parentheses)	
	Fisher	Klotz
$\ell n(W_C/W_M)_{t-1}$	-7.09	-7.0
	(-2.19)	(-1.4)
$\ell n(1-A/P)_t$	-312.	-444.
	(-7.61)	(-7.40)
$\ell n(1-U)_{t-1}$	-8.91	-17.
	(-.87)	(-1.89)
R^2	.90	.70
Durbin-Watson Statistic	.31	2.14

see that he estimated his voluntary enlistment elasticity by simply dividing his total enlistment elasticity, -0.46, by .62, the estimated proportion of enlistees who are true volunteers derived from a 1964 DOD survey.[21] In addition, he should have adjusted his reduced form coefficient of $\ell n(W_C/W_M)_{t-1}$ by the factor $\frac{1}{1+\beta_3}$ to derive an estimate of the structural coefficient of $\ell n(W_C/W_M)_{t-1}$.

To summarize Klotz's empirical contribution, by excluding seasonal variables from his estimating equation, Klotz allows all seasonal factors to be captured by the included variables with large seasonal components, $\ell n(1-A/P)_t$ and $\ell n(1-U)_{t-1}$. The result is a dramatic increase in the absolute size and statistical significance of the reduced form coefficients of these variables. Moreover, the larger absolute value of the reduced form parameters of $\ell n(1-A/P)_t$ results in further increases in the voluntary enlistment rate elasticities of (W_C/W_M) and $(1-U)$ that are independent of any changes that might have been produced in their reduced form coefficients.

Results with Original Data Base

The Klotz experiment is replicated using the same data base used in replicating Fisher's results. Since Fisher's reported findings could be duplicated with this data base, there should be no reason why one should not be able to duplicate Klotz's findings. Unfortunately, simply dropping the seasonal dummies from the estimating equation summarized in Table 7 as FISH B does not produce the same results reported by Klotz. Therefore, an experiment with alternative samples allows one to come closest to reproducing Klotz's results when using the sample period 2:58-4:65. Results of these experiments are summarized in Table 20 as KLOTZ A and

Table 20

COMPARISON OF ENLISTMENT SUPPLY PARAMETERS DERIVED FROM KLOTZ MODEL USING ORIGINAL DATA BASE AND ALTERNATIVE SAMPLE PERIODS

	Estimated Coefficient (t-statistic in parentheses)			
	Klotz (AER)[a]	KLOTZ A[b]	KLOTZ B[c]	KLOTZ C[d]
$n(W_C/W_M)_{t-1}$	-7.0	-4.996	-7.054	-5.332
	(-1.4)	(-0.9)	(-1.2)	(-1.2)
$n(1-A/P)_t$	-444.[e]	-452.[e]	-450.[e]	-449.[e]
	(-7.4)	(-7.7)	(-7.3)	(-7.3)
$n(1-U)_{t-1}$	-17.	-22.9	-17.2	-16.3
	(-1.9)	(-1.9)	(-1.2)	(-1.1)
R^2	.70	.69	.68	.68
Durbin-Watson Statistic	2.14	2.05	2.05	2.08

[a]Reported in Klotz, op. cit., p. 971.

[b]Computed from the sample 4:57-4:65 using Fisher's estimate of W_C/W_M.

[c]Same as KLOTZ A, only sample is 2:58-4:65.

[d]Same as KLOTZ B, only W_C/W_M is estimated using nominal values for both W_C and W_M.

[e]t-statistic > $|3.0|$

KLOTZ B. This table also summarizes the results of an experiment in which the seasonals are dropped from the results summarized in Table 7 as FISH C (i.e., the estimating equation in which both W_M and W_C are measured in nominal terms). These results are described as KLOTZ C.

Qualitatively, the results are about the same as the reported results. Dropping the seasonal dummies increases the absolute values of the coefficients of $\ell n(1\text{-}A/P)_t$ and $\ell n(1\text{-}U)_{t-1}$. However, in KLOTZ B the coefficient of $\ell n(1\text{-}U)_{t-1}$ is one-third larger (in absolute value than that reported by Klotz, while the coefficient of $\ell n(W_C/W_M)_{t-1}$ is almost one-third lower (in absolute value). In KLOTZ C, the coefficients are about the same as those reported by Klotz, but the t-statistic for the coefficient of $\ell n(1\text{-}U)_{t-1}$ is -1.2 compared to the -1.9 reported by Klotz. As in FISHC, the coefficient of $\ell n(W_C/W_M)_{t-1}$ is notable lower in absolute terms when military and civilian pay are both measured in nominal terms. The coefficient falls from -7.054 in KLOTZ B to -5.332 in KLOTZ C. The explanatory power of the estimating equations hover near an R^2 of 0.7 and there is no significant serial correlation in the residuals.

These results alter the findings with respect to the effects of the draft and of (1-U) on enlistment behavior (Table 21). Estimates of the proportion of enlistees in Mental Groups 1-3 who were draft-motivated hover around one-third, much closer to the 39 percent derived from the 1964 DOD survey. Estimates of the elasticity of (1-U) both with and without a draft exceed unity by substantial margins. Finally, the smaller proportion of voluntary enlistees tends to further increase the voluntary supply elasticities of W_C/W_M and (1-U) relative to those derived from the Fisher estimating equation.

Table 21

COMPARISON OF ENLISTMENT SUPPLY ELASTICITIES WITH AND WITHOUT A DRAFT AND PROPORTION OF ENLISTEES IN MENTAL GROUPS 1-3 WHO WERE DRAFT MOTIVATED DERIVED FROM KLOTZ MODEL USING ORIGINAL DATA BASE AND ALTERNATIVE SAMPLE PERIODS

	Klotz (AER)	KLOTZ A	KLOTZ B	KLOTZ C
Elasticities (with draft)[a]				
W_C/W_M	-.82	-.60	-.83	-.63
1-U	-1.99	-2.73	-2.03	-1.92
Proportion draft motivated[b]	.33	.34	.33	.33
Elasticities (without draft)[c]				
W_C/W_M	-1.22	-.91	-1.24	-.94
1-U	-2.97	-4.14	-3.03	-2.87

[a]Estimated from the following formula:

$$\varepsilon_i = \frac{\beta_i}{1+\beta_3} \cdot \frac{1}{\bar{e}},$$

where: ε_i = total enlistment elasticity of the ith variable; β_i = regression coefficient of ith variable; β_3 = regression coefficient of $\ell n(1-A/P)_t$; $\bar{e}$ = average enlistment rate for the sample period.

[b]Estimated from the formula:

$$D = \frac{\beta_3}{1+\beta_3} \cdot \frac{\bar{I}}{\bar{e}},$$

where: $\bar{I}$ = average induction rate for the sample period; D = proportion of enlistees in Mental Groups 1-3 who are draft motivated.

[c]Estimated from the formula:

$$V_i = \frac{1}{1-D} \cdot \varepsilon_i$$

where: V_i = voluntary enlistment elasticity of ith variable.

While the magnitude of the parameters summarized in Table 21 are more in accord with a priori reasoning and with evidence derived from related sources, the larger degree of imprecision of the parameters of the Klotz estimating equation (signified by their low t-values) is disappointing and may be reflecting the lack of statistical variation in W_M/W_C and (1-U) discussed earlier in this chapter.

FOOTNOTES, CHAPTER IV

1. See Kim et al., pp. 196-197, for details of how this variable was estimated.

2. Ibid., pp. 187-198 for details of the estimating procedure.

3. Presumably, this includes inductees and enlistees in Mental Group 4.

4. Assumed to have the standard stochastic properties.

5. For elaboration of the adjustment model and how one dervies these estimates, see Fechter, op. cit., pp. 5-6.

6. The accession rate was unusually high during the period of the Berlin crisis because of unusually high draft calls. Thus, the major enlistment effect of this crisis appears to have been the increase in draft-motivated enlistments prompted by the higher draft calls.

7. The Viet Nam period is assumed to begin in 3:65, the quarter that the United States Congress passed the Tonkin Gulf resolution.

8. Results derived from the absolute pay models of enlistment behavior are somewhat different from those summarized in Table 6 for relative pay models. The estimated coefficient of the Viet Nam dummy is consistently positive, but never statistically significant.

9. Fechter's comparable Army relative pay elasticity with a draft was .87. However, Fechter's estimate is derived from a model comparable to FISH C in which both military and civilian pay are both estimated in real (or nominal) dollars. Recall that the relative pay elasticity estimated from FISH B was about 41.1 percent larger in absolute terms than the relative pay elasticity estimated from FISH C. Adjusting Fechter's relative pay elasticity to reflect this difference produces a comparable elasticity of 1.23.

10. Recall that Cook's Air Force relative pay elasticity was 2.23.

11. The Cook results for absolute pay models were not reported in the original Cook study. They were found in unpublished tabulations of regression results that were provided by Cook. The author is indebted to him for his generosity and assistance.

12. Richard S. Toikka, "A Note on Estimating the Fisher Model of Military Enlistments," unpublished manuscript, undated.

13. Similar experiments were performed using absolute pay models but the results were essentially the same as those summarized in Table 7; the military pay variable had a significantly negative coefficient, contrary to theoretical expectations, in all the equations that were fit.

14. The simple correlations between $\ell n(1-A/P)$ and the seasonal dummies for the period 4:57-4:65 were statistically significant.

15. The simple correlations between $\ell n(1-U)$ and the seasonal dummies for the period 4:57-4:65 were statistically significant.

16. Klotz, op. cit., p. 972.

17. Fisher, op. cit., p. 249.

18. See supra, equation 1'.

19. Klotz, op. cit., p. 972, especially footnote 6.

20. It is assumed in this exercise that I equals .0147.

21. Fisher, op. cit., p. 249.

Chapter V

POOLED TIME-SERIES CROSS-SECTIONAL ANALYSIS

by

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PURPOSE AND SCOPE OF THE ANALYSIS

The focus of this chapter is on combined time series and cross-sectional models for estimating the supply levels of Army and Navy volunteers. The strength of the models being tested rests in their ability to adequately account for the variance in enlistment patterns across the nine census regions and variance due to longitudinal changes in the enlistment data. The major purpose of this analysis, however, is to determine the rates of change in the supply parameters given an effective pooled time series cross-sectional model for different quality enlistment groups.

The models tested by regression analysis fall into two main categories: (1) strictly linear models and (2) log-linear models. The general form of the models is given below.

$$(1) \quad E_i = \chi_i\alpha+\epsilon_i \quad , i = 1, 2, \ldots, 9$$

where:

E_i is a 5 x 1 vector representing the number of enlistees in region i during the five years 1970-1974.

χ_i is a 5 x K matrix of K independent variables in the analysis pertaining to region i.

α is a K x 1 vector representing the coefficients of the independent variables.

ϵ_i is an error term (5 x 1) for the equation representing regional variation from the linear fit

It should be noted that the general form of the model assumes a test of the hypothesis of constant slopes and intercepts in the regressions. This means that with a single set of coefficients α supply parameter estimates will be made for each region. Hence regional variance exhibited by ε_i becomes extremely important in determining the adequacy of a model for use in regional forecasting. It is also important that the models tested contain as many as possible of the independent variables which may be significant variates in each region's enlistment pattern.

The set of variables used in the study are presented and explained in detail in Appendix C. Brief descriptions of these variables are given below.

o E/QMA = the ratio of 17-21 year old male volunteer enlistees to a branch of service to the number of 17-21 year old qualified military availables as the dependent variable.

o PAYR = the ratio of average military pay and benefits to average civilian income for 17-21 year old males working full-time the year round.

o UNEMP = the unemployment rate for 17-21 year old males who were in the full-time labor force.

o ADV = the amount of money spent in print media advertising and related advertising funds.

o REC/QMA = the ratio of recruiters assigned by a branch of service to the number of 17-21 year old qualified military availables.

o PROB = the proportion of 17-21 year old black military availables to the total number of 17-21 year old military availables.

The necessary tests of significance for all of these variables to be included in the original model were performed in studies performed prior to this analysis.[1] Equation (1) could therefore be written in the following manner for strictly linear models:

$$(2) \quad (E/QMA)_i = \alpha_0 + \alpha_1 PAYR_i + \alpha_2 UNEMP_i + \alpha_3 ADV_i + \alpha_4 (REC/QMA)_i + \alpha_5 PROB_i + \varepsilon_i , \qquad i = 1, 2, \ldots, 9$$

The volunteer enlistee groups studied in this analysis were limited to high school graduates of Mental Category 1, 2 and 3 and non-high school graduates of the same mental categories. Regression results presented here are for the following separate classes as dependent variables:

1. Category 1-2 high school graduate enlistment rates
2. Category 3 high school graduate enlistment rates
3. Category 1-3 total volunteer enlistment rates
4. Category 1-3 non-high school graduate enlistment rates

The method for determining service volunteers prior to July 1973 is described in Appendix D. The traditional method for determining the mental category groups is presented in Appendix D.

The analysis of Army enlistment rates included regressions on groups 1, 2, and 3 described above. The analysis of Navy enlistments included groups 1, 2, and 4 above as dependent variables in regressions. An analysis of Mental Category 4 volunteer enlistees was not attempted since enlistment levels for the group are usually determined by service policies regarding quotas.

METHODOLOGY

The rationale for the use of models with constant slopes versus random component models varies with different studies of time series cross-sectional structures. One major reason for constant slopes in this analysis is the fact that the principal exogenous variables available for enlistment supply estimation are policy variables such as pay, recruiting and advertising. It is important to test the likelihood of estimating all parameters under the assumption of the same directional response to policies across all regions. Other reasons for the use of a fixed parameter model includes the desire on the part of analysts to measure the effectiveness of nationwide policies -- or policies in the aggregate.

To test the pooled time series cross-sectional models for significance and adequacy, the following procedure was employed:

1. First, perform regressions by ordinary least squares techniques with both linear and log-linear forms on only the exogenous variables most highly correlated with the dependent variable.

2. Perform regressions including all exogenous variables.

3. Perform regressions with all variables and include dummy variables as proxies for regional components of variance.

4. Perform a two-stage analysis regressing the dependent variable on the exogenous variables as in 2 above and then regress the residuals on the regional dummy variables.

This approach to the analysis provided important and varied information at each step. From steps 1 and 2, a determination was made on just how well the national scale model performs without regional component proxies. Steps 3 and 4 were employed to show the adequacy of the models as forecasting tools.

The linear model employed in steps 1 and 2 will be referred to as Model 1 in this chapter; the log-linear model for steps 1 and 2 is Model

2. The respective models used in step 3 are Model 1-R and Model 2-R. The step 4 process was designed as a check on residual variation.

RESULTS OF ARMY REGRESSIONS

The Army's concern for obtaining a sufficient supply of high school graduate enlistees has prompted the analysis of Mental Category 1, 2 and 3 high school graduate enlistment levels in most recent studies. In this analysis, regressions were run against Mental Category 1-2 high school graduate enlistment levels separate from those run against Category 3 high school graduate enlistments. This was done in order to estimate variable elasticities pertaining to the most supply-limited group of volunteer enlistees -- Mental Category 1-2 high school graduates.

The results of the regressions are presented in Tables 22-25. Tables 22 and 24 show the results of regressions which did not include regionals (regional dummy variables); Tables 23 and 25 show results on regressions with regionals.[2] The results for all variables in the same table were obtained from the same regression. The following information is provided in each table.

- The coefficient estimated via the regression.
- The significance level based on a t-statistic for that variable.
- The elasticity or rate of change for the dependent variable per unit change in the independent variable.

Regression results for Army Mental Category 1-2 high school graduate volunteer rates indicate no statistically significant response to increases in military pay across regions.

Table 22

ARMY ENLISTMENT SUPPLY PARAMETER ESTIMATES
REGRESSION MODEL M1

Independent variables	Category 1-2 High School Graduates			Category 3 High School Graduates			Category 1-3 Total		
	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity
Pay ratio	.0017	a	.179	.0019	a	.154	-.0028	a	-.062
Advertising	.000013	a	.014	.00031	.05	.254	.00049	a	.109
Unemployment	-.00012	a	-.088	-.0001	a	-.057	-.0008	a	-.123
Recruiters/QMA	3.7835	.01	.747	3.596	.05	.551	14.966	.05	.624
% Black	-.0203	.01	-.204	.0184	.05	.144	.0288	a	.061

[a]The t-statistic for this variable was less than the critical value at the .10 level.

Table 23

ARMY ENLISTMENT SUPPLY PARAMETER ESTIMATES
REGRESSION MODEL M1-R

Independent variables	Category 1-2 High School Graduates			Category 3 High School Graduates			Category 1-3 Total		
	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity
Pay ratio	.00056	a	.059	-.0030	a	-.241	-.0193	.10	-.429
Advertising	.000087	.10	.092	.00027	.01	.224	.00056	.05	.124
Unemployment	-.00019	a	-.139	.00003	a	.018	-.00036	a	-.055
Recruiters/QMA	1.6657	.10	.329	1.6037	a	.246	5.8788	.10	.245
% Black	-.0109	a	-.109	-.0408	.05	-.319	-.0904	.10	-.192
Regionals:									
NE	0.			0.			0.		
MA	-.0004	a		.0038	.10		.0073	a	
ENC	.0003	a		.0036	.10		.0095	a	
WNC	.0049	.01		.0059	.01		.0233	.01	
SA	.0018	a		.0174	.01		.0422	.01	
ESC	-.0001	a		.0158	.01		.0421	.01	
WSC	.0007	a		.0112	.01		.0314	.01	
MD	.0061	.01		.0044	.01		.0246	.01	
PD	.0047	.01		.0071	.01		.0232	.01	

[a]The t-statistic for this variable was less than the critical value at the .10 level.

Table 24

ARMY ENLISTMENT SUPPLY PARAMETER ESTIMATES
REGRESSICN MODEL M2

Independent variables	Category 1-2 High School Graduates			Category 3 High School Graduates			Category 1-3 Total		
	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity
Pay ratio	.269	a	.269	.227	a	.227	.034	a	.034
Advertising	-.002	a	-.002	.119	.05	.119	.170	a	.170
Unemployment	.003	a	.003	-.077	a	-.077	-.120	a	-.120
Recruiters/QMA	.628	.05	.628	.531	.05	.531	.620	.05	.620
% Black	-.159	.01	-.159	.086	.10	.086	.042	a	.042

[a]The t-statistic for this variable was less than the critical value at the .10 level.

Table 25

ARMY ENLISTMENT SUPPLY PARAMETER ESTIMATES
RECRESSION MODEL M2-R

Independent variables	Category 1-2 High School Graduates			Category 3 High School Graduates			Category 1-3 Total		
	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity
Pay ratio	.105	a	.105	.181	a	.181	-.355	a	-.355
Advertising	.065	a	.065	.159	.01	.159	.067	.05	.067
Unemployment	-.047	a	-.047	.062	a	.062	.003	a	.003
Recruiters/QMA	.104	a	.104	.060	a	.060	.104	a	.104
% Black	-.132	a	-.132	-.118	a	-.118	-.089	a	-.089
Regionals:									
NE (Control)	0.			0.			0.		
MA	.0306	a		.2054	a		.1230	a	
ENC	.1336	a		.1886	a		.1889	a	
WNC	.5425	.01		.5215	.01		.5355	.01	
SA	.2735	a		.8248	.01		.6635	.01	
ESC	.1362	a		.8834	.01		.7544	.01	
WSC	.2462	a		.7080	.01		.6238	.01	
MD	.5232	.01		.3857	.01		.5135	.01	
PD	.5296	.01		.5529	.01		.5068	.01	

[a]The t-statistic for this variable was less than the critical value at the .10 level.

The recruiting variable appeared in all regressions without regional components for Army enlistment rates with a coefficient significantly positive at the .05 level or better. Inclusion of regional components reduced the significance of the recruiting variable substantially. In fact, the regional components are, in general, of much higher statistical significance than any of the exogenous variables as defined in this analysis. It is also seen that including regional components significantly reduced the recruiter elasticities from .60-.75 without regionals to a range of .10-.30 with regional components.

The amount of advertising by the Army appears to have a significant effect on only one enlistment group — the Category 3 high school graduates. Advertising elasticities for this group range from .16 to .25 at the .05-.01 significance level. Other volunteer groups appear to be rather less responsive to increases in advertising expenditures.

The unemployment rate variable did not appear in significant proportions in any of the regressions. The results for this variable do not immediately rule out its importance for enlistment projections where post 1974 unemployment data can be used in the analysis.

The variable used in the analysis to represent the proportion of black male youth in the population enters the high school graduate regressions in significant proportions. The percentage of blacks in a region, however, may be only a partial proxy variable in effect for attitudes toward the military or a particular branch of service. This is reinforced by the occurrence of a significantly positive coefficient in regressions without regional dummy variables and the occurrence of negative coefficients in regressions with regional dummy variables with a decidedly reduced significance for the black proportion variable.

The overall significance of the Army pooled time series cross-sectional models for Category 1-2 high school graduates is summarized in Table 26. The amount of variation in the enlistment data explained by the models with no regional components is slightly greater than 50 percent. The standard error of the regression ranges between 6 and 23 percent on the mean of the dependent variable. With the inclusion of regional proxy variables in all regressions, the variables explain an additional 28 to 30 percent of the variance and the standard error of the regression is decreased. R^2 values for Model 1-R and Model 2-R are .80 and .82, respectively. R^2 and standard error measures are almost identical for regressions on the other mental category groups.

Pay ratio elasticities are generally the same for both sets of models tested in the regressions; however, the recruiting variable elasticities tend to be sensitive to the inclusion of dummy variables representing regional components. The apparent reason for this occurrence is the attempt by the Army to place recruiters where larger elements of the military available population are to be found. Based on this assumption, the recruiting variable elasticity ranges between .10 and .33 for Category 1-2 high school graduates.

Regional dummy variables for census regions 4, 8 and 9 entered the regressions at a significance level of .01 consistently. These proxy variables represent variation in enlistment patterns for the West North Central, Mountain and Pacific census regions. Appendix E contains a list of the census subdivisions or regions for population estimation and the states contained within these regions.

Table 26

REGRESSIONS FOR ARMY CATEGORY 1-2 HIGH SCHOOL GRADUATES
REGIONAL VARIATION

	R^2	Standard Error of the Regression	Black Prop. Elasticity	Recruiter Elasticity
Model 1 Regressions				
With regionals[a]	.835	.0017 (15%)	-.109	.329
Without regionals	.517	.0026 (24%)	-.204	.747
Model 2 Regressions				
With regionals	.813	.160 (3.5%)	-.132	.104
Without regionals	.490	.236 (5.2%)	-.159	.628

[a]Regionals = regional dummy variables.

RESULTS FOR NAVY ENLISTMENTS

Regressions were performed against three different Navy enlistment levels: these were Mental Category 1-2 high school graduate, Mental Category 3 high school graduate and Mental Category 1-3 non-high school graduate volunteer rates. The results of the analysis showed a dominance of the recruiters per QMA variable in explaining the variance in the pooled data. The only other significant variable in the analysis was the percentage black youth population variable when no regional components were considered.

The results of regressions on Navy enlistment levels are summarized in Tables 27-30. It is not possible to estimate the true range of elasticity measurements from the results for the pay ratio, advertising and unemployment variables. However, since no pair of the exogenous variables exhibited high multicollinearity, the implication is that the effect of changes in military pay, the youth unemployment rate and the Navy advertising effort are only supportive at best to the service's recruiting effort.

The recruiting variable elasticity is measured at about .75-.90 for Category 1-2 high school graduates and .85-1.20 for Category 3 high school and Category 1-3 non-high school graduates. The predominance of the recruiting variable in all the regressions was tested by examining different forms of the recruiter variable in regressions for Model 1 and Model 2. The number of recruiters was always a significant variable in the regressions.

Another exogenous variable of importance in the Navy regression results is the proportion of black youth in the population (PROB). This variable enters most log-linear regressions with a negative sign and is significant at the .01 or .05 level when no regional components are included. The

Table 27

NAVY ENLISTMENT SUPPLY PARAMETER ESTIMATES
REGRESSION MODEL M1

Independent variables	Category 1-2 High School Graduates			Category 3 High School Graduates			Category 1-3 Non-High School Graduates		
	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity
Pay ratio	.00045	a	.041	-.00360	a	-.298	-.00480	a	-.438
Advertising	.000012	a	.005	-.00012	a	-.047	.00003	a	.013
Unemployment	-.00011	a	-.069	.00006	a	.034	-.00044	a	-.277
Recruiters/QMA	5.9746	.01	.908	7.6206	.01	1.057	6.7750	.01	1.035
% Black	-.0130	a	-.112	-.0107	a	-.084	-.0132	a	-.115

[a]The t-statistic for this variable was less than the critical value at the .10 level.

Table 28

NAVY ENLISTMENT SUPPLY PARAMETER ESTIMATES
REGRESSION MODEL M1-R

Independent variables	Category 1-2 High School Graduates			Category 3 High School Graduates			Category 1-3 Non-High School Graduates		
	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity
Pay ratio	-.480E(-2)[b]	a	-.436	-.401E(-2)	a	-.332	-.102E(-1)	a	-.932
Advertising	.487E(-4)	a	.021	.862E(-4)	a	.034	.157E(-3)	a	.068
Unemployment	.930E(-4)	a	.058	-.482E(-4)	a	-.027	.121E(-3)	a	.076
Recruiters/QMA	5.665	.01	.861	7.833	.01	1.087	7.843	.01	1.198
% Black	-.895E(-2)	a	-.077	-.754E(-2)	a	-.059	-.234E(-1)	a	-.203
Regionals:									
NE (Control)	0.			0.			0.		
MA	-.445E(-2)	a		-.384E(-2)	a		-.567E(-2)	a	
ENC	-.324E(-2)	a		-.256E(-2)	a		-.383E(-2)	a	
WNC	.265E(-2)	a		-.220E(-2)	a		-.259E(-2)	a	
SA	-.117E(-2)	a		-.291E(-2)	a		-.142E(-2)	a	
ESC	-.015E(-2)	a		-.039E(-2)	a		.069E(-2)	a	
WSC	-.478E(-2)	a		-.280E(-2)	a		-.138E(-2)	a	
MD	-.388E(-2)	a		-.414E(-2)	a		-.129E(-1)	:10	
PD	-.298E(-2)	a		-.049E(-2)	a		-.599E(-2)	a	

[a]The t-statistic for this variable was less than the critical value at the .10 level.
[b]$E(-x) = 10^{-x}$.

Table 29

NAVY ENLISTMENT SUPPLY PARAMETER ESTIMATES
REGRESSION MODEL M2

Independent variables	Category 1-2 High School Graduates			Category 3 High School Graduates			Category 1-3 Non-High School Graduates		
	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity
Pay ratio	-.1455	a	-.1455	-.4160	a	-.4160	-.2088	a	-.2088
Advertising	.0195	a	.0195	.0267	a	.0267	.0157	a	.0157
Unemployment	-.0939	a	-.0939	-.1348	a	.1348	-.2595	a	-.2595
Recruiters/QMA	.7576	.01	.7576	.9175	.01	.9175	.8488	.01	.8488
% Black	-.2005	.01	-.2005	-.1646	.01	-.1646	-.2008	.05	-.2008

[a]The t-statistic for this variable was less than the critical value at the .10 level.

Table 30

NAVY ENLISTMENT SUPPLY PARAMETER ESTIMATES
REGRESSION MODEL M2-R

Independent variables	Category 1-2 High School Graduates			Category 3 High School Graduates			Category 1-3 Non-High School Graduates		
	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity	Coefficient	Significance level	Elasticity
Pay ratio	-.2418	a	-.2418	-.3326	a	-.3326	-.3540	a	-.3540
Advertising	.0036	a	.0036	.0164	a	.0164	.0390	a	.0390
Unemployment	-.1336	a	-.1336	-.1692	a	-.1692	-.0614	a	-.0614
Recruiters/QMA	.6330	.01	.6330	.8984	.01	.8984	.9438	.01	.9438
% Black	-.2208	a	-.2208	-.2324	a	-.2324	-.2117	a	-.2117
Regionals:									
NE (Control)	0.			0.			0.		
MA	-.0975	a		-.0178	a		-.3741	a	
ENC	-.0484	a		.0328	a		-.2685	a	
WNC	.2126	a		-.0594	a		-.4077	a	
SA	.0366	a		-.0479	a		-.4425	a	
ESC	.3981	a		.3857	a		.0156	a	
WSC	-.5211	a		-.1060	a		-.2667	a	
MD	-.0859	a		-.1723	a		-.7264	.10	
PD	.1312	a		.2573	a		-.4092	a	

[a]The t-statistic for this variable was less than the critical value at the .10 level.

results indicate a possible .2 percent decrease in the quality group enlistment levels for 1 percent increase in the proportion of black youth in the regional populations.

The regional dummy variables in the analysis were almost uniformly not statistically significant for the Navy accessions groups analyzed here. The overall results for the analysis on Navy enlistments are presented in Table 31. The R^2 measurements show that a significant proportion of the variance in the enlistment data for Mental Category 1-2 high school graduates is explained by the Model 1 equation alone. The addition of the regional components in Model 1-R account for only an additional 4 or 5 percent of the variance.

SUMMARY OF FINDINGS

The pooled time series cross-sectional models used in this analysis provide for an evaluation of the Army and Navy pay and recruiting policies across the diverse nine census regions. The results obtained from testing these models may be summarized with the following observations:

1. The number of recruiters per QMA variable is the most significant factor in the pooled models. Recruiter elasticities range from .08 to .68 for Army results and from .70 to 1.20 for Navy results. It must be considered, though, that this observation of recruiter impact may be due to a very successful tracking of enlistment potential supply by service recruiters and not simply a productivity measure.

2. The pay ratio variable defined in this analysis elicits no significant response in regressions on Army volunteers nor in regressions on Navy volunteers. Since the standard error of regression is smaller for Army

Table 31

REGRESSION RESULTS FOR NAVY CATEGORY 1-2 HIGH SCHOOL GRADUATES
REGIONAL VARIATION

	R^2	Standard Error of the Regression	Black Prop. Elasticity	Recruiter Elasticity
Model 1 Regressions				
With regionals[a]	.901	.0035 (27%)	-.077	.861
Without regionals	.860	.0037 (29%)	-.112	.908
Model 2 Regressions				
With regionals	.846	.315 (6.9%)	-.201	.633
Without regionals	.733	.370 (8.5%)	-.201	.758

[a]Regionals = regional dummy variables.

regressions than for Navy regressions, it is possible that more conclusive results for the Navy could be obtained via a different method, perhaps separate regional time series models.

3. The youth unemployment rate does not appear to have a significant effect on enlistment patterns across geographic regions. The unemployment rate variable enters regressions with a non-significant and usually negative coefficient. A regional time series analysis could possibly dispute the measured elasticities from this analysis provided the most recent trends (1974-1975) in unemployment rates are considered.

4. The percentage of blacks in the male youth population enters regressions on volunteer levels in statistically significant proportions. This indicates a sensitivity of Army and Navy enlistment patterns to the black population distribution. The presence of this variable in the regressions is assumed to be that of a proxy variable for service attitudes or attitudes toward the two branches of service. The variable occurs with greater significance in Navy regressions than in Army regressions. This variable also performs as a camouflage for quality levels within the services.

5. The amount of advertising appears to be an important factor in the volunteer supply level of Army Category 3 high school graduates. It is not, however, of significant determination for other enlistee groups.

6. The tradeoffs between the use of type 1 and type 2 models are not always apparent. In most regressions, the standard errors of the regressions as well as the standard errors of the coefficients are less than 20 percent, generally about 14 percent for statistically significant coefficients. R^2 values are greater for Model 1 type regressions than for

Model 2 types. Because there is also a lack of complexity in the structure of the type 1 equation with regional component proxies, the use of Model 1-R is considered the best choice as a forecasting tool among the pooled models considered in this study.

Appendix A

ANALYSIS OF FISHER'S METHOD OF ESTIMATING W_C AND W_M

Fisher's method of estimating W_C using data from Census Current Population Reports[1] for year-round, full-time male workers, age 14-19 (Y14-19) and 20-24 (Y20-24) is replicated in this analysis. Fisher's estimating equation was:

$$W_C = 2Y_{14-18} + Y_{20-24}$$

The results of GRC calculations and Fisher's estimates (as reported in several publications)[2] are summarized in Table A1, which also contains an estimate of the Consumer Price Index derived from monthly data. Following the description of Fisher's estimating procedure described in Kim, Farrell and Clague, it was assumed that the statistics reported in the Current Population Report for a given year were for the third quarter of that year.

Comparison of the rates of W_C estimated from the CRP to W_C reported by Fisher (column 3) with the Consumer Price Index for that quarter produces strikingly similar numbers. This comparison constitutes strong evidence in support of the notion that Fisher used the real value of W_C as his estimate of the returns to not enlisting.

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[1] U.S. Bureau of the Census, Current Population Reports, Series R60, Consumer Income.

[2] Kim, et al., pp. 200-201; Hause-Fisher, pp. 131-132.

Table A1

ESTIMATES OF W_C DERIVED FROM CPR AND REPORTED BY FISHER, 1957-1965

Quarter and year	W_c CPR (1)	Fisher (2)	Relative difference[a] (3)	CPI[b] (4)
3:57	6,859	6,999	98.0	n.a.
3:58	7,128	7,078	100.7	100.8
3:59	7,201	7,095	101.5	101.4
3:60	7,860	7,628	103.0	103.2
3:61	7,921	7,602	104.2	104.0
3:62	8,457	8,024	105.4	105.3
3:63	8,671	8,127	106.7	106.6
3:64	9,067	8,388	108.1	108.0
3:65	10,854	9,876	109.9	110.2

[a] Relative difference = [Col. (1) ÷ Col. (2)] x 100.

[b] For the month of June; 1957-59 = 100.

Sources:

Col. (1): Kim, et al., pp. 200-201 (1958-1965); Hause-Fisher, pp. 131-132 (1957).

Col. (2): Unpublished calculations derived from U.S. Bureau of Census, Current Population Reports, Series P-60, Consumer Income, Nos. 33, 35, 37, 39, 41, 43, 47, 51 and 53.

Table A2 summarizes Fisher's estimate and is shown below. Since nominal military pay rates are subject to Congressional control, they would be expected to display a fairly discrete pattern of movement over time, changing only when Congress approved changes in military pay rates. There, one would expect to find an estimate of nominal W_M to display a step-line pattern of movement with the steps occurring at the time of Congressionally approved pay increases. Until recently, the Congress has not been inclined to award nominal pav increases to first-term enlistees. Fisher's estimates, contained in Table A1 display a step-function pattern of movement that is consistent with the notion that his estimates of nominal (as opposed to real) returns to enlisting. W_M remains extremely stable in the earlier quarter and begins to rise more rapidly in the later quarter. This pattern is consistent with Congressional neglect of nominal first-term W_M in the earlier years of Fisher's analysis. Presumably an estimate of real W_M would have shown some decline in these early years because of inflation.

Table A2

ESTIMATES OF W_M REPORTED BY FISHER, 1957-1965

Quarter and year	W_M	Quarter and year	W_M
3:58	6869	1:62	6933
4:58	6869	2:62	6933
1:59	6933	3:62	6933
2:59	6933	4:62	6933
3:59	6933	1:63	7074
4:59	6933	2:63	7074
1:60	6933	3:63	7074
2:60	6933	4:63	7276
3:60	6933	1:64	7276
4:60	6933	2:64	7276
1:61	6933	3:64	7276
2:61	6933	4:64	7307
3:61	6933	1:65	7307
4:61	6933	2:65	7307
		3:65	7884
		4:65	7884

Source: Hause-Fisher, p. 132; see also Kim, et.al., pp. 200-201.

Appendix B

COMPARISON OF FISHER DATA BASE WITH COMMON DATA BASE

Fisher's data base consists of six variables: enlistments, accessions, population, unemployment rates, military pay, and civilian pay. Enlistments and accessions are generated on the same data source by both the Fisher data base and the common data base. They therefore should not differ from one another by any substantial amount. Table B1 compares Fisher's enlistment and accession estimates with those derived for this study. With the exception of four quarters, 1:58, 3:58, 3:43 and 1:65, the estimates differ by no more than three percent. Fisher estimates his population variable to include the male civilian population in the age range 17 through 20. Fisher reports that he derives his estimates from the Current Population Reports, which present annual observations. Quarterly estimates of population are derived from these annual observations by linear interpolation. The Current Population Reports include two possible sources of population statistics. One source reports population estimates as of July; the other source, derived from school enrollment reports, reports population figures as of October. Unfortunately, Fisher does not report which of these two series he uses to generate his population estimates. The common data base uses the population series generated from the October figures.

Estimates of the male civilian noninstitutional population age 17-20 were derived for the common data base from the Current Population Reports. These estimates were obtained by linear interpolation of the

Table B1

COMPARISON OF DOD ENLISTMENTS AND ACCESSIONS, MENTAL GROUPS 1-3 (IN THOUSANDS), 1957-1965

	Enlistments (in thousands)			Accessions (in thousands)		
	(1) Fisher	(2) Common Data Base	(3) (2) ÷ (1)	(4) Fisher	(5) Common Data Base	(6) (5) ÷ (4)
1957:3	68.9					
:4	38.8					
1958:1	60.5	60.7	1.003	101.4	106.3	1.048
:2	55.2	55.5	1.005	96.6	98.8	1.022
:3	81.2	77.5	.954	124.3	121.8	.980
:4	54.8	55.0	1.004	96.6	97.8	1.012
1959:1	63.4	63.6	1.003	93.7	94.9	1.012
:2	54.4	54.6	1.004	72.6	73.4	1.011
:3	80.4	80.9	1.006	109.5	111.1	1.014
:4	58.8	59.0	1.003	87.0	88.0	1.011
1960:1	69.5	69.6	1.001	90.4	91.4	1.011
:2	67.7	67.8	1.001	97.6	98.8	1.012
:3	97.0	97.5	1.005	130.1	131.5	1.011
:4	63.7	63.8	1.001	94.0	95.2	1.013
1961:1	76.0	76.3	1.004	92.7	93.5	1.009
:2	67.4	67.6	1.003	73.8	74.1	1.004
:3	108.9	109.6	1.006	158.8	161.8	1.018
:4	76.3	76.8	1.007	139.0	141.2	1.015
1962:1	83.7	84.1	1.005	122.8	124.9	1.017
:2	66.1	66.6	1.008	89.1	90.1	1.010
:3	90.0	91.8	1.020	113.9	114.0	1.001
:4	57.3	58.4	1.019	76.7	78.4	1.022
1963:1	67.3	69.3	1.030	91.5	92.6	1.012
:2	57.8	59.6	1.031	88.3	88.7	1.004
:3	91.4	91.4	1.	130.6	130.4	.998
:4	61.3	61.3	1.	115.3	115.3	1.000
1964:1	77.6	77.6	1.	129.6	129.4	.999
:2	65.3	65.3	1.	95.8	95.7	.999
:3	90.1	90.1	1.	112.9	112.6	.998
:4	51.3	51.3	1.	78.7	78.6	.999
1965:1	63.5	57.5	.906	89.6	83.7	.935
:2	66.7	66.7	1.	120.3	120.1	.999
:3	113.9	113.9	1.	188.0	186.6	.993
:4	102.7	102.7	1.	220.0	216.0	.982

Source: (1), (4) Hause-Fisher, p. 129

October statistics that are reported by the Census annually. Estimates for 17-20 year olds were derived from the following equation:

$$P1720 = 1/2\ P1617 + P1819 + 1/2\ P2021,$$

where P1720 is the 17-20 year old civilian, noninstitutional population, P1617 is the 16-17 year old civilian noninstitutional population, P1819 is the 18-19 year old civilian noninstitutional population, and P2021 is the 20-21 year old population. The equation is modified for the years 1957 and 1958 to accomodate the fact that 20 year olds are reported with 20-24 year olds (20-21 year olds are not reported separately). For these years, 1/2 P2021 is replaced by 0.2 P2024, where P2024 is the 20-24 year old population.

Fisher reports th. . he derives his statistics from the Current Population Reports, but does not provide any further details as to how his series is constructed. Table B2 compares the two population series, for the period 3:57-4:65. On average, the series developed for the common data base is one percent higher than the series reported by Fisher. This overstatement is quite pronounced in the early part of the series, where the population estimates developed for the common data base are as much as seven percent above those developed by Fisher. This discrepancy narrows to less than two percent by late 1959 and remains within two percent (in absolute terms) thereafter, except for one quarter, when it reaches 2.5 percent.

Fisher's unemployment rate is the rate for all males age 18 and 19. His series is taken from unpublished Bureau of Labor statistics figures. Table B3 compares Fisher's unemployment rate estimates with those derived for this study. The estimates are exactly the same in all but one quarter (2:65).

Table B2

COMPARISON OF ESTIMATES OF THE 17-20 YEAR OLD MALE CIVILIAN NONINSTITUTIONAL POPULATION, 1957-1965

	(1) Fisher	(2) Common Data Base	(3) (2) ÷ (1)
1957:3	3633	3890	1.071
:4	3717	3935	1.060
1958:1	3787	3979	1.051
:2	3861	4025	1.042
:3	3935	4069	1.034
:4	3997	4118	1.030
1959:1	4058	4168	1.027
:2	4119	4219	1.024
:3	4180	4271	1.022
:4	4261	4327	1.015
1960:1	4342	4396	1.012
:2	4423	4465	1.009
:3	4503	4533	1.007
:4	4565	4594	1.006
1961:1	4628	4637	1.002
:2	4691	4683	.998
:3	4754	4727	.994
:4	4787	4762	.995
1962:1	4820	4780	.992
:2	4853	4797	.984
:3	4886	4814	.985
:4	4900	4850	.990
1963:1	4913	4922	1.002
:2	4927	4994	1.013
:3	4940	5067	1.025
:4	5042	5136	1.019
1964:1	5145	5204	1.011
:2	5248	5270	1.004
:3	5351	5337	.997
:4	5492	5426	.998
1965:1	5634	5555	.986
:2	5776	5685	.984
:3	5918	5813	.982
:4	5973	5912	.990

Source: (1) Hause-Fisher, p. 130.
(2) U.S. Bureau of the Census, Current Population Reports, Series P-20, Population Characteristics, School Enrollments.

Table B3

COMPARISON OF ESTIMATES OF THE 18-19 YEAR OLD MALE UNEMPLOYMENT RATES, 1957-1965

	(1) Fisher	(2) Common Data Base
1957:3	.098	.098
:4	.124	.124
1958:1	.204	.204
:2	.182	.184
:3	.173	.173
:4	.157	.157
1959:1	.170	.169
:2	.145	.146
:3	.138	.139
:4	.146	.146
1960:1	.162	.162
:2	.150	.150
:3	.135	.135
:4	.159	.159
1961:1	.203	.203
:2	.176	.176
:3	.136	.136
:4	.146	.146
1962:1	.170	.170
:2	.139	.139
:3	.119	.119
:4	.131	.132
1963:1	.178	.178
:2	.178	.178
:3	.140	.140
:4	.144	.144
1964:1	.160	.160
:2	.161	.161
:3	.133	.132
:4	.135	.135
1965:1	.140	.140
:2	.157	.165
:3	.106	.107
:4	.099	.099

Source: (1) Hause-Fisher, p. 133.

Fisher's estimate of civilian pay is generated as a weighted sum of the year-round, full-time earnings of male workers in the age groups 14 to 19 and 20 to 24. (Fisher, p. 247. He assumes enlistment at age 18.) Table B4 compares Fisher's estimates to those generated from the common data base compiled for this study. With the exception of one quarter, 4:65, the estimates are within three per cent of each other; they are within one percent of each other in all but six quarters, 64:3 - 65:4.

The most difficult variable to replicate was Fisher's estimate of military pay. Recall that Fisher's estimate included basic pay, quarters and subsistence allowances, and an imputed value for medical services (Hause-Fisher, p. 64). He reports that his estimate of basic pay and allowances is taken from schedules giving base pay by pay grade that were in force at the time of the observation. These estimates are used together with estimates of average time in grade provided by the services to produce average pay for the first three-year enlistment in the military. Fisher reports an estimate of military pay for 1957-58 that is equivalent to a monthly average of approximately $193 for basic pay, allowances, and medical benefits. Medical benefits are reported valued at $253 per year (Kim, et al., p. 198). Thus we can estimate Fisher's measure of average basic pay and allowances to be about $171. Using reasonable assumptions about promotion rates and basic pay scales for 1957-63, the average monthly value of basic pay over a three-year enlistment should have ranged between $102 and $115.[1] Deducting this estimate of basic pay from Fisher's estimate

[1] Two promotion rates (slow and rapid) were used. The time in grade assumptions associated with these rates, together with the relevant value of basic pay, are described below:

Table B4

COMPARISON OF ESTIMATES OF W_C, 1958-1965

	(1) Common Data Base	(2) Fisher	(3) (2) ÷ (1)
1958:1	7016.00	7039.00	1.00328
1958:2	7033.76	7059.00	1.00359
1958:3	7077.38	7078.00	1.00009
1958:4	7095.24	7082.00	.998134
1959:1	7114.09	7086.00	.996052
1959:2	7111.77	7090.00	.996939
1959:3	7128.68	7095.00	.995275
1959:4	7255.13	7228.00	.996260
1960:1	7416.42	7361.00	.992527
1960:2	7535.47	7494.00	.994497
1960:3	7625.00	7628.00	1.00039
1960:4	7594.41	7622.00	1.00363
1961:1	7607.90	7615.00	1.00093
1961:2	7615.01	7609.00	.999210
1961:3	7629.31	7602.00	.996420
1961:4	7750.24	7707.00	.994421
1962:1	7863.42	7813.00	.993588
1962:2	7953.42	7918.00	.995546
1962:3	8017.03	8024.00	1.00087
1962:4	8052.88	8050.00	.999642
1963:1	8088.60	8076.00	.998443
1963:2	8122.30	8102.00	.997501
1963:3	8126.98	8127.00	1.000000
1963:4	8197.39	8192.00	.999342
1964:1	8272.30	8257.00	.998150
1964:2	8349.72	8322.00	.996680
1964:3	8509.70	8338.00	.979824
1964:4	8897.79	8760.00	.984514
1965:1	9274.31	9132.00	.984655
1965:2	9614.75	9504.00	.988481
1965:3	9776.77	9876.00	1.01015
1965:4	9512.20	10248.0	1.07735

Source: (2) Hause-Fisher, p.

of basic pay and allowances, it appears that Fisher valued the allowances at $56-$69 per month. Since the quarters allowance for enlisted men was $55.20 per month at the time, it is assumed that Fisher added the value of the quarters allowance to his estimates of basic pay (using a liberal promotion assumption) and his imputed value of medical services. Table B5 compared our estimates of W_M with those reported by Fisher. The estimates range within 3 percent of each other, suggesting that the assumptions newly adopted in generating an estimate of military pay for the common data base were reasonable proxies for those employed by Fisher.[2]

Pay Grade	Time-in-grade (in months) Slow Promotion	Rapid Promotion	Monthly Basic Pay Less than 2 years	2 years, less than 3 years
E1	4	4	$78.00	
E2	8	8	$85.80	
E3	24	6	$99.37	$124.00
E4		18	$122.30	$150.00
3 year monthly average base pay	$102	$115		

[2]The relatively large differences occurring in the third and fourth quarters of 1965 are the result of an error in Fisher's estimates. He assumed the 1965 base pay increase occurred at the beginning of the third quarter, when it actually became effective in September, the end of the third quarter. Adjusting Fisher's estimates for this error brings his estimates to within three dollars of our estimates.

Table B5

COMPARISON OF ESTIMATES OF W_M, 1958-1965, FISHER VS. COMMON DATA BASE

	(1) Common Data Base	(2) Fisher	(3) (2) - (1) Difference
1958:1	6755.02	6869.00	113.980
1958:2	6755.02	6869.00	113.980
1958:3	6755.02	6933.00	177.980
1958:4	6755.02	6933.00	177.980
1959:1	6755.02	6933.00	177.980
1959:2	6755.02	6933.00	177.980
1959:3	6755.02	6933.00	177.980
1959:4	6755.02	6933.00	177.980
1960:1	6755.02	6933.00	177.980
1960:2	6755.02	6933.00	177.980
1960:3	6755.02	6933.00	177.980
1960:4	6755.02	6933.00	177.980
1961:1	6755.02	6933.00	177.980
1961:2	6755.02	6933.00	177.980
1961:3	6755.02	6933.00	177.980
1961:4	6755.02	6933.00	177.980
1962:1	6755.02	6933.00	177.980
1962:2	6755.02	6933.00	177.980
1962:3	6755.02	6933.00	177.980
1962.4	6755.02	6933.00	177.980
1963:1	6895.42	7074.00	178.580
1963:2	6895.42	7074.00	178.580
1963:3	6895.42	7074.00	178.580
1963:4	7255.42	7276.00	20.580
1964:1	7255.42	7276.00	20.580
1964:2	7255.42	7276.00	20.580
1964:3	7273.42	7276.00	2.57996
1964:4	7305.42	7307.00	-2.42004
1965:1	7309.42	7307.00	-2.42004
1965:2	7309.42	7307.00	-2.42004
1965:3	7555.95	7884.00	328.053
1965:4	8049.00	7884.00	-165.00

Source: (2) Hause-Fisher, p. 132

Appendix C

DESCRIPTIONS OF VARIABLES USED IN THE POOLED TIME SERIES CROSS SECTIONAL MODEL REGRESSIONS

The specifications of variables used in the pooled time series cross sectional regressions were the result of an expansive effort to obtain the best estimates of the primary factors in the analysis. No reliance was made on published estimates of civilian pay, youth unemployment rates, the educational stratification of the population, nor the supply of accessions from different geographical locations. All data variables except recruiting and advertising were derived in-house by GRC staff.

The methods for obtaining the population variable estimates included the processing of Bureau of the Census "Current Population Survey" tapes for the years 1970 through 1974. The 1975 tape was not available to GRC analysts. Also derived from the CPS tapes were the estimates of the 17-21 year old male unemployment rates and the civilian pay received by 17-21 year old males in the diverse regional locations.

True volunteer accession variables were derived from the USAREC[1] files for 1970-1975 and were extracted by region, race, mental category and educational group. The method for determining the enlistee mental categories is explained in Appendix D. More prospective dependent variables were derived than were in the analysis due to limitations of the independent variables. The data were extracted on a monthly basis, then summed for yearly estimates.

[1] United States Army Recruiting Command.

Although the 17-21 year old male civilian population, the military availables, was derived from census tapes, estimates of the physically and mentally qualified population had to be obtained from different sources. The use of disqualification rates against population variables is assumed to provide a better measure of the potential enlistees to a branch of service. A set of disqualification rates was obtained for this study from the HumRRO[2] master file of the mental category and physical acceptance rate distributions of pre-inductees for the year 1972.

Disqualification rates were extracted at GRC from the HumRRO file by race and by state of residence of the pre-inductee.[3] State qualification rates were grouped according to regional location and a set of nine qualification factors was derived by weighted averaging. The disqualification rates per region were assumed constant over time.

The military pay variables for the Army and Navy were derived by a sequence of averaging processes on the total of basic pay, quarters and subsistence allowances and the tax advantage on those allowances for pay grades E1-E6 by years of service. The Army pay variable assumes expected income by a new recruit is for three years of service, whereas the Navy pay variable assumes a four-year expectation of pay. The military pay variable does not change from region to region in the series; however, the civilian pay variable had both regional and yearly differences. This provided variation in the pay ratio variable attributable to regional differences.

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[2]Human Resources Research Organization.

[3]"FY77 Qualified Military Available (QMA) Inventory," prepared for Headquarters, US Army Recruiting Command, General Research Corporation, 25 November 1975.

The recruiting and advertising data were obtained from the services. Yearly estimates were derived from monthly Army recruiting data and from Navy fiscal year data. Army recruiting data were broken out by recruiting command regions and state distributions had to be calculated at GRC. The state recruiting data were then grouped into regional data.

The variables used in the pooled time series cross sectional analysis are listed on the following pages and have the structural form shown below.

Data Series Structure

	1970	Value for 1970	1
	1971	Value for 1971	2
Region 1	1972	.	3
	1973	.	4
	1974	.	5
	1970	Value for 1970	6
	1971	.	7
Region 2	1972	.	8
		.	
		.	
		.	
	1970	Value for 1970	41
	1971	.	42
Region 9	1972	.	43
	1973	.	44
	1974	.	45

QUALIFIED MILITARY AVAILABLES

1	87191.764
2	90155.193
3	112555.57
4	122718.31
5	129154.44
6	220439.49
7	344283.48
8	294185.59
9	319433.64
10	361299.26
11	334565.93
12	439473.45
13	477443.93
14	472243.38
15	525249.96
16	99913.859
17	118925.41
18	139395.45
19	191593.33
20	230491.84
21	213375.45
22	253974.39
23	296159.14
24	352449.86
25	357341.54
26	93029.693
27	92759.969
28	114365.65
29	132474.44
30	145949.46
31	135717.53
32	154225.00
33	167551.30
34	194470.00
35	224592.00
36	35194.343
37	93395.753
38	95697.843
39	96197.329
40	95517.303
41	199041.91
42	217575.48
43	256992.46
44	247732.15
45	274719.93

1

MILITARY AVAILABLES

1	126758.00
2	170105.00
3	212369.00
4	226827.00
5	243689.00
6	423920.00
7	508249.00
8	565551.00
9	614232.00
10	694737.00
11	567061.00
12	693955.00
13	809227.00
14	803412.00
15	899254.00
16	167515.00
17	201399.00
18	235755.00
19	324717.00
20	353375.00
21	417795.00
22	497983.00
23	561114.00
24	710686.00
25	721454.00
26	153732.00
27	189304.00
28	234585.00
29	270355.00
30	297854.00
31	271435.00
32	308452.00
33	335132.00
34	388857.00
35	449184.00
36	53543.000
37	143546.00
38	153014.00
39	152907.00
40	152700.00
41	337359.00
42	368772.00
43	435394.00
44	419885.00
45	465627.00

1

	ARMY ADVERTISING EXPENDITURES (IN THOUSANDS)	AVERAGE PAY FOR ARMY ENLISTEES
1	1409.9500	4392.1000
2	9474.6500	5345.8000
3	14574.2500	5929.9000
4	13755.4500	6456.1000
5	19077.7000	6803.4000
6	1409.9500	4392.1000
7	9474.6500	5345.8000
8	14574.2500	5929.9000
9	13755.4500	6456.1000
10	19077.7000	6803.4000
11	1409.9500	4392.1000
12	9474.6500	5345.8000
13	14574.2500	5929.9000
14	13755.4500	6456.1000
15	19077.7000	6803.4000
16	1409.9500	4392.1000
17	9474.6500	5345.8000
18	14574.2500	5929.9000
19	13755.4500	6456.1000
20	19077.7000	6803.4000
21	1409.9500	4392.1000
22	9474.6500	5345.8000
23	14574.2500	5929.9000
24	13755.4500	6456.1000
25	19077.7000	6803.4000
26	1409.9500	4392.1000
27	9474.6500	5345.8000
28	14574.2500	5929.9000
29	13755.4500	6456.1000
30	19077.7000	6803.4000
31	1409.9500	4392.1000
32	9474.6500	5345.8000
33	14574.2500	5929.9000
34	13755.4500	6456.1000
35	19077.7000	6803.4000
36	1409.9500	4392.1000
37	9474.6500	5345.8000
38	14574.2500	5929.9000
39	13755.4500	6456.1000
40	19077.7000	6803.4000
41	1409.9500	4392.1000
42	9474.6500	5345.8000
43	14574.2500	5929.9000
44	13755.4500	6456.1000
45	19077.7000	6803.4000

ARMY RELATIVE PAY VARIABLE

1	.9939905
2	1.0340947
3	1.1773345
4	1.2144658
5	1.1770599
6	.9319698
7	.9901613
8	1.1945585
9	1.2269289
10	1.0337947
11	.7923477
12	1.0394512
13	1.0373263
14	1.0715519
15	1.0311306
16	1.2365917
17	1.2459792
18	1.6932995
19	1.3595589
20	1.6233357
21	1.0532614
22	1.1511209
23	1.3229789
24	1.1579907
25	1.2153269
26	1.2916166
27	1.2734159
28	1.4932991
29	1.3133161
30	1.2320141
31	1.0950791
32	1.0599019
33	1.3456241
34	1.3851012
35	1.2716636
36	1.0333584
37	1.7315314
38	1.2313545
39	.8853675
40	1.1637445
41	.8549757
42	.9743817
43	1.1339946
44	1.0370827
45	.9924604

ARMY MENTAL CATEGORY 1-3 TOTAL VOLUNTEERS PER QMA

1	.0377988
2	.0394352
3	.0376314
4	.0329570
5	.0455898
6	.0374477
7	.0343293
8	.0396951
9	.0354564
10	.0451494
11	.0423551
12	.0453431
13	.0358838
14	.0333157
15	.0435831
16	.0588096
17	.0563323
18	.0678531
19	.0425340
20	.0458339
21	.0430662
22	.0544994
23	.0586021
24	.0609304
25	.0576819
26	.0479690
27	.0537938
28	.0577924
29	.0587783
30	.0675719
31	.0479351
32	.0562997
33	.0506129
34	.0566698
35	.0654229
36	.0628507
37	.0544319
38	.0577741
39	.0518534
40	.0911465
41	.0521532
42	.0600224
43	.0553548
44	.0512622
45	.0564277

3

ARMY MENTAL CATEGORY 3 HIGH SCHOOL GRADUATE VOLUNTEERS PER QMA

1	.0099058
2	.0099599
3	.0103579
4	.0094099
5	.0113507
6	.0077784
7	.0100993
8	.0127593
9	.0115532
10	.0117717
11	.0079485
12	.0113736
13	.0107392
14	.0099344
15	.0103444
16	.0135265
17	.0163139
18	.0179592
19	.0137807
20	.0115407
21	.0097225
22	.0147377
23	.0190653
24	.0135333
25	.0221312
26	.0096923
27	.0160525
28	.0243111
29	.0225942
30	.0222293
31	.0110243
32	.0141573
33	.0190255
34	.0195777
35	.0175423
36	.0125157
37	.0128225
38	.0144713
39	.0170595
40	.0144302
41	.0124643
42	.0163222
43	.0157133
44	.0193255
45	.0179655

ARMY MENTAL CATEGORY 1-2 HIGH SCHOOL GRADUATE VOLUNTEERS PER QMA

1	[illegible]
2	.0091659
3	.0071721
4	[illegible]
5	[illegible]
6	[illegible]
7	.0064736
8	.0076445
9	[illegible]
10	.0107443
11	.0075537
12	[illegible]
13	[illegible]
14	[illegible]
15	[illegible]
16	[illegible]
17	[illegible]
18	[illegible]
19	[illegible]
20	[illegible]
21	[illegible]
22	[illegible]
23	[illegible]
24	[illegible]
25	[illegible]
26	[illegible]
27	[illegible]
28	[illegible]
29	[illegible]
30	[illegible]
31	[illegible]
32	[illegible]
33	[illegible]
34	[illegible]
35	[illegible]
36	[illegible]
37	[illegible]
38	[illegible]
39	[illegible]
40	[illegible]
41	[illegible]
42	[illegible]
43	[illegible]
44	[illegible]
45	[illegible]

ARMY RECRUITERS PER QMA

1	.0014637
2	.0026975
3	.0022547
4	.0019662
5	.0019114
6	.0016325
7	.0024327
8	.0022736
9	.0019374
10	.0014833
11	.0017076
12	.0022655
13	.0019077
14	.0018737
15	.0019133
16	.0021155
17	.0022099
18	.0017665
19	.0027475
20	.0021727
21	.0021736
22	.0025472
23	.0022967
24	.0019568
25	.0017423
26	.0015313
27	.0029777
28	.0023537
29	.0023715
30	.0017466
31	.0017991
32	.0027035
33	.0025291
34	.0022749
35	.0019497
36	.0023356
37	.0024269
38	.0024535
39	.0025123
40	.0027257
41	.0015131
42	.0025029
43	.0024335
44	.0025619
45	.0023121

1

PROPORTION OF BLACKS IN THE YOUTH POPULATION

1	.0467137
2	.0172561
3	.0313961
4	.0175111
5	.0109712
6	.1573675
7	.1101177
8	.0775469
9	.0759516
10	.1017457
11	.1636190
12	.1373356
13	.1353236
14	.0962415
15	.0377777
16	.0774196
17	.0515747
18	.0427163
19	.0755327
20	.0527745
21	.2316690
22	.2303727
23	.2913673
24	.2557699
25	.2312515
26	.2373457
27	.1711293
28	.1744231
29	.1975472
30	.1945407
31	.1992553
32	.1973446
33	.1757599
34	.1120357
35	.0913935
36	.0145520
37	.0219447
38	.0279607
39	.0122671
40	.0515297
41	.0921995
42	.0957573
43	.0563298
44	.0525528
45	.0556583

1

AVERAGE PAY FOR NAVY ENLISTEES

1	4387.8000
2	5349.1000
3	5978.5000
4	6603.3000
5	6768.3000
6	4387.8000
7	5349.1000
8	5578.5000
9	6603.3000
10	6768.3000
11	4387.8000
12	5349.1000
13	5578.5000
14	6603.3000
15	6768.3000
16	4387.8000
17	5349.1000
18	5578.5000
19	6603.3000
20	6768.3000
21	4387.8000
22	5349.1000
23	5578.5000
24	6603.3000
25	6768.3000
26	4387.8000
27	5349.1000
28	5578.5000
29	6603.3000
30	6768.3000
31	4387.8000
32	5349.1000
33	5578.5000
34	6603.3000
35	6768.3000
36	4387.8000
37	5349.1000
38	5578.5000
39	6603.3000
40	6768.3000
41	4387.8000
42	5349.1000
43	5578.5000
44	6603.3000
45	6768.3000

3

NAVY RECRUITERS

1	91.000000
2	91.000000
3	97.000000
4	124.000000
5	113.000000
6	111.000000
7	419.000000
8	419.000000
9	433.000000
10	669.000000
11	573.000000
12	625.000000
13	411.000000
14	411.000000
15	440.000000
16	614.000000
17	626.000000
18	646.000000
19	301.000000
20	301.000000
21	333.000000
22	417.000000
23	396.000000
24	412.000000
25	266.000000
26	266.000000
27	346.000000
28	442.000000
29	457.000000
30	443.000000
31	153.000000
32	153.000000
33	234.000000
34	232.000000
35	240.000000
36	233.000000
37	214.000000
38	214.000000
39	291.000000
40	374.000000
41	392.000000
42	379.000000
43	98.000000
44	98.000000
45	135.000000

2

NAVY MENTAL CATEGORY 1-2 HIGH SCHOOL GRADUATE VOLUNTEERS PER QMA

1	.0086710
2	.0130111
3	.0059724
4	.0095358
5	.0119469
6	.0056569
7	.0058610
8	.0093298
9	.0089871
10	.0082551
11	.0110023
12	.0079183
13	.0040875
14	.0083830
15	.0084630
16	.0437350
17	.0465553
18	.0304539
19	.0075008
20	.0133316
21	.0143406
22	.0051264
23	.0050786
24	.0056808
25	.0029177
26	.0223922
27	.0225307
28	.0199745
29	.0225402
30	.0189382
31	.0025120
32	.0053360
33	.0050931
34	.0050254
35	.0054009
36	.0179581
37	.0128967
38	.0219708
39	.0254638
40	.0236452
41	.0126607
42	.0103688
43	.0032255
44	.0067549
45	.0068458
	4

NAVY ADVERTISING EXPENDITURES (IN MILLIONS)

1	.8000000
2	.5000000
3	1.6000000
4	11.5000000
5	13.4000000
6	.8000000
7	.5000000
8	1.6000000
9	11.5000000
10	13.4000000
11	.8000000
12	.5000000
13	1.6000000
14	11.5000000
15	13.4000000
16	.8000000
17	.5000000
18	1.6000000
19	11.5000000
20	13.4000000
21	.8000000
22	.5000000
23	1.6000000
24	11.5000000
25	13.4000000
26	.8000000
27	.5000000
28	1.6000000
29	11.5000000
30	13.4000000
31	.8000000
32	.5000000
33	1.6000000
34	11.5000000
35	13.4000000
36	.8000000
37	.5000000
38	1.6000000
39	11.5000000
40	13.4000000
41	.8000000
42	.5000000
43	1.6000000
44	11.5000000
45	13.4000000
	1

NAVY MENTAL CATEGORY 3 HIGH SCHOOL GRADUATE VOLUNTEERS PER QMA

1	.0076821
2	.0110522
3	.0097044
4	.0091373
5	.0104361
6	.0163147
7	.0155330
8	.0095667
9	.0101663
10	.0076262
11	.0105654
12	.0091550
13	.0039462
14	.0077951
15	.0037562
16	.0475102
17	.0421795
18	.0357309
19	.0053417
20	.0031178
21	.0015151
22	.0080318
23	.0070055
24	.0050352
25	.0030271
26	.0156935
27	.0322323
28	.0243486
29	.0266617
30	.0223024
31	.0032620
32	.0154552
33	.0079569
34	.0078849
35	.0081481
36	.0167653
37	.0213423
38	.0214061
39	.0291171
40	.0327680
41	.0392223
42	.0136550
43	.0021357
44	.0049095
45	.0054725
	3

NAVY MENTAL CATEGORY 1-3 NON-HIGH SCHOOL GRADUATE VOLUNTEERS PER QMA

1	.0122532
2	.0059653
3	.0050396
4	.0101777
5	.0172739
6	.0031400
7	.0074399
8	.0072833
9	.0075984
10	.0105409
11	.0188244
12	.0057348
13	.0045128
14	.0045859
15	.0065947
16	.0444110
17	.0591793
18	.0184693
19	.0062721
20	.0058936
21	.0020114
22	.0035939
23	.0021665
24	.0025652
25	.0029363
26	.0174567
27	.0280165
28	.0283326
29	.0400455
30	.0055956
31	.0037606
32	.0051411
33	.0070716
34	.0076052
35	.0109755
36	.0101841
37	.0153713
38	.0142683
39	.0221971
40	.0259507
41	.0180867
42	.0023781
43	.0025675
44	.0026162
45	.0038912
	3

	NAVY RECRUITERS PER QMA	YOUTH UNEMPLOYMENT RATE
1	.0013545	6.2000000
2	.0010094	13.400000
3	.0008618	10.200000
4	.0010315	4.5000000
5	.0008517	7.2000000
6	.0005039	8.2000000
7	.0015354	12.200000
8	.0014248	12.600000
9	.0013557	9.2000000
10	.0019268	10.200000
11	.0017037	5.5000000
12	.0015265	6.3000000
13	.0008608	13.700000
14	.0008703	8.6000000
15	.0008377	7.5000000
16	.0062[illegible]24	5.5000000
17	.0052682	5.9000000
18	.0046443	6.2000000
19	.0015711	5.4000000
20	.0014437	5.3000000
21	.0015628	3.9000000
22	.0016419	8.0000000
23	.0013839	7.0000000
24	.0011367	5.1000000
25	.0007229	7.3000000
26	.0029546	3.5000000
27	.0037301	4.2000000
28	.0038453	7.0000000
29	.0034497	4.4000000
30	.0030698	7.2000000
31	.0011273	6.9000000
32	.0009921	6.7000000
33	.0013966	7.3000000
34	.0011932	3.6000000
35	.0010686	5.4000000
36	.0043129	5.4000000
37	.0026622	12.500000
38	.0024974	14.000000
39	.0032603	7.1000000
40	.0043737	10.200000
41	.0019694	12.700000
42	.0017419	15.300000
43	.0003815	13.100000
44	.0003956	10.100000
45	.0003822	11.200000
	1	4

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NAVY RELATIVE PAY VARIABLE

1	.9999116
2	1.0367145
3	1.1455912
4	1.2621559
5	1.1773467
6	.9339958
7	.9937566
8	1.1367459
9	1.2563331
10	1.3991511
11	.7915919
12	1.0393779
13	1.1354764
14	1.0359974
15	1.0259139
16	1.2556536
17	1.2677609
18	1.7371477
19	1.6337969
20	1.6149605
21	1.0372939
22	1.1919775
23	1.3199733
24	1.1365199
25	1.2370569
26	1.2937519
27	1.2762320
28	1.5325132
29	1.3679537
30	1.1358127
31	1.0153169
32	1.3704523
33	1.3514771
34	1.6364577
35	1.2351028
36	1.0990617
37	1.9796769
38	1.7311997
39	.9355563
40	[illegible]
41	.8517906
42	.9949779
43	1.1179076
44	1.0766835
45	.9777719

CIVILIAN PAY VARIABLE

1	4437.0000
2	5324.0000
3	5038.0000
4	5316.0000
5	5780.0000
6	4870.0000
7	5454.0000
8	5006.0000
9	5262.0000
10	6220.0000
11	5614.0000
12	5701.0000
13	5431.0000
14	6025.0000
15	6558.0000
16	3435.0000
17	4247.0000
18	3592.0000
19	4714.0000
20	4191.0000
21	4170.0000
22	4604.0000
23	4452.0000
24	5528.0000
25	5594.0000
26	3477.0000
27	4194.0000
28	3979.0000
29	4917.0000
30	5660.0000
31	4044.0000
32	4997.0000
33	4397.0000
34	4695.0000
35	5350.0000
36	4227.0000
37	4107.0000
38	4592.0000
39	7292.0000
40	5964.0000
41	5093.0000
42	5911.0000
43	5446.0000
44	6133.0000
45	6925.0000

Appendix D

DETERMINATION OF TRUE VOLUNTEERS AND MENTAL CATEGORY GROUPS

TRUE VOLUNTEERS

Monthly estimates of true volunteer enlistments for the months prior to July 1973 were made from an analysis of the lottery numbers of accessions for each month. Volunteer accessions were calculated from the following formula:

$$E_k = \frac{365}{365-240} \left(\sum_{i=241}^{365} a_{ik} \right) + a_o$$

where

E_k = true volunteers for month k.

a_{ik} = total accessions with lottery number i in month k.

a_o = total accessions who enlisted before the lottery draw for their age group.

By use of this formula, it is assumed that youth with lottery numbers greater than 240 felt no draft pressure and are true volunteers. The formula was derived by an analysis of typical distributions of enlistees by lottery number. Enlistees who enter a branch of service prior to publication of their lottery numbers and enlistees still 17 years of age are assumed to be true volunteers.

MENTAL CATEGORY GROUPS

Classifications of enlistees by mental category are made by use of the Armed Forces Qualifications Test (AFQT) scores present on each enlistee's USAREC accession file. The scores for the separate categories are listed below:

Mental Category Group	AFQT Score Percentile
1	93-100
2	66-92
3	31-65
4	11-30
5	0-10

Appendix E

GEOGRAPHIC REGIONS

There are four census regions which are distributed into nine geographic divisions. These nine divisions and the states included within them are presented below.

Northeast Region

1. New England Division (NE)

   Connecticut
   Maine
   Massachusetts
   New Hampshire
   Rhode Island
   Vermont

2. Middle Atlantic Division (MA)

   New Jersey
   New York
   Pennsylvania

North Central Region

3. East North Central Division (ENC)

   Illinois
   Indiana
   Michigan
   Ohio
   Wisconsin

4. West North Central Division (WNC)

   Iowa
   Kansas
   Minnesota
   Missouri
   Nebraska
   North Dakota
   South Dakota

South Region

5. South Atlantic Division (SA)

   Delaware
   District of Columbia
   Florida
   Georgia
   Maryland
   North Carolina
   South Carolina
   Virginia
   West Virginia

South Region (continued)

6. East South Central Division (ESC)

   Alabama
   Kentucky
   Mississippi
   Tennessee

7. West South Central Division (WSC)

   Arkansas
   Louisiana
   Oklahoma
   Texas

West Region

8. Mountain Division (MD)

   Arizona
   Colorado
   Idaho
   Montana
   Nevada
   New Mexico
   Utah
   Wyoming

9. Pacific Division (PD)

   Alaska
   California
   Hawaii
   Oregon
   Washington