SMU ECONOMICS & STATISTICS
WORKING PAPER SERIES

Non-market Leadership Experience and Labor Market Success: Evidence From Military Rank

Myoung-Jae Lee, Yip Chun Seng

May 2005

Paper No. 12-2005

Non-market Leadership Experience and Labor Market Success: Evidence From Military Rank.

Myoung-Jae Lee and **Chun Seng Yip***

Singapore Management University

May 31, 2005

Abstract

There has been much recent interest in the effects of pre and non-market skills on future labor market outcomes. This paper examines one such effect: the effect on future wages of military leadership experience among "Vietnam generation" American men. We study rank, not just veteran status. We argue that rank is a good measure of pre-market leadership skills because of the clear military hierarchy and the primarily youth experience of Vietnam service. Two sources of selection bias are accounted for: non-random military entry and eventual rank attained. We apply a modified 2-stage parametric sample selection method. The rank premia on future wages are estimated using the parametric selection corrections and a propensity score matching with two indices. We find evidence of a leadership premium, but not a veterans' premium. It is the rank that matters. If one joins the military believing that military service commands a future wage premium, he had better become an NCO or an officer. **JEL**: J24, J10.

*Corresponding author: Chun Seng Yip, Sch. of Economics and Social Sciences, Singapore Management University, 90 Stamford Road, Singapore 178903. **Email: yipcs@smu.edu.sg.** This research is funded by the SMU Office of Research.

I am certain that ...[universal military] training ... will not only make our youth better equipped to serve their country, but better *mentally, morally*, and physically.

President Harry S Truman, 1947 Commencement Address, Princeton University.

1 Introduction

In recent years, there has been growing interest among economists in the relation between labor market outcomes and various forms of non-schooling human capital. Specifically we are interested in intangible qualities not measurable by classroom learning attainment, cognitive tests, receipt of a diploma/degree, or acquisition of specific job skills through training. Some examples are qualities such as discipline, responsibility, leadership, and motivation — qualities that are hard to quantify or observe, and for which data is often lacking. In the economics literature, these qualities are sometimes referred to as "non-cognitive" skills (e.g. Heckman (2000)).

Why might looking at such non-schooling qualities be a profitable direction to explain labor market success? We can think of three reasons. One reason is the emerging consensus that "traditional" predictors of labor market outcomes, such as own and parental schooling, labor market experience, occupation and industry characteristics and so on, account for less than half of observed wage variation. Much of the observed wage variation remains relegated to the realm of "unobserved heterogeneity" (Bowles, Gintis and Osborne (2001) and Abowd et al. (1999), for example). A second reason is that recent evidence from employer surveys finds employers caring more about qualities such as "attitude" and "motivation" than about schooling attainment in hiring decisions. (See, e.g. Green et al., (1998)). A third reason is evidence that some aspects of adolescent experience matter for future labor market success. Participation in high school sports and experience in a high-school leadership position have been found to be related to higher future wages, while deviant and delinquent behavior are correlated with lower future wages. (See Kuhn and Weinberger (2005), Anderson (2000), Cawley et al. (2000), and Barron et al., (2000)).

This paper seeks to understand the nature of the relationship between "non-cognitive" skills and labor market success, using military rank as the measure of a quality such as "leadership" and "discipline". In particular, we use data on Vietnam-era youths (including veterans and non-veterans) to study whether the highest attained military rank has further labor market impact in subsequent civilian life.

Why Study Leadership, Military and Rank? We consider rank and veteran experience particularly suitable measures of non-cognitive factors. Our motivation arises from the observation that leadership qualities are highly valued in all sectors of work, public or private. While this appears to be conventional wisdom, we have not encountered much work that relates leadership to what we think of as human capital. The literature cited above represents the two related views that (1) non-cognitive human capital and (2) pre-labor market factors such as late-adolescent experiences are significant determinants of subsequent labor market success. Military experience, we believe, lies somewhere between these two views. The military provides training in discipline and motivation that is more rigorous than high school leadership experience or atheletics. Furthermore, rank is an easily observable measure that is likely to be highly correlated with actual leadership experience. Higher rank implies greater responsibilities and leadership. Finally, military experience, at least in the sample we consider, is largely constrained to late teenage, post-high school years. From this, we argue that by considering the military experience of young men, we essentially consider the effect of pre-labor-market experience.

Empirical Approach We use data from the National Longitudinal Surveys' 1966 Young Men Cohort (NLSYM). This data covers the experiences of the very cohorts known as the "Vietnam generation" and contains substantial information regarding individuals' experiences with the military and the draft system during that time. A well known empirical challenge is that of non-random selection into the military. To that, our research question adds the selection problem of the eventual rank attained. These challenges arise from two sources: on the one hand, there is heterogeneity in military ability or interest, while on the other hand, it is well known that loopholes in the draft system allowed some eligible males to avoid conscription (Baskir and Strauss (1978) and Foley (2003) provide a good background to the draft system and avoidance behavior).

In addition to the economic contribution, we also make a methodological contribution. We present ways to account for the above two sources of selectivity: (1) selection into the military; and (2) selection on the eventual rank attained. We do so by developing and estimating a simultaneous probit and ordered probit for military service and rank respectively. From this we derive parametric selection correction techniques and two-dimensional propensity-score matching techniques to address the question of the rank premium.

Summary of Results Our findings are twofold. One methodological, the other substantive. We find that parametric selection correction fails to display evidence of a premium on rank. We find also that the selection correction terms are jointly not significant. This suggests that a least squares regression of a Mincer type wage equation augmented to include rank dummies, may suffer from less selection bias than originally thought. Our matching method delivers more promising results. In a variety of matching techniques, we find that the rank wage premium is absent among privates, but is positive and significant among corporals and sergeants. Thus we take this as evidence that veteran service *per se*, is *not* enough to generate a future wage premium; one needs to have been a veteran *and* a leader. We consider this our main finding, supporting the view that non-cognitive skills matter for the labor force.

Discussion of Related Work This paper draws together two strands of empirical work in labor economics. The first strand is that of the growing literature on non-cognitive skills that has been discussed earlier. The other literature is that of the veterans' wage premium, elaborated in the next section. The early papers mostly used ordinary least squares approaches to ask whether veteran status affects future wages. Subsequent work applied various methods such as instrumental variables and matching methods to account for the nonrandom selection. There have not been too many studies in this latter vein, and hence it has not been easy to draw firm conclusions.

The closest work to ours are Kuhn and Weinberger (2005) and Hirsch and Mehay (2003). Kuhn and Weinberger present evidence that teenage leadership experience relates to future labor market success. Hirsch and Mehay do matching to compare active duty servicemen with reserve component servicemen. They find a small veterans' premium of around 3 percent. However their matching method is restricted to a limited set of covariates, namely age and race, and they do not control for selection into eventual rank. Interestingly they report a larger veteran's premium among officers suggesting some evidence of greater returns to leadership and responsibility.

While our paper relates to the previous works, it is different in several ways. Firstly, whereas veteran status is a binary state in previous work, rank is a measure of leadership. We argue, if there is a wage premium among veterans, it is rank (and thus a non-cognitive skill such as leadership), and not veteran status *per se*, that generates heterogeneity in military returns. Secondly, we fill a gap in the literature by applying propensity score matching and thus making use of a higher dimensioned set of covariates, including longitudinal information.

The rest of the paper is laid out as follows. Section 2 presents the background and motivation. Section 3 shows our parametric approach. Section 4 presents our two-dimensional propensity-score matching approach. Section 5 discusses the data and provides the results. Section 6 discusses and concludes.

2 Background and Motivation

Our motivation arises from the observation that leadership qualities are valued in civilian labor markets. For example, a search of articles in the Harvard Business Review with the keyword "leadership" yielded 316 hits. A search at the websites of the top ranked American Business Schools show that leadership courses increasingly form part of the core curriculum. Interestingly Kuhn and Weinberger (2005) cite examples of top business schools incorporating Marine "boot camps" as part of the MBA education. This suggests belief that business leadership and military leadership skills are correlated.

American society at large appears to value military service. A distinguished service record opens lucrative civilian job offers, while questionable service records can imply character flaws that can haunt the individual many years. It is revealing that individuals seeking political office are often scrutinized for their military experience. For instance, of the eleven American presidents since WWII, eight performed wartime service. The two most recent Presidents Bill Clinton and George W. Bush were frequently dogged by calls to account for their failure to serve in Vietnam.

Previous work has found that participation in high school athletics and leadership of clubs and societies have a positive effect on future wages (Kuhn and Weinberger (2005), Eide and Ronan (2000), Barron et al. (2000) and Postlewaite and Silverman (2005)). It is argued that better future labor market outcomes are related to the non-cognitive skills such as leadership, personal motivation, discipline developed through these activities. Relatedly Heckman and Carneiro (2004) report that early childhood intervention and teenage mentoring programs often uncover substantial program impact on non-cognitive skills that are stronger than effects on mesurable cognitives. Taken together, we see a growing body of evidence pointing to the significance of pre-market non-cognitive skills in human capital.

Many studies on military service suggest the importance of military experience for future wages, but few have approached it from the viewpoint of non-cognitive skills. Most work comes from America's Vietnam experience, while some are from WWII. The results appear mixed. The earlier work applies OLS to wage regressions incorporating a dummy variable for veteran status (For example De Tray (1982), Berger and Hirsch (1983) and Goldberg and Warner (1986).). On average the results suggest a positive wage premium for World War II veterans, but somewhat weaker evidence on Vietnam veterans. However as pointed out by Hirsch and Mehay (2003) the problem of selection bias makes it hard to draw clear conclusions. The more recent studies, notably by Angrist (1989, 1990), Angrist and Krueger (1994), Hirsch and Mehay (2003) that provide more robust controls for military selection, find generally small benefits to veteran status and on occasion finds wage penalties. Angrist (1990) used the draft lottery number as an instrument for military service. While he tries a variety of methods, his most notable contribution is to separate the population into two groups: the "at risk" group consisting of those whose draft number falls below the announced threshold, and those not at risk, i.e. with numbers above the threshold. Thus he constructs a "Wald" estimator of the military effect, with a binary instrumental variable. Hirsch and Mehay (2003) adopt a different approach. They perform a matching estimation to estimate the effect of treatment on the treated (TT). However rather than using propensity score methods, they deliberately keep matching to a limited set of covariates, namely by age and race. This, they argue, is due to the fact that they use a data set consisting of reservists, and in that sample, age is the major way in which veterans and nonveterans differ.

A dummy variable for veteran status seems inadequate to capture the military experience. Many different military specializations and the echelons of leadership and responsibility suggest that the military experience is far from homogeneous. Consider the case of rank. The experience of a private is different than that of a commissioned officer or NCO. The private is primarily a follower, whereas the NCO and officer are leaders responsible for increasing numbers of troops. A corporal is often in charge of around 5 soldiers. A sergeant would normally lead upwards of 10 troops, a lieutenant would be a leader of 40 troops, and a captain would lead 150 to 200 troops. Clearly it is harder to be a lieutenant than a private. Those promoted to the higher ranks would normally have displayed the potential to assume greater responsibility. Being exposed to higher levels of responsibility in turn affords many opportunities to further develop non-cognitive skills of the sort under investigation.

If our view is correct, then we should detect premia not simply to veteran status, but to leadership experience (i.e. rank in our case). Considering that high school leadership has significant effects on future wages, we should expect military leadership experience, which is more challenging, to have at least as much effect on future wages.

Our main challenge is that we face a non-standard sample selection problem for which a solution needs to be developed. It is that we must account for entry into the military, as well as the rank attained. This is what we turn to in the next two sections. We present a parametric approach and a non-parametric approach. The former is based upon parametric sample selection, with a non-standard first stage; we develop parametric selection correction terms for the second stage. The nonparametric approach is propensity-score matching using the two indices estimated via the first stage.

3 Two-Stage Parametric Approach

We adopt a two-stage parametric approach. In the first stage, selection is based upon two events. The first selection denotes military entry, and the second selection is an ordered discrete response (ODR) representing rank finally attained. In the second stage, we estimate a log-wage equation of the usual form.

Let d_{i1} be a binary variable denoting military entry for individual i, and d_{i2} is ordered response denoting the rank such that $\{1, 2, 3, 4\}$ corresponds to {private, corporal, sergeant, officer}. Define indicator function $1[A] = 1$ if A holds and 0 otherwise. The empirical model is as follows:

$$\begin{aligned}
\text{First Selection} \quad &: \quad d_{i1}^* = w_i'\alpha_1 + \varepsilon_{i1}, \quad d_{i1} = 1[d_{i1}^* > 0], \\
\text{Second Selection} \quad &: \quad d_{i2}^* = w_i'\alpha_2 + \varepsilon_{i2}, \quad d_{i2} = \sum_{r=1}^{3} 1[d_{i2}^* \geq \gamma_r], \quad \gamma_1 = 0, \\
\text{Log-wage} \quad &: \quad y_i = x_i'\beta + u_i, \quad E(\varepsilon_1) = E(\varepsilon_2) = E(u) = 0
\end{aligned} \tag{1}$$

where the regressor w may include x and the second selection equation is ODR with unknown thresholds γ_2 and γ_3. What is observed is

$$(d_{i1},\ d_{i1}d_{i2},\ w_i',\ y_i)', i = 1, ..., N,\ iid;$$

d_{i2} is observed only when $d_{i1} = 1$. In view of the iid assumption, we will often omit the subscript

i in the following. Defining $\varepsilon \equiv (\varepsilon_1, \varepsilon_2)'$, further assume

$$\text{Normality of } \varepsilon \quad : \quad \varepsilon \text{ follows } N(0, \Omega), \ \Omega \equiv \begin{bmatrix} 1 & \rho_{12} \\ \rho_{12} & 1 \end{bmatrix}, \text{ independently of } w,$$

$$\text{Linear Projection } u|\varepsilon \quad : \quad E(u|w, \varepsilon) = (\theta_1, \theta_2)\varepsilon, \text{ where } (\theta_1, \theta_2) = E(u\varepsilon') \cdot \{E(\varepsilon\varepsilon')\}^{-1}.$$

The joint estimation of the parameters with MLE is difficult because we have to estimate three correlation parameters, ρ_{12}, $\rho_{1u} \equiv COR(\varepsilon_1, u)$, and $\rho_{2u} \equiv COR(\varepsilon_2, u)$; estimating covariance matrix is often the source of non-convergence in numerical optimization. Also the MLE requires the joint normality of $(\varepsilon_1, \varepsilon_2, u)$ that is stronger than the combination of the normality of ε and the linear projection assumption. Instead, we can proceed in the following two stages. First, estimate $(\alpha_1', \alpha_2', \rho_{12}, \gamma_2, \gamma_3)$ with the MLE for $(d_1, d_1 d_2)$; denote the estimator as $(a_1', a_2', r_{12}, g_2, g_3)$. Second, do a Heckman type least squares estimation of y on x and the selection correction terms that are known functions of two indices $w'a_1$ and $w'a_2$.

For this procedure, we need to derive the likelihood function for the first step and the selection correction terms for the second step. The main complication here is the selection problem of the ODR d_2 observed only when $d_1 = 1$. Although an analogous selection problem with a binary d_2 appeared in Van de Ven and Van Praag (1981) and Dubin and Rivers (1989), our type of the selection problem has not appeared in the literature as far as we are aware of. As shown in the following, the first step is relatively straightforward, but the closed form formulas are difficult to obtain for some selection correction terms; for these, we will use numerically computed selection correction terms.

3.1 Likelihood Function for the First Stage

Define $\psi(\varepsilon_1, \varepsilon_2, \rho_{12})$ as the standard joint normal density function with correlation ρ_{12}, and

$$\Psi(\varepsilon_1, \varepsilon_2, \rho_{12}) \equiv \int_{-\infty}^{\varepsilon_2} \int_{-\infty}^{\varepsilon_1} \psi(t_1, t_2, \rho_{12}) dt_1 dt_2.$$

The log-likelihood function to maximize for $(\alpha_1, \alpha_2, \rho_{12}, \gamma_2, \gamma_3)$ consists of five terms corresponding to the five cases: $d_1 = 0$, $(d_1 = 1, d_2 = 0)$, $(d_1 = 1, d_2 = 1)$, $(d_1 = 1, d_2 = 2)$, and $(d_1 = 1, d_2 = 3)$.

The log-likelihood function is

$$\sum_{i=1}^{N}[\ (1-d_{i1})\cdot\ln\Phi(-w_i'\alpha_1) \tag{2}$$
$$+d_{i1}1\left[d_{i2}=0\right]\cdot\ln P\left(-w_i'\alpha_1<\varepsilon_{i1},\ \varepsilon_{i2}<-w_i'\alpha_2\right)$$
$$+d_{i1}1\left[d_{i2}=1\right]\cdot\ln P\left(-w_i'\alpha_1<\varepsilon_{i1},\ -w_i'\alpha_2<\varepsilon_{i2}<-w_i'\alpha_2+\gamma_2\right)$$
$$+d_{i1}1\left[d_{i2}=2\right]\cdot\ln P\left(-w_i'\alpha_1<\varepsilon_{i1},\ -w_i'\alpha_2+\gamma_2<\varepsilon_{i2}<-w_i'\alpha_2+\gamma_3\right)$$
$$+d_{i1}1\left[d_{i2}=3\right]\cdot\ln P\left(-w_i'\alpha_1<\varepsilon_{i1},\ -w_i'\alpha_2+\gamma_3<\varepsilon_{i2}\right).$$

A detailed explanation of the derivation of this likelihood function is in the appendix.

3.2 Correction Term for the Second Stage

The selection correction term for the case $d_1 = 0$ is easy to derive. As is well known, with $\sigma_u \equiv SD(u)$,

$$E(y|w, d_1 = 0) = x'\beta + E(u|w, \varepsilon_1 < -w'\alpha_1) = x'\beta - \rho_{1u}\sigma_u\frac{\phi(-w'\alpha_1)}{\Phi(-w'\alpha_1)}. \tag{3}$$

More difficult to derive are the selection correction terms for the $d_1 = 1$ cases. We use the linear projection assumption of $u|\varepsilon$. The correction terms are presented here while the details are left to the appendix. Define the correction term in (3) as $\lambda_c(\alpha_1, \alpha_2, \rho_{12}, \gamma_2, \gamma_3)$, and $\lambda_{jk}(\alpha_1, \alpha_2, \rho_{12}, \gamma_2, \gamma_3)$, $j = 0, 1, 2, 3$ (denoting rank), and $k = 1, 2$ (denoting first or second correction) for $d_1 = 1$ such that

non veterans : $E(u|w, d_1 = 0) = -\rho_{1u}\sigma_u \cdot \lambda_c = -\sigma_{1u}\lambda_c \ where\ \sigma_{1u} \equiv COV(\varepsilon_1, u)$,

privates : $E(u|w, d_1 = 1, d_2 = 0) = \theta_1\lambda_{01} + \theta_2\lambda_{02}$,

corporals : $E(u|w, d_1 = 1, d_2 = 1) = \theta_1\lambda_{11} + \theta_2\lambda_{12}$,

sergeants : $E(u|w, d_1 = 1, d_2 = 2) = \theta_1\lambda_{21} + \theta_2\lambda_{22}$,

officers : $E(u|w, d_1 = 1, d_2 = 3) = \theta_1\lambda_{31} + \theta_2\lambda_{32}$;

where for privates,

$$\lambda_{01} = E\left(\varepsilon_1|w, -w'\alpha_1 < \varepsilon_1, \varepsilon_2 < -w'\alpha_2\right) = \frac{\int_{-\infty}^{-w'\alpha_2}\int_{-w'\alpha_1}^{\infty}\varepsilon_1\psi\left(\varepsilon_1, \varepsilon_2, \rho_{12}\right)d\varepsilon_1 d\varepsilon_2}{\int_{-\infty}^{-w'\alpha_2}\int_{-w'\alpha_1}^{\infty}\psi\left(\varepsilon_1, \varepsilon_2, \rho_{12}\right)d\varepsilon_1 d\varepsilon_2}, \tag{4}$$

$$\lambda_{02} = E\left(\varepsilon_2 | w, -w'\alpha_1 < \varepsilon_1, \varepsilon_2 < -w'\alpha_2\right) = \frac{\int_{-\infty}^{-w'\alpha_2} \int_{-w'\alpha_1}^{\infty} \varepsilon_2 \psi\left(\varepsilon_1, \varepsilon_2, \rho_{12}\right) d\varepsilon_1 d\varepsilon_2}{\int_{-\infty}^{-w'\alpha_2} \int_{-w'\alpha_1}^{\infty} \psi\left(\varepsilon_1, \varepsilon_2, \rho_{12}\right) d\varepsilon_1 d\varepsilon_2}; \tag{5}$$

for corporals,

$$\lambda_{11} = E\left(\varepsilon_1 | w, -w'\alpha_1 < \varepsilon_1, -w'\alpha_2 < \varepsilon_2 < -w'\alpha_2 + \gamma_2\right) = \frac{\int_{-w'\alpha_2}^{-w'\alpha_2+\gamma_2} \int_{-w'\alpha_1}^{\infty} \varepsilon_1 \psi\left(\varepsilon_1, \varepsilon_2, \rho_{12}\right) d\varepsilon_1 d\varepsilon_2}{\int_{-w'\alpha_2}^{-w'\alpha_2+\gamma_2} \int_{-w'\alpha_1}^{\infty} \psi\left(\varepsilon_1, \varepsilon_2, \rho_{12}\right) d\varepsilon_1 d\varepsilon_2;}, \tag{6}$$

$$\lambda_{12} = E\left(\varepsilon_2 | w, -w'\alpha_1 < \varepsilon_1, -w'\alpha_2 < \varepsilon_2 < -w'\alpha_2 + \gamma_2\right) = \frac{\int_{-w'\alpha_2}^{-w'\alpha_2+\gamma_2} \int_{-w'\alpha_1}^{\infty} \varepsilon_2 \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2}{\int_{-w'\alpha_2}^{-w'\alpha_2+\gamma_2} \int_{-w'\alpha_1}^{\infty} \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2}; \tag{7}$$

for sergeants,

$$\lambda_{21} = E\left(\varepsilon_1 | w, -w'\alpha_1 < \varepsilon_1, -w'\alpha_2 + \gamma_2 < \varepsilon_2 < -w'\alpha_2 + \gamma_3\right) = \frac{\int_{-w'\alpha_2+\gamma_2}^{-w'\alpha_2+\gamma_3} \int_{-w'\alpha_1}^{\infty} \varepsilon_1 \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2}{\int_{-w'\alpha_2+\gamma_2}^{-w'\alpha_2+\gamma_3} \int_{-w'\alpha_1}^{\infty} \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2}, \tag{8}$$

$$\lambda_{22} = E\left(\varepsilon_2 | w, -w'\alpha_1 < \varepsilon_1, -w'\alpha_2 < \varepsilon_2 < -w'\alpha_2 + \gamma_2\right) = \frac{\int_{-w'\alpha_2+\gamma_2}^{-w'\alpha_2+\gamma_3} \int_{-w'\alpha_1}^{\infty} \varepsilon_2 \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2}{\int_{-w'\alpha_2+\gamma_2}^{-w'\alpha_2+\gamma_3} \int_{-w'\alpha_1}^{\infty} \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2.}; \tag{9}$$

and for officers,

$$\lambda_{31} = E\left(\varepsilon_1 | w, -w'\alpha_1 < \varepsilon_1, -w'\alpha_2 + \gamma_3 < \varepsilon_2\right) = \frac{\int_{-w'\alpha_2+\gamma_3}^{\infty} \int_{-w'\alpha_1}^{\infty} \varepsilon_1 \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2}{\int_{-w'\alpha_2+\gamma_3}^{\infty} \int_{-w'\alpha_1}^{\infty} \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2}, \tag{10}$$

$$\lambda_{32} = E\left(\varepsilon_2 | w, -w'\alpha_1 < \varepsilon_1, -w'\alpha_2 + \gamma_3 < \varepsilon_2\right) = \frac{\int_{-w'\alpha_2+\gamma_3}^{\infty} \int_{-w'\alpha_1}^{\infty} \varepsilon_2 \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2}{\int_{-w'\alpha_2+\gamma_3}^{\infty} \int_{-w'\alpha_1}^{\infty} \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2}. \tag{11}$$

3.3 Second-Stage OLS Estimation

We next define five dummy variables $\delta_c, \delta_0, ..., \delta_3$ for the five categories (e.g., δ_c is for $d_1 = 0$, δ_0 is for ($d_1 = 1$, $d_2 = 0$),...), and its category number

$$\delta \equiv \delta_c + 2\delta_0 + 3\delta_1 + 4\delta_2 + 5\delta_3.$$

Define $\tilde{m}_i$ and β such that

$$x_i = (\delta_{ci}, \delta_{0i}, \delta_{1i}, \delta_{2i}, \delta_{3i}, \tilde{m}_i')' \quad \textit{and } \beta = (\beta_c, \beta_0, \beta_1, \beta_2, \beta_3, \beta_{\tilde{m}}')';$$

here $\tilde{m}$ does not include the unity but it may include interaction terms between $\delta_c, \delta_0, ..., \delta_3$ and some covariates. Rewrite the outcome equation as

$$y = \delta_c\beta_c + \delta_0\beta_0 + \delta_1\beta_1 + \delta_2\beta_2 + \delta_3\beta_3 + \tilde{m}'\beta_{\tilde{m}}$$
$$+ \left\{ \delta_c E(u|w, \delta_c = 1) + \sum_{j=0}^{3} \delta_j E(u|w, \delta_j = 1) \right\} + \left[u - \delta_c E(u|w, \delta_c = 1) - \sum_{j=0}^{3} \delta_j E(u|w, \delta_j = 1) \right].$$

The part in $\{\cdot\}$ is the 'selection correction' term, and the part in $[\cdot]$ is the error term. The selection correction term can be written as

$$-\delta_c\sigma_{1u}\lambda_c + \sum_{j=0}^{3} \delta_j(\theta_1\lambda_{j1} + \theta_2\lambda_{j2}) = -\sigma_{1u} \cdot \delta_c\lambda_c + \theta_1 \sum_{j=0}^{3} \delta_j\lambda_{j1} + \theta_2 \sum_{j=0}^{3} \delta_j\lambda_{j2}.$$

Further define

$$v \equiv u - \delta_c E(u|w, \delta_c = 1) - \sum_{j=0}^{3} \delta_j E(u|w, \delta_j = 1)$$

to rewrite the outcome equation as

$$y = \delta_c\beta_c + \delta_0\beta_0 + \delta_1\beta_1 + \delta_2\beta_2 + \delta_3\beta_3 + \tilde{m}'\beta_{\tilde{m}} - \sigma_{1u}\delta_c\lambda_c + \theta_1 \sum_{j=0}^{3} \delta_j\lambda_{j1} + \theta_2 \sum_{j=0}^{3} \delta_j\lambda_{j2} + v.$$

Observe that

$$E(v|w, \delta = j) = E(v|w, \delta_j = 1) = E(u|w, \delta_j = 1) - E(u|w, \delta_j = 1) = 0 \quad \textit{for } j = c, 0, 1, 2, 3,$$

which justifies OLS estimation for the outcome equation. The identified parameters are

$$\beta', \ -\sigma_{1u}, \ \theta_1, \theta_2.$$

For our empirical analysis, instead of the five dummies $\delta_c, \delta_0, ..., \delta_3$, we will use only the last four $\delta_0, ..., \delta_3$ along with the unity. In this case, substituting $\delta_c = 1 - \delta_0 - \delta_1 - \delta_2 - \delta_3$ into the above outcome equation, we get

$$\begin{aligned} y &= \delta_0(\beta_0 - \beta_c) + \delta_1(\beta_1 - \beta_c) + \delta_2(\beta_2 - \beta_c) + \delta_3(\beta_3 - \beta_c) + m'\beta_m \\ &\quad -\sigma_{1u}\delta_c\lambda_c + \theta_1 \sum_{j=0}^{3} \delta_j\lambda_{j1} + \theta_2 \sum_{j=0}^{3} \delta_j\lambda_{j2} + v \\ \textit{where } m &\equiv (1, \tilde{m}')' \ \textit{ and } \ \beta_m = (\beta_c, \beta'_{\tilde{m}})'. \end{aligned} \tag{12}$$

Interaction terms between δ_c and covariates in $\tilde{m}$ are also removed by substituting $\delta_c = 1-\delta_0-\delta_1-\delta_2-\delta_3$; i.e., m includes interaction terms between covariates and $\delta_0, ..., \delta_3$ only. The parameters for $\delta_0, ..., \delta_3$ show the 'military rank premium' relative to the 'base' civilian case.

4 Program Evaluation Approach

In this section, we describe our matching approach. Our outcome of interest is post-military civilian wage. In the data we use wage in 1980 and 1981, and our parameter of interest is the effect of treatment on the treated (TT). We have four treatments (private, corporal, sergeant, and officer) relative to no treatment (no military service).

The matching approach is based on a simple idea. For each subject in the four treated pools, we find a set of controls (untreated) closest in terms of the two indices that are the two linear functions in the first stage MLE. The matched controls are used as the 'comparison group' to construct the counter-factual no-treatment response of the treated. In the following, for the sake of expostion, we will use a binary variable $d = 0, 1$ to denote treatment and no treatment. Since we will be comparing each treatment (there are four) to no treatment, this simplification does no harm.

Let y_1 and y_0 be two potential responses, treated and untreated, respectively. The TT effect is

$$E\left(y_1 - y_0 | d = 1\right) = E\left(y_1 | d = 1\right) - E\left(y_0 | d = 1\right)$$

where the second term on the r.h.s. is a counter-factual to be constructed. If we have some covariates x and if $E\left(y_0 | d, x\right) = E\left(y_0 | x\right)$, then the following construction of the counterfactual is valid:

$$E\left(y_0 | x, d = 0\right) = E\left(y_0 | x, d = 1\right).$$

In other words, for a treated subject with x_i, we find a group of untreated controls with $x = x_i$. Since their covariates are similar, then presumably so is y_0. It thus follows that

$$E\left(y_1 - y_0 | d = 1\right) = \int \left\{E\left(y | x, d = 1\right) - E\left(y | x, d = 0\right)\right\} f\left(x | d = 1\right) dx.$$

When d takes values $0, 1, 2, 3, 4$, we can think of $E(y_2 - y_0 | d = 2)$ and so on.

An important issue is how to choose the control groups. If the dimension of x is large, it is advantageous to replace x by its propensity score $\Pr(d = 1|x)$ as proposed by Rosenbaum and Rubin (1983). Since $P(d = 1|x)$ in our data depends on two events of joining the military and to attain a particular rank, there are two linear indices determining $P(d = 1|x)$. We thus use the two indices obtained in section 3, instead of four propensity scores; the issue of how to do matching in multiple treatment cases is discussed in some detail in Lee (2005). For a given treated case i, we use the following metric, known as the 'Mahalanobis distance' to measure similarity of a control to case i :

$$D_{i,m} = (\pi_i - \pi_m)' C_N^{-1} (\pi_i - \pi_m) \tag{13}$$

where $\pi_i = (w_i' a_1, w_i' a_2)'$, m indexes the subject in the control group, C_N is the sample covariance matrix of π. With $D_{i,m}$ we implement several different matching criteria, namely pairwise, M-nearest neighbor (M-NN), and fixed caliper matching. Let N_T be the number of treated cases, N_i be the number of successful matches to treated subject i, C_i be the set of controls for treated subject i, and y_{mi} be the wage of the m-th control that is matched to treated subject i. An estimator for the effect of treatment on the treated is

$$TT_N = \frac{1}{N_T} \sum_{i \in T} \left[y_i - \frac{1}{N_i} \sum_{m \in C_i} y_{mi} \right] \tag{14}$$

where '$i \in T$' means that i is in the treatment group, and '$m \in C_i$' means that m is in the matched controls for the treated subject i. In the case of pairwise matching, this is

$$TT_N = \frac{1}{N_T} \sum_{i \in T} (y_i - y_{mi}). \tag{15}$$

In the case of M-NN matching this is

$$TT_N = \frac{1}{N_T} \sum_{i \in T} \left[y_i - \frac{1}{M} \sum_{m \in C_i} y_{mi} \right]. \tag{16}$$

In the case of fixed caliper matching, C_i in equation (14) is the set $\{m \in \text{control group} : D_{i,m} \leq K\}$ where K is the caliper. In caliper matching there will be some unmatched treated units if K is small.

To evaluate matching success, we check how "balanced" π is across the two groups at the aggregate level, by calculating the following statistic for each linear index $j = 0, 1$:

$$M_{(j)} = \frac{\sum_{i \in T} \left(\pi_{ij} - N_i^{-1} \sum_{m \in C_i} \pi_{mj}\right) \cdot 1\left[C_i \neq \phi\right]}{\sum_{i \in T} 1\left[C_i \neq \phi\right]}. \tag{17}$$

For each index j we calculate the average distance between the treated subjects and their respective matched controls. This number should be close to zero if the matching is successful.

5 Empirical Results

5.1 Data

The NLSYM began in 1966 and ended in 1981. The first five waves and last two waves were collected annually, and remaining waves were collected biannually. This data consists of 5225 young men aged 14 to 24 in 1966 and it is unique for its coverage of the very cohorts that faced the draft during the Vietnam years - the young men known as the "Vietnam generation". Compared to other data sets, the NLSYM contains a substantial amount of information regarding military service and experience with the draft system. Questions pertaining to military and draft experience were asked in the 1966, 1969, 1971, 1976, and 1981 waves. They included questions on veteran status, the branch of service, the rank he held, whether he had enlisted or been drafted, the duration of active duty etc. Those who did not serve were asked questions pertaining to their draft eligibility. This data forms a rich body of wartime information available in a longitudinal survey. Unfortunately, one drawback is its high attrition rates. By the end of the survey in 1981, only 65 percent of the 1966 respondents were contacted. Most of the remainder were refusals or dropped after two consecutive non-interviews; 139, or 2.7 percent of the sample, were deceased - a high mortality rate to be sure, but not significant enough relative to attrition.

We start by presenting the summary statistics of the variables in table 1.

table 1 *about here*

About 32% of the sample are veterans. An asterisk indicates that the difference in means between veterans and nonveterans is significant at the 5% level. Veterans are significantly more likely to have higher socio-economic status, live in the south, have lower draft numbers, have a male presence in youth, live in a metropolitan area, and earn higher wages. Turning to the rank attained. In the US Military there are nine grades of enlisted ranks (E1 to E9), five grades of warrant officers (W1-W5) and eleven grades of officer ranks (O1 to O11). However the overwhelming majority of the servicemen in this sample, due to a relatively short average spell of service, would only have achieved the lower grades: enlisteds would rarely achieve a rank much higher than sergeant (E5), and officers would rarely be promoted beyond captain (O3). There are virtually no warrant officers. As a result, it is convenient to partition the rank distribution into four categories: private, corporal, sergeant, and officer. "Private" refers to servicemen who attained the ranks of private, private first class, or lance corporal (corresponding to enlisted ranks E1 to E3). "Corporal" refers only to servicemen who attained that rank (E4). "Sergeant" refers to servicement attaining all ranks of sergeant up to sergeant major or in the case of the Navy, petty officer (corresponding to enlisted ranks E5 to E9).[1] "Officer" refers to all warrant officers and commissioned officers. In actuality, officers largely consist of lieutenants and captains. To summarize, the data was recoded as follows:

Recoded Rank	Actual Rank
Private	private, pvt 1st class, lance corporal (E1-E3)
Corporal	Corporal (E4)
Sergeant	Sergeant to Sergeant Major (E5-E9)
Officer	Warrant Officer, Lieutenant to Major (W1-W5,O1-O5)

From table 1 we see that the among the veterans, 32% attained the rank of private, 38% corporal, 26% sergeant, and 5% officer. Observe also that wages increase steadily with rank. The increase is

[1]For a comparison of the ranks across the services army, navy, air force and marines, refer to http://usmilitary.about.com/library/milinfo/rankchart/blenlistedrank.htm

slight from non-veterans to privates, and then much more steeply as rank progresses.

5.2 Estimation Results

5.2.1 First Stage

We begin by estimating the first stage bivariate probit-ordered probit pair of equations in (1) by MLE. We selected a variety of covariates that reflected "initial conditions", household and individual characteristics, and geographical characteristics. For initial conditions we use variables at the beginning of the survey in 1966. These include geographical information such as measures of local level unemployment, labor market condition, and labor force, metropolitan status, information on region. These variables are meant to capture whether labor market conditions may affect individuals' willingness to enlist. We also include family background characteristics such as whether the parents were alive in the initial year, whether there was a paternal presence during adolescence, and measures of family socio-economic status in the initial survey.[2] For individual characteristics, we focus on anthropometric measures such as height. We avoid using variables such as schooling to minimize issues of endogeneity; the usage of parental socio-economic status proxies fairly well for schooling. Cohort effects were also included via the use of an orthogonal polynomial of birth year (up to the fifth degree). We experimented also with the use of cohort dummies but there were little differences in the results. Following Angrist (1990) we use the draft number of cohorts facing the draft lottery. In view of this, we consider two samples. The first sample consists only of respondents born between 1944 and 1952 — the cohorts that faced the draft lottery. The second, larger sample in addition uses cohorts born before 1944. After eliminating missing observations we are left with 2856 observations in the draft lottery sample, and 3554 observations in the full sample.

[2] Socio Economic Index is a composite measure of parents' schooling, parents' occupation, education of eldest sibling and indicator of reading materials at home (eg, newspapers).

table 2 about here

In table 2 we present the results of estimating the first stage equation that jointly determines military entry and subsequent rank. The first three columns refer to the military entry equation, while the last three columns refer to the rank equation. From the estimates in the first three columns, we see that race, socio-economic status, the draft lottery number, and family background characteristics are significant predictors of military entry. Blacks are less likely to enter the military. The presence of the father at age 14 has a positive effect of military entry, but if the parents are alive in 1966, the individual is much less likely to serve in the military. However in the rank equation, only height and socio-economic status of parents are significant predictors of rank. Both are positive. The finding that height is significant is in line with other research that finds a wage premium to height (e.g. Sargent and Blanchflower (1994) and Persico et al. 2004). Here, height does not affect military entry, but affects rank, suggesting that tall individuals acquire more leadership skills only after entry into the military. There is concavity in the effect of socioeconomic status of the family, which suggests that those with high and low levels of socioeconomic status are less likely to serve. This effect is reversed when we look at the rank equation, where rank increases with family background. As is often the case with estimating bivariate probit models, we are unable to estimate the correlation parameter ρ_{12} with much precision. Though not reported here, we conducted a simulation study which confirmed that ρ_{12} is not well estimated whereas α_1 and α_2 are. Performing the same estimation on the larger full sample yields very similar results as shown in table A1.

Our main interest in the first stage is to obtain the two linear indices $w'\alpha_1$ and $w'\alpha_2$. We calculate the parametric selection correction terms according to equations (3) to (11) . The first equation is the selection term corresponding to those who did not enter the military, and the latter eight equations are parametric selection terms corresponding to the terms $\lambda_{01}, \lambda_{02}, \lambda_{11}, \lambda_{12}, \lambda_{21}, \lambda_{22}, \lambda_{31}$, and λ_{32}. There are two expressions per rank, which correspond to the first correction (military entry) and second correction (rank). For instance, $(\lambda_{01}, \lambda_{02})$ correspond to the entry and rank correction terms of privates, $(\lambda_{11}, \lambda_{12})$ for corporals and so on. We calculate these expressions using numerical integration routines such as those found in Press et al. (1992). To have an idea of what the indices look like, refer to figure 1 which presents kernel density estimates of the two indices for both the lottery and the full samples.

figure 1 *here*

5.2.2 Parametric Selection Wage Regressions

Table 3 (lottery sample) presents the second stage wage regressions, equation (12). We compare several specifications of the wage regression including and not including the selection correction terms. We try these regressions using wages in two years. The sample consists of only those individuals who reported positive earnings. The sample size is around 1900 to 2000 observations in the draft lottery sample. We also consider the full sample which consists of around 2350 to 2450 observations in table A2.

table 3 *about here*

The first column presents OLS regression of ln(wage in 1981) on a unity, polynomial in cohort, a dummy for metropolitan standard area (MSA), years of schooling, dummies for black and rank (the omitted category being non-veterans). As can be seen, the rank coefficients for corporals and sergeants are significant. Column 2 presents the same specification including the parametric selection correction terms λ_c, λ_1 and λ_2. Immediately we see that the rank premium on corporals and sergeants is no longer significant. At the same time, none of the correction terms are significant as well. As years of schooling are likely to be endogenous, we also present instrumental variable estimates of the same equation in column 3.[3] Likewise, neither the rank terms nor the selection

[3]The instruments used are : central MSA, noncentral MSA, indicators of local unemployment rate and labor force

correction terms are significant. Columns 4 to 6 in **table 3** present the same results with 1980 wages. We see that the results are similar, in that the coefficients on the ranks of corporal and sergeant are significant under OLS, but when we add the selection correction terms and instrument for years of schooling, we find that neither the rank dummies nor the selection correction terms are significant. We also perform F-tests to check for joint significance of the selection correction terms and for the rank coefficients. The tests fail to reject the null hypotheses in all but one case. The results of the full sample turn out to be similar (see appendix). Although far from definitive, we take this as supporting evidence that our parametric selection correction procedure does not invalidate our use of OLS estimates of the rank premium.

To delve deeper into the lack of significance when the selection correction terms are included, we check for correlations among the regressors. We found that correlations between several of the rank dummies and the selection correction terms were quite high: $cor\left(\delta_c\lambda_c, \sum_j \delta_j\lambda_{j1}\right) = -0.77$, $cor\left(\delta_0, \sum_j \delta_j\lambda_{j2}\right) = -0.75$, and $cor\left(\delta_1, \sum_j \delta_j\lambda_{j1}\right) = 0.54$. This suggests that near-multicollinearity affecting some key variables could be the main reason why neither the selection correction terms nor the rank coefficients are significant. It also appears that the high correlations are more severe around the lower levels of military participation (i.e. non-veterans, and privates). Unfortunately there does not appear to be much we can do about it.

5.2.3 Matching Estimates

The preceding results indicate that the selection problem would not be severe and the OLS would be valid. But OLS assumes the linear functional form and the OLS results may be biased if the functional form is misspecified. To avoid this problem, we now try matching estimates of the rank wage premium. Considering each rank separately, we have four treatments vs no treatment (non-veteran). For each treated subject we find a group of controls closest in terms of two indices that are

size in 1960, dummies for whether father and mother and respondent were born in the US, dummies for whether parents were alive in 1966, and dummy for whether father was present in the household when the respondent was 14.

the two linear indices in the first stage MLE. The metric we use is the Malahanobis distance as per equation (13). Table 4 presents the matching estimates of the TT effect using M nearest-neighbor estimation with $M = 1, 5, 10$. The TT effect corresponds to equation (15) in the case of pairwise matching, and equation (16) in the case of general M. Table 5 presents the estimates using caliper matching with $K = 0.007$, 0.01, and 0.03. In each case bootstrap standard errors were calculated using 2000 repetitions, except the case of $M = 1$ where the formula for asymptotic variance is straightforward treating each pair as one unit. To assess the overall match, we also calculate the average measure of covariate balance as per equation (17). As before, our calculations use 1980 wages and 1981 wages, and we also perform the analysis using the lottery sample and the larger full sample.

We first focus on table 4, nearest neighbor matching. Observe that as the number of neighbors increase, across all ranks, the standard error of the estimated TT decreases. There is no discernible pattern for the mean difference. The net result is some gain in precision of the estimated rank premium. It ranges between 8% and 14% for corporals, 9% and 12% for sergeants, and 15% and 34% for officers. The most consistent results in terms of significance levels are for corporals, followed by officers (for the latter case we should note the smaller sample size). Indeed the standard errors for corporals is the lowest across the four ranks. Turning next to average imbalance of the indices, as expected, average imbalance increases as we increase the number of neighbors; but also as would be expected, with larger numbers of neighbors in the control group, the rank premium emerges as significant. Fortunately this is achieved without significant drop in covariate balance. From this table, it begins to emerge that veteran status is not enough to command a wage premium (observing the wage premium for privates would confirm this) but it is rank that carries with it a premium.

table 4, 5, 6, 7 *about here*

Nearest neighbor matching may have the disadvantage that some treated subjects could be poorly matched due to lack of close enough controls. Indeed this might be true if covariate balance were at worrying levels. To examine whether this might be a factor, we turn to consider caliper matching, using a variety of different-sized calipers. Refer to table 5. As before, observe that with an increase in the caliper size, more controls are included in the control group, and thus we observe a fall in the standard errors of the estimated TT. There is no discernible pattern in the mean difference, and covariate balance levels are maintained at fairly low levels compared with the M-NN. With a caliper size of 0.007, the percentage of unmatched treated subjects ranges from 11% to 29%. This decreases to 2.7% to 7.1% when caliper size is 0.03. Despite the smaller sample size of officers, covariate balance does not change dramatically even at larger caliper sizes. The levels of significance are somewhat less than nearest neighbor estimates. However if one is willing to admit significance levels at the 10% level, then we still observe significant wage premiums to corporals, sergeants and officers at larger caliper sizes. As with the nearest neighbor estimates, corporals appear to have the most consistent wage premiums.

We also perform the same analysis with a larger sample that includes the earlier cohorts (that is, before the draft lottery was implemented). The first stage results are found in appendix in table A1. However we present the propensity score matching results here in tables 6 and 7. The addition of the earlier cohorts increases the sample size by around 20%. The increased sample sizes result in lower standard errors all round, and hence higher levels of significance. As before, there is no premium to being a private, but in this larger sample, we find that the rank premium increases with rank. For instance, looking at table 6, the 1981 wages of corporals were 10% to 13% higher than comparable non-servicemen, but 14% to 17% higher for sergeants. In 1980 the wages were 8% to 10% for corporals, 13% to 17% for sergeants. We find the same pattern holding when we perform caliper matching in **table 7**. There is a premium to corporals of 9.5 to 14%, a premium to sergeants of 16 to 18%, and a premium to officers of 18 to 33%.

For comparison, the bottom rows of tables 4 to 7 present the average of the index 1 and index 2 (before matching) in each group. However these rows also reveal some interesting findings in themselves. Notice from table 4 that the average of index 2 for non-veterans is 0.381, which is in between that of sergeants (0.289) and officers (0.672). We find the same pattern in the other tables as well. This implies that among the Vietnam era youth, there were many eligible among those who *did not* serve in the military. In fact, they *could* have become sergeants or officers had they done so!

The overall finding, therefore, is that among the Vietnam era youth cohorts, men born between 1940 and 1952 (full sample), there is a positive and significant wage premium among veterans above private, which increases monotonically in rank. When we restrict to the draft lottery cohorts (those born 1944 and after), we find slightly weaker results, but in those cases the rank premium is most consistent among corporals, followed by officers. It is not significant among sergeants.

Why Corporals? It seems surprising that in the lottery sample, there is stronger evidence of a corporal's wage premium over a sergeant's wage premium. An understanding of NCO training may provide a clue. Typically, NCOs are promoted from the lower ranks. The rank of Sergeant could only be attained after 4 to 6 years of service. High casualty rates in the early years of the Vietnam War led to attrition of the NCO ranks. In later years of the conflict, the military sought to replenish the depleted NCO ranks by establishing NCO training schools. Promising candidates would be selected from among enlistees or draftees at an early stage (usually after basic training), be trained for about 12 weeks, and instantly promoted to sergeants, bypassing the corporal rank. The program was controversial, was widely thought to have produced "shake and bake" sergeants or "instant NCOs" who were untested and lacked the experience and skill needed to lead men who were their peers. By contrast, promotions to corporal could have come more slowly than promotions to sergeant, and consequently may have been a clearer indicator of skill, leadership or ability.

Early Cohorts and Later Cohorts. We also observe that inclusion of the early cohorts born before 1944 significantly improved the matching results. On the one hand it could be due to sample size. On the other hand, it is well documented that public opinion turned against the Vietnam war in the later years (1968 onwards). The later cohorts may have faced more of the negative repercussions associated with an taking part in an increasingly unpopular war (for instance, greater incidence of post-traumatic stress disorder and discrimination). The stronger anti-war sentiment faced by the lottery cohorts (draftees in 1969 to 1972) may also explain why the rank premium is weaker among the younger cohorts.

6 Conclusion

This paper set out to establish whether there is indeed a premium to leadership experience in the military. We measure this premium using the highest attained rank in the military. To our knowledge, our attempt at quantifying the impact of military leadership is a first in the literature. Our paper applies a parametric method to account for the possibility of non-random selection into the military and selection into the final rank attained. We then applied a nonparametric matching method to avoid parametric "regression function" misspecification.

Our results from the parametric selection correction procedure indicate that after applying the correction terms, the coefficients on the rank dummy variables were no longer significant. But then, we failed to reject the null hypothesis that all correction terms are irrelevant. When we turn to the nonparametric matching estimates, we found more convincing results. Limiting ourselves only to lottery-era cohorts, we found the rank premium strongest among veterans who were corporals. The rank premium for sergeant was less precisely estimated, and the premium of officers was significant and of large magnitude, although the sample size was small. We also found that extending our sample to include pre-lottery cohorts improved the matching estimates significantly. While it could be due to sample size, it could also be that the results reflect the negative repercussions associated stronger anti-war sentiment faced by the later lottery cohorts of 1969 to 1972.

From our findings, we argue that it is rank, not veteran status, that commands a wage premium in the labor market. This is seen from the fact that the wage premium for privates is consistently small and insignificant, but for corporals or above it is positive and significant. We see two implications of this. Firstly, we interpret this as a return to leadership skills, and thus regard this as contributing evidence to there being returns to non-cognitive skills. Secondly, it implies that the veterans' premium (if there is one) is primarily one rewarding leadership, rather than participation in the military. If one joins the military expecting a future wage premium but fails to get promoted in rank, he would be disappointed.

This paper is related to a growing body of work focusing on non-schooling characteristics which have an effect on labor market success. This is a contribution to the existing literature documenting the effects of non-schooling, non-cognitive characteristics on labor market success. We stress in particular the role played by leadership in the military, and not just veteran status. This emerging body of evidence on the role of non-cognitive skills suggests there is much to uncover in the area of pre-market human capital effects.

7 Appendix

7.1 Derivation of Likelihood

Observe that the second through the last terms of the log-likelihood function in equation (2) are the likelihood contributions corresponding to being a private, corporal, sergeant and officer respectively:

$$\begin{aligned} P\left(-w'\alpha_1 < \varepsilon_1, \varepsilon_2 < -w'\alpha_2\right) &= P\left(\varepsilon_2 < -w'\alpha_2\right) - P\left(\varepsilon_1 < -w'\alpha_1, \varepsilon_2 < -w'\alpha_2\right) \\ &= \Phi\left(-w'\alpha_2\right) - \Psi\left(-w'\alpha_1, -w'\alpha_2, \rho_{12}\right) \equiv p_0; \end{aligned}$$

$$\begin{aligned} & P\left(-w'\alpha_1 < \varepsilon_1, -w'\alpha_2 < \varepsilon_2 < -w'\alpha_2 + \gamma_2\right) \\ = \;& P\left(\varepsilon_2 < -w'\alpha_2 + \gamma_2\right) - P\left(\varepsilon_1 < -w'\alpha_1, \varepsilon_2 < -w'\alpha_2 + \gamma_2\right) - P\left(-w'\alpha_1 < \varepsilon_1, \varepsilon_2 < -w'\alpha_2\right) \\ = \;& \Phi\left(-w'\alpha + \gamma_2\right) - \Psi\left(-w'\alpha_1, -w'\alpha_2 + \gamma_2, \rho_{12}\right) - p_0 \equiv p_1; \end{aligned}$$

$$
\begin{aligned}
& P\left(-w'\alpha_1 < \varepsilon_1, -w'\alpha_2 + \gamma_2 < \varepsilon_2 < -w'\alpha_2 + \gamma_3\right) \\
= \; & P\left(\varepsilon_2 < -w'\alpha_2 + \gamma_3\right) - P\left(\varepsilon_1 < -w'\alpha_1, \varepsilon_2 < -w'\alpha_2 + \gamma_3\right) - P\left(-w'\alpha_1 < \varepsilon_1, \varepsilon_2 < -w'\alpha_2 + \gamma_2\right) \\
= \; & \Phi\left(-w'\alpha_2 + \gamma_3\right) - \Psi\left(-w'\alpha_1, -w'\alpha_2 + \gamma_3, \rho_{12}\right) - \left(p_0 + p_1\right);
\end{aligned}
$$

$$
\begin{aligned}
& P\left(-w'\alpha_1 < \varepsilon_1, -w'\alpha_2 + \gamma_3 < \varepsilon_2\right) \\
= \; & P\left(-w'\alpha_1 < \varepsilon_1\right) - P\left(\varepsilon_2 < -w'\alpha_2 + \gamma_3\right) + P\left(\varepsilon_1 < -w'\alpha_1, \varepsilon_2 < -w'\alpha_2 + \gamma_3\right) \\
= \; & \Phi\left(w'\alpha_1\right) - \Phi\left(-w'\alpha_2 + \gamma_3\right) + \Psi\left(-w'\alpha_1, -w'\alpha_2 + \gamma_3, \rho_{12}\right).
\end{aligned}
$$

The above is made clear by referring to the following diagram, which pictures, for an individual with characteristics w, the thresholds of ε_1 and ε_2 that determine his veteran status and rank.

figure A1 about here

The shaded area below the horizontal line $-w\alpha_1$ consists of non-veterans. Entry requires that $\varepsilon_1 \geq -w\alpha_1$. Among entrants into the military, the thresholds, $-w\alpha_2$, $-w\alpha_2 + \gamma_2$, and $-w\alpha_2 + \gamma_3$ denote the ε_2 thresholds separating private from corporal, corporal from sergeant, and sergeant from officer respectively. For instance an individual i would be a private if $-w_i\alpha_1 + \varepsilon_{1i} \geq 0$ and $-w_i\alpha_2 + \varepsilon_{2i} < 0$, and so on.

7.2 Second Stage Correction Term

The selection correction terms are derived using the linear projection assumption of $u|\varepsilon$. To ease exposition, we will deal with the fourth and first terms, for which closed forms are obtainable. The middle terms will then be examined, for which the correction terms are computed numerically.

Observe, for the fourth term.

$$
\begin{aligned}
E\left(u|w, d_1 = 1, d_2 = 3\right) \; = \; & \theta_1 E\left(\varepsilon_1 | w, \; -w'\alpha_1 < \varepsilon_1, \; -w'\alpha_2 + \gamma_3 < \varepsilon_2\right) \\
& + \theta_2 E\left(\varepsilon_2 | w, \; -w'\alpha_1 < \varepsilon_1, \; -w'\alpha_2 + \gamma_3 < \varepsilon_2\right).
\end{aligned}
$$

Using the moment formulas for truncated bivariate normal distributions (e.g., Maddala (1983, p.368)), the selection correction term is

$$\theta_1 \Psi_{12}^{-1} \left\{ \phi(-w_1'\alpha_1)\Phi\left(t_1\right) + \rho_{12}\phi\left(-w'\alpha_2+\gamma_3\right)\Phi\left(t_2\right) \right\}$$
$$+\theta_2 \Psi_{12}^{-1} \left\{ \phi\left(-w'\alpha_2+\gamma_3\right)\Phi\left(t_2\right) + \rho_{12}\phi\left(-w'\alpha_1\right)\Phi\left(t_1\right) \right\},$$

where, because the distribution of $(\varepsilon_1, -\varepsilon_2)$ is the same as that of $(\varepsilon_1, \varepsilon_2)$ except that ρ_{12} is replaced by $-\rho_{12}$,

$$\begin{aligned}
\Psi_{12} &\equiv P(-w'\alpha_1 < \varepsilon_1,\ -w'\alpha_2+\gamma_3 < \varepsilon_2) = P(-\varepsilon_1 < w'\alpha_1,\ -\varepsilon_2 < w'\alpha_2 - \gamma_3) \\
&= \Psi(w'\alpha_1, w'\alpha_2-\gamma_3, -\rho_{12}), \\
t_1 &\equiv \frac{-w'\alpha_2+\gamma_3+\rho_{12}w'\alpha_1}{\sqrt{1-\rho_{12}^2}}, \quad t_2 \equiv \frac{-w'\alpha_1+\rho_{12}(w'\alpha_2-\gamma_3)}{\sqrt{1-\rho_{12}^2}}.
\end{aligned}$$

Analogously, for the first term, we get

$$\begin{aligned}
& E\left(u|w, d_1=1, d_2=0\right) \\
= \;& \theta_1 E\left(\varepsilon_1 | w,\ -w'\alpha_1 < \varepsilon_1,\ \varepsilon_2 < -w'\alpha_2\right) + \theta_2 E\left(\varepsilon_2 | w,\ -w'\alpha_1 < \varepsilon_1,\ \varepsilon_2 < -w'\alpha_2\right) \\
= \;& \theta_1 E\left(\varepsilon_1 | w,\ -w'\alpha_1 < \varepsilon_1,\ w'\alpha_2 < -\varepsilon_2\right) + \theta_2 E\left(\varepsilon_2 | w,\ -w'\alpha_1 < \varepsilon_1,\ w'\alpha_2 < -\varepsilon_2\right).
\end{aligned}$$

The selection correction term is

$$\theta_1 \Psi_{12}^{-1} \left\{ \phi(-w_1'\alpha_1)\Phi(t_1) - \rho_{12}\phi(w'\alpha_2)\Phi(t_2) \right\} + \theta_2 \Psi_{12}^{-1} \left\{ \phi(w'\alpha_2)\Phi(t_2) - \rho_{12}\phi(-w'\alpha_1)\Phi(t_1) \right\}$$

where

$$\begin{aligned}
\Psi_{12} &\equiv P\left(-w'\alpha_1 < \varepsilon_1,\ w'\alpha_2 < -\varepsilon_2\right) = P\left(-\varepsilon_1 < w'\alpha_1,\ \varepsilon_2 < -w'\alpha_2\right) \\
&= \Psi\left(w'\alpha_1, -w'\alpha_2, -\rho_{12}\right), \\
t_1 &\equiv \frac{w'\alpha_2-\rho_{12}w'\alpha_1}{\sqrt{1-\rho_{12}^2}}, \quad t_2 \equiv \frac{-w'\alpha_1+\rho_{12}w'\alpha_2}{\sqrt{1-\rho_{12}^2}}.
\end{aligned}$$

Turning to the second term, observe

$$E(u|w, d_1 = 1, d_2 = 1)$$
$$= \theta_1 E(\varepsilon_1|w,\ -w'\alpha_1 < \varepsilon_1,\ -w'\alpha_2 < \varepsilon_2 < -w'\alpha_2 + \gamma_2)$$
$$+ \theta_2 E(\varepsilon_2|w,\ -w'\alpha_1 < \varepsilon_1,\ -w'\alpha_2 < \varepsilon_2 < -w'\alpha_2 + \gamma_2).$$

Differently from the previous two terms, the closed form for this seems difficult to get. Instead, we get the selection correction terms numerically. Observe

$$E(\varepsilon_1|w, -w'\alpha_1 < \varepsilon_1, -w'\alpha_2 < \varepsilon_2 < -w'\alpha_2 + \gamma_2) = \frac{\int_{-w'\alpha_2}^{-w'\alpha_2+\gamma_2} \int_{-w'\alpha_1}^{\infty} \varepsilon_1 \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2}{\int_{-w'\alpha_2}^{-w'\alpha_2+\gamma_2} \int_{-w'\alpha_1}^{\infty} \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2},$$
$$E(\varepsilon_2|w, -w'\alpha_1 < \varepsilon_1, -w'\alpha_2 < \varepsilon_2 < -w'\alpha_2 + \gamma_2) = \frac{\int_{-w'\alpha_2}^{-w'\alpha_2+\gamma_2} \int_{-w'\alpha_1}^{\infty} \varepsilon_2 \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2}{\int_{-w'\alpha_2}^{-w'\alpha_2+\gamma_2} \int_{-w'\alpha_1}^{\infty} \psi(\varepsilon_1, \varepsilon_2, \rho_{12}) d\varepsilon_1 d\varepsilon_2};$$

which are equations (6) and (7) respectively. For the given estimate for $(\alpha_1', \alpha_2', \rho_{12}, \gamma_2, \gamma_3)$ from the first stage, these integrals can be obtained numerically for each observation in the second stage using a variety of quadrature routines.

As for the third term, observe

$$E(u|w, d_1 = 1, d_2 = 2) = \theta_1 E(\varepsilon_1|w,\ -w'\alpha_1 < \varepsilon_1,\ -w'\alpha_2 + \gamma_2 < \varepsilon_2 < -w'\alpha_2 + \gamma_3)$$
$$+ \theta_2 E(\varepsilon_2|w,\ -w'\alpha_1 < \varepsilon_1,\ -w'\alpha_2 + \gamma_2 < \varepsilon_2 < -w'\alpha_2 + \gamma_3).$$

The integrals (8) and (9) then follow. For consistency we use numerical evaluations for all the integrals. Thus we derive the numerical integrals for the first and fourth terms in equations (4), (5), (10) and (11) in the same way.

References

Abowd, J. M., F. Kramarz, and D. N. Margolis (1999): "High Wage Workers and High Wage Firms," *Econometrica*, 67(2), 251–333.

Anderson, D. J. (2000): "If You Let me Play: The Effects of Participation in High School Athletics on Students' Behavior and Economic Success," Mimeo, Cornell University.

Angrist, J. D. (1989): *Using the Draft Lottery to Measure the Effect of Military Service on Civilian Labor Market Outcomes*vol. 10 of *Research in Labor Economics*. JAI Press, Inc., Greenwich.

——— (1990): "Lifetime Earnings and the Vietnam Era Draft Lottery: Evidence from Social Security Administrative Records," *American Economic Review*, 80(3), 313–36.

——— (1998): "Estimating the Labor Market Impact of Voluntary Military Service Using Social Security Data on Military Applicants," *Econometrica*, 66(2), 249–288.

Angrist, J. D., and A. Krueger (1994): "Why Do World War II Veterans Earn More Than Nonveterans?," *Journal of Labor Economics*, 12(1), 74–97.

Arvey, R., M. Rotundo, and W. Johnson (2002): "The Determinants of Leadership: The Role of Genetic, Personality and Cognitive Factors," Mimeo, Carlson School of Management, Minneapolis.

Arvey, R., M. Rotundo, W. Johnson, and M. McGue (2004): "The Determinants of Leadership: The Role of Genetic, Personality and Cognitive Factors," Mimeo, University of Minnesota.

Barron, J. M., B. Ewing, and G. Waddell (2000): "The Effects of High School Athletic Participation on Education and Labor Market Outcomes," *Review of Economics and Statistics*, 82(3), 1–13.

Baskir, L. A., and W. A. Strauss (1978): *Chance and Circumstance: the Draft, the War, and the Vietnam Generation*. Alfred A. Knopf, New York.

Berger, M. C., and B. T. Hirsch (1983): "The Civilian Earnings Experience of Vietnam-Era Veterans," *Journal of Human Resources*, XVIII(4), 455–479.

Bowles, S., H. Gintis, and M. Osborne (2001): "The Determinants of Earnings: Skills, Preferences and Schooling," *Journal of Economic Literature*, 39(4), 1137–1176.

Cawley, J., J. J. Heckman, and E. Vytlacil (2001): "Three Observations on Wages and Measured Cognitive Ability," *Labour Economics*, 8, 419–442.

DeTray, D. (1982): "Veteran Status as a Screening Device," *American Economic Review*, 72, 133–42.

Dubin, J. A., and D. Rivers (1989): "Selection bias in linear regression," *Sociological Methods and Research*, 18, 360–390.

Eccles, J., and B. Barber (1999): "Student Council, Volunteering, Basketball, or Marching Band: What Kind of Extracurricular Involvement Matters?," *Journal of Adolescent Research*, 14(1), 10–43.

Eide, E., and N. Ronan (2000): "Is Participation in High School Athletics an Investment or Consumption Good? Evidence from High School and Beyond," Mimeo, Brigham Young University.

Foley, M. S. (2003): *Confronting the War Machine: Draft Resistance During the Vietnam War*. The University of North Carolina Press, Chapel Hill and London.

Goldberg, M. S., and J. T. Warner (1986): "Military Experience, Civilian Experience, and the Earnings of Veterans," *Journal of Human Resources*, XXII(1), 62–81.

Green, F., S. Machin, and D. Wilkenson (1998): "The Meaning and Determinants of Skills Shortages," *Oxford Bulletin of Economics and Statistics*, 60(2), 165–188.

Hamermesh, D., and J. Biddle (1994): "Beauty and the Labor Market," *American Economic Review*, 84(5), 1174–94.

Heckman, J. J. (2000): "Policies to Foster Human Capital," *Research in Economics*, 54, 3–56.

Heckman, J. J., and P. Carneiro (2004): *Human Capital Policy,*chap. 2, pp. 77–239, Inequality in America: What Role for Human Capital Policies? MIT Press.

Hirsch, B. T., and S. L. Mehay (2003): "Evaluating the Labor Market Performance of Veterans Using a Matched Comparison Group Design," *Journal of Human Resources*, 38(3), 673–700.

Kuhn, P., and C. Weinberger (2005): "Leadership Skills and Wages," Forthcoming, Journal of Labor Economics.

Lee, M.-J. (2005): *Micro-econometrics for Policy, Program and Treatment Effects*. Oxford University Press.

Maddala, G. S. (1983): *Limited-dependent and qualitative variables in econometrics*. Cambridge University Press.

Mueller, U., and A. Mazur (1996): “Facial Dominance of West Point Cadets as a Predictor of Later Military Rank,” *Social Forces*, 74(3), 823–850.

Persico, N., A. Postlewaite, and D. Silverman (2004): “The Effect of Adolescent Experience on Labor Market Outcomes: The Case of Height,” *Journal of Political Economy*, 112(5), 1019–53.

Postlewaite, A., and D. Silverman (2005): “Social Isolation and Inequality,” Mimeo, University of Pennsylvania.

Press, W., S. A. Teukolsky, W. T. Vetterling, and B. P. Flannery (1992): *Numerical Recipes in Fortran*, vol. 1. Cambridge University Press, 2 edn.

Rosenbaum, P., and D. Rubin (1983): “The central role of the propensity score in observational studies for causal effects,” *Biometrika*, 70, 41–55.

Sargent, J. D., and D. G. Blanchflower (1994): “Obesity and Stature in Adolescence and Earnings in Young Adulthood: Analysis of a British Birth Cohort,” *Archives Pediatric and Adolescent Medicine*, 148, 681–87.

Tarr, C. W. (1981): *By the Numbers: the Reform of the Selective Service System 1970-1972*. National Defense University Press, Fort Lesley J. McNair Washington, DC.

Van de Ven, W., and B. V. Praag (1981): “The demand for deductibles in private health insurance,” *Journal of Econometrics*, 17, 229–252.

Wong, L., P. Bliese, and D. McGurk (2003): “Military Leadership: A Context Specific Review,” *The Leadership Quarterly*, 14, 657–692.

Table 1: Summary Statistics

	Non-Veterans		Veterans		
Variable	**Mean**	**SD**	**Mean**	**SD**	
Panel A: Summary Statistics					
Central	0.32	0.47	0.34	0.47	
Noncentral	0.32	0.47	0.35	0.48	
Black	0.24	0.43	0.16	0.37	*
Height	70.62	3.34	70.67	3.31	
Socio Economic Index 1966	10.03	2.39	10.29	1.96	*
South Residence 1966	0.41	0.49	0.36	0.48	*
Male Labor Market Index 1966	0.52	0.50	0.54	0.50	
Draft Number *100	1.55	1.19	1.37	1.16	*
Unemployment Rate 1960	5.23	1.68	5.17	1.71	
Labor Force Size 1960	0.57	1.06	0.55	1.02	
Father Present at Age 14	0.63	0.48	0.67	0.47	*
Respondent Born US	0.73	0.44	0.75	0.43	
Father Alive 1966	0.92	0.28	0.91	0.28	
Mother Alive 1966	0.97	0.16	0.97	0.18	
Log wage 1980	9.74	0.75	9.88	0.64	*
Log Wage 1981	9.70	0.84	9.88	0.69	*
Years of Schooling	13.17	3.00	13.32	2.09	
SMSA Residence in 1981	0.67	0.47	0.73	0.44	*
Private			0.32	0.46	
Corporal			0.38	0.49	
Sergeant			0.26	0.44	
Officer			0.05	0.22	
Number of obs	2605		1171		

	1980		1981	
Panel B: Wages by Rank	**Mean**	**SD**	**Mean**	**SD**
Nonveteran	9.74	0.75	9.70	0.84
Private	9.77	0.80	9.74	0.80
Corporal	9.86	0.50	9.88	0.59
Sergeant	9.97	0.59	9.97	0.67
Officer	10.25	0.59	10.26	0.62

* Significantly different at 5% level.

Table 2: First Stage Selection Model

	Military Entry Equation			Rank Equation		
	Coeff	**Std err**	**T-ratio**	**Coeff**	**Std err**	**T-ratio**
Constant	1.231	0.755	1.63	-1.166	1.229	-0.95
Central MSA	0.107	0.073	1.47	0.090	0.273	0.33
Noncentral MSA	0.084	0.071	1.18	-0.039	0.225	-0.17
Black	-0.205	0.074	-2.75	-0.021	0.508	-0.04
Height	0.001	0.007	0.08	0.024	0.013	1.84
South Resident in 1966	-0.044	0.062	-0.72	0.197	0.132	1.50
Male Labor Market index in 1966	-0.048	0.064	-0.76	-0.125	0.145	-0.86
Draft Sequence Number * 100	-0.086	0.024	-3.58	-0.025	0.207	-0.12
Unemployment Rate 1960	-0.025	0.015	-1.67	-0.023	0.062	-0.38
Size Of Labor Force 1960	-0.035	0.029	-1.21	-0.037	0.093	-0.40
Father In The Household At Age 14	0.333	0.068	4.88	0.251	0.794	0.32
Respondent Born In The Us	0.196	0.064	3.06	-0.080	0.467	-0.17
Father Alive In 1966	-0.364	0.104	-3.50	-0.134	0.862	-0.16
Mother Alive In 1966	-0.399	0.162	-2.47	0.018	0.932	0.02
Socio-Economic Index	0.024	0.032	0.75	0.193	0.084	2.29
Socio-Economic Index Squared	-0.176	0.030	-5.85	0.063	0.425	0.15
gamma2				1.074	0.052	20.53
gamma3				2.340	0.104	22.44
rho12				-0.006	3.352	0.00
Cohort Polynomial Included?						yes
Log Likelihood						-2801.1
Number of Observations						2867

Table 3: Second Stage of Parametric Selection Model

	OLS		IV	OLS		IV
	(1)	(2)	(3)	(4)	(5)	(6)
	lw81	lw81	lw81	lw80	lw80	lw80
years of schooling	0.065***	0.063***	0.084**	0.056***	0.054***	0.082**
	(0.007)	(0.007)	(0.036)	(0.006)	(0.006)	(0.032)
In MSA 1981	0.247***	0.250***	0.236***	0.245***	0.247***	0.226***
	(0.038)	(0.038)	(0.049)	(0.034)	(0.034)	(0.043)
Black	-0.302***	-0.306***	-0.276***	-0.293***	-0.287***	-0.243***
	(0.043)	(0.046)	(0.072)	(0.040)	(0.043)	(0.066)
Private	0.008	-0.309	-0.257	-0.014	-0.252	-0.116
	(0.061)	(0.284)	(0.327)	(0.055)	(0.252)	(0.300)
Corporal	0.179***	-0.029	-0.016	0.143***	0.031	0.114
	(0.053)	(0.228)	(0.246)	(0.048)	(0.203)	(0.227)
Sergeant	0.123**	-0.011	-0.037	0.109**	0.082	0.108
	(0.061)	(0.204)	(0.207)	(0.055)	(0.182)	(0.188)
Officer	0.136	0.151	0.031	0.139	0.292	0.185
	(0.148)	(0.247)	(0.290)	(0.129)	(0.220)	(0.249)
lambdac		-0.019	-0.069		0.058	0.046
		(0.156)	(0.163)		(0.143)	(0.145)
lambda1		0.111	0.107		0.047	0.014
		(0.152)	(0.159)		(0.135)	(0.143)
lambda2		-0.158	-0.105		-0.188*	-0.114
		(0.119)	(0.152)		(0.106)	(0.134)
Observations	2010	2010	2001	1914	1914	1908
Adjusted R-squared	0.14	0.14	0.136	0.147	0.147	0.139
P value of F-test ranks=0		0.14	0.24		0.05	0.14
P-value of F-test lambdas=0		0.59	0.82		0.34	0.83

Standard errors in parentheses
* significant at 10%; ** significant at 5%; *** significant at 1%
All equations include cohort polynomials.

Table 4: Pairwise and M-NN Matching, Lottery Sample

	1981					1980				
	Non-vet	**Private**	**Corporal**	**Sergeant**	**Officer**	**Non-vet**	**Private**	**Corporal**	**Sergeant**	**Officer**
# Treated	1369	179	252	182	28	1375	183	257	187	31
Panel A: Pairwise										
Mean ln(wage) difference		0.012	0.082	0.098	0.151		-0.013	0.118	0.054	0.182
Std Err		(0.088)	(0.061)	(0.082)	(0.148)		(0.079)	(0.061)	(0.069)	(0.143)
T-ratio		0.14	1.35	1.21	1.03		-0.16	1.93	0.79	1.27
Cov. Balance:										
Index 1		0.0012	0.0003	0.0043	0.0022		-0.0002	0.0019	0.0062	-0.0031
Index 2		-0.0011	-0.0012	0.0044	0.0005		-0.0024	-0.0012	0.0060	-0.0040
Panel B: M=5										
Mean ln(wage) difference		-0.021	0.143	0.090	0.341		-0.022	0.096	0.081	0.347
Std Err		(0.076)	(0.060)	(0.071)	(0.164)		(0.076)	(0.052)	(0.062)	(0.143)
T-ratio		-0.27	2.38	1.26	2.07		-0.29	1.85	1.30	2.43
Cov. Balance:										
Index 1		0.0025	0.0052	0.0113	0.0020		0.0030	0.0052	0.0112	0.0014
Index 2		-0.0019	0.0025	0.0073	0.0138		-0.0021	0.0026	0.0084	0.0131
Panel C: M=10										
Mean ln(wage) difference		-0.010	0.127	0.120	0.329		-0.021	0.086	0.121	0.347
Std Err		(0.071)	(0.052)	(0.065)	(0.151)		(0.070)	(0.044)	(0.056)	(0.132)
T-ratio		-0.14	2.43	1.84	2.17		-0.30	1.96	2.17	2.63
Cov. Balance:										
Index 1		0.0057	0.0081	0.0136	0.0186		0.0057	0.0076	0.0139	0.0104
Index 2		-0.0009	0.0040	0.0072	0.0245		-0.0018	0.0036	0.0079	0.0200
Balance Before Matching										
Index 1	-0.655	0.170	0.263	0.361	0.343	-0.647	0.154	0.241	0.345	0.305
Index 2	0.381	0.052	0.154	0.289	0.672	0.384	0.037	0.136	0.281	0.646

NB: Balance before matching: the number in the "nonvet" column refers to the average of the index among non-veterans.
The next four columns after that refer to difference of the average from non-veterans. For instance, -0.655 is the average value of index 1 among nonvets. For privates, the average value of index 1 is 0.17 higher, and so on.

Table 5: Caliper Matching - Lottery Sample

	1981					1980				
	Non-vet	**Private**	**Corporal**	**Sergeant**	**Officer**	**Non-vet**	**Private**	**Corporal**	**Sergeant**	**Officer**
# Treated	1369	179	252	182	28	1375	183	257	187	31
Panel A: Caliper = 0.007										
Mean ln(wage) difference		-0.054	0.130	0.062	0.147		-0.083	0.086	0.086	0.183
Std Err		(0.094)	(0.071)	(0.091)	(0.188)		(0.089)	(0.066)	(0.075)	(0.204)
T-ratio		-0.58	1.84	0.68	0.78		-0.93	1.30	1.15	0.90
Cov. Balance:										
Index 1		-0.0006	0.0005	0.0002	-0.0021		-0.0007	0.0007	0.0009	-0.0030
Index 2		-0.0024	-0.0001	0.0005	0.0004		-0.0033	-0.0001	0.0008	0.0006
Nonmatched		0.129	0.111	0.176	0.286		0.148	0.128	0.150	0.290
Panel B: Caliper = 0.01										
Mean ln(wage) difference		-0.055	0.145	0.082	0.243		-0.052	0.110	0.102	0.247
Std Err		(0.089)	(0.063)	(0.080)	(0.225)		(0.089)	(0.059)	(0.071)	(0.211)
T-ratio		-0.61	2.30	1.02	1.08		-0.58	1.85	1.43	1.17
Cov. Balance:										
Index 1		-0.0006	0.0008	0.0009	0.0027		-0.0022	0.0007	0.0014	-0.0009
Index 2		-0.0019	-0.0010	0.0008	0.0054		-0.0017	-0.0010	0.0009	0.0035
Nonmatched		0.084	0.071	0.126	0.179		0.087	0.090	0.107	0.226
Panel C: Caliper = 0.03										
Mean ln(wage) difference		-0.009	0.137	0.115	0.206		-0.021	0.074	0.143	0.323
Std Err		(0.077)	(0.054)	(0.066)	(0.157)		(0.076)	(0.044)	(0.056)	(0.156)
T-ratio		-0.12	2.57	1.74	1.31		-0.28	1.67	2.56	2.08
Cov. Balance:										
Index 1		0.0014	0.0017	0.0030	0.0024		0.0002	0.0010	0.0039	-0.0010
Index 2		-0.0005	-0.0001	0.0018	0.0049		-0.0014	-0.0011	0.0030	0.0037
Nonmatched		0.039	0.028	0.033	0.071		0.038	0.027	0.027	0.065
Balance Before Matching										
Index 1	-0.655	0.170	0.263	0.361	0.343	-0.647	0.154	0.241	0.345	0.305
Index 2	0.381	0.052	0.154	0.289	0.672	0.384	0.037	0.136	0.281	0.646

NB: Balance before matching: the number in the "nonvet" column refers to the average of the index among non-veterans.
The next four columns after that refer to difference of the average from non-veterans. For instance, -0.655 is the average value of index 1 among nonvets. For privates, the average value of index 1 is 0.17 higher, and so on.

Table 6: Pairwise and M-NN Matching, Full Sample

	1981					1980				
	Non vet	**Private**	**Corporal**	**Sergeant**	**Officer**	**Non vet**	**Private**	**Corporal**	**Sergeant**	**Officer**
# Treated	1665	234	302	222	32	1680	240	309	224	34
Panel A: Pairwise										
Mean ln(wage) difference		0.001	0.104	0.161	0.297		-0.020	0.102	0.172	0.389
Std Err		(0.073)	(0.051)	(0.067)	(0.180)		(0.070)	(0.052)	(0.071)	(0.182)
T-ratio		0.01	2.02	2.41	1.65		-0.28	1.95	2.43	2.14
Cov. Balance:										
Index 1		0.0010	0.0012	0.0037	0.0064		0.0007	0.0021	0.0040	0.0047
Index 2		-0.0008	0.0015	0.0024	0.0107		-0.0002	0.0033	0.0021	0.0084
Panel B: M=5										
Mean ln(wage) difference		-0.0166	0.1297	0.1404	0.2796		-0.018	0.100	0.132	0.307
Std Err		(0.067)	(0.051)	(0.063)	(0.156)		(0.064)	(0.046)	(0.059)	(0.136)
T-ratio		-0.25	2.54	2.23	1.79		-0.28	2.18	2.23	2.25
Cov. Balance:										
Index 1		0.0026	0.0022	0.0055	0.0034		0.0022	0.0029	0.0044	0.0087
Index 2		-0.0006	0.0017	0.0037	0.0117		0.0002	0.0028	0.0019	0.0126
Panel C: M=10										
Mean ln(wage) difference		-0.019	0.099	0.174	0.368		-0.028	0.077	0.146	0.311
Std Err		(0.061)	(0.045)	(0.056)	(0.151)		(0.057)	(0.040)	(0.053)	(0.123)
T-ratio		-0.31	2.20	3.12	2.44		-0.49	1.95	2.74	2.53
Cov. Balance:										
Index 1		0.00	0.01	0.01	0.01		0.00	0.00	0.01	0.01
Index 2		0.00	0.00	0.01	0.02		0.00	0.00	0.01	0.02
Balance Before Matching										
Index 1	-0.619	0.174	0.226	0.312	0.361	-0.610	0.163	0.209	0.296	0.330
Index 2	0.369	0.038	0.116	0.244	0.594	0.372	0.026	0.093	0.243	0.590

NB: Balance before matching: the number in the "nonvet" column refers to the average of the index among non-veterans. The next four columns after that refer to difference of the average from non-veterans. For instance, -0.619 is the average value of index 1 among nonvets. For privates, the average value of index 1 is 0.174 higher, and so on.

Table 7: Caliper Matching - Full Sample

	1981					1980				
	Nonvet	**Private**	**Corporal**	**Sergeant**	**Officer**	**Nonvet**	**Private**	**Corporal**	**Sergeant**	**Officer**
# Treated	1665	234	302	222	32	1680	240	309	224	34
Panel A: Caliper = 0.007										
Mean ln(wage) difference		-0.015	0.095	0.167	0.187		-0.0282	0.1169	0.1955	0.2279
Std Err		(0.079)	(0.055)	(0.069)	(0.167)		(0.071)	(0.060)	(0.076)	(0.170)
T-ratio		-0.19	1.72	2.40	1.12		-0.40	1.96	2.57	1.34
Cov. Balance:										
Index 1		-0.0005	0.0005	0.0012	0.0020		-0.0002	0.0004	0.0015	0.0026
Index 2		-0.0010	0.0000	0.0001	0.0042		0.0006	0.0006	0.0001	0.0044
Nonmatched		0.120	0.099	0.077	0.125		0.146	0.123	0.085	0.177
Panel B: Caliper = 0.01										
Mean ln(wage) difference		0.001	0.139	0.167	0.232		-0.040	0.096	0.187	0.274
Std Err		(0.072)	(0.054)	(0.066)	(0.165)		(0.070)	(0.052)	(0.075)	(0.169)
T-ratio		0.02	2.57	2.53	1.41		-0.57	1.85	2.51	1.62
Cov. Balance:										
Index 1		-0.0001	0.0017	0.0015	0.0008		0.0000	0.0014	0.0012	0.0011
Index 2		-0.0005	0.0012	0.0012	0.0037		0.0000	0.0017	0.0007	0.0037
Nonmatched		0.077	0.060	0.050	0.125		0.096	0.078	0.054	0.177
Panel C: Caliper = 0.03										
Mean ln(wage) difference		-0.001	0.128	0.185	0.326		-0.020	0.088	0.159	0.329
Std Err		(0.062)	(0.044)	(0.060)	(0.163)		(0.060)	(0.039)	(0.051)	(0.144)
T-ratio		-0.02	2.87	3.11	2.00		-0.34	2.24	3.08	2.28
Cov. Balance:										
Index 1		0.0016	0.0024	0.0024	0.0024		0.0017	0.0022	0.0027	0.0033
Index 2		0.0007	0.0021	0.0029	0.0029		0.0007	0.0024	0.0026	0.0049
Nonmatched		0.043	0.013	0.014	0.031		0.029	0.013	0.013	0.088
Balance Before Matching										
Index 1	-0.619	0.174	0.226	0.312	0.361	-0.610	0.163	0.209	0.296	0.330
Index 2	0.369	0.038	0.116	0.244	0.594	0.372	0.026	0.093	0.243	0.590

NB: Balance before matching: the number in the "nonvet" column refers to the average of the index among non-veterans.
The next four columns after that refer to difference of the average from non-veterans. For instance, -0.619 is the average value of index 1 among nonvets. For privates, the average value of index 1 is 0.174 higher, and so on.

Table A1: First Stage Selection Model - Full Sample

	Military Entry Equation			Rank Equation		
	Coeff	**Std err**	**T-ratio**	**Coeff**	**Std err**	**T-ratio**
Constant	-0.206	0.511	-0.40	-1.509	3.346	-0.45
Central MSA	0.174	0.065	2.67	0.014	0.422	0.03
Noncentral MSA	0.145	0.063	2.29	-0.106	0.362	-0.29
Black	-0.237	0.066	-3.56	-0.034	0.588	-0.06
Height	-0.003	0.006	-0.42	0.026	0.013	1.96
South Resident in 1966	-0.035	0.055	-0.64	0.178	0.113	1.57
Male Labor Market index in 1966	-0.043	0.057	-0.76	-0.135	0.129	-1.05
Draft Sequence Number * 100	-0.088	0.024	-3.67	-0.029	0.213	-0.14
unemployment rate 1960	-0.009	0.013	-0.65	-0.012	0.028	-0.41
size of labor force 1960	-0.029	0.025	-1.16	-0.025	0.080	-0.31
Father in the household at age 14	0.296	0.058	5.10	0.180	0.714	0.25
Respondent born in the US	0.201	0.056	3.57	-0.079	0.489	-0.16
Father alive in 1966	-0.233	0.086	-2.72	-0.070	0.566	-0.12
Mother alive in 1966	-0.313	0.138	-2.27	-0.066	0.757	-0.09
socio-economic index	0.045	0.028	1.58	0.214	0.129	1.66
socio-economic index squared	-0.179	0.026	-6.82	0.070	0.444	0.16
lottery sample	0.238	0.193	1.23	0.335	0.622	0.54
gamma2				1.046	0.043	24.05
gamma3				2.293	0.085	27.06
rho12				-0.001	3.406	0.00
cohort polynomial included?						yes
log likelihood						-3538.9
Number of observations						3554

Table A2: Second Stage of Parametric Selection Model - Full Sample

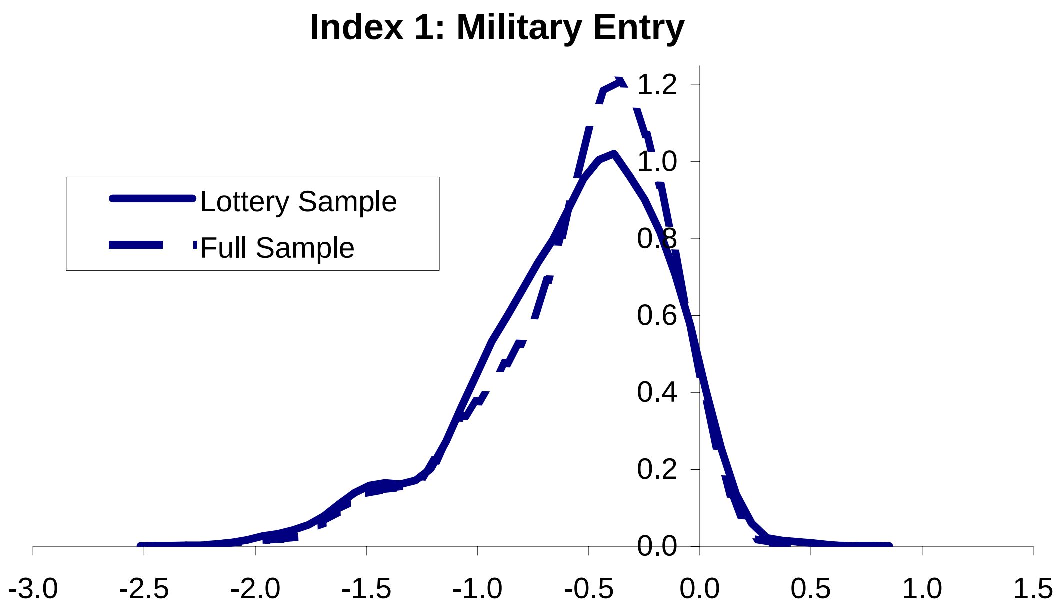
	OLS		IV	OLS		IV
	(1)	(2)	(3)	(4)	(5)	(6)
	lw81	lw81	lw81	lw80	lw80	lw80
years of schooling	0.068***	0.067***	0.091***	0.058***	0.056***	0.097***
	(0.006)	(0.006)	(0.029)	(0.005)	(0.005)	(0.026)
cohort1	-0.134***	-0.136***	-0.144***	-0.125***	-0.126***	-0.136***
	(0.015)	(0.018)	(0.020)	(0.014)	(0.017)	(0.018)
cohort2	-0.015	-0.014	-0.018	-0.02	-0.018	-0.02
	(0.015)	(0.019)	(0.019)	(0.014)	(0.017)	(0.017)
cohort3	0.008	0.005	0.002	0.015	0.012	0.01
	(0.017)	(0.018)	(0.018)	(0.015)	(0.017)	(0.017)
In MSA 1981	0.250***	0.254***	0.234***	0.263***	0.269***	0.230***
	(0.032)	(0.033)	(0.043)	(0.030)	(0.030)	(0.039)
black	-0.324***	-0.330***	-0.297***	-0.325***	-0.330***	-0.266***
	(0.037)	(0.041)	(0.059)	(0.035)	(0.038)	(0.056)
private	-0.008	-0.273	-0.219	-0.007	-0.269	-0.116
	(0.051)	(0.237)	(0.255)	(0.047)	(0.214)	(0.237)
corporal	0.156***	-0.015	-0.017	0.116***	-0.026	0.035
	(0.046)	(0.192)	(0.195)	(0.042)	(0.174)	(0.181)
sergeant	0.127**	0.007	-0.036	0.108**	0.024	0.011
	(0.053)	(0.180)	(0.183)	(0.048)	(0.164)	(0.166)
officer	0.15	0.124	-0.021	0.126	0.159	-0.029
	(0.131)	(0.221)	(0.269)	(0.118)	(0.201)	(0.234)
lambda_c		-0.028	-0.091		-0.04	-0.099
		(0.141)	(0.155)		(0.130)	(0.137)
lambda_1		0.112	0.107		0.073	0.036
		(0.134)	(0.136)		(0.121)	(0.125)
lambda_2		-0.112	-0.051		-0.144	-0.038
		(0.097)	(0.122)		(0.088)	(0.110)
Observations	2455	2455	2444	2349	2349	2341
Adjusted R-squared	0.175	0.175	0.169	0.18	0.18	0.161
P value of F-test ranks=0		0.15	0.32		0.14	0.41
P value of F-test lambdas=0		0.66	0.8		0.39	0.83

Standard errors in parentheses

* significant at 10%; ** significant at 5%; *** significant at 1%

All equations include cohort polynomials.

Figure 1: Indices From first Stage



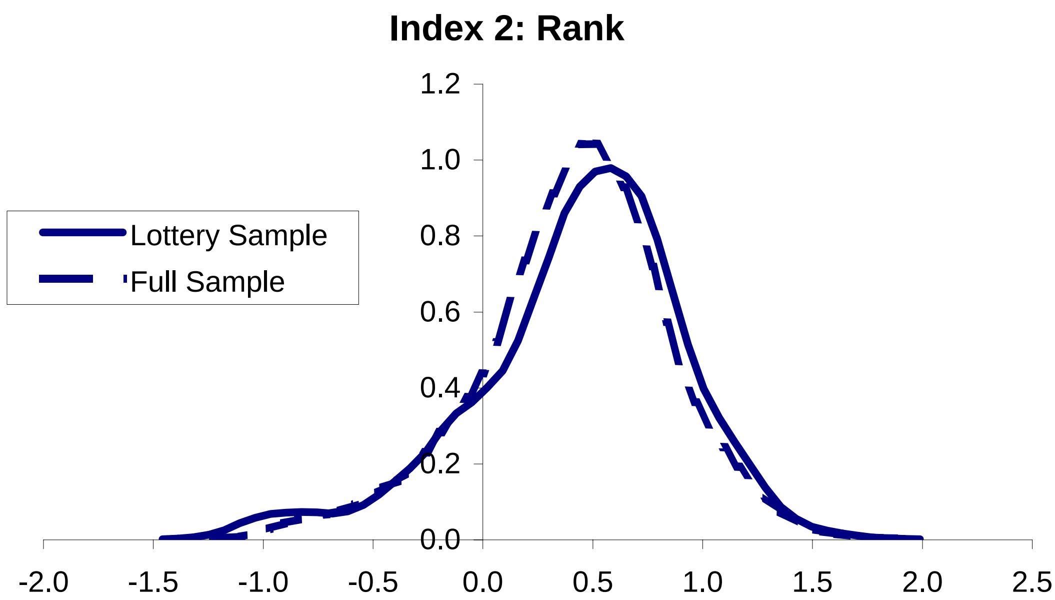


Figure A1: Rank by Region in (ε_1,ε_2) Space.