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Chris Rohlfs, *Syracuse University*

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The Economic Cost of Conscription and an Upper Bound on the Value of a Statistical Life: Hedonic Estimates from Two Margins of Response to the Vietnam Draft

Chris Rohlfs

Abstract

This study estimates the cost of the Vietnam draft by applying hedonic methods to the decision to attend college and the decision to voluntarily enlist. In 2009 dollars, the estimated cost of the draft is roughly \$115,000 for the marginal military recruit. For the marginal college student, the estimated cost is only \$30,000 and probably understates the true amount because men were credit-constrained and college required an upfront cost. Supposing that the costs other than fatality risk were positive, our preferred specifications produce an upper bound on the Value of a Statistical Life ranging from \$7 million to \$12 million.

KEYWORDS: conscription, draft, hedonic, Vietnam, value of a statistic life, deferment, enlistment, military

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1. INTRODUCTION

The majority of countries in the world procure military personnel through some form of draft (Mulligan and Shleifer, 2005). Economists have long argued against conscription in favor of an all-volunteer army (United States President's Commission on an All-Volunteer Armed Force, 1970). Recent evidence highlights the many long-term negative effects that conscription has on young men's lives (Dobkin and Shabani, 2009; Eisenberg and Rowe, 2009; Lindo and Stoecker, 2010; Rohlf's, 2010; Conley and Heerwig, in press). The draft is also a recent policy issue, and policymakers in Germany and Sweden recently moved to end conscription in those countries (Dempsey, 2010, Wallis, 2010). This study aims to provide a composite measure of the cost that the Vietnam draft imposed on young men.

Policies affecting national security are notoriously difficult to analyze in a cost-benefit framework, and the current study applies that framework to one important defense policy. For a country conducting a cost-benefit analysis of the switch from a draft to an all-volunteer force, it is fairly straightforward to determine the degree to which the switch would increase military personnel costs. It is less clear, however, how one might measure the dollar value of the utility gain to young men and women who would no longer be required to serve in the military. The current study provides estimates of the dollar value of that utility gain for the case of conscription in the U.S. during the Vietnam Era. To the extent that this utility benefit exceeds the cost savings to the military, switching to an all-volunteer force passes the cost-benefit test and would increase social welfare.

The estimates provided in this study can also help to evaluate the U.S. government's controversial "stop-loss" policy, which allows the military to extend the contract of any active duty service member without that service member's consent. The authority to impose involuntary extensions to service members' contracts was granted by the U.S. Military Congress in 1984. The number of soldiers working involuntarily through this policy at any given time ranged from 8,540 to 15,758 from November 2004 to May 2009 (Henning, 2009). In early 2009, the U.S. military took steps to reduce its use of this policy (Shanker, 2009). Estimates of the economic cost of mandatory service in the U.S. military provide a rough dollar-denominated measure of the utility cost of stop-loss to military personnel, a cost that the U.S. military can compare in a cost-benefit framework to the alternative cost of increasing reenlistment bonuses.

In addition to providing insights into the decision to end a draft, this study expands the domain of revealed preference studies to understand an important action that affects the lives of many people around the world. These estimates can also help to inform a growing literature measuring the value of different types of freedom. Previous studies have estimated the willingness-to-pay for freedom from

jail or from slavery (Cole, 2005; Abrams and Rohlfs, 2011); the current study on mandatory military service provides evidence on the economic cost of another important form of coercion. These estimates also provide insights into the Value of a Statistical Life (VSL). Being drafted exposes men to fatality risk and imposes additional economic, health, and utility costs. Assuming that these non-fatality related costs are positive, it is possible to construct an upper bound on the VSL that addresses an important obstacle that researchers have pointed out in previous VSL estimates – that of obtaining plausibly exogenous variation across agents in their choice sets (Ashenfelter and Greenstone, 2004; Ashenfelter, 2006). Examining the decision to avoid the draft helps to illustrate the broad range of behaviors from which VSL estimates can be generated and the contexts in which dollar-fatality tradeoffs are relevant.

The ideal experiment for estimating the cost of the draft would randomly assign different values of a “buyout price” across draft-age men, where paying the buyout price would exempt one from service. The distribution of willingness to pay for draft avoidance could then be estimated by measuring the fraction buying out at different price levels.

Although no buyout option existed during the Vietnam Era, young men took many costly actions to avoid being drafted. The current study focuses on two draft avoidance strategies: attending college and voluntarily enlisting. Previous work has found that increases in the risk of being drafted led to increases in the fraction of men engaging in these activities (Angrist, 1991; Card and Lemieux, 2001). This study converts these previous estimates into dollar-denominated measures of the “statistical cost of the draft” using hedonic methods developed in a theoretical study by Rohlfs (2011a). The statistical draft is defined in a way that is similar to the VSL. If a young man is willing to pay c dollars to avoid a 10 percentage point increase in the risk of being drafted, then the *statistical cost of the draft* is $\frac{c}{0.10}$ and is an estimate of the value of reducing one’s risk of being drafted from 100% to zero.

The intuition for this estimation approach is illustrated in Figure 1. During the Vietnam Era, a young man could reduce his risk of being drafted by enrolling in college and obtaining a student deferment. The two curves in Figure 1A plot the demand for college with and without this benefit. The vertical difference in demand measures the value of a deferment, which is estimated in Section 2 as the effect of this benefit on the number attending college (shown along the horizontal axis) divided by negative one times the slope of the demand curve for college attendance, where slope estimates are taken from Dynarski (2003). This ratio can be interpreted as the amount of tuition assistance that would be required to increase attendance by as much as a shield from the draft did.

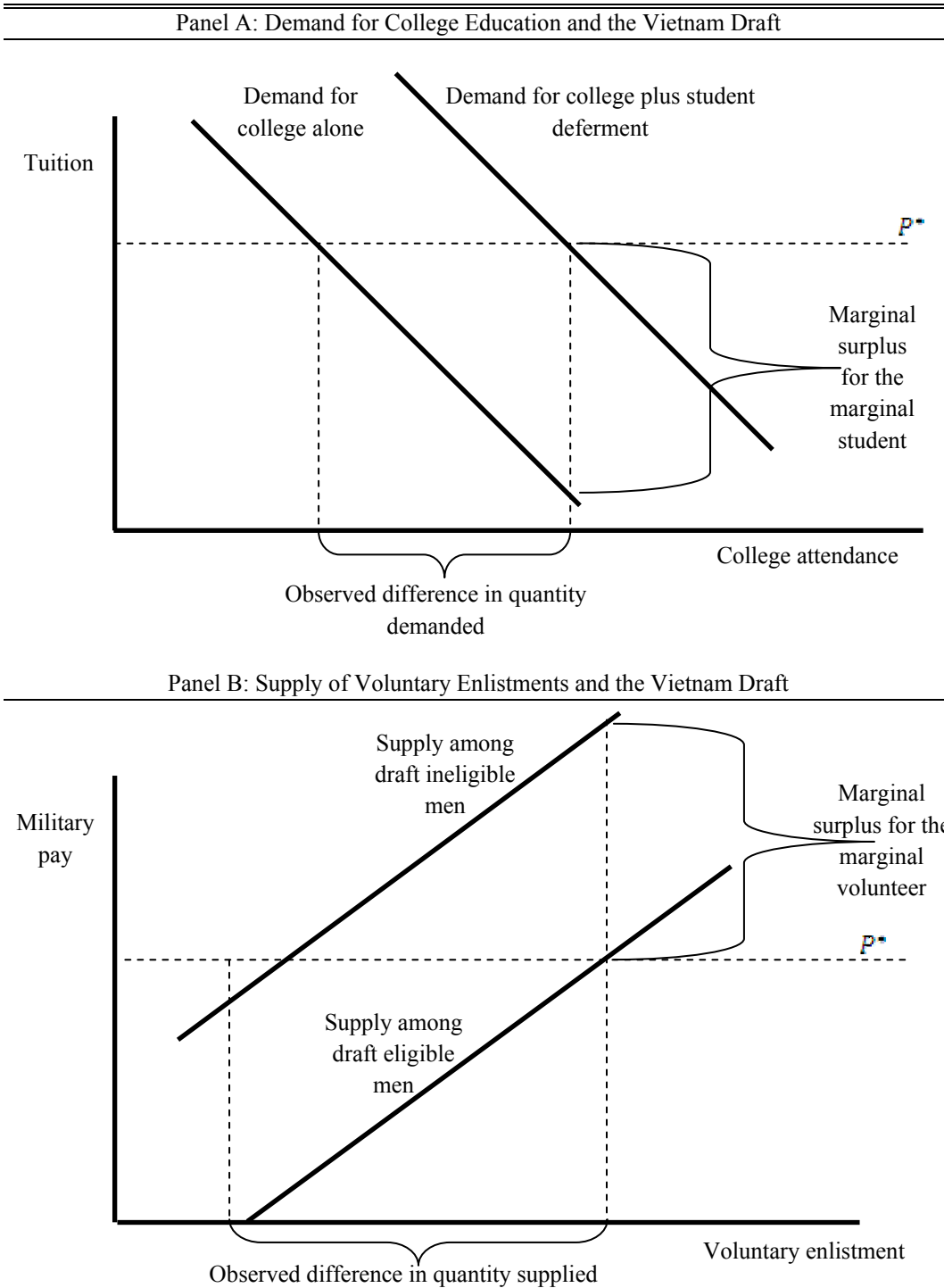
The two curves in Figure 1B plot the supply of military recruits among draft eligible and ineligible men. For draft eligible men, working in the civilian

sector entailed a risk of having that career interrupted with a call to military service. This risk increased the desirability of voluntary enlistment which provided a certain timeline and the opportunity to obtain a safer military posting. The effect of draft eligibility on the fraction joining the military is shown on the horizontal axis and is estimated in Section 3 using data from Angrist (1991) on the effect of lottery-induced conscription risk on voluntary enlistment. The slope of the supply curve for military labor is taken from a 1983 enlistment bonus experiment by Polich, Dertouzos, and Press (1986). The vertical difference in the supply curves measures the cost of the draft for the marginal recruit – the bonus that would be required to increase enlistment by as much as draft eligibility did – and is estimated by dividing the quantity difference by the slope of the supply curve.

The vertical differences in Figure 1 — which are labeled “marginal surplus” — are comparable to marginal willingness to pay in standard hedonic models. The method used here is transparent and does not require the standard hedonic model’s restrictions of perfect competition, “thick” markets, or separability of demand or supply between the product of interest and all other goods. We require for consistency that the quantity differences and the slopes of the price responses are identified and that Dynarski’s (2003) estimated effect of tuition on college attendance among children of deceased fathers in the early 1980s and Polich, Dertouzos, and Press’s (1986) estimates of the slope of the military labor supply function in 1983 both generalize to young men in the Vietnam Era.

Figure 1: Graphical Illustration of Estimation Strategy

(A) Demand for college education and the Vietnam draft (B) Supply of voluntary enlistments and the Vietnam draft



The combination of findings obtained in this study is at first puzzling and helps to provide insights into the cost of the draft. Evidence from the voluntary enlistment margin indicates that the statistical cost of the draft was between \$103,000 and \$131,000 in 2009 dollars. Hence, the cost of conscription appears to have been substantial, even for an individual on the margin of voluntary service. Interestingly, the college attendance margin produces lower estimates of the statistical cost of the draft, ranging from \$23,000 to \$36,000. One likely explanation for this discrepancy is that potential draftees were credit constrained, and due to the upfront cost of college, the combined demand for college and draft avoidance was less than the demand for either good provided individually. The lower estimates based upon college attendance are probably most appropriate for measuring actual willingness to spend money for draft avoidance and the potential demand for a draft buyout, which would involve a large upfront cost. The higher numbers that do not reflect the effects of credit constraints are probably most appropriate for evaluating the cost of the draft from the perspective of an optimal social planner.¹

These costs of conscription estimates are used to construct an upper bound on the VSL in Section 4. For the college attendance margin, this upper bound estimate is approximately \$3 million, which is low relative to existing estimates of the VSL from that era, possibly due to the interaction between credit constraints and the upfront costs of schooling. In our preferred specifications, which are taken from the voluntary enlistment margin, the implied upper bound on the VSL varies from \$7.28 million to \$11.9 million.

2. COLLEGE ATTENDANCE AS A MARGIN OF RESPONSE

2.1 CONCEPTUAL FRAMEWORK

Let the fraction attending college among a set of individuals with characteristics \mathbf{x} and sex $s \in \{f, m\}$ from cohort c be denoted $Q_{sc}^{college}(\tau, \delta, \mathbf{p}, \mathbf{x})$, where τ is a tuition subsidy that can be applied to any college, δ is the reduction in conscription risk that is provided by attending college, \mathbf{p} is a vector of prices of

¹ This difference is related to the well-documented difference between “willingness to pay” and “willingness to accept,” where the value of a good being sold is generally larger than the value of a good being purchased, possibly due in part to liquidity constraints for buyers (cf., Horowitz and McConnell, 2002). Attending college and enlisting in the military were costly actions that young men had to take to avoid conscription; hence, both estimates can be thought of as “willingness to pay” estimates; however, the requirement for liquidity was substantially larger for college than for voluntary enlistment.

colleges and other relevant goods, and \mathbf{x} is a vector of individual characteristics that influence the demand for college and the ability to avoid the draft. Suppose that this function has the following log-linear form:

$$\ln \left(Q_{sc}^{college}(\tau, \delta, \mathbf{p}, \mathbf{x}) \right) = -\alpha_{\tau} * \tau + \alpha_{\delta s} * \delta + \mathbf{p}'\boldsymbol{\alpha}_{\mathbf{p}}(\mathbf{x}) + \alpha_c(\mathbf{x}) + \alpha_s(\mathbf{x}) + \varepsilon_{sc}(\mathbf{x}), \quad (1)$$

where α_{τ} , $\alpha_{\delta s}$ and $\boldsymbol{\alpha}_{\mathbf{p}}(\mathbf{x})$ represent the demand effects of the tuition subsidy, conscription risk, and the price vector, $\alpha_c(\mathbf{x})$ and $\alpha_s(\mathbf{x})$ represent cohort- and sex-specific fixed effects, and $\varepsilon_{sc}(\mathbf{x})$ is random error that is mean zero for all \mathbf{x} . Hence, the effects of sex, cohort, and prices may vary depending on the individual characteristics \mathbf{x} . We require that the effect of tuition is constant with respect to \mathbf{x} , so that an estimate of α_{τ} can be taken from a separate data source. We also assume for simplicity of interpretation that the effect of conscription risk is constant with respect to \mathbf{x} . Finally, we require for our difference-in-differences strategy that the effects of tuition, other prices, and cohort are common across sexes.

Aggregate demand is observed separately by sex for a non-draft cohort $c = 0$ and a cohort $c = 1$ that is subject to the draft. Let $\tau_c(\mathbf{x})$ and $\mathbf{p}_c(\mathbf{x})$ denote the tuition subsidy and price vector faced by an individual with characteristics \mathbf{x} from cohort c ; these values are assumed to be constant across sexes. Additionally, let $\delta_{sc}(\mathbf{x})$ denote the reduction in risk that college would provide for an individual with characteristics \mathbf{x} from sex s in cohort c . We apply the normalization that, $\tau_0(\mathbf{x}) = \tau_1(\mathbf{x}) = 0 \forall \mathbf{x}$ so that any differences in the cost of attendance across individuals and cohorts are reflected in the price vectors $\mathbf{p}_0(\mathbf{x})$ and $\mathbf{p}_1(\mathbf{x})$. The risk of being drafted is always zero for all women, so that $\delta_{f0}(\mathbf{x}) = \delta_{f1}(\mathbf{x}) = 0 \forall \mathbf{x}$, and the distribution $f(\mathbf{x})$ of characteristics \mathbf{x} defined over the support \mathbf{X} is the same for men and women.

As in a standard supply and demand setting, the relative quantity $Q_{sc}^{college}$ of people attending college and the various prices $\mathbf{p}_c(\mathbf{x})$ of different types of college and complements and substitutes to college are all potentially endogenous and may be affected by exogenous changes in the number of males being drafted. The increased demand for college among men from draft cohorts may lead to higher tuitions, overcrowding, or tougher admissions standards for these cohorts, and these effects may have been larger at some types of colleges than at others. By assumption, these effects are common across men and women, so that women can serve as a control group to capture the effects of cohort-level changes in tuition, college quality, and admissions standards.

Potential violations of the assumption that $\boldsymbol{\alpha}_{\mathbf{p}}(\mathbf{x})$ is constant across sexes include differences across sexes in the elasticity of college attendance with

respect to prices or college quality, changes in aspects of college quality that are valued differently by sex (such as the male-female sex ratio), a draft-induced reduction in unskilled male labor supply resulting in a sex-specific increase in the wages of men relative to women (making college less attractive for men in the short run), and sympathy for men by admissions boards (so that admissions standards were relatively tough for women in draft years). The practice of going to college to avoid the draft is well documented (e.g., Baskir and Strauss, 1978), and draft avoidance is probably a first-order effect that was large relative to these other factors. Nevertheless, these other factors probably did affect the demand for college to varying degrees, and it is important to keep these potential biases in mind when interpreting the estimates from this portion of the study.

Given the assumption that $\alpha_p(\mathbf{x})$ is constant across sexes, the conditional expectation of the log of the male-female attendance ratio can now be written as:

$$E[\ln(Q_{mc}^{college}/Q_{fc}^{college}) | \mathbf{x}] = (\alpha_m(\mathbf{x}) - \alpha_f(\mathbf{x})) + \alpha_{\delta m} * \delta_{mc}(\mathbf{x}). \quad (2)$$

Assuming that the unobservable determinants of male versus female taste for education are uncorrelated with the change in conscription risk, the expected difference in the log male-female attendance ratio between cohorts one and zero is $\alpha_{\delta m} * (\delta_{m1}(\mathbf{x}) - \delta_{m0}(\mathbf{x}))$. Averaging this difference-in-differences over \mathbf{x} and dividing by $\int_{\mathbf{x}} (\delta_{m1}(\mathbf{x}) - \delta_{m0}(\mathbf{x}))f(\mathbf{x})d\mathbf{x}$ produces $\alpha_{\delta m}$, the effect of student deferments on the demand for college, which corresponds to the difference in quantity shown in Figure 1A.²

² Strictly speaking, because the log-linear functional form is assumed conditional on \mathbf{x} , identifying $\alpha_{\delta m}$ requires taking the log of the geometric mean over \mathbf{x} of the exponential of each \mathbf{x} -specific difference-in-differences. For ease of interpretation and due to the structure of the data, an arithmetic mean is instead used for the empirical results in this paper.

The risk of being drafted varied across men due to differences in the ability to obtain deferments or exemptions and the ability to pass the entrance exams. Supposing that $\alpha_{\delta m}$ is constant with respect to these abilities, the difference-in-differences strategy described here will identify this constant effect. The economic cost of the draft can be expressed by rewriting Eq. (1) in terms of inverse demand:

$$-\tau(Q, \delta, \mathbf{p}, \mathbf{x}) = -\frac{\alpha_{\delta s}}{\alpha_{\tau}} * \delta - \frac{1}{\alpha_{\tau}} (\ln(Q) + \mathbf{p}'\alpha_{\mathbf{p}}(\mathbf{x}) + \alpha_c(\mathbf{x}) + \alpha_s(\mathbf{x}) + \varepsilon_{sc}(\mathbf{x})) \quad (3)$$

Holding other factors constant, the effect of a one-unit change in δ on the surplus obtained from college is $-\alpha_{\delta s}/\alpha_{\tau}$, which corresponds to the vertical difference in demand shown in Figure 1A. This vertical difference can be interpreted as the amount of the tuition subsidy that would be required to increase college attendance by as much as the risk of being drafted did. Supposing that estimates of α_{τ} and $\int_{\mathbf{x}} (\delta_{m1}(\mathbf{x}) - \delta_{m0}(\mathbf{x}))f(\mathbf{x})d\mathbf{x}$ can be measured from other sources, the marginal surplus shown in Figure 1A is then estimated as $-1/\left[\alpha_{\tau} * \int_{\mathbf{x}} (\delta_{m1}(\mathbf{x}) - \delta_{m0}(\mathbf{x}))f(\mathbf{x})d\mathbf{x}\right]$ times the difference-in-difference in the log of college attendance. This estimation strategy is an application of the “placebo consumers estimator” developed in Rohlfs (2011a, p. 34).

To the extent that $\alpha_{\delta m}$ varies with \mathbf{x} , the difference-in-differences in college attendance identifies a Local Average Treatment Effect (LATE; Angrist and Imbens, 1994), which can be viewed as a weighted average of $\alpha_{\delta m}(\mathbf{x})$, with greater weight placed on the observations for which the change $(\delta_{m1}(\mathbf{x}) - \delta_{m0}(\mathbf{x}))$ in risk was higher. Hence, the effect of the draft on college attendance is measured for a population of men who were relatively unable to avoid the draft through other means.

2.2 STUDENT DEFERMENTS AND THE RISK OF BEING DRAFTED

To measure the reduction in conscription risk provided by a student deferment, the fraction drafted is estimated separately by birth year for men who did not obtain exemptions or deferments and those who obtained student deferments. The data source for these fractions is the National Vietnam Veterans Readjustment Study (NVVRS; Research Triangle Institute, 1989, described in detail in Rohlfs, 2010). Men who volunteered for service are excluded from the sample. Owing to measurement problems in the survey data, administrative sources are used to calculate conscription risk for the main lottery cohorts (1950 and later) for non-

students.³ Following Card and Lemieux, a 3-year moving average is used; for 1949, the last birth year primarily affected by the pre-lottery rules, a 2-year average of 1948 and 1949 is used.

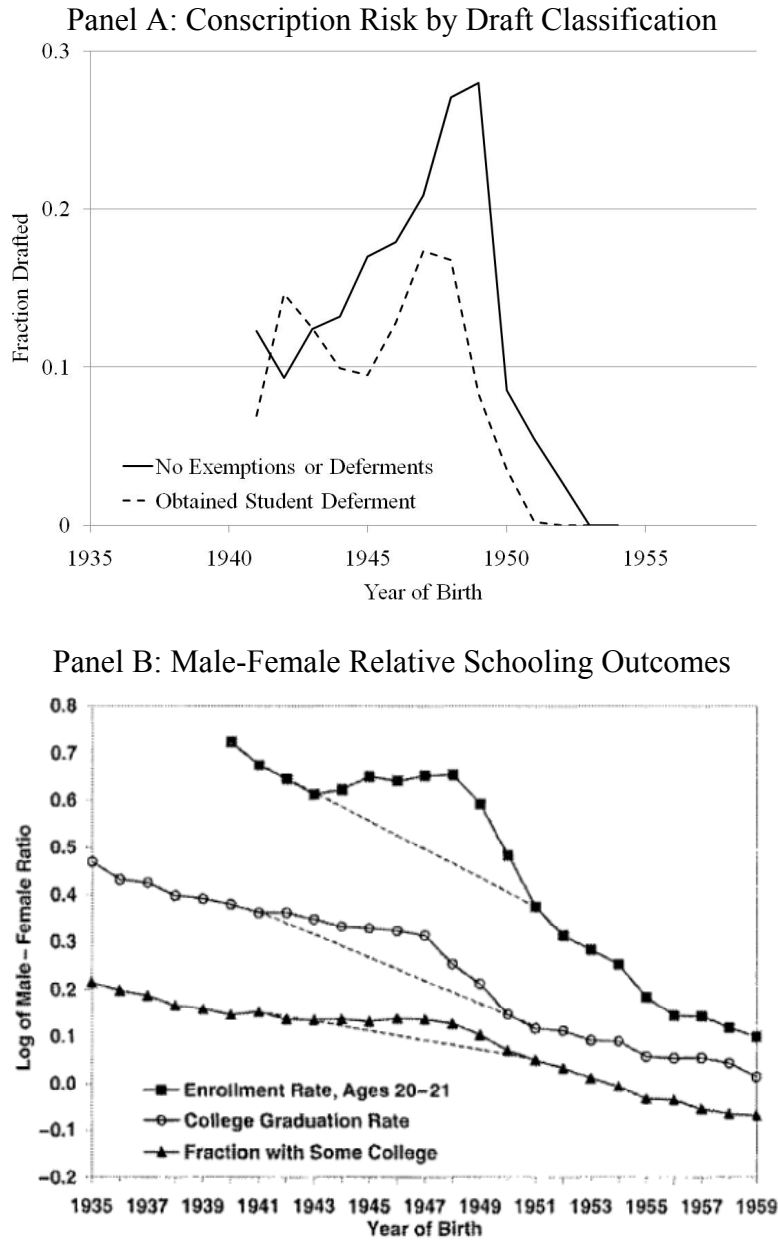
By directly estimating conscription risk for those who did and did not obtain student deferments, the risk measure used here more accurately estimates the tradeoffs faced by the marginal college student than does Card and Lemieux's measure, which supposes that the risk of being drafted is zero for college students and equal to the number of conscriptions per capita for non-students. Nevertheless, we must assume that the marginal college student could not avoid the draft through other means such as conscientious objector status, a divinity or farm occupation, or fatherhood and had the average likelihood of passing the military's physical and cognitive exams (56.8% over 1965–1973; United States Census Bureau, 1974). The biases introduced by these assumptions partially counteract each other, as the marginal student probably had alternative deferment options and above average likelihood of passing the exams.

Figure 2A shows the fraction drafted by birth year among those men who did not obtain exemptions or deferments and those who obtained student deferments. The likelihood of being drafted was the largest for men born in 1947 and 1948, and the difference in the fraction drafted between those without and with student deferments was also largest for these cohorts. The measure of conscription risk used in this study is this difference in these two fractions, and it is supposed that attending college reduced one's risk of being drafted from the level shown on the solid line to the dashed one, a difference that varied by cohort.

³ The fraction drafted among men born in 1950 and 1951 is unusually high in the NVVRS relative to other datasets, probably due to small non-veteran sample sizes and lottery-induced enlistment being misclassified as conscription. Additionally, unlike the schooling decision, the enlistment decision was often made after lottery numbers were known, so that the population who did not volunteer is not comparable between the pre-lottery and lottery periods (cf. Angrist, 1991; Angrist and Krueger, 1992). For these reasons, conscription risk for non-students in the lottery periods is computed from administrative data as the ratio of draftees to those examined (from U.S. Census Bureau, 1972, 1974; U.S. Selective Service System, 1971–1973) times the fraction of lottery numbers that were called for examination (195/365, 125/365, and 95/366 for the 1950, 1951, and 1952 cohorts and zero for later cohorts). Excluding these cohorts from the sample has negligible effects on the final estimates.

Figure 2: Fraction Drafted and Log Male-Female Schooling Ratio by Cohort

- (A) Conscription risk by draft classification
- (B) Male-female relative schooling outcomes



Notes to Figure 2: Panel A is constructed from the NVVRS and administrative sources as described in the text. The unusual pattern of higher risk among deferrers in 1941–1942 is driven by very small sample sizes among the earliest birth cohorts. Panel B is reproduced with permission from Card and Lemieux (2001, p. 100).

2.3 STUDENT DEFERMENTS AND COLLEGE ATTENDANCE

Figure 2B, reproduced from Card and Lemieux's (2001) influential study, illustrates the difference-in-differences strategy used to measure the effect of the draft on college attendance. The log of the ratio of men to women attending college is plotted along the vertical axis, and year of birth is shown on the horizontal axis. The top curve shows the log of the male-female ratio of the fraction attending college at ages 20–21 years in the October Current Population Survey (CPS; Bruno and Curry, 1995). The middle and lower curves show the log of the male-female ratio of the fraction that had graduated college and the fraction that had completed some college by 1990, as estimated from the U.S. Census. All three curves show declines in the log male-female ratio in successive birth cohorts, as college attendance became increasingly common among women. Additionally, all three curves show departures from these downward trends for the cohorts most affected by the Vietnam draft, changes that Card and Lemieux attribute to draft avoidance behavior among men.

The CPS data on fraction enrolled at ages 20–21 years produce much larger estimated effects of the draft than do the Census data on the fraction who graduated or completed some college by 1990. Card and Lemieux attribute this difference to current servicemen being excluded from the CPS data but included as veterans in the Census data. The relevant population in this study is those who did not volunteer, some of whom were drafted; hence, the appropriate sample population is something between what is found in these two datasets.

2.4 TUITION SUBSIDIES

Dynarski's (2003) estimates of the slope of the demand curve for college attendance use a difference-in-differences strategy that focuses on the elimination in 1982 of a ~\$8,300 (measured in 2009 dollars) Social Security tuition benefit for the children of deceased fathers. The effect used here is the coefficient from the regression with covariates from that paper (0.219, from p. 283) divided by \$8,300, giving a per dollar effect of a tuition subsidy on the probability of attending college by age 23 years of -2.64×10^{-5} with a standard error of 1.33×10^{-5} . As Dynarski shows, children of deceased fathers are significantly more likely than average to be black and to have parents who did not attend college. The fraction attending college among children of deceased fathers was 0.63 among cohorts receiving the subsidy and 0.32 among those not receiving the subsidy. If the relatively poor population examined by Dynarski is more price sensitive than the marginal college student in the Vietnam Era, then the procedure used here will understate the true cost of the draft. As Kane (2004, pp. 340–341) discusses,

however, recent studies have failed to find an interaction effect between tuition and parental income on college attendance.

2.5 THE COST OF THE DRAFT

Table 1 presents the results from the schooling portion of the current study. Columns 1 and 2 show estimates in which schooling is measured as the fraction attending college at ages 20–21 years, as in the top curve in Figure 2B. Columns 3 and 4 measure schooling as the fraction that completed some college by 1980.⁴ Columns 1 and 3 estimate the effect of the draft on schooling using Card and Lemieux’s linear interpolation approach that assumes that, in the absence of the draft, the log male-female schooling ratio would have declined linearly according to the dashed line in Figure 2B. Columns 2 and 4 estimate this effect using Card and Lemieux’s alternative approach in which the log of the male-female schooling ratio is regressed on cohort-specific conscription risk (in this case the difference in the two curves shown in Figure 2A), controlling for a time trend.

The first row of Table 1 shows the log difference in schooling between the birth years with the highest conscription risk and those with zero risk. For the linear interpolation method, this difference is computed as the distance between the actual and interpolated enrollment ratios shown in Figure 2B, averaged over 1947–1949. For the regression-based method, this difference is computed as the average over the 1947–1949 cohorts of the risk reduction provided by student deferments multiplied by the coefficient on that risk reduction in the schooling regression. The risk reduction provided by deferments is computed as the difference between the two curves in Figure 2A, averaged over the 1947–1949 cohorts (hence, a 3-year average over an already smoothed variable). Computed directly from the microdata, this weighted mean difference is 0.112 with a standard error of 0.026.

Row 2 presents estimates of the statistical cost of the draft, computed as row 1 divided by the 0.112 risk reduction, then divided by Dynarski’s (2003) estimate of the per dollar effect of a tuition subsidy. Two methods are used to compute the standard error for this ratio, both of which take into account the imprecision in the time series schooling regressions, the risk measure, and the tuition effect.⁵ They do not, however, take into account the imprecision and

⁴ The estimates in this table use the 1980 Census (Ruggles *et al.*, 2004) rather than the 1990 Census used by Card and Lemieux because the microdata samples of the 1980 Census include both age and quarter of birth. The quarter of birth variable helps to avoid misclassification of birth years, as those with birthdays before the April 1 survey date (in quarter one) have birth years of 1979 minus age rather than 1980 minus age.

⁵ White heteroskedasticity-consistent standard errors are used for the time series regressions, which include 20 observations for the 1941–1960 birth years. Dynarski’s reported standard error is used for the tuition effect. Enrollment and conscription risk are assumed to be independent across

possible inaccuracy associated with applying Dynarski's slope estimate from the population of students with deceased fathers in 1982 to the population of draft-age men in the Vietnam Era. The standard errors shown in parentheses are computed using the delta method. The 95% confidence interval (CI) shown in brackets below it replaces the imprecise linear approximation that the delta method imposes for a ratio of two coefficients with the formula from Hinkley (1969) for the Cumulative Distribution Function (CDF) of the ratio of two normally distributed random variables.

Table 1 College attendance and the statistical cost of the draft

| | 1 | 2 | 3 | 4 |
|--|--|--------------------------------|---|--------------------------------|
| | Effect on Schooling estimated using... | | | |
| | Fraction enrolled at ages 20–21 years | | Fraction completed some college by 1980 | |
| | Linear interpolation | Conscription risk as regressor | Linear interpolation | Conscription risk as regressor |
| 1. Log difference in enrollment between highest and lowest draft years | 0.187 (0.092)** | 0.119 (0.038)** | 0.024 (0.017) | 0.017 (0.006)** |
| 2. Estimated statistical cost of the draft | \$63,515 (45,553) | \$40,306 (20,693)* | \$8,247 (7,120) | \$5,930 (3,142)* |
| Hinkley 95% CI | [-204,345, 475,088] | [-139,562, 292,948]* | [-26,153, 65,203] | [-20,310, 43,238]* |

Notes to Table 1: Row 1 shows the log difference in the male-female college attendance ratio between the peak years of the draft (1947–1949) and the lowest draft years. For columns 1 and 3, the lowest draft years are estimated as a linear interpolation between the 1941 and 1951 cohorts, as in Figure 2B. For columns 2 and 4, this difference is estimated by regressing the log male-female

cohorts. For the linear interpolation, a constant standard deviation is assumed for the time series observations and is estimated from the variation in the log of the male-female enrollment ratio across the 1941–1951 cohorts. The variance of the linear interpolation is then computed as the variance of the difference between the average from 1947 to 1949 and the predicted value (a weighted average of the 1941 and 1951 values). For conscription risk, a separate standard deviation is measured from the NVVRS microdata for each of the students' and non-students' moving averages for each of 1947, 1948, and 1949. The variance of the effect of a student deferment on risk is then computed as the variance of the difference between students and non-students in these 3-year averages, averaged over 1947–1949.

attendance ratio on the difference in risk shown in Figure 2A and computing the effect on the log male-female attendance ratio of moving from zero risk difference to the average risk difference from 1947 to 1949. Columns 1 and 2 show results from the CPS, and columns 3 and 4 show results from the 1980 U.S. Census. Row 2 divides the numbers in row 1 by 2.64×10^{-5} , Dynarski's (2003) estimated per dollar effect of a tuition subsidy on log enrollment, and then divides by 0.112, the average difference in conscription risk from 1947 to 1949. The standard errors in parentheses in row 2 are computed using the delta method. The 95% confidence intervals in brackets are computed by treating the numerator and the denominators of the ratios as independent normal random variables and using Hinkley's (1969) formula for the CDF of the ratio of two independent normal random variables. ** or * denotes that zero falls outside the 95% or 90% confidence interval. Additional details in the text.

All four estimated statistical costs of the draft are imprecisely estimated; the two using conscription risk as a regressor in columns 2 and 4 are marginally significant using both the delta method and the Hinkley 95% CI, and the two estimates using the linear interpretation are nonsignificant using both the delta method and the Hinkley 95% CI. Averaging across the estimates from the CPS and Census datasets, the linear interpolation method in columns 1 and 3 produces an estimated statistical cost of the draft of \$35,800. The same average for the regression approach in columns 2 and 4 produces an estimated cost of the draft of \$23,100.

3. VOLUNTARY ENLISTMENT AS A MARGIN OF RESPONSE

3.1 CONCEPTUAL FRAMEWORK

Estimating marginal surplus for the marginal military recruit uses a similar strategy as for the marginal college student, as shown in Figure 1 B. Among men with characteristics \mathbf{x} and randomly assigned draft eligibility $l \in \{0,1\}$, let $Q_l^{enlist}(b, \delta, \mathbf{p}, \mathbf{x})$ denote the fraction of men voluntarily joining the military. This quantity depends on the enlistment bonus b , the risk δ of being drafted if one does not voluntarily enlist, the vector \mathbf{p} of the prices of all relevant goods and the vector \mathbf{x} of individual characteristics. Assuming a log-linear functional form produces:

$$\ln \left(Q_l^{enlist}(b, \delta, \mathbf{p}, \mathbf{x}) \right) = \beta_b * b + \beta_\delta * \delta + \mathbf{p}'\boldsymbol{\beta}_p(\mathbf{x}) + \beta(\mathbf{x}) + \varepsilon_l(\mathbf{x}), \quad (4)$$

w where β_b , β_δ , and $\boldsymbol{\beta}_p(\mathbf{x})$ are the supply effects of the enlistment bonus, conscription risk, and all other prices, $\beta(\mathbf{x})$ is a constant term, and $\varepsilon_l(\mathbf{x})$ is mean zero random error that is uncorrelated with eligibility. The constant term and the

effects of prices may vary depending on \mathbf{x} . The effect of the enlistment bonus is assumed to be constant, so that estimates of this effect can be taken from a different data source. As in the college attendance case, the effect of conscription risk is assumed to be constant with respect to \mathbf{x} for simplicity of interpretation.

Symmetrically to the college attendance case, let $b(\mathbf{x})$ and $\mathbf{p}(\mathbf{x})$ denote the enlistment bonus and price vector faced by an individual with characteristics vector \mathbf{x} ; these values are assumed to be constant across draft eligibility categories. Additionally, let $\delta_l(\mathbf{x})$ denote the risk of being drafted for an individual with characteristics \mathbf{x} from eligibility category l . We will apply the normalization that $b(\mathbf{x}) = 0 \forall \mathbf{x}$, so that any differences in military pay across people are reflected in prices $\mathbf{p}(\mathbf{x})$. The distribution $f(\mathbf{x})$ of characteristics \mathbf{x} defined over the support \mathbf{x} is the same for draft eligible and ineligible men.

For men with characteristics \mathbf{x} , the expected log difference in enlistment between draft eligible and ineligible men is $\beta_\delta * (\delta_1(\mathbf{x}) - \delta_0(\mathbf{x}))$, and the cost of a one-unit change in conscription risk can be expressed as $1 / \left[\beta_b * \int_{\mathbf{x}} (\delta_1(\mathbf{x}) - \delta_0(\mathbf{x})) f(\mathbf{x}) d\mathbf{x} \right]$ times this difference, where β_b and $\int_{\mathbf{x}} (\delta_1(\mathbf{x}) - \delta_0(\mathbf{x})) f(\mathbf{x}) d\mathbf{x}$ are estimated from other sources. To the extent that β_δ varies with \mathbf{x} , our estimated effect of draft eligibility on enlistment can be interpreted as a LATE, so that, as in the schooling case, our estimated effect measures the effect of the draft on recruitment for a subpopulation who was relatively unable to avoid it.

3.2 THE DRAFT LOTTERY AND CONSCRIPTION RISK

To measure the effect of conscription risk on voluntary enlistment, Angrist (1991) examines the effects of the draft lottery drawings for men born in 1951 and 1952. Through these drawings, each man born in 1951 and 1952 received a random number from 1 to 365 or 366 based upon his birthday, with lower lottery numbers corresponding to higher risk of conscription. Lottery numbers were drawn in the summer (July 1, 1970 and August 5, 1971), but draft eligible men were not called to report to their local draft offices for inspection until the following year. Men with lower lottery numbers were called first, and the "ceilings" (i.e., the highest lottery numbers to be called for inspection) of 125 and 95 were not announced until October 1971 and September 1972 (Angrist, 1991; Angrist and Krueger, 1992).

Men could voluntarily enlist at any time before their numbers were called. Voluntarily enlisting provided one the opportunity to select one's branch of service (Army, Navy, Air Force, or Marine Corps), obtain specialty training, and possibly enter as an officer. Conscription involved a 2-year term, generally in a non-specialized occupation in the Army, and volunteering typically involved a 3-

year term in the Army or a 4-year term in any of the other branches. In the enlisted ranks over the 1971–1973 fiscal years, 40.7% of first enlistments were into the Army. Over the same period, the 36.7% of active duty officers were in the Army (Computed from Tables 507 and 515, United States Census Bureau, 1974). Among those joining the military between the 1971 and 1973 fiscal years, 10.3% entered as officers. For the purposes of this study, as in Angrist (1991), enlisting is a general term that includes those voluntarily joining the enlisted ranks and those commissioned into the officer ranks.

Following Angrist (1990, 1991), we compare conscription risk and enlistment rates between draft eligible men whose lottery numbers equaled or fell below the ceilings and draft ineligible men whose lottery numbers exceeded the ceilings. Unlike the perceived risk of being drafted, which probably varied continuously with draft lottery numbers, Angrist (1991) shows that the fraction actually drafted was relatively flat among draft eligible men. The probability of being drafted that is used in this study is the number of men who were actually drafted divided by the size of the at-risk population (the men from that eligibility group who did not enlist). This risk is computed separately by eligibility category. Only conscriptions and enlistments occurring in the lottery year and the year afterward are counted in the numbers who enlisted or were drafted. Over the year during which the cut-off for draft eligibility was unknown, men with lottery numbers just above the cut-off probably believed that they had positive risks of being drafted. Consequently, the effect of eligibility status on the perceived likelihood of being drafted is probably smaller than the effect estimated in this study. For this reason, the statistical cost of the draft as estimated in this section may be biased toward zero, and the true cost per probability unit may be larger than what is presented here.

The numbers of men who were drafted and who voluntarily enlisted by birth cohort and eligibility category are taken from data reported in Angrist (1991). Two data sources are used to construct alternative sets of estimates of the numbers of American men by eligibility group and cohort. The first data source is a special extract requested by Angrist of the Social Security Administration's Continuous Work History Sample (CWHS). These data, used in Angrist (1990) and provided on the author's website, are tallies from a 1% sample of all individuals with Social Security numbers.⁶ The second data source is the total

⁶ Angrist (1990) also estimates the fraction who served in the military using data from the Survey of Income and Program Participation (SIPP). These SIPP data produce considerably higher estimates of the fraction who served in the military than the administrative data do. These estimates are probably especially high because (a) the fractions are 3-year moving averages across cohorts (so that the 1951 average is calculated from the 1950–1952 cohorts), (b) when self-reporting veteran status, some men may have misrepresented service in the reserves as active duty service, and (c) the SIPP measure service over the lifetime, whereas the administrative data used in this study only count entry into the military in the year of the lottery or the year afterward. For

number of American-born men surviving infancy by cohort, estimated as male births minus male infant deaths by cohort (1.86 million and 1.91 million for the 1951 and 1952 cohorts), obtained from the Vital Statistics of the United States (VSUS; United States Centers for Disease Control and Prevention, 1951–1952) and used in Rohlf's (2010). The fraction within each eligibility category is assumed to be proportional to the number of birthdays in that eligibility group. Hence, the numbers of draft eligible men from the 1951 and 1952 cohorts are estimated as $125/365$ and $95/366$ times the VSUS population totals, and the numbers of ineligible men from those cohorts are estimated as $(365-125)/365$ and $(366-95)/366$ times those same totals.

One limitation with the conscription and enlistment data used by Angrist (1991) and obtained from the Defense Manpower Data Center (DMDC) is that a small fraction of volunteers opted for the 2-year contract given to conscripts. As Angrist notes, these volunteers were classified in the DMDC system as conscripts. For this reason, 6.3% of men from the 1951 cohort who were not eligible for the draft and entered the military in the lottery year or the year afterward were classified as draftees in the DMDC data. For the 1952 cohort, this rate is 3.0%. To estimate the true numbers who were drafted and enlisted, it is supposed here that the fraction of voluntary recruits who opted for the draftee terms of service and were incorrectly classified is the same between draft eligible and ineligible men. The enlistment counts are then scaled up by 6.3% and 3.0% for the 1951 and 1952 cohorts, and those numbers of enlistees are subtracted from the conscription totals, both for draft eligible and ineligible men.

After making this adjustment, the fraction drafted among draft eligible men who did not enlist is shown in Figure 3A. The black columns show the fractions using CWHS population data, and the gray columns use the VSUS data. These fractions are shown separately for the 1951 and 1952 cohorts. For the 1951 cohort, this fraction is 0.069 using the CWHS data and 0.083 using the VSUS. For the 1952 cohort, this fraction is 0.067 using the CWHS data and 0.084 using the VSUS.

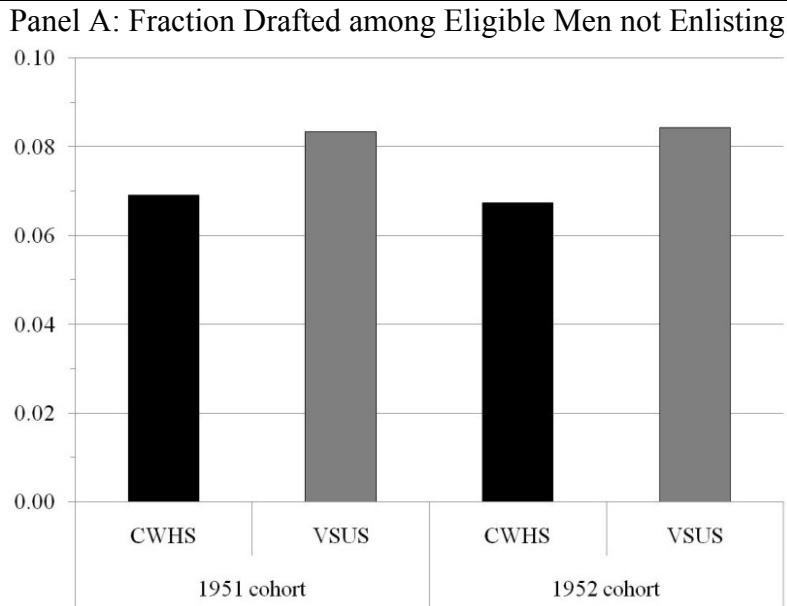
these reasons, and because the SIPP does not differentiate between conscription and voluntary enlistment, the SIPP data are not used in this study.

3.3 THE DRAFT LOTTERY AND VOLUNTARY ENLISTMENT

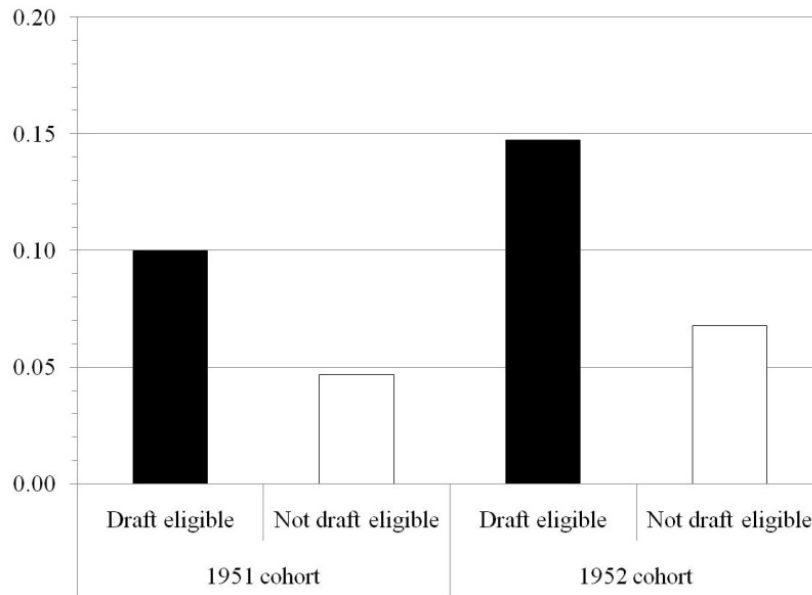
Figure 3B and C show the effect of lottery-induced draft eligibility on the fraction enlisting. As in Figure 3A, only enlistments in the lottery year and the year after the lottery are counted in the totals. Figure 3B shows results using the population size from the CWHS in the denominator; Figure 3C uses the VSUS data for population size. In both panels, the black columns show the fraction enlisting among draft eligible men, and the white columns show the fraction enlisting among ineligible men. Estimates for the 1951 birth cohort appear on the left, and estimates for the 1952 cohort appear on the right.

As in Angrist (1991), there appear to be large positive effects of draft eligibility on enlisting behavior in both cohorts, and in all four cases, the fraction enlisting is more than twice as high for draft eligible as for ineligible men. For the 1951 cohort, eligibility increased the fraction enlisting by five to six percentage points; for the 1952 cohort, the effect of eligibility on enlistment is higher, at 8.0 percentage points using the CWHS data and 9.5 percentage points using the VSUS.

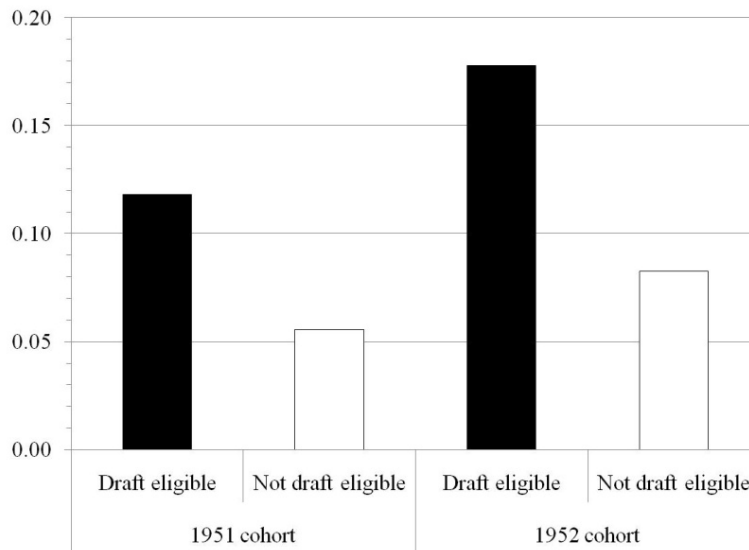
Figure 3: Fraction Drafted and Fraction Enlisting, Adjusted for Misclassification
 Panel (A): Fraction drafted among eligible men not enlisting
 Panel (B:) Fraction enlisting, CWHS population data
 Panel (C:) Fraction enlisting, VSUS population data



Panel B: Fraction Enlisting, CWHS Population Data



Panel C: Fraction Enlisting, VSUS Population Data



Notes to Figure 3: Numbers drafted and voluntarily enlisting by eligibility category taken from Angrist (1991) and only include new conscriptions and enlistments in the lottery year and the year afterward. The size of the draft eligible population is estimated using two alternative data sources, the CWHS (as in Angrist, 1990) and the VSUS (as in Rohlf, 2010a), as described in the text. In all three panels, the numbers who were drafted and who voluntarily enlisted are corrected for the 6.3% and 3.0% of enlisted men in the 1951 and 1952 cohorts who opted for the draftee terms of service and were incorrectly classified as draftees in the data (hence, enlistments are scaled upward and that fraction of enlistments is subtracted from the numbers of draftees).

3.4 ENLISTMENT BONUSES

The effect of enlistment bonuses on recruiting is taken from an experiment conducted by Polich, Dertouzos, and Press (1986) for Rand Corporation in 1983. Through the experiment, the authors randomly assigned recruiters in different geographic areas of the United States to offer different enlistment bonus packages. These bonus packages were made available to “high quality” recruits (i.e., men with high school diplomas and above-median standardized test scores) who agreed to specialize in high need areas. The experiment consisted of a control group that provided the same set of bonuses as was available in previous years, a treatment group that was offered an additional \$8,616 bonus (in 2009 dollars) for signing up for a 4-year contract, and a second treatment group that was offered the \$8,616 bonus for a 4-year contract and a \$6,462 bonus for a 3-year contract. Recruiters in the treatment groups were also given advertising money to promote the bonuses.

Supposing that the marginal recruit would opt for the longer 4-year commitment offered by the service branches other than the Army, the slope of the supply curve is the effect of a bonus for a 4-year contract, which is estimated as the log difference in high quality, high need enlistments between the control group and the first treatment group. This effect of the \$8,616 bonus on enlistment is 0.46, or 5.3×10^{-5} log points per dollar. Supposing that the marginal recruit would opt for the shorter 3-year commitment available from the Army, the slope of the supply curve is the effect of a bonus for a 3-year contract, which is estimated as the log difference in high quality, high need enlistments between the second and the first treatment groups. This effect of the \$6,462 bonus on enlistment is 0.57 log points, or 8.8×10^{-5} log points per dollar. Both of these differences were computed from the figures provided in Polich, Dertouzos, and Press (1986).⁷ For the calculations in this study, a weighted average of $0.40 \times 8.8 \times 10^{-5} + 0.60 \times 5.3 \times 10^{-5} = 6.7 \times 10^{-5}$ is used, where the weight of 0.40 is the ratio of new Army enlistments to new enlistments to the military overall.

One key concern in using the estimated effects in the enlistment bonus experiment is the extent to which the effect of bonuses on military labor supply among high quality recruits in 1983 generalizes to the marginal recruit during the Vietnam Era. No experimental data are available for the period being studied; however, non-experimental estimates of the price elasticity of military labor supply are available from a variety of contexts. These studies generally examine time series and geographic variation in civilian earnings and express the military supply effects in units of log change in enlistment per log change in military-

⁷ Both differences also appear in Rohlfs (2011a) in a hedonic estimation of the marginal recruit’s valuation of educational benefits.

civilian relative earnings. It is not straightforward to compare the units of these estimated effects of a possibly temporary change in local average wages to those of the experimentally estimated effects of a lump sum bonus offered to an individual; however, it is informative to see how these estimated elasticities vary across different studies.

Literature reviews in Asch and Heaton (2008) and Warner, Simon, and Payne (2001), mostly of studies using data from the 1980s and 1990s, find elasticities around one, as does Brown (1985) using a state-by-quarter panel from 1975 to 1982, shortly before the time of the enlistment bonus experiment.⁸ Estimates from the Vietnam Era also vary widely and center around one; however, it is generally agreed upon that, due to the rise in military pay that coincided with the end of the draft, the wage elasticity of enlistment was somewhat lower just before the end of the draft than it was in the years afterward.⁹ Researchers are also in agreement that the wage elasticity of enlistment is especially low for high earning populations such as high quality recruits (cf., Altman, 1969, p. 45; Ash, Udis, and McNown, 1983). These two differences between high quality recruits in 1983 and the marginal recruit in 1970–1972 have countervailing effects, so that it is not clear which population would have a higher elasticity. Additionally, both differences arise because the supply of recruits is a less concave function of relative wages than the constant elasticity formulation predicts, and the linear formulation used in the current study should help to address that concern.

⁸ While varying widely across studies, the 16 estimates shown in Table 1 of Asch and Heaton average to 1.07; the 45 estimates in Table B.2 of Warner, Simon, and Payne average to 0.941. The 12 estimates presented in Table 1 of Brown's study produce an average elasticity of 1.06.

⁹ The United States President's Commission on an All Volunteer Force (1970) surveys a variety of studies and finds estimates that generally exceed one and center around 1.25. Ash, Udis, and McNown (1983) and Brown (1985), in surveying the literature, find estimates that center slightly below one. Ash, Udis, and McNown note that the differences in these estimates across studies are generally attributable to differences in model specification. The functional forms that many authors find best fit the data have elasticities that increase with military pay and the resulting higher numbers of "true volunteers" who would enlist in the absence of the draft (cf., Oi, 1967; Altman, 1969; Fisher, 1969; Cooper, 1977). Using data on enlistments of high school graduates from just before and after the end of the draft, Cooper finds elasticities of 0.95 and 0.97 at the pay levels of the draft and 1.19 to 1.23 at the 30% higher first-term pay levels of the all-volunteer force (Cooper, 1977, p. 168).

Table 2 Voluntary enlistment and the statistical cost of the Vietnam draft

| | 1 | 2 | 3 | 4 |
|---|-----------------------------------|-------------------------|-----------------------------------|-------------------------|
| | Cohort size measured from CWHS | | Cohort size measured from VSUS | |
| | 1951 Cohort | 1952 Cohort | 1951 Cohort | 1952 Cohort |
| 1. Effect of eligibility on conscription risk | 0.069 (0.003)** | 0.067 (0.004)** | 0.083 (0.000)** | 0.084 (0.000)** |
| 2. Effect of eligibility on ln(enlistment rate) | 0.757 (0.051)** | 0.776 (0.042)** | 0.756 (0.005)** | 0.766 (0.004)** |
| 3. Estimated statistical cost of the draft | \$124,983 (27,989)** | \$131,273 (29,142)** | \$103,394 (21,614)** | \$103,574 (21,649)** |
| Hinkley 95% CI | [88,554, 217,526]** | [90,588, 228,828]** | [73,333, 175,118]** | [73,467, 175,425]** |

Notes to Table 2: Row 1 shows the effect of lottery-induced draft eligibility on conscription risk among men who did not voluntarily enlist, using data from Angrist (1990, 1991). These means are the same as the bars in Figure 3A. Row 2 shows the effect of draft eligibility on the log of the probability of enlistment. These figures are the log difference between the black and white bars shown in Figure 3B and C. To estimate the statistical cost of the draft, row 3 divides the log difference in enlistments in row 2 by 6.7×10^{-5} , the effect of an enlistment bonus on log enlistments, and then divides that ratio by the change in conscription risk shown in row 1. Delta method standard errors and Hinkley 95% confidence intervals are calculated in the same way as in Table 1. Additional details in the text.

3.5 THE COST OF THE DRAFT

Table 2 shows the main results from the enlistment portion of this study. In columns 1 and 2, population size is measured from the CWHS; in columns 3 and 4, it is measured from the VSUS. Columns 1 and 3 show estimates for the 1951 birth cohort, and columns 2 and 4 show estimates for the 1952 cohort. The first row shows the probability of being drafted among draft eligible men who did not

enlist. These effects are the same as appear in Figure 3A. The standard errors are calculated using the formula of $p * (1 - p)/n$ for the variance of a sample mean estimated from n draws from a binomial random variable with success probability p . These standard errors are smaller for the VSUS than for the CWHS because the CWHS uses a 1% sample, and the VSUS is a population total. Row 2 shows the effect of draft eligibility on the log of the probability of enlisting. This effect is the same as the log of the difference in the rates for draft eligible and ineligible men that is shown in Figure 3B and C. The standard error for this log difference is computed by applying the formula for the variance of a binomial random variable to the enlistment probability and using the delta method to compute the standard error for the log of this probability. Row 3 presents the estimated statistical cost of the draft, which is computed as row 2 divided by row 1, all divided by the 6.7×10^{-5} per dollar effect of enlistment bonuses. As in Table 1, the numbers in parentheses in row 3 are standard errors computed using the delta method; the Hinkley 95% CI appears in brackets below. Both the standard error and the CI row 3 take into account the imprecision in the estimates of the draft avoidance effect, the enlistment bonus effect, and the risk of being drafted, but do not take into account the imprecision associated with applying the slope of military labor supply from the 1983 experiment to draft-age men during the Vietnam Era.

The statistical cost of the draft estimates in Table 2 range from \$125,000 to \$131,000 using the CWHS data and from \$103,000 to \$104,000 using the VSUS data. The primary difference between the two sets of estimates is that size of the eligible population is smaller using the VSUS data, leading to a larger probability of conscription among draft eligible men. All four estimated costs of the draft are statistically significant using both the delta method and the Hinkley CIs.

4. AN UPPER BOUND ON THE VSL

This section relates the cost of being drafted to the risk of death. Fatality risk was one of the many costs of serving in the military. In addition to the risk of death, being drafted entailed career interruption and lost earnings, deleterious health effects including the possibility of injury, relatively harsh living conditions, a loss of freedom, and low social status as a low-ranking member in the hierarchy of the military. Here, we consider the extreme case in which the only cost associated with mandatory service is the risk of death. Under the assumption that, for the marginal college student or military recruit, these other costs outweigh the benefits of the positive feeling associated with patriotic service and fulfilling an obligation, the military training and job experience, and the novelty of traveling to

Southeast Asia, the estimates presented in this section constitute upper bounds on the VSL.¹⁰

These upper bounds differ from conventional VSL estimates in two important ways. First, traditional VSL estimates involve direct comparisons between dollars and risk, as in the decision to work in a risky job or to purchase an automobile air bag. In the current setting, a costly action – rather than an explicit dollar cost – is compared to the risk of death. Second, traditional VSL estimates are based upon Ordinary Least Squares (OLS) regressions of price or wage on the level of risk associated with an automobile or job (cf., Viscusi, 1993). As previous researchers have noted, this level of risk is correlated with many unobserved factors about cars, drivers, workers, and jobs, and these statistical associations are probably biased estimates of the true causal relationships (cf., Ashenfelter and Greenstone, 2004; Ashenfelter, 2006). A recent study by Rohlfs (2011b) uses a quasi-experiment based upon automobile air bag regulations to obtain causal estimates of the VSL; in the same vein, the current study aims to construct causal estimates of the VSL through exogenous shifts in conscription risk.

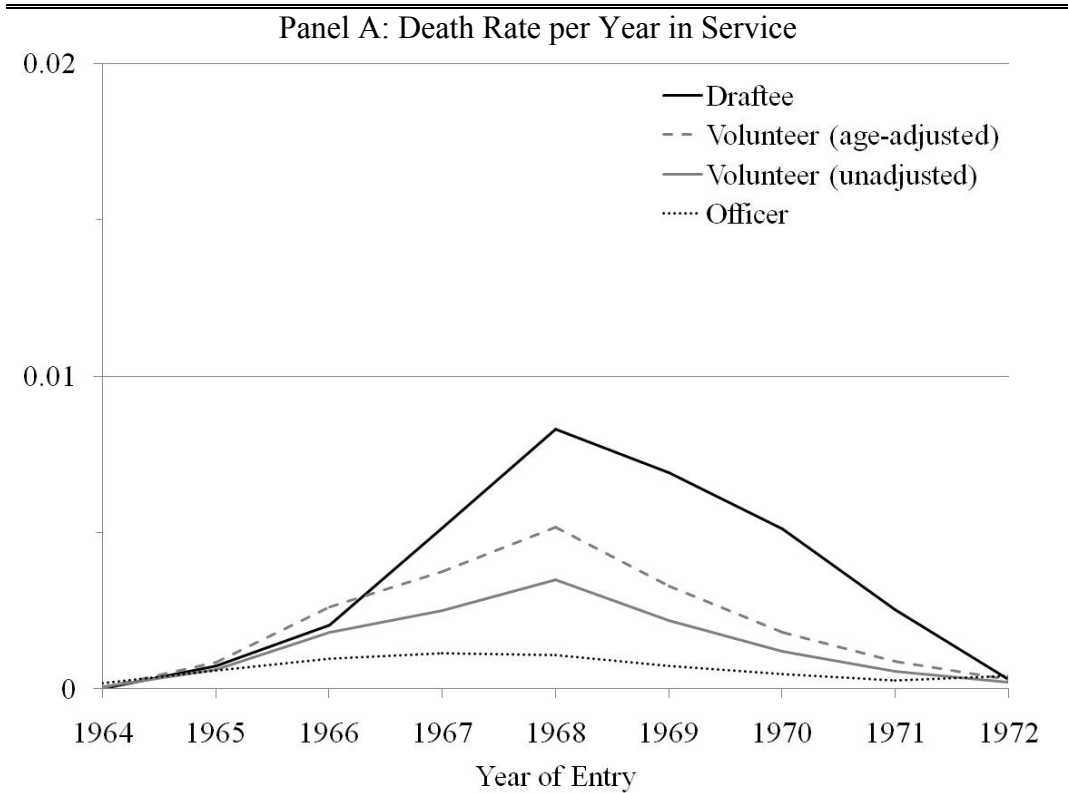
For the approach used in this study to produce valid estimates of the VSL, it is essential that draft-age men from the Vietnam Era had reasonably accurate estimates of the risk of death associated with serving in the military. Little data exist from that period on the beliefs that young men had about the risks of serving. Nevertheless, in a 1968 survey, 57% of Americans reported paying a “good deal” of attention to the Vietnam issue, 69% mentioned it as one of the top three problems facing the country, and 78% listed it as “very important” in their votes for president (Lau, Brown, and Sears, 1978). The amount of attention paid to the Vietnam War and to deaths there was probably higher for draft-age men than for the average American. The war and military deaths were highly publicized in newspapers and on television. Given the large amounts of available information and attention paid to the problem, it is reasonable to expect that draft eligible men had accurate impressions of what service in Vietnam was like.

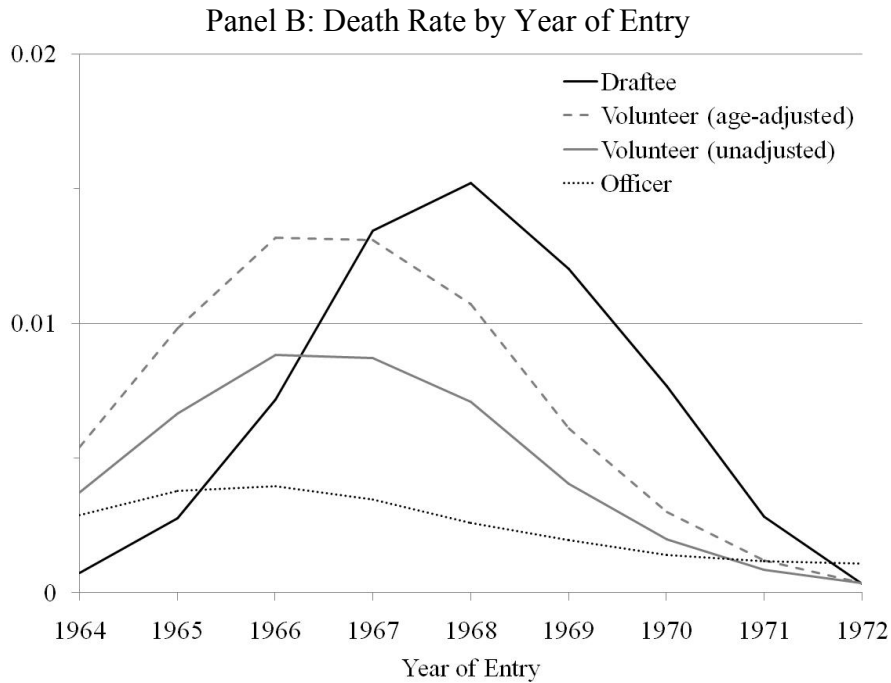
The likelihood of death among draftees, volunteers (including enlisted and officer ranks), and officers is plotted for different years in Figure 4. These rates are computed from microdata on U.S. military deaths in Vietnam combined with published statistics on the numbers of each type of soldier serving in the military. The rates in Figure 4A show the number of deaths divided by the total number of active duty personnel, annually by personnel type from 1964 to 1975. The age-

¹⁰ The phrase “upper bound” is somewhat imprecise, because a true upper bound would take into account the imprecision in the estimates. Owing to this imprecision, the true VSL for young men could potentially be larger than what is estimated here (and might be as high as the 95% upper bound in the CI). The use of the phrase is meant to convey the point that the VSL estimate presented here is biased upwards, and that this bias could be large.

adjusted death rates for enlisted men measure the death rates for enlisted men of draft age (25 years or under). Some of the numbers of active duty personnel are imputed using assumptions described in the footnotes to the figure. Figure 4B shows estimated death rates by year of entry into the military. These rates are computed by assuming that a draftee served 2 years (the year of entry and the year afterward), a volunteer served 3.5 years, beginning at the year of entry, and an officer served 4 years. In each case, the soldier is assumed to experience the average risk of death for his personnel type for each year in service.

Figure 4: Estimated Risk of Death in Vietnam for Draftees, Volunteers, and Officers





Notes to Figure 4: Numbers of draftees by year are administrative data provided by the Selective Service System (D. Card and T. Lemieux, unpublished data); numbers of draftees in service in a given year is computed by assuming that each draftee served 2 years. The volunteer category includes both enlisted men and officers. Numbers of volunteers and officers by year are taken from the United States Census Bureau (1966, 1976). Deaths by year, type of entry (draftee/volunteer), rank (enlisted or officer), and age are computed from microdata on U.S. military deaths in Vietnam (United States Department of Defense, 2005). The fraction shown in (A) of military personnel who died in Vietnam each year is computed separately by enlistment category by dividing these death totals by the number in service that year. The probability of death by entry year shown in (B) is then computed separately by supposing that a draftee entering in a given year experienced the draftee fatality rate for 2 years, a volunteer entering in a given year experienced the volunteer fatality rate for 3.5 years, and an officer experienced the officer fatality rate for 4 years. The age-adjusted risk of death for volunteers 25 years or under is computed using estimates by Oi (1967, Table 2), obtained from Census data, that 46.7% of military personnel were above 25 years in 1960. The number of military personnel above 25 years is assumed to have remained constant over this period, so that all of the changes in the military population were driven by changes in the numbers of draft-age men in service. Deaths in Vietnam are then calculated separately by year for volunteers 25 years and under and divided by the estimated number of volunteers 25 years and under who were in service each year. Additional details in the text.

As the estimates from Figure 4A show, the annual rate of death tended to be considerably higher for draftees than for volunteers. With the exception of the death rate for officers, which was essentially flat, all of these rates peaked in 1968. When the number of years in service is taken into account in Figure 4B, we observe

earlier peaks for enlisted men, both with and without the age adjustment, so that the enlisted men with the highest fatality rates were those entering service in 1966 and 1967. The death rate peaks later, in 1968, for draftees, due to their shorter term of service. Consequently, the death rate was higher for volunteers than for draftees among those entering between 1964 and 1966. In both Figure 4A and B, but especially in Figure 4B, restricting the sample of volunteers to draft-age men shrinks the gap between draftees and volunteers considerably but does not eliminate it.

Table 3 presents estimated upper bounds on the VSL. Part (A) shows estimates using the college attendance margin, and part (B) shows estimates for the voluntary enlistment margin. Within each panel, the column headings are the same as in Tables 1 and 2, and the estimates are shown for the same specifications and datasets as in Tables 1 and 2. The two rows in part (A) show estimated upper bounds on the VSL using two alternative models of the perceived likelihood of death if drafted. In both cases, fatality risk is estimated as an average of the risk for draftees entering in certain years, as in the black curve from Figure 4B. Row 1 supposes that the marginal college student could correctly predict the fatality risk for the peak years of the war and uses the fatality risk averaged across years of entry from 1967 to 1969. Row 2 supposes that the marginal college student could not predict the peak years of the war and uses the risk averaged across the years of entry from 1964 to 1972, the complete range of the war.

The two rows in part (B) show estimated upper bounds on the VSL from the enlistment margin using two alternative models of the difference in the likelihood of death between draftees and volunteers. This difference is shown in Figure 4B as the difference between the solid black curve and the dashed gray curve. Row 3 shows this difference using the unadjusted fatality rate for volunteers; this difference appears in Figure 4B as the difference between the solid black curve and the solid gray curve. Row 4 shows this difference using the age-adjusted rate for volunteers who are aged 25 years or younger. In both rows, the marginal recruit is assumed to use the fatality rate for two recent cohorts of draftees, those who entered in 1969 and 1970. Owing to the end of the war, the actual risk of death for draftees entering in 1971 and 1972 was much lower than the 1969–1970 rates (and the implied upper bound on the VSL would be much higher). The 1969–1970 years are used because it is these deaths that young men would have learned about in press accounts as they made their enlistment decisions.¹¹

¹¹ Additionally, the estimates from Table 2 produced very similar estimates of the draft avoidance behavior for the 1951 and 1952 cohorts, despite the sharp decline in fatality risk (down to practically zero) for the 1952 cohort. This finding is consistent with the view that young men could not foresee high frequency changes in the risk of death from service.

Table 3 Estimated upper bounds on the value of a statistical life

| | 1 | 2 | 3 | 4 |
|---|--|--|--|--|
| (A) College attendance as a margin of response | | | | |
| | Fraction enrolled at ages 20–21 years | | Fraction completed some college by 1980 | |
| Risk measured as... | Linear interpolation | Conscription risk as regressor | Linear interpolation | Conscription risk as regressor |
| 1. Rate of death among draftees in 1967–1969 [95% CI in millions] | \$4.68 million (3.48 million) [–15.1, 35.0] | \$2.97 million (1.91 million) [–10.3, 21.6]* | \$0.61 million (0.54 million) [–1.93, 4.81] | \$0.29 million (0.44 million) [–1.50, 3.19]* |
| 2. Rate of death among draftees in 1964–1972 [95% CI in millions] | \$9.08 million (6.74 million) [–29.2, 67.9] | \$5.76 million (3.70 million) [–20.0, 41.9]* | \$1.18 million (1.04 million) [–3.74, 9.33] | \$0.85 million (0.56 million) [–2.91, 6.18]* |
| (B) Voluntary enlistment as a margin of response | | | | |
| | Cohort size measured from CWHS | | Cohort size measured from VSUS | |
| | 1951 Cohort | 1952 Cohort | 1951 Cohort | 1952 Cohort |
| 3. Unadjusted difference, 1969–1970 [95% CI in millions] | \$8.80 million (2.52 million)** [5.55, 19.5]** | \$9.24 million (2.64 million)** [5.87, 20.5]** | \$7.28 million (2.00 million)** [4.72, 15.8]** | \$7.29 million (2.01 million)** [4.73, 15.8]** |
| 4. Age-adjusted difference, 1969–1970 [95% CI in millions] | \$11.3 million (3.64 million)** [6.82, 29.3]** | \$11.9 million (3.80 million)** [7.21, 30.9]** | \$9.36 million (2.91 million)** [5.80, 23.8]** | \$9.37 million (2.92 million)** [5.81, 23.8]** |

Notes to Table 3: These upper bounds on the VSL are computed by dividing the estimates of the statistical cost of the draft (in Tables 1 and 2) by the effect of being drafted on fatality risk. For the college attendance margin in part (A), fatality risk is measured as the risk among draftees by year of entry (as in Figure 4), averaged across the years of entry listed. For the voluntary enlistment margin, fatality risk is measured as the difference between the risk among draftees and the risk among enlisted men, averaged across years of entry. Standard errors and CIs are calculated as in Tables 1 and 2. Additional details in the text.

It is useful to have benchmark VSL estimates to understand how the upper bounds in Table 3 compare to VSL estimates from the literature. Relatively few estimates of the VSL are available from the Vietnam Era; however, Costa and Kahn (2004) provide labor market OLS estimates of the VSL for 18- to 30-year-olds using data from the U.S. Census from multiple decades including 1970. For later decades, their estimates are similar to prevailing estimates (e.g., Viscusi, 1993). For a young man in 1970, Costa and Kahn estimate VSLs ranging in 2009 dollars from \$4.58 million to \$8.10 million; the estimates those authors cite from previous studies vary widely but are generally consistent with this range.

As with the statistical cost of the draft, using college attendance as a margin of response produces somewhat small and nonsignificant implied upper bounds on the VSL. When averaged across the CPS and Census estimates as in Table 2, using the higher 1967–1969 measure of fatality risk, the implied upper bound on the VSL from the college attendance margin is \$2.64 million using the linear interpolation approach and \$1.63 million using the regression approach. When the lower 1964–1972 fatality risk measure is used, the implied upper bound on the VSL when averaged across the CPS and Census estimates is \$5.13 million using the linear interpolation and \$3.30 million using the regression. These upper bound VSL estimates are somewhat lower than Costa and Kahn's VSL estimates for a young man from 1970, suggesting that fatality risk accounted for essentially all of the cost of being drafted and that the other costs associated with conscription were negligible or possibly negative. However, as with the cost of the draft, these estimates are consistent with the interpretation that young men were credit constrained and the upfront cost of college reduced their valuations of both the fatality risk and the other costs of service.

Examining voluntary enlistment as the margin of response, so that the estimates are not affected by credit constraints, produces considerably higher upper bounds on the VSL than are obtained from the college attendance margin. Using the higher 1969–1970 measure of fatality risk and not adjusting for age in row 3, the estimated upper bound on the VSL ranges from \$7.3 million to \$9.2 million. If the age adjustment is applied, so that the fatality risk for enlisted men is calculated from enlisted men who were 25 years or under, the implied upper bound on the VSL is higher, ranging across samples from \$9.4 million to \$11.9 million.

If the marginal recruit was able to predict the end of the war and the large decline in risk among draftees and enlisted men, then it may be appropriate to use the 1971–1972 measure of fatality risk, as in rows 5 and 6. Using this lower risk measure produces considerably higher (and consequently less informative) upper bounds on the VSL. Without adjusting for age, this estimated upper bound ranges across samples from \$47.3 million to \$60.0 million. After adjusting for age, the upper bound on the VSL ranges from \$70.0 million to \$88.9 million.

5. DISCUSSION AND CONCLUSION

This study presents new dollar-denominated estimates of the utility cost that the draft imposed on young men in the Vietnam Era. Estimates of this cost of the draft are key ingredients that one could use to evaluate the cost-effectiveness of conscription in other countries and the “stop-loss” policy in the U.S. military. This study also applies revealed preference estimation to an important set of behaviors that have not been previously examined from this perspective.

Two forms of draft avoidance behavior in the Vietnam Era are examined to produce two sets of estimates of the economic cost that the draft imposed on young American men. The first draft avoidance behavior is college attendance, which reduced one’s risk of being drafted through the institution of college draft deferments. The second draft avoidance behavior examined in this study is voluntary enlistment. During the Vietnam Era, men who were concerned about the risk of being drafted could voluntarily enlist in the military to reduce uncertainty (so that they could better plan their careers) and to obtain safer postings.

The estimates from the market for higher education indicate that, in 2009 dollars, the statistical cost of the draft for the marginal college student was approximately \$30,000. Considerably higher estimates are obtained from the market for voluntary enlistment, indicating that statistical cost of the draft for the marginal military recruit was approximately \$115,000. The explanation proposed in this paper for this difference in the estimated valuations is that young men were credit constrained, and the demand for a combined package of college and draft avoidance was lower than the demand for the two goods individually. Of these two sets of estimates, those based upon the voluntary enlistment margin are probably more appropriate for measuring the social cost of the draft, as the optimal social planner is not credit constrained.

Relative to the voluntary enlistment estimates, those based on the college attendance margin more closely approximate demand for draft avoidance (willingness to exchange dollars for draft avoidance) as it would typically be modeled in a revealed preference setting. Such estimates could be used to forecast the potential demand for a draft buyout, which, like college, would require an upfront cost. Additionally, the extent to which credit constraints influence schooling decisions is a topic of considerable debate (cf., Becker, 1975; Card, 1995, 1999; Carneiro and Heckman, 2002; Cameron and Taber, 2004; Kane, 2004). In addition to providing information on the cost of the draft, the estimates from the college attendance margin provide evidence to support the view that credit constraints played an important role in young men’s college attendance decisions in the 1960s.

Supposing that conscription would be costly in the absence of fatality risk, so that the overall value of the earnings, health, and utility effects of conscription was negative for the marginal college student or military recruit, the estimates of the statistical cost of the draft produced in this study can be used to measure an upper bound on the VSL. For the college attendance margin, this upper bound ranges across the preferred specifications from \$1.6 million to \$5.1 million, which is on the low end of VSL estimates for young men in that era. This finding is consistent with the view that the estimates using college attendance as a margin of response are downward biased due to credit constraints. For the voluntary enlistment margin, this upper bound ranges across specifications from \$7.3 million to \$11.9 million. These figures are on the upper range of Costa and Kahn's (2004) estimates of the VSL in that era. A large portion of the demand for draft avoidance was probably due to the loss of freedom and the substantial health and economic costs of conscription. Taking these other costs into account, the VSL implied by the voluntary enlistment margin is probably at or below the lower end of Costa and Kahn's range of VSL estimates.

The Vietnam War involved the conscription of 1.83 million men (United States Selective Service System, 1965–1973). If we suppose that the cost to the average draftee of being drafted was approximately \$115,000, the estimated cost for the marginal military recruit, then the loss in surplus caused by the draft was \$210 billion. By comparison, Daggett (2010) estimates that U.S. government expenditures attributable to the Vietnam War totaled \$720 billion in 2009 dollars. Hence, the estimates from this study indicate that the economic cost of the Vietnam Draft was approximately 29% as large as the dollar cost of the war effort.

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