Paying to Avoid Recession: Using Reenlistment to Estimate the Cost of Unemployment

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Appendix Tables and Figures

A1 Earnings Estimation: Reenlistment Selection-correction

We only observe earnings outcomes for those servicemembers who exit the military, creating a potential selection problem. Specifically, as reenlistment rises and fewer servicemembers exit the military during recessions, the characteristics of veterans who exit the military may change. Our earnings regressions will produce unbiased estimates under the assumption that we can capture these changes with our covariates. This selection-on-observables assumption may be plausible, considering our rich set of controls; however, we go one step further and make use of variation in reenlistment rates at the military occupation level to account directly for selection on unobservable characteristics. This strategy follows Card and Rothstein (2007), which uses the share of students taking the SAT in a high school to correct for selection in the outcome distribution; see the Appendix to that paper for a derivation.

We define $O(o)$ to be the first digit of the first observed primary occupation code (the military MOS code) and use movements in occupation-level reenlistment to construct a control function approach to Heckman's two-step selection correction Heckman (1979).

In the first stage, we estimate a linear probability model of reenlistment, adding fixed effects for the interaction of $O(o)$ and t to the model. We then include the predicted probability of exit from the first stage in the main earnings equation, using a semi-parametric transformation suggested by Newey (2009):

$$\mathbb{E}[y_{iostk}|D_{iost}=0] = w_{O(o)k} + a_{sk} + b_{j(i)tk} + d_k UR_{st} + c\mathbf{X}_{istk} + \Lambda_k(\widehat{p_{iost}}) \tag{1}$$

$$\widehat{p_{iost}} = z_{O(o)t} + f_s + g_{j(i)t} + hUR_{st} + X_{ist}. \tag{2}$$

Here, $\widehat{p_{iost}} \equiv \mathbb{P}[D_{iost}=0]$ is the predicted probability of exit,

$$\Lambda_k(\widehat{p_{iost}}) \equiv \sum_{l=1}^{4} \lambda_l^k (2\Phi(\widehat{p_{iost}})-1)^l$$

is the semi-parametric transformation and $\Phi(\bullet)$ is the cumulative distribution function.[1] X_{ist} and X_{istk} contain the same variables as the above specifications. The selection-correction is identified by the exclusion of the $z_{O(o)t}$ cross-effect between military occupation and the period of discharge from the earnings equation, where we include only the independent effects of occupation o and year of eligibility t in Equation 1. The occupation-by-year effects have strong predictive power; we reject the null of $z_{1t} = ... = z_{O(o)t} = 0$ at $p < 0.001$. These $O(o)t$ effects reflect the average reenlistment rate in occupation o at time t, conditional on changes in the composition

[1] A linear probability model is used to estimate $\widehat{p_{iost}}$. We set extreme values to 1 and 99%; this is less than 0.0003% of observations.

of servicemembers with contracts ending in the period. Identification of the selection function with this exclusion restriction requires changes in state-level economic conditions and occupation-level demand to be uncorrelated.[2]

In practice, there is limited scope for selection to influence our estimates. Since close to half of servicemembers will reenlist, the percentage change in reenlistment is relatively small, even for large movements in the unemployment rate. To take an example, a one p.p. innovation in the unemployment rate (an average shock) and a one p.p. response per point of unemployment (approximately our estimate of β), will raise the average exit rate from 60 to 61%. With this relatively small shift, the characteristics of servicemembers on the margin of reenlisting (compliers) would have to be dramatically different from the average servicemember who always exits (always takers) in order for this (average) shift in reenlistment behavior to exert a meaningful change in the average characteristics, and particularly, expected earnings.

Appendix Figure A4 allows us to assess the magnitude adjustment to our estimates due to the selection-correction. For reference, we include the effect on civilian earnings. The figure displays the earnings losses that result from higher unemployment at discharge

[2] As in the main text, the analysis was conducted in two stages, with the individual-level covariates, now containing the (transformed) probabilities of exit, included in the estimation at SSA, with cell-level residuals computed at the state by quarter of exit level. In principle, we could have used the bonuses from the previous section; however, the SSA earnings analysis was completed before the bonus data was accessed.

for the full sample of young veterans with initial contracts ending in 1993-2009. We plot three sets of estimates: total earnings losses, total losses corrected for selection out of the military, and civilian earnings losses. Total earnings drop nearly 4% in the year following separation for each percentage point increase in the longer sample. Table A3 reports average earnings, the d_k regression coefficients, their standard errors and the implied value of horizon k losses in 2010 dollars, where we use the CPI to deflate nominal income. Negative effects on total earnings remain statistically significant through the 8th year, and point estimates return to zero between the ninth and tenth years. Summing the stream of earnings losses, young veterans lose $4875 over the next 10 years for each point increase in home-state unemployment; with a discount factor of 0.9, the present discounted value of the earnings losses is $3450. We plot effects without the selection correction for reenlistment, and as expected, correcting for selection increases the size of the effect, but the increase is economically negligible.

A1.1 Estimation: Average Value of Bonus Payment

In Section 2 and Section 4.3, we discuss the estimation of mitigating mechanisms as the difference in the present discounted value of earnings losses and bonus payments. In order to translate our estimates from the model parameters in the reenlistment estimation to the dollar value of losses, we require an estimate of Δy^{bonus}, the average PDV of a unit movement in our bonus variable. We use the rules of bonus payments to calibrate this for different values of the discount rate, taking into account the size and timing of the intital bonus payment, and the possibility of persistence in the bonus between the end of the first contract and end of the second contract.

In terms of size, the bonus pays the multiple of the length of reenlistment, the bonus multiplier and the base pay of servicemembers. To account for timing, we decompose the value of a bonus payment into three payment streams. First, half of the bonus is paid as a lump-sum on signing. Second, the remaining half of the bonus is spread over the course of the next contract. Third, the bonus offered at the end of the first contract is predictive of the bonus offered at the end of the second contact (this effect is small, as discussed below). To address this persistence of offered bonuses, we calibrate our estimates using the averages of second-term contract length and reenlistment. Second-term contracts average a four year commitment (average second contract length is 49 months between 1993 and 2004). A regression of the bonus at the end of the second contract on the bonus at the end of the first contract reveals that 25% of the bonus persists; at this next reenlistment point, 55% of servicemembers will continue in the military.[3]

Pulling together these components, we estimate the dollar value

[3] Appendix Figure A5 displays the histogram of second contract lengths. To estimate the persistence of the bonus, we use the framework of the reenlistment model section to predict the bonus offered at the conclusion of the second contact with the the bonus offered at the end of the first contract.

of the bonus as:

$$\Delta y^{bonus} = y_t^{mil}(2\% + \frac{1}{2}0.5\% + \rho 0.5\% + \rho^2 0.5\% + \rho^3 0.5\% + \frac{1}{2}\rho^4 0.5\%)$$
$$+ \rho^4(0.55)(0.25)y_{t+4}^{mil}(2\% + \frac{1}{2}0.5\% + \rho 0.5\% + \rho^2 0.5\% + \rho^3 0.5\% + \frac{1}{2}\rho^4 0.5\%). \quad (3)$$

The four units of bonus, reflecting the four additional years of new service commitment (an average of 49 months in the data), are paid as two units in year zero, and half a unit per year served after that. Since the average reenlistment occurs midway through the year, and we use an annual approximation to the payment stream, the servicemember receives 2.25 of the four bonus units in the year of reenlistment, half a unit for the next three years, and a quarter of a unit in the last year of the contract. We set y_t^{mil} = \$32,000, average reenlistment-year base military earnings in 2010 dollars, meaning that each bonus point, B_{iot}, equates to a stream of payments (over the course of a four year contract) of \$640 and \$80 in year zero, \$160 in years one through three and \$80 in year four. After the second contract is concluded, we assume that continuation occurs at the average rate, and those who continue receive bonus offers that reflect the average persistence of bonuses. This implies that 55% will continue at the second contract, with 25% of the unexpected component of the first contract bonus persisting. With a discount factor of 0.9, these values imply that the increase in expected second contract bonus comprises around 10% of the value of a movement in a the first contract bonus. Following the second contract, we assume the servicemember exits. This is consistent with the reduced-form evidence showing that the increase in military income (in response to home-state unemployment) is statistically significant for nine years following the end of the first contract.

A2 Extensions to the Reenlistment Model

A2.1 Independent Variation in Bonuses and State Unemployment Rates

We demonstrate independent variation in the two variables of interest in the reenlistment analysis by reconstructing Figure 3 with the addition of fixed effects to remove the effect of the other series. The results appear in Appendix Figure A2. In the upper panel, we add occupation by year fixed effects to the controls in the state unemployment rate figure. The estimated effect of the unemployment rate changes from 0.87 in Figure 3 to 0.90 in Appendix Figure A2. For the bonus, the inclusion of state by year fixed effects does not change the estimate from 0.33 in the main text.

A2.2 Logit Transformation

The linear probability model used in the reenlistment analysis is likely to perform well in this setting, where the mean probability of reenlistment is far from zero or one for any group of servicemembers. Figure 3 shows that the linear model fits the data. In this section, we demonstrate robustness to a logit functional form. The non-linear logit regression is challenging to estimate with the large number of observations and fixed effects. To facilitate computation, we use Berkson's transformation to a regression on aggregate probabilities, following the discussion in Cameron and Trivedi (2005).

To implement this strategy, we first form groups of state by two-digit occupation by year of eligibility for reenlistment. We impose the $N_{ot} > 15$ to maintain connection to the estimates in the main text while minimizing the number of groups with all successes or all failures at the group level. The group-level probability of reenlistment is transformed to logit, $log(\frac{p}{1-p})$. We then regress this transformed probability on a restricted set of variables: the state, year, and two-digit occupation fixed effects, along with the unemployment rate and bonus measures (none of which vary within cells). The regressions are weighted by the estimated within-cell variance. Note that we are dropping individual-level controls from this regression as in Figure 3; however, our empirical estimates in the main text were largely insensitive to their inclusion.

Appendix Table A4 reports the results transformed into marginal effects that are comparable to the main estimates. As predicted, the point estimates mirror the linear specification, and are particularly close to the grouped regression estimates reported in Panel B of Table 4.

A2.3 Instrumental Variables Estimates

The reenlistment model uses variation in bonuses across occupations to estimate the responsiveness of reenlistment to financial incentives. Variation in the bonus may arise for either demand or supply-related reasons. Ideally, we would like to isolate pure demand-side variation, in which the military changes bonuses in response to (exogenous) shocks to the target-level of enlistment in an career field, holding constant working conditions, civilian labor market opportunities and other aspects of broadly-defined compensation. The 1993-2004 era presents a promising era for this type of variation, as the structure of the military changed a great deal in the years following the Cold War and Gulf War of 1990-1991, and again in the initial stages of the Afghanistan and Iraq Wars (initially predicted to be short engagements). Consistent with the choice of era and sample, we found the estimates to be insensitive to a range of controls for the most likely confounds (such as exposure to deployment and hostile fire) and insensitive to the inclusion of the average characteristics of servicemembers who do reenlist (which would be the case if the military changes reenlistment eligibility standards).

To investigate further the potential endogeneity of bonuses, we estimated an instrumental variables (IV) model based on career field-specific shocks to military labor demand. To isolate labor demand movements, we use the flow of new recruits within a career field. This flow may reflect a superior measure of labor demand within the career field, as it is unconstrained by previous assignments of servicemembers to occupations, and unrelated to reenlistment choices (the outcome in the model). On the other hand, flows of new recruits may be endogenous itself, if the military has problems filling those same occupations which are targeted by the bonus, or places more recruits in occupations with reenlistment shortfalls. Thus, the estimates from this instrumental variables strategy can inform the effect of bonuses, but may introduce other sources of bias.

Specifically, we instrument the bonus with the flow of new recruits to the career field in the reenlistment year t, L_{ot}, conditional on the flow at the time of entry, L_{ot-j}.[4] The first-stage equation in the IV model is:

$$B_{iot} = \psi_o + \sigma_{j(i)t} + \phi L_{ot} + \pi L_{ot-j(i)} + \tau B_{iot-j(i)} + \xi \mathbf{X}_{iot} + \upsilon_{iot}, \quad (4)$$

where ψ_o indicates military career field, $\sigma_{j(i)t}$ indicates year of eligibility for discharge, L_{ot} is the labor demand shock, the size of the entering class in the o career field, $L_{ot-j(i)}$ and b_{iot-j} control for the size of the career field cohort and bonus at entry, and $\mathbf{X}_{iost}$ contains the full set of controls required to maintain consistency with the previous analysis. Our estimates will capture the causal effect of bonus payments on reenlistment under the assumption that the same underlying career-field demand shock drives both the target level of new

[4]To the best of our knowledge, this is the first published paper to employ this strategy. We know of at least one internal analysis at RAND that explored a similar IV strategy using occupation-specific staffing authorizations, but was never released. Results in that study were similar to those reported here.

recruits in the field and the bonus level for re-enlistees, and that the bonus represents the military's sole control variable in reenlistment.[5] As the military desires a stable tenure pyramid and does not permit mid-career entry, the flow of new recruits into career fields should also reflect an increased demand for higher tenured workers of the same type that cannot be met except through reenlistment. New recruits do not appear to substitute for higher tenure servicemembers, as the size of the entry cohort does not predict speed of promotion.

Table A5 contains the results of the IV estimation. The upper panel reports the first-stage results, including Kleibergen-Paap (heteroskedasticity-robust) F-statistics for the significance of the excluded regressor. In each case, the first stage F-statistics narrowly exceeds 10. This suggests the size of the entry cohort in the career field has predictive power over the bonus offer made to those up for reenlistment, but not so much that we can completely exclude a bias from weak instruments. We do not report results for the longer sample (1993-2009), as the first-stage did not approach the rule of thumb F-statistic of 10. This likely reflects recruitment and reenlistment shortfalls that occurred in 2005-2008.

The lower panel applies the IV model to the 1993-2004 sample. Instrumenting the bonus increases the reenlistment response by a factor of approximately 1.5 to 3 compared to the OLS esti-

[5] We also assume that military occupation-specific demand is uncorrelated with local labor market conditions. We test this by regressing the the bonus and the recruitment flow on unemployment rates, finding no significant effect.

mates, implying a corresponding reduction in the value of unemployment. Note that the unemployment rate coefficient (β) remains robust across the OLS and IV specifications: the gap between the OLS and IV estimates does not arise from the avoidance behavior (higher reenlistment in response to arises from the response to pay), but rather from the price of this avoidance behavior (the increase in reenlistment in response to bonus pay). The coefficient in the first column of 0.93 means that a 10% increase in pay results in a 9.3% increase in reenlistment. Moving across the columns, we cannot reject the equality of bonus coefficients in the 1993-2004 sample period, although the estimates are larger when we control for (average) in-service experiences and smaller for those with longer initial contracts. Off a base 38% reenlistment rate, these point estimates imply an elasticity between 2.5 and 3 in the earlier sample period. This elasticity is on the higher end or above what has been found in previous studies of reenlistment.

We find some evidence of bias in the first-stage, and for this reason have placed the results in an Appendix. Specifically, the largest estimates of the bonus response in the IV model come when we include post-enlistment controls. We found that these had no effect on the estimates in the main text; however, they seem to increase the response to the bonus by 30-40% in the IV model. This suggests the flow of new recruits in the occupation is correlated with μ_{ot}, violating the exclusion restriction. Despite this increase, the IV estimates allow us to test and reject the presence of full insurance against movements in end-of-contract unemployment in all specifications. Table 5 allows the reader to replace our estimates of β/γ with these results, if so desired.

A3 Behavioral Responses to Unemployment and Effect on Estimates

In this appendix section, we describe several analyses of behavioral responses to unemployment at the end of the first contract, with the goal of relating these changes to the results presented in the main text. We also discuss measurement issues in the primary variables. These calculations are referenced in Section 5.1.

We require alternative data sources for some of this analysis, as we cannot observe outcomes beyond earnings and college benefit usage in the administrative data. First, we present results on the effect of unemployment at separation, which may not occur in the month the initial contract ends, even for those who do not reenlist. Next, we summarize our analysis of young veterans' migration. Finally, we detail the response of college enrollment and benefits usage. In each case, we find evidence that our estimates represent lower bounds on the true effect of interest. In sum, effects of unemployment in the state of residence at the precise time of separation appear to be 24% to 32% larger than the effect of unemployment in the state of enlistment at the end of contract.

A3.1 Timing of Separation

Servicemembers have the opportunity to seek early discharge, and short extensions that would not meet our definition of reenlistment. Appendix Figure A6 displays the c.d.f. of month of separation for non-reenlisting servicemembers with different initial contract lengths. The figure is truncated at 100 months, however, it is clear that the separation hazard is nearly zero at this point. Removing the truncation would reveal that almost all separations of those who we classify as not reenlisting occurs within the 100-month window.

For a policy maker interested in the welfare of young veterans, the relevant estimate would likely be that of the effect of unemployment at the actual time of discharge. Using the end-of-contract unemployment rate effectively results in a mis-measured separation unemployment rate, correction for which requires an increase in the coefficient by the inverse of the linear prediction of the separation rate by the end of contract unemployment rate, the Wald estimator. A regression in the framework from above of the unemployment rate at actual separation on the rate at the end of the first contract reveals a coefficient of 0.845 in 1993-2004, and 0.843 in 1993-2009; the relationship between the end of contract and separation date appears to be relatively stable over time. In either case, the inverse of the regression coefficient implies an increase of 17-18% would result from the use of the end of contract rate as an instrument for the separation rate, with little loss of precision. While we continue to use the end of contract rate throughout the paper, estimated effects can be scaled by this factor to reach the effect at separation.

A3.2 Migration

Veterans are a highly mobile population, and when we correct for the resulting mis-measurement of the relevant unemployment rate, the earnings effects above are magnified by as much as 20%.

There are two elements of the migration margin that may enter our calculation. First, some veterans will migrate over the course of their first contract, resulting in measurement error from the use of the home-state unemployment rate; this occurs independent of the period. Second, veterans may respond to changes in economic conditions by migrating at higher or lower rates, or to target better performing labor markets. Data is a serious challenge, because we require a panel that surveys veterans both before and after service. We draw evidence from the Current Population Survey, American Community Survey and 1979 and 1997 National Longitudinal Surveys of Youth. We summarize the analysis here, the details of which are available upon request from the authors.

Volunteer-era veterans in the Census and ACS migrate at higher rates than non-veterans, measured both as probability of residing in their state of birth, and the probability of recent migration. In the NLSY79 military oversample, 38% of veterans will have migrated from their home state, measured at age 14, when we observe them in the first full year following separation. A good portion of this migration will be to states near their home state, where unemployment rate movements will be similar. So, the reduction in the precision of our measure of the unemployment rate will be less than the implied 61% increase in the estimates that would occur if the unemployment rate in the destination state was entirely unrelated to that of the home state. We can take 38% as an upper bound, and attempt to correct the estimates for three elements of slippage: migration between age 14 and enlistment (38 to 30%); the decline in migration since the early 1980s, (which appears of similar percentage terms for veterans as with non-veterans, 30 to 20%); and, the correlation between the unemployment rate in the home state and the unemployment rate in the destination state (20% to 10%). This calculation implies that we should divide our earnings losses by 0.9, or multiply by 11%, to correct for in-service migration.

The second issue is increased migration in response to unemployment rate shocks themselves. We regress a dummy for living outside of the age 14 state in the year following separation on the unemployment rate in the nation and in the age 14 state unemployment rate. We estimate that an increase of one point in the state unemployment rate leads to 10% (t=1.8) of veterans in the NLSY79 not returning to their home state following separation; the estimate using the national unemployment rate (dropping year effects) is 13% (t=2). If economic conditions in the destination state are unrelated to the source state, we should adjust our estimates upwards by a similar amount

(11-15%) as we did to account for the timing of separation. Correcting for the directedness of migration (i.e. migrating to better labor market) using the 2006-2011 estimate of Yagan (2013) leads us to reduce this to 8-13%.[6] We can further adjust these for the decline in migration—although, whether migration in response to unemployment rate shocks has fallen is an open question, as far as we know. Taking the decline in responsiveness to be proportional to the overall decline, a one-third reduction in migration responsiveness would imply we increase our estimated earnings effects by 6-10%.

Together, the effects of average migration and migration in response to unemployment rate shocks imply an approximate 20% increase in the estimated effects of the unemployment rate on earnings and reenlistment. We fully acknowledge the many rough elements of this calculation. we've made here. For this reason, we report the "reduced-form" effects in our main tables, and the "first-stage" results as scaling factors.

Finally, some portion of the persistence of the shock likely reflects migration away from the state of enlistment in the years following discharge. In other words, the initial unemployment rate loses much of its relevance to labor market outcomes as young veterans flow out of their state of enlistment over the years following discharge. Somewhere between 8 and 14% of veterans ages 22-30

[6] Yagan (2013) uses IRS tax records to document the response of migration to local labor market shocks in 2001-2011, and finds that destination states received 25% of the shock received by source states, the estimate of directedness in this period. This estimate uses e-pop instead of the unemployment rate; these two measures of business cycle conditions generally move in tandem.

in the ACS report migrating between states in the preceding year, suggesting this explains a reasonable portion of the decay in the effect. To the degree that this reflects differential migration in response to labor market conditions, this migration may permit young veterans to improve the outcome on the earnings dimension, while suffering costs on other dimensions. As we lack detailed evidence on these effects, and believe our current estimates reflect the primary economically- and policy-relevant effects of interest, we leave these questions regarding veterans migratory behavior for future work.

A3.3 Education Benefits

The final margin of behavioral response that plays a role in our analysis is the use of G.I. Bill benefits. We have no quantitative estimate of the bias arising from selection into benefit usage, but the evidence supports the conclusion that it is small.

G.I. Bill benefits pay up to 36 months of tuition and fees, materials, and a monthly stipend to veterans who enroll in degree-granting programs within 10 years of separation. These benefits are not taxed, and will not show up in the SSA or military earnings. College benefit usage during the 1991-2005 period has been analyzed by Simon et al. (2010), who find that benefit usage rises during times of national unemployment. Estimated earnings responses do not change when we drop those who use their college benefits. The response of earnings for those who do not use their college benefits closely mirrors the average effect in the population, suggesting differential use of the benefits over the business cycle does not affect average earnings outcomes. Estimates for the 1993-2004 cohorts (not shown) look similar, with slightly larger losses for college attendees in the first two years. When we analyze the response of college benefit usage to state unemployment in our framework we find smaller responses than Simon et al. (2010) (who analyze national unemployment). Specifically, a one point increase in unemployment at discharge in 1993-2004 is associated with a 0.49 percentage point increase in the probability of college benefit usage (with a standard error of 0.14 percentage points); this is a 1.3% increase in the probability of benefits usage. This definition of benefit usage sets the threshold for usage at a minimum of 3 months and $2000 of total payment. Using a lower threshold of any usage would result in an estimate of 0.39 percentage points, while a higher threshold of 15 months and $5000 would generate an estimate of 0.40 percentage points. These regressions use the same specification as in the earnings analysis.

The selection effect of increased college benefit usage on our earnings estimates are likely to be small for the same reasons that the selection correction for reenlistment had little effect: namely, the composition of the no-reenlistment, no-college benefit usage pool of young veterans is not changed in a meaningful way by removing a relatively small fraction of people. We chose to report estimates for those who do not use their college benefit, as the college benefit margin may affect the overall estimates through an earnings-composition effect, since earnings are significantly lower for those who use their college benefit. However, these “losses” come from the choice of servicemembers to go to school rather than work. Therefore, they should not directly affect welfare—in the language of our model, college benefit usage is a mitigating mechanism.

We take the evidence as suggesting that G.I. Bill benefits offer, at best, very minor protection from local economic conditions for the average service member. We hope to explore G.I. Bill benefit usage further in future work, however, its role as insurance against unemployment shocks does not appear important enough to affect our estimates for the population of young veterans.

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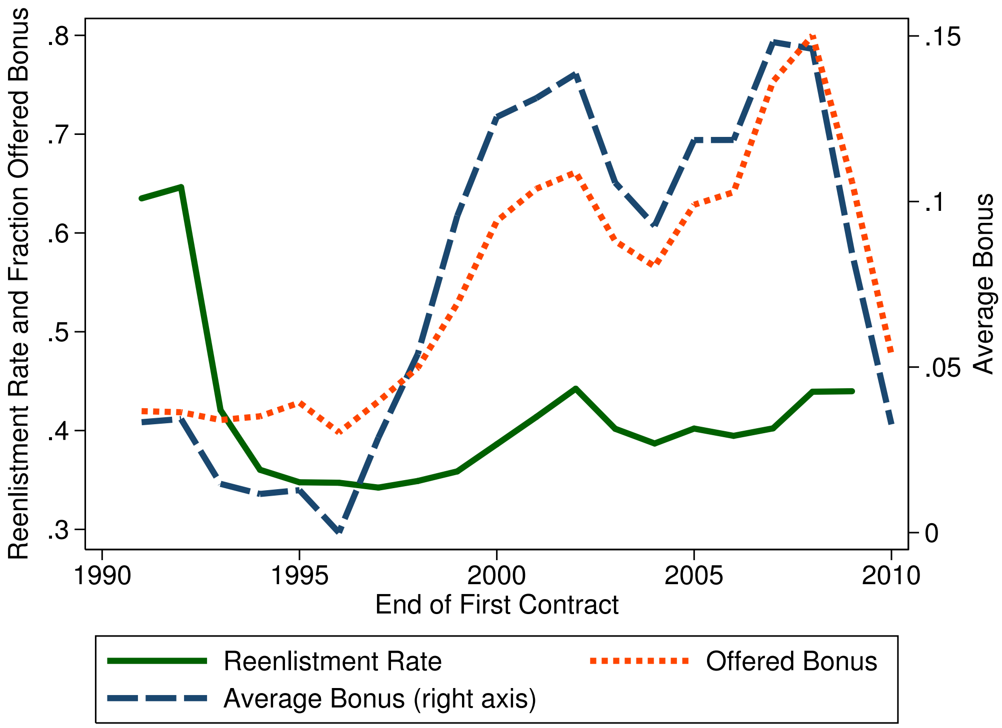
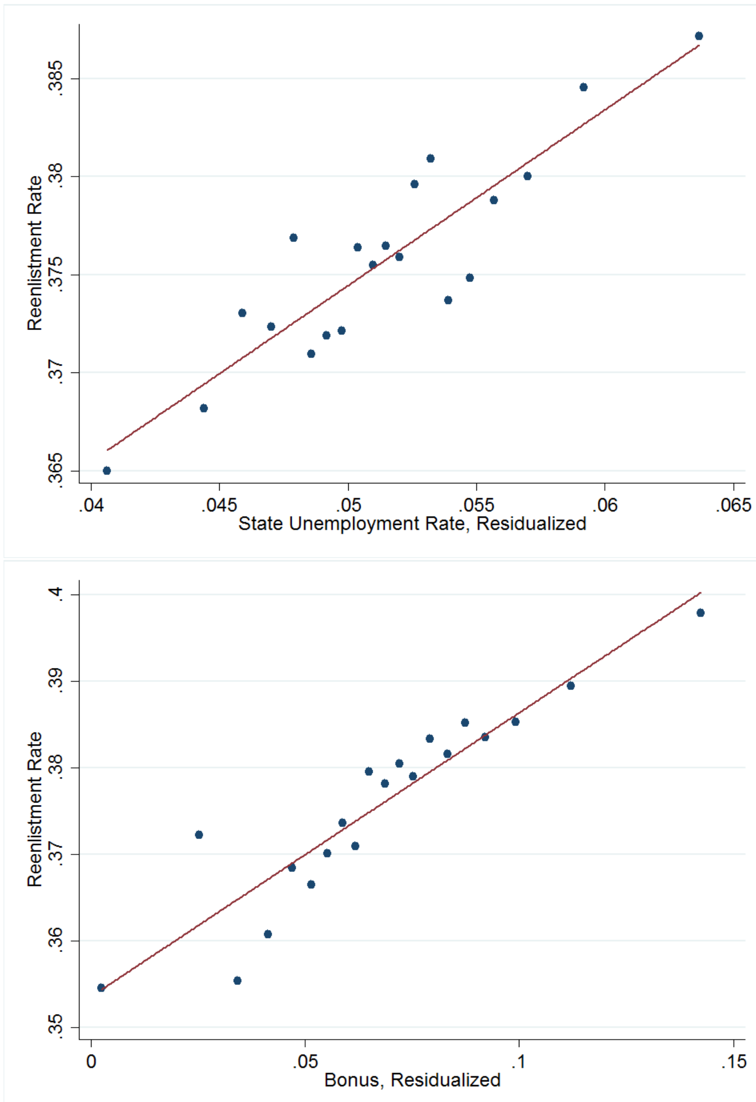
Figure A1: Bonuses and Reenlistment



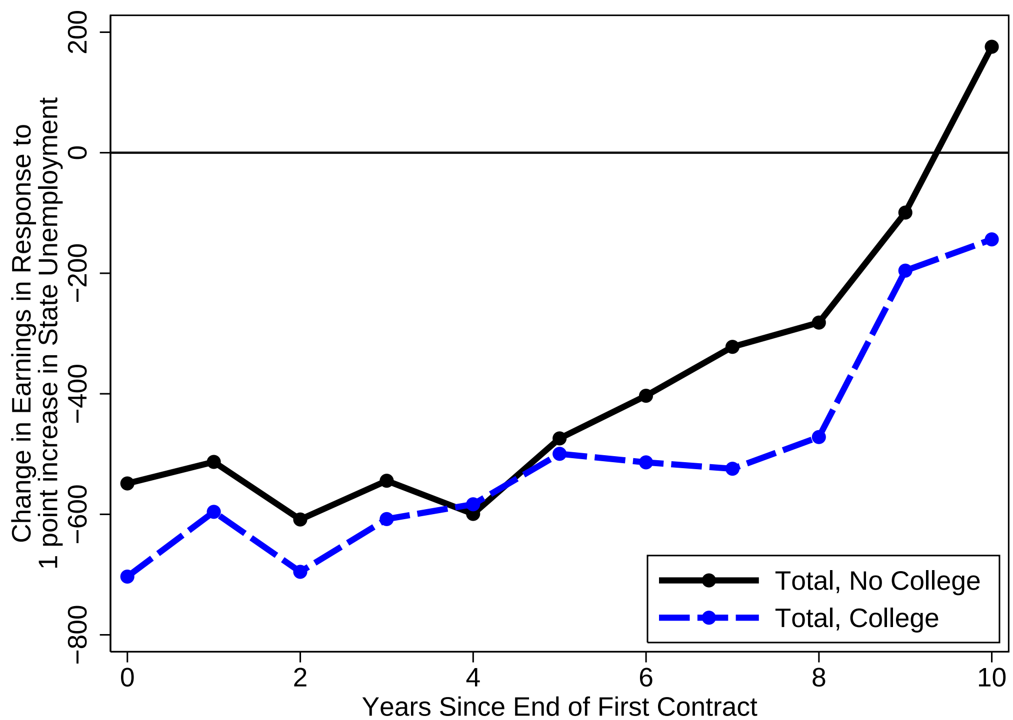
Notes: Bonuses grew in importance over the 1991-2010 period, increasing in economic expansions (particularly the late 1990s) and wars, and falling during military drawdowns (for example, in the mid-1990s and at the end of the sample) and recessions. Reenlistment shortfalls (relative to goals) occurred in 2005-2007. Sample comprised of all enlisted servicemembers at end of first contract. The bonus formula = months reenlisted (T^*) x bonus multiplier (b_{ot}) x monthly basic salary (y_{it}^{mil}).

Figure A2: Reenlistment Rises with State Unemployment and Bonus Offers, 1993-2004



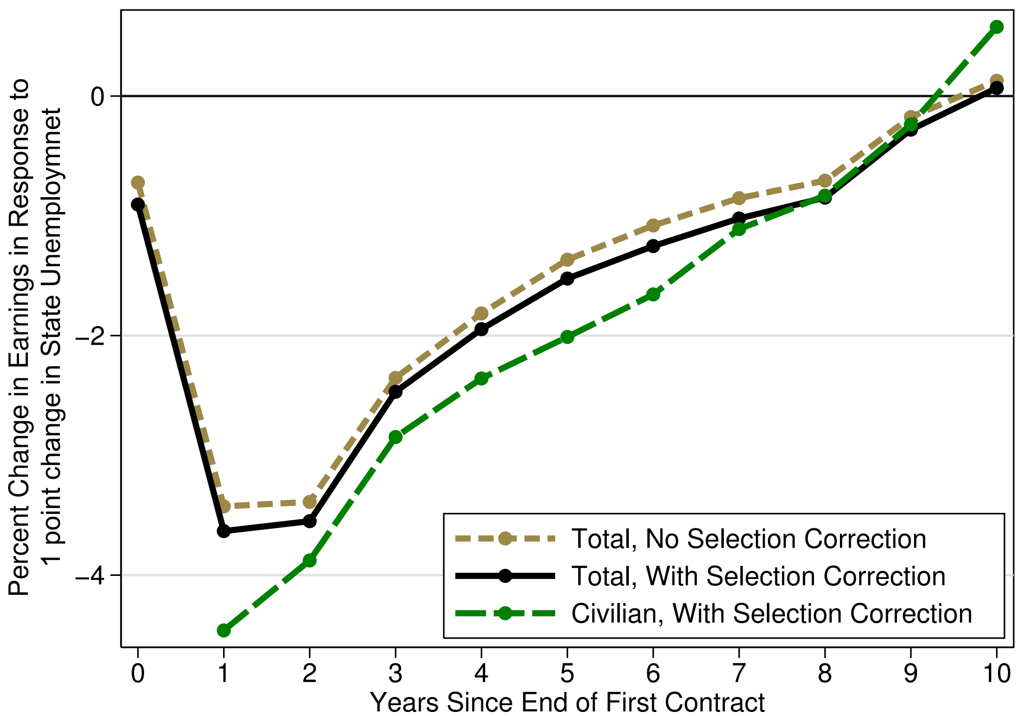
Notes: The figures plot the mean reenlistment rate within 20 equally sized bins of the independent variable, after controlling for state, occupation, and year by term effects in both regressions, and with the addition of occupation by year fixed effects in the state unemployment rate figure and state by year fixed effects in the bonus figure. The slope of the fit line in the upper figure is 0.90 and in the lower panel the slope is 0.33.

Figure A3: Effect of Unemployment on Earnings by College Benefit Usage, 1993-2004 Cohorts



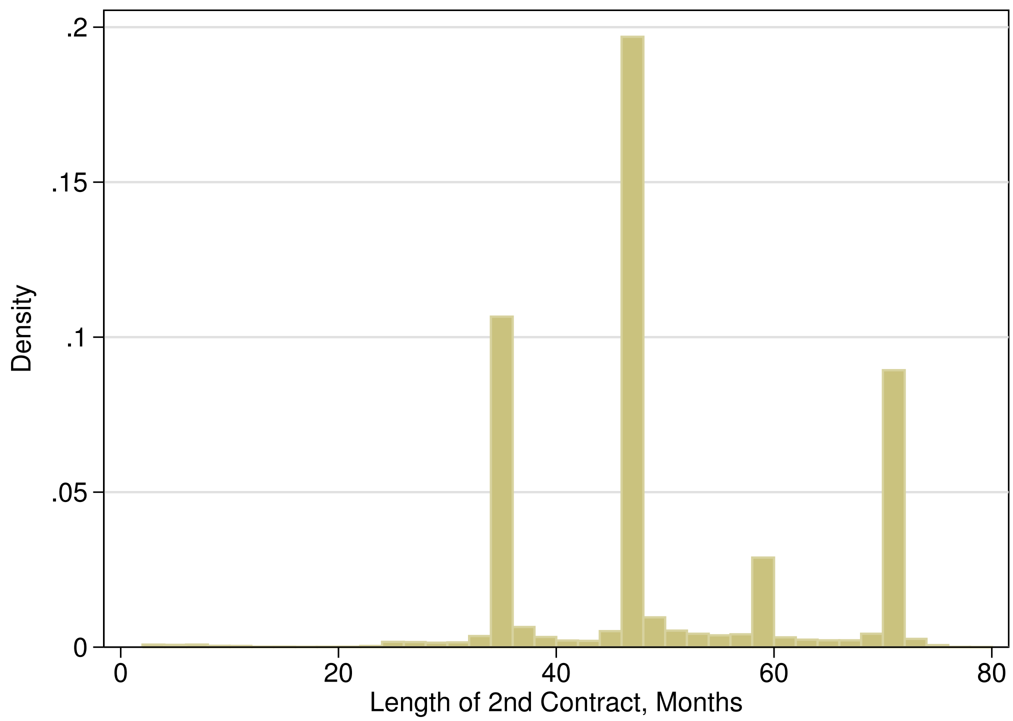
Notes: Figure plots the δ_k regression coefficients separately for all young veterans and those who do and do not use their college benefits.

Figure A4: Effect of Unemployment on Earnings, Selection and Civilian Earnings, 1993-2009 Cohorts



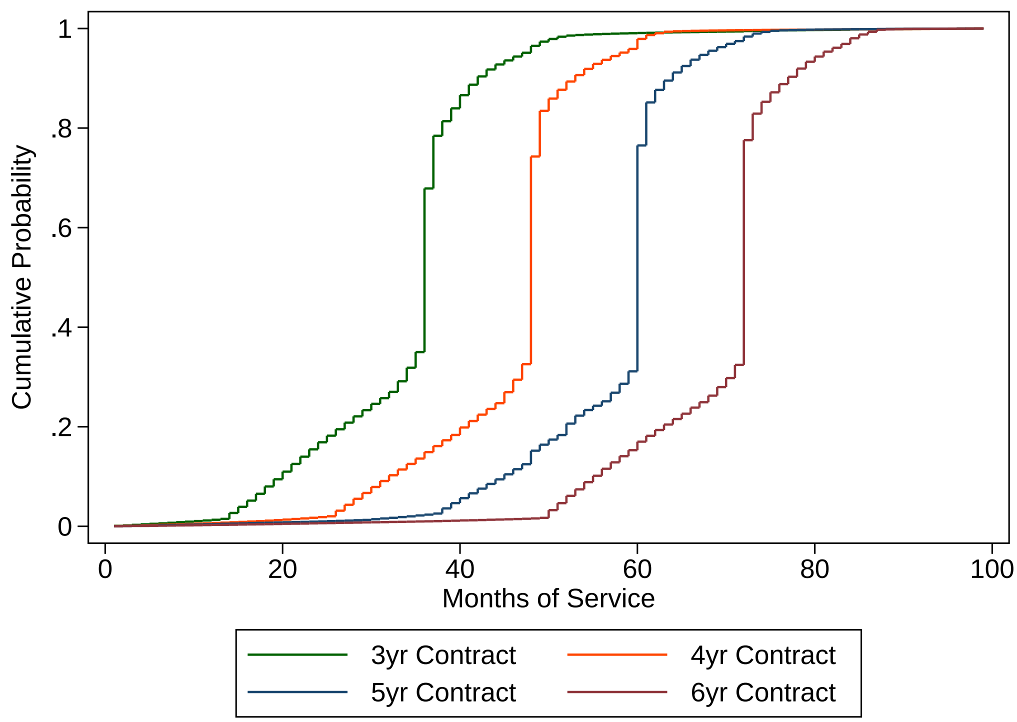
Notes: Figure plots coefficients from a regression of earnings *k* years after the end of the initial enlistment contract on state unemployment in the quarter of the end of the contract for our primary specification (Total, with Selection Correction) and two alternative specifications: the first alternative specification drops the selection correction; the second alternative specification plots the effect on civilian earnings, defined as SSA earnings minus military income. Civilian earnings show a large proportional effect in year zero, however, this comes on a small base in income; we omit the year zero civilian estimate from the figure to maintain scale. See text for additional details on regression specification, and Table A3 for average earnings, dollar value of losses, and standard errors.

Figure A5: Length of New Contract



Notes: Sample is all enlisted servicemembers who do not separate within three months following the end of their first contract. Two-month bins. In our empirical specifications, we define reenlistment as an increase in months of service committed of 24 or more months within three months of the end of the first contract. As can be seen, this captures virtually all second contracts.

Figure A6: Months of Service, No Reenlistment by End of First Contract



Notes: The figures plot the cumulative distribution of exits from the military relative to the end of the first contract. In the Appendix, we estimate the relationship between the realized unemployment rate at separation to the unemployment rate at the end of the first contract.

Table A1: Summary Statistics

	All Enlisted Men: 1992-2009				Young Veterans: 1992-2009			
	Mean	SD	p10	p90	Mean	SD	p10	p90
Eligible Age	24.06	2.39	22	27	23.92	2.26	22	27
More than HS	0.09	0.29			0.09	0.28		
Single at Entry	0.88	0.32			0.90	0.30		
AFQT Percentile	58.94	17.89	35	84	58.89	17.92	35	84
White	0.67	0.47			0.70	0.46		
Black	0.16	0.37			0.14	0.35		
Hispanic	0.10	0.30			0.10	0.30		
Initial Term	4.03	0.69			4.00	0.66		
N	1,098,377				673,260			

Notes: Descriptive statistics for All Enlisted Men with initial enlistment contracts ending between 1993 and 2009, and the subgroup of Young Veterans, defined as enlisted men who do not reenlist by the end of the first contract. AFQT, education and marital status measured at entry.

Table A2: Career Fields and Occupation Examples

Service/Career Field	Abbr.	3-digit Occupations
Army		
Infantry	11	Infantryman (11B), Mortarman (11C)
Aviation	15	Air Traffic Control Operator (15Q), Aircraft Electrician (15F)
Signal Corps	25	Radio Operator (25C), Cyber Network Defender (25D)
Adjunct General	42	Human Resources Specialist (42A), Musician (42R)
Navy		
Naval Aircrewman	AW	Aircrewman Helicopter (AWS), Aircrewman Avionics (AWV)
Steelworker	SW	no 3-digit sub-classifications
Culinary Specialist	CS	Culinary Specialist-Submarine (CSS)
Cryptologic Technician	CT	Maintenance (CTM), Collection (CTR)
Marines		
Personnel and Admin.	01	Personnel Clerk (0121), Career Retention Specialist (0143)
Infantry	03	Rifleman (0311), Mortarman (0341)
Utilities	11	Electrician (1141), Refrigeration and AC Technician (1161)
Ground Ordinance Maint.	21	Small Arms Repair (2111), Machinist (2161)
Air Force		
Aircrew Operations	1A	Aircraft Loadmaster (1A2X1), Flight Attendant (1A6X1)
Command and Control	1C	Air Traffic Control (1C1X1), Airfield Management (1C7X1)
Aerospace Maint.	2A	Aircraft Fuel Sys. (2A6X4), Aircraft Struct. Maint. (2A7X2)
Civil Engineering	3E	HVAC and Refrigeration (3E1X1), Fire Protection (3E7X1)

Table A3: Earnings: 1993-2009 Cohorts, Civilian Earnings and Selection

	Civilian Earnings, all		Total Earnings, all			
Year (k)	Av. Earn.	d_k	Av. Earn.	d_k	d_k, no $\Lambda(\cdot)$	N
0	\$9225	-7.31***	\$23409	-0.91***	-0.72**	348,092
		(1.15)		(0.32)	(0.33)	
1	\$19192	-4.46***	\$21878	-3.63***	-3.43***	353,264
		(0.66)		(0.47)	(0.48)	
2	\$23112	-3.88***	\$25222	-3.55***	-3.39***	335,130
		(0.69)		(0.57)	(0.58)	
3	\$25466	-2.85***	\$28098	-2.47***	-2.35***	314,538
		(0.84)		(0.72)	(0.73)	
4	\$27523	-2.36***	\$30477	-1.95***	-1.81**	294,804
		(0.81)		(0.71)	(0.71)	
5	\$29490	-2.01***	\$32675	-1.52**	-1.37**	272,422
		(0.73)		(0.64)	(0.63)	
6	\$31405	-1.66***	\$34770	-1.25***	-1.08**	248,844
		(0.57)		(0.48)	(0.47)	
7	\$33221	-1.11***	\$36739	-1.02***	-0.85***	227,524
		(0.42)		(0.34)	(0.33)	
8	\$34890	-0.83**	\$38517	-0.85***	-0.71**	208,301
		(0.36)		(0.33)	(0.32)	
9	\$36289	-0.24	\$39997	-0.28	-0.18	189,920
		(0.30)		(0.26)	(0.25)	
10	\$37664	0.58	\$41419	0.07	0.13	169,253
		(0.56)		(0.46)	(0.45)	

Notes: Table reports average total earnings and regression coefficients on state unemployment in the quarter of eligibility for separation (d_k, from Equation 1). Sample comprised of an unbalanced panel of all enlisted men who separate at the end of their first contract. In Columns 2 and 3 we report results on the civilian component of earnings, in Columns 4-5 we report the full sample result (with the selection correction), in Column 6 we report the effects without the selection correction (i.e. excluding $\Lambda(\cdot)$ from the estimation). Real 2010 \$ (CPI). s.e. clustered by state of enlistment. Significance levels: *** p<1%, ** p<5%, * p<10%.

Table A4: Effect of Bonus and Unemployment Rate on Reenlistment: Logit Estimation

	1993-2004			1993-2009
	All	$j \geq$4y	Late	All
Bonus (γ)	0.43***	0.39***	0.42***	0.49***
	(0.07)	(0.07)	(0.08)	(0.04)
UR (β)	1.03***	0.88***	0.98***	0.80***
	(0.07)	(0.07)	(0.08)	(0.04)
β/γ	2.39***	2.22***	2.32***	1.62***
	(0.75)	(0.69)	(0.81)	(0.43)
Career Fields	226	225	220	231
N (state x year x occ)	45,985	39,080	40,114	64,696
N (individuals)	661,545	494,474	558,599	960,731

Notes: Table reports results of regressions of reenlistment on bonus incentives and state unemployment rates , where the probability of reenlistment is in the logit form. The ratio of regression coefficients, β/γ, represents the willingness-to-pay (as a percentage of earnings on the next contract) for reenlistment in response to changes in home state economic conditions. We drop small occupations (N_{ot} $\geq$15) and include state, year and occupation fixed effects. Standard errors are clustered by career field and state of enlistment. Significance levels: *** p<1%, ** p<5%, * p<10%.

Table A5: Effect of Bonus and Unemployment Rate on Reenlistment: IV Model, 1993-2004

	OLS	IV	IV	IV	IV	IV	IV	IV
		All	All	All	$j \geq 4$	$j \geq 4$	$N_{ot} > 15$	Late
*ln*EntryCohort		0.010***	0.010***	0.010***	0.011***	0.012***	0.010***	0.010***
		(0.002)	(0.003)	(0.002)	(0.004)	(0.003)	(0.003)	(0.003)
1st stage F-stat		12.60	12.92	13.44	10.56	10.62	12.74	13.25
	OLS	IV	IV	IV	IV	IV	IV	IV
Bonus(γ)	0.37***	0.93***	0.88***	1.16***	0.66***	0.94***	0.85***	0.80***
	(0.04)	(0.30)	(0.29)	(0.34)	(0.27)	(0.31)	(0.29)	(0.29)
UR (β)	0.85***	0.84***	0.84***	0.83***	1.06***	1.06***	0.90***	0.95***
	(0.19)	(0.18)	(0.18)	(0.17)	(0.18)	(0.17)	(0.17)	(0.20)
β/γ	2.27***	0.91**	0.96**	0.71**	1.62**	1.13***	1.06**	1.17**
	(0.63)	(0.37)	(0.40)	(0.27)	(0.71)	(0.42)	(0.45)	(0.51)
Mean AFQT, BMI	x		x	x	x	x	x	x
In-service X_i				x		x		
Career Fields	323	303	303	291	299	281	212	294
N	632,261	625,755	625,755	622,855	491,893	490,162	619,210	562,144

Notes: Table reports results of IV regressions of reenlistment on bonus incentives and state unemployment rates with a rich set of controls (see text for details). The ratio of regression coefficients, β/γ, represents the willingness-to-pay (as a percentage of earnings on the next contract) for reenlistment in response to changes in home state economic conditions. Sample comprised of all enlisted servicemembers at end of first contract between 1993 and 2004 with other sample restrictions noted in column headers: all contracts, initial contracts of at least four years ($j \geq$4y), dropping small occupations ($N_{ot} \geq$15) and restricting to reenlistments occurring in the last year of the contract (Late). s.e. clustered by career field and state of enlistment. Kleibergen-Paap F-statistics reported for 1st stage. Significance levels: *** p<1%, ** p<5%, * p<10%.