Does Cognitive Ability Reduce Psychological Bias?

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Abstract

A burgeoning literature in economics argues that bounded cognition can explain many observed empirical deviations from rationality. Consistent with this hypothesis, we show that individuals with greater cognitive ability behave more closely in accordance with economic decision theory. However, even the most cognitively skilled individuals display significant biases. In two laboratory studies, one conducted with Harvard undergraduates and one with Chilean high school students, we find that individuals with greater cognitive ability are more patient over short-term trade-offs and less risk-averse over small-stakes gambles. In both studies, mathematical ability seems to be more predictive of normative decision-making than verbal ability. In the sample of Chilean students, achievement in elementary school (grades 1-6) is strongly predictive of decisions made at the end of secondary school. Drawing on the National Longitudinal Survey of Youth (NLSY), we show that, even after controlling carefully for labor income, more cognitively skilled individuals are more likely to participate in financial markets, are more knowledgeable about their pension plans, accumulate more assets, and are more likely to have tax-deferred savings. These findings persist when we use sibling relationships to identify models using within-family variation in cognitive ability. Finally, various institutional measures of school quality are predictive of sophisticated decision-making, suggesting a role for human capital policy in reducing the impact of psychological biases.

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1 Introduction

Real people do not live up to the ideal of infinite intelligence generally assumed in economic models. An enormous literature in psychology argues that cognitive constraints force people to rely on fallible heuristics (Kahneman, Slovic and Tversky, 1982; Kahneman 2003), resulting in systematic deviations from normative[1] decision-making. A burgeoning literature in economics argues that bounded cognition can explain many observed empirical deviations from rationality, such as low stock market participation rates, failure to achieve diversified portfolios, and inability to smooth consumption in response to transitory income shocks (e.g., Thaler 1992). A natural implication of this hypothesis is that individuals with greater cognitive ability—those whose cognition is "less bounded"—will behave in a way that is more consistent with standard economic models.[2] In this paper, we test whether sophisticated economic decision-making is correlated with cognitive ability.

Our first set of tests addresses whether cognitive ability predicts greater risk-neutrality over small-stakes gambles and greater patience over short-term trade-offs. We use data from two laboratory studies in which we measured preferences using decisions with real stakes and collected information on participants' standardized test scores. Participants in the first study were Harvard students with SAT scores ranging from the 76th to the 99th percentile. Participants in the second study were Chilean high school students whose standardized test scores range from the 45th percentile to the 99th percentile. Rabin's (2000) calibration theorem implies that participants who rejected the gambles we offered them could not plausibly have been behaving in accordance with standard expected utility theory, the normative benchmark. If a participant's risk-aversion in our studies were truly explained by expected utility theory, then that participant's utility function would have to be so concave that the person would turn down a gamble that gives a 50% chance of losing $5 and a 50% chance of winning an infinite amount of money! Similarly, impatient choices by participants in our studies cannot be explained by time-consistent exponential discounting at a

[1] We use the term "normative" throughout this paper. To be clear, by "normative" decision theory, we mean exponentially discounted, expected utility with parameter values that imply relative patience over long horizons and reasonable risk-aversion over large stakes—the standard economic model that is often taught prescriptively in business and public policy policy schools.

[2] Some existing evidence suggests that more cognitively able individuals tend to be less susceptible to biases in judgment and decision-making (such as the sunk-cost fallacy) in hypothetical choice scenarios. See Stanovich (1999) and Stanovich and West (1998) for a review. For earlier work on the connection between cognitive ability and judgment, see Jepson, Krantz, and Nisbett (1983) and Larrick, Nisbett, and Morgan (1993).

plausible discount rate (less than 50% per year). Therefore, risk-averse and impatient choices in our studies imply deviations from standard prescriptive models.

In both the Harvard and the high school population, we find an economically large and statistically significant positive relationship between cognitive ability and sophisticated decision-making. For example, in the second study we calculate that a one-standard-deviation increase in measured mathematical ability is associated with an increase of about 13 percentage points in the probability of behaving in a risk-neutral fashion over small stakes (as against a mean probability of about 10%) and an increase of about 18 percentage points in the probability of behaving patiently over short-run trade-offs (with a mean of about 28%).[3] Results from both studies also indicate that mathematical ability is more predictive of sophisticated decision-making than verbal ability. Finally, in the second study we show that elementary-school (grades 1 through 6) grade point averages in mathematics strongly predict decision-making at the end of secondary school (grade 12), and may be even more predictive than contemporaneous measures of cognitive ability.

Given these striking results from the laboratory, we turn next to the question of whether the added biases of less skilled individuals have meaningful effects on real-world economic behavior. Using data from the National Longitudinal Survey of Youth (NLSY), we show that individuals with greater cognitive skills make consumption and personal financial decisions that more closely resemble the predictions of economic theory. Individuals with greater cognitive ability are more likely to participate in financial markets, are more knowledgeable about their pension plans, accumulate more wealth, and are more likely to participate in tax-deferred savings programs. These results survive detailed controls for household labor income. They also remain important when we take advantage of sibling groups in the NLSY to control for family fixed-effects. As in our laboratory studies, we find that mathematical ability is more predictive of sophistication than measured verbal skills. Finally, in light of potentially important policy implications, we also show that individuals attending schools with superior staff and resources perform better on a test of cognitive ability and are subsequently more likely to accumulate financial assets and invest them wisely.

A classic defense of the assumption that rationality prevails in markets is that individual heterogeneity in decision-making skill allows the most highly skilled to sort into occupations in which

[3] See also Shoda, Mischel and Peake (1990), who show that preschoolers' ability to delay gratification is positively related to subsequent high school performance on the SATs.

decision-making skill is most important (Friedman, 1953; Glaeser, 2004). Our results suggest that this view is partially right and partially wrong. On the one hand, since more cognitively skilled individuals make better decisions, sorting is possible. On the other hand, a large fraction of even the most cognitively skilled individuals in the population display substantial biases. For example, among the participants in our study of Harvard undergraduates with perfect math SAT scores, only $\frac{1}{3}$ behaved in a manner fully consistent with risk-neutrality over small gambles, and only $\frac{2}{3}$ behaved patiently over short-run intertemporal trade-offs, and a majority of NLSY respondents in the upper decile of cognitive ability do not own stocks or have an IRA.

The remainder of the paper is organized as follows. In Section 2, we first review recent arguments that small-stakes risk-aversion and short-term impatience are non-normative. We then present new laboratory evidence showing that small-stakes risk-aversion and short-horizon time discounting are both lower in individuals with greater cognitive ability. In Section 3, we ask whether this relationship extends to real-world economic behavior. We use data from the NLSY to investigate the correlation between IQ and anomalies such as low levels of financial market participation and limited knowledge about pension plans. Section 4 presents additional evidence from the NLSY suggesting that human capital policy can affect financial sophistication by raising cognitive ability. Section 5 interprets our results and concludes with welfare and policy implications.

2 Cognition and Preferences

In this section, we ask whether preferences—usually taken as exogenous by economists—are related to cognitive ability. To address this issue, we conducted two laboratory studies. In both, we measured participants' (small-stakes) risk preferences and (short-term) time preferences by having participants make decisions with real stakes. We also collected information on participants' standardized test scores. Participants in the first study were Harvard students, whose SAT scores ranged from the 76th to the 99th percentile (conversions from raw scores to population percentiles based on College Board, 2001). To show that our results are not restricted to the high end of the cognitive ability distribution, we replicated our results in a second study with Chilean high school students, whose standardized test scores range from the 45th percentile to the 99th percentile (conversions from raw scores to population percentiles based on Universidad de Chile, 2004). Moreover, in the second study we obtained data on participants' elementary-school grade point averages, allowing us to investigate the role of heterogeneity in cognitive ability arising early in life.

2.1 Calibrating Preferences

In the studies reported below, we will investigate the connection between cognitive ability and behaviors such as small-stakes risk-aversion and short-term impatience. We will interpret our results as a test of the hypothesis that such behaviors represent a "mistake" from the perspective of economic decision theory (Rabin, 2000; Rabin and Thaler, 2001; Rabin, 2002). In this subsection, we describe the decision problems we used in our laboratory studies, and we use existing calibration theorems to argue that any risk-aversion or discounting displayed in the laboratory cannot be a result of agents maximizing standard preferences with reasonable parameters.

Small-Stakes Risk Preference. We elicited risk preferences with five questions of the following form:

> Please circle either Choice A or Choice B.
> (A) You get \$0.50 for sure.
> (B) If the die comes up 1, 2, or 3, you get X. If the die comes up 4, 5, or 6, you get nothing.

For the five questions, X was \$0.95, \$1.05, \$1.15, \$1.25, and \$1.35. In each case, (A) is the safe bet, and (B) is the risky bet.

In this problem, any reasonable expected-utility preferences imply perfectly risk-neutral behavior: choosing the safe bet for $X = \$0.95$ and choosing the risky bet for all other values of X. To understand why any other behavior is extreme, consider the risk-averse choice of the safe bet when $X = \$1.05$. Rabin's (2000) calibration theorem implies that an expected-utility maximizer who would make this choice at any level of initial wealth would also turn down a gamble that gives a 50% chance of losing \$5 and a 50% chance of winning an infinite amount of money. (Put another way, an expected-utility maximizer with a constant relative risk-aversion utility function and lifetime wealth of \$1 million would make this decision only if the coefficient of relative risk-aversion exceeded 100,000!) The reason is that risk-aversion over such small stakes requires non-negligible local concavity that, when extrapolated, leads to extraordinary risk-aversion over larger stakes. Because this behavior over larger-stakes gambles is obviously implausible, the risk-averse choice of the safe bet when $X = \$1.05$ is quantitatively inconsistent with standard expected utility theory. Making the risk-averse choice of the safe bet when X is larger than \$1.05 implies even more extreme risk-aversion over larger stakes and is therefore even more questionable.[4]

Short-Term Time Preference. We measured time preferences with six questions of the form:[5]

> Please circle either Choice A or Choice B.
> (A) You get \$5.00 right now.
> (B) You get X a week from now.

For the six questions, X was \$5.05, \$5.35, \$5.55, \$5.75, \$5.95, and \$6.15. The impatient choice is always (A), and the patient choice is (B).

Taking time-consistent exponential discounting as the normative standard, an impatient choice is non-normative in the sense of implying an absurdly high discount rate. For example, a person indifferent between receiving \$5.00 now and \$5.05 in one week is, under constant discounting, implicitly discounting cash flows at a rate of $52 \times \ln\left(\frac{\$5.05}{\$5.00}\right) \approx 52\%$ per annum![6] Someone who makes the impatient choice for larger values of X is implicitly discounting at an even higher rate. Such high discount rates imply virtually no regard for the future, which seems unlikely to be the correct explanation of a participant's impatient choice in this decision problem.

These arguments suggest that short-run impatience and small-stakes risk-aversion are inconsistent with sensibly calibrated models of decision-making—expected utility maximization and exponential discounting—that have been argued to have normative force. In the next two subsections, we will test the hypothesis that these deviations from normative benchmarks result from bounded

[4] The risk-seeking choice of the risky bet for $X = \$0.95$ is similarly suspect. Even if risk-loving preferences are considered acceptable over larger stakes, they are inappropriate in this context of small stakes. Assuming a convex utility function, an argument analogous to Rabin's (2000) calibration theorem would show that the risk-seeking choice of the risky bet for $X = \$0.95$ is quantitatively inconsistent with standard expected utility theory. Such a choice for an expected-utility maximizer would imply a counterfactual degree of risk-seeking over somewhat larger stakes. Therefore, we take risk neutrality as the normative benchmark in our studies.

[5] Our approach is a standard laboratory tool for measuring time preference (Frederick, Loewenstein and O'Donoghue, 2002). However, according to economic theory, questions involving monetary rewards should not measure impatience, since subjects can borrow or lend money at the market rate of interest regardless of how they discount future utility (Fuchs, 1982). Nevertheless, time preference elicited in a manner similar to the one we employ here has been shown to correlate with behaviors such as drug addiction (Kirby, Petry and Bickel, 1999), cigarette smoking (Fuchs, 1982), and rapid exhaustion of food stamps (Shapiro, 2005).

[6] Imputing an exact discount rate over utility flows from this indifference requires assumptions about the utility function. However, Arrow (1971) and Rabin (2000) imply that, for participants with a reasonable amount of lifetime wealth, utility should be approximately linear over the small-stakes choices we offer. Within the standard model, the discount rate over small cash flows therefore approximates the discount rate over utility flows.

cognition by estimating the relationship between cognitive ability and preferences expressed in the laboratory.

2.2 Study 1: Harvard College Undergraduates

2.2.1 Participants

Participants were 60 undergraduates (virtually all Harvard students), recruited through on-campus posters and e-mail solicitations. We promised students $5 for participating in a 45-minute experiment, with the possibility to earn more "depending on your responses [in the experiment]." We allowed only non-economics majors to participate because we were concerned that economics students would be familiar with our preference elicitation procedures. In total, we held six sessions, on February 8, 21 and 22 and May 14, 15 and 16, 2004.

2.2.2 Procedure

Each session began with an experimenter handing out a questionnaire booklet to each participant in the classroom. The participants moved through each page of the questionnaire in unison, at a pace determined by the experimenter. The experimenter read aloud instructions throughout the session. The questionnaire was divided into sections (with neutral labels such as "Choices" and "More Choices"), each of which elicited a type of preference, as described below. The order of the sections differed across sessions; this order had no effect on the results, so we do not discuss it further.

At the end of the questionnaire, after all the preference elicitation questions, we asked participants for their major, year in school, and gender. Most importantly for our analysis later, we asked participants for their highest Math and Verbal SAT I scores. We also asked participants for ACT or other standardized test scores, but virtually all participants had taken the SAT, and virtually none had taken the ACT or other standardized test. We therefore focus our analysis on SAT scores, which are available for 57 of our 60 participants. As a check on the accuracy of the self-reported SAT scores, we included a section of SAT-like math questions near the end of the questionnaire (after participants had responded to all the preference-elicitation sections). The correlation between performance on the SAT-like math questions and self-reported Math SAT score was 0.33, which is significantly different from 0 ($p = 0.012$). Given that our math test contained six questions and the Math SAT contains about 50, this correlation is not especially low and seems to indicate that students' self-reported scores contain valid information.

We paid all participants \$5 in cash for their participation immediately at the completion of the session. We also paid participants (by check) for their choices so that our preference-elicitation procedures were incentive-compatible.[7] As described below, for some sections we paid participants immediately after the experiment. For other sections, we mailed checks to participants within a week of their participation.

Small-Stakes Risk Preference. The section of the questionnaire that elicited risk preferences comprised the five questions described above. Each question offered the choice between \$0.50 for sure and a gamble that gives a 50% chance of winning \$0 a 50% chance of winning X, where X takes the values \$0.95, \$1.05, \$1.15, \$1.25, and \$1.35. The questionnaire contained all five questions on the same page, with X in ascending order. This presentation made salient to participants the strategy of choosing (A) (the safe bet) for small X and (B) (the risky bet) for large X. In fact, 51 out of 60 gave monotonic responses, choosing (A) below some threshold value of X and (B) above it. For these participants, the threshold measures the level of risk-aversion.[8] Seventeen participants behaved in a way that is consistent with perfect risk-neutrality, rejecting the bet for $X = \$0.95$ and accepting the bet for all other values of X.

To make sure that participants understood the choices they were making, we gave them an example question in the instructions for this section. We also informed participants that they would answer five questions of the above form. Finally, we gave participants the opportunity to ask any questions about the instructions.

[7] During each of the preference-elicitation sections of the questionnaire, participants heard a CD of piano notes while they filled out the section. In each section, half the participants were required to remember the number of times they heard a specific sequence of musical tones. This kind of distraction ("cognitive load") may reduce the cognitive resources available to participants (e.g., Gilbert and Silvera, 1996). To incentivize participants in the cognitive load condition to pay attention to the tones, we made payment for that section contingent on correct recall of the number of repetitions of the sequence of tones. The cognitive load manipulation severely hurt participants' performance on the Math SAT-like questions, but it had no effect on preferences. We note here the fact that we conducted this manipulation because it affected how much participants were paid. In particular, participants who incorrectly recalled the number of repetitions of the sequence of tones were not paid for the corresponding preference-elicitation section of the questionnaire. In Study 2, we did not manipulate cognitive load.

[8] In all cases, we present statistical results for the whole sample of subjects but note that our results are substantively unchanged if we restrict attention to only those subjects whose responses were monotonic.

Participants had 15 seconds to answer the five questions. Participants knew from the instructions that after they made a selection for each question, we would roll a die five times to determine their payment for this section. We asked a participant to roll the die to maximize our credibility. We paid participants for this section in the check that we mailed within a week of the participant's participation.

Short-Term Time Preference. We measured time preferences with the six questions described above. For each question, the participant chose between \$5.00 today and X a week from today. All six questions were on the same page, with X in ascending order: \$5.05, \$5.35, \$5.55, \$5.75, \$5.95, and \$6.15. We ordered the questions this way to make obvious to participants the strategy of choosing (A) (the immediate payoff) for small X and (B) (the delayed payoff) for large X. As it turned out, 53 out of 60 participants chose (A) below some threshold value of X and (B) above it. Among the 57 participants for whom SAT scores are available, 6 did not answer all of the time preference questions. These participants have been omitted from our analysis.

In the instructions for this section, the experimenter gave participants an example question, informed them that they would have 15 seconds to answer the six actual questions, told them that a die roll would select the question to be implemented, and gave them a chance to ask questions. We asked a participant to roll the die. The instructions explained that participants would receive a check to pay them for this section. That check would be post-dated by a week if the participant had chosen (B) for the relevant question. We gave participants this check immediately after the session.

2.2.3 Results

In this subsection we present evidence on the connection between cognitive ability and preferences expressed in the laboratory setting. Our primary measure of cognitive ability is the participant's math SAT score. As Figure 1 illustrates, participants' math SAT scores were all between 600 and 800, with a mean of about 750 and a median of 760. Almost a quarter of our respondents reported a perfect score of 800. The nationwide average math SAT score was about 500 during the time period in which our participants took the exam (College Board, 2001), suggesting that our sample of (mostly) Harvard undergraduates is not representative of the general population. Since we are only observing the upper tail of the distribution of scores, our sample selection is likely to bias us against finding evidence of a role for cognitive ability in determining preferences.

Figure 1 also serves to illustrate that the distribution of math SAT scores is highly non-normal, which may present a problem for a small-sample study. To simplify the statistical analysis, we therefore conduct most of our analysis by comparing participants with below-median math and verbal SATs to participants with above-median scores. When we instead use specifications that are linear in SAT score, the estimated coefficients are comparable in sign and magnitude, but less statistically precise and in general not statistically distinguishable from zero.

Small-stakes risk preference. As explained in more detail above, if people with greater cognitive ability reason more normatively, then we should expect them to be essentially risk-neutral over the small-stakes gambles we offered in our study.[9] We coded participants as risk-neutral if they accepted all gambles for more than $1.00 and rejected the gamble for $0.95.

Column (1) of Table 1 presents a probit analysis comparing the risk-neutrality of participants who are below and above the median SAT math score. While only 21% of below-median scorers made choices fully consistent with risk-neutral expected value maximization, 43% of those reporting an above-median SAT math score made such choices. This difference is statistically significant at the 10% level. As column (2) shows, including controls for gender and year in school reduces the coefficient only slightly, and it remains significant at the 10% level. In column (3), we simultaneously estimate the effects of math and verbal performance, and find a positive effect of an above-median math SAT that is statistically significant at the 5% level. In contrast, the estimated effect of an above-median verbal SAT score is negative and statistically imprecise.

Short-term time preference. As discussed above, if people with greater cognitive ability reason more normatively, then we should expect them to exhibit perfect patience in the choices we offered them. We coded participants as perfectly patient if they always took the (larger) delayed reward over the immediate reward.

Column (4) reports our findings on the share of respondents whose behavior is perfectly patient.

[9] Prospect theory (Kahneman and Tversky, 1979) can explain small-stakes risk aversion by assuming that people have locally concave utility over monetary gains. Our finding that people with greater cognitive ability are less risk-averse over small-stakes gambles suggests that these people behave more in accordance with expected utility theory. This finding–and even more so the finding on loss-aversion in the Chile study discussed next–is consistent with evidence reviewed in Stanovich (1999) that people with greater cognitive ability are less subject to gain-loss framing effects.

While the results go in the expected direction—patience is greater among the high SAT math scorers—the difference is not statistically significant. When we add demographic controls in column (5), the estimated effect of math ability increases but remains statistically insignificant. Finally, column (6) shows that controlling for mathematical ability, verbal performance is estimated to have a negative (but statistically imprecise) effect on patience.

The statistical weakness of these findings is driven largely by the fact that many high scorers were not willing to accept a delay in payment (i.e., a post-dated check) at any price, and therefore chose the current payment in all cases. Columns (7) through (9) therefore re-estimate the models of columns (4) through (6), restricting attention to those participants who accepted at least some delay in payment. As column (7) shows, the positive effect of an above-median math SAT score is statistically significant at the 10% level. When we add basic controls in column (8), the coefficient increases and becomes statistically significant at the 5% level. Finally, column (9) shows that when we include both mathematical and verbal performance, mathematical performance has a large positive effect (significant at the 10% level), whereas the effect of verbal ability is quite small (though statistically imprecise).

Two patterns emerge from this study. First, individuals with greater cognitive ability do appear to make choices that are more consistent with normative standards. This is consistent with the view that cognitive ability reduces the influence of psychological biases in economic decision-making. Second, mathematical ability is more strongly related to decision-making than is verbal ability. One attractive, though tentative, explanation for this finding is the observation that math SAT scores are much more correlated than verbal SAT scores with cognitive scientists' measures of "general cognitive ability" (Frey and Detterman, 2004).

While the results of this study are suggestive, it is difficult to generalize the findings out of sample due to the relatively small amount of variation in cognitive ability among participants. We therefore turn next to a population with a wider range of measured ability.

2.3 Study 2: Chilean High School Students

Study 2 closely mirrored Study 1, except that participants were seniors at a semi-private high school in Santiago, Chile. To make sure the questionnaire from Study 2 was comparable to analogous parts of the Study 1 questionnaire, those parts were translated by a native Spanish speaker from Santiago.[10]

At the end of their senior year, Chilean high school students take a national standardized test, the Prueba de Selección Universitaria (PSU), which has three obligatory sections—Math, Verbal, and Chilean History and Geography—as well as specific subject-area sections. The math section is very much like the SAT I Math Section, while the verbal section covers literary concepts, reading comprehension, logical paragraph organization, and vocabulary. For many Chilean universities, the PSU score together with GPA are the sole determinants of admission. Because performance on the exam is so important, seniors at this school take monthly practice tests. We obtained from the school for each participant 5 practice test scores for April through August, 2004.

Most participants entered the school for kindergarten at age 4 or 5. Some students were admitted because older siblings had attended, but most were admitted on the basis of adequate performance on an entry exam. Most students (more than 80%) had received their entire formal education at the school. Therefore, these participants had had a similar schooling experience. The school gave us grade point averages for grades 1 through 11 for all students participating in our study for whom such data were available.

2.3.1 Participants

Participants were the 92 out of 160 members of the senior class of a Chilean high school who turned in parental consent forms. None had received any formal training in economics. We held a single 30-minute session on August 24, 2004, with participants sitting in widely-separated desks in the school gym.

2.3.2 Procedure

As in Study 1, an experimenter guided participants through the questionnaire in unison by reading instructions aloud. The questionnaire contained a section that elicited small-stakes risk preferences, a section that elicited short-term time preferences, another section that specifically examined loss aversion, and a final section that asked a few demographic questions.

[10] To be sure that the translation was accurate, we asked a different native Spanish speaker to back-translate the questionnaire into English. The back-translated questionnaire closely matched the original.

Participants were paid in cash for their choices in the risk-preferences and loss-aversion sections, as well as paid a participation fee of 1250 pesos (about \$2.00 at the then-exchange rate of 632 pesos/\$), during lunch break the following day. Participants who chose to be paid "now" in the time-preference section were also paid in cash for that section at the same time. Participants who chose to be paid "a week from now" in the time-preference section were paid in cash during lunch break one week after the experiment.

Small-Stakes Risk Preference. The procedure for measuring risk preferences was the same as in Study 1, except that option (A) (the safe bet) paid 250 pesos, and option (B) (the risky bet) paid X, where X was 400, 550, 700, 850, and 1000 pesos. We gave participants an example question in the instructions and the opportunity to ask questions. There was no stated time limit for answering the questions, but we waited about 6 minutes for all participants to finish before moving on. We then rolled a die five times to determine their payment for this section. 70 out of 92 participants gave monotonic responses, choosing (A) below some threshold value of X and (B) above it.[11]

Short-Term Time Preference. We measured time preferences the same way as in Study 1. After an example question and an opportunity to ask the experimenter about the instructions, participants chose between 500 pesos now and X a week from now, where X was 450, 550, 650, 750, 850, and 950 pesos. After giving participants enough time (about 6 minutes) to answer the six questions, the experimenter rolled a die to determine which question would be implemented. 87 out of 92 participants had a threshold value, taking the 500 pesos now for X below the threshold and taking the money in a week from now for X above the threshold.

2.3.3 Results: Preferences and Cognitive Ability

In this section, we present our findings regarding the connection between cognitive ability and economic preferences. For each participant, we calculated the average score on the math and verbal sections of the five practice PSU exams for which we have data. We then standardized these measures by dividing each by the standard deviation in the entire Chilean test-taking population so that coefficients can be interpreted as marginal effects of a one-standard-deviation increase in

[11] In our study of Chilean high school students, we also collected information on loss aversion. Since the results for small-stakes loss aversion closely parallel those for risk-aversion, we present our findings on loss aversion in Appendix Table 2 and focus here on risk and time preferences.

the independent variable.

We will in each case use as our dependent variable a dummy for whether the participant's choices were perfectly consistent with normative predictions (i.e., perfectly risk-neutral or perfectly patient).[12] Roughly 11% of participants were perfectly risk-neutral, considerably lower than the 33% of Harvard undergraduates who displayed risk-neutrality. Similarly, 28% of the high school students exhibited perfect patience, which is again considerably lower than the 53% of Harvard undergraduates who display no short-term discounting. Since average cognitive ability (relative to the overall population) is higher in the sample of Harvard students, these facts seem supportive of our core hypothesis that cognitive ability reduces deviations from normative economic decision-making. However, despite our efforts to ensure comparability, differences in responses between participant populations may be contaminated by small differences in context and experimental procedure. We therefore turn in Table 2 to a within-population study of Chilean high school students.

Small-Stakes Risk Preference. Columns (1) of Panel A presents probit estimates of the effects of mathematical ability on the participant's propensity to display perfect risk-neutrality. Coefficients can be interpreted as marginal effects evaluated at the mean of the independent variables. We estimate that a one-standard-deviation increase in measured mathematical ability is associated with a 14 percentage point greater likelihood of risk-neutrality, a statistically and economically significant effect given the base rate of 11%. As column (2) shows, this effect is robust to the inclusion of dummies for age, gender, and the respondent's municipality of residence.

Column (3) presents joint estimates of the effects of mathematical and verbal ability. Because the two are highly related (with a correlation coefficient of 0.46), it is difficult to separately identify the effects of each of these measures. The point estimates do suggest that mathematical ability plays a larger role, although the large standard errors on the effect of verbal ability make it impossible to statistically reject the equality of these coefficients.

Short-Term Time Preference. Results are similar when we examine the relationship between cognitive ability and patience in columns (4) through (6). As column (4) shows, there is an economically large and statistically significant positive effect of mathematical ability on the propensity

[12] Appendix Table 1 shows that our results are similar when we use categorical measures of the number of normative choices made by the subject as the dependent variable.

to display perfect patience. A one-standard-deviation increase in mathematical performance raises the propensity to be patient by 18 percentage points, relative to a base of 28%. As column (5) shows, this effect actually becomes somewhat stronger when we include demographic controls. The model in column (6) estimates the effects of mathematical and verbal ability jointly. As with risk preference, there is insufficient power to distinguish the effects statistically, but the point estimates indicate a much larger effect of mathematical ability than of verbal ability.

Overall, our findings in this population are quite consistent with those from Harvard undergraduates, which suggests that the connection between cognitive ability and normative decision-making is not limited to the very top of the ability distribution. Also consistent with our findings from the Harvard study is a greater effect of mathematical than of verbal ability.

2.3.4 Results: Preferences and Early-life Achievement

Thus far we have shown that individuals with greater measured cognitive ability are more risk-neutral over small-stakes gambles and more patient over short-run trade-offs than those with less cognitive ability. We have not identified, however, whether the relevant differences in cognitive ability arise early in life or rather develop later as a result of differential behaviors or investments. As a step toward answering this question, we ask now whether student achievement in elementary school predicts more normative decision-making among the twelfth-graders in our study.

For each student, we calculate the mean grade point average (GPA) in mathematics and language over all years in elementary school.[13] Among the 84 participants for whom we can make this calculation, the correlation between the average elementary-school GPA in mathematics and our measure of current mathematical ability is 0.65. The analogous correlation between GPA in language and verbal ability is 0.57. Thus early-life achievement is strongly, but not perfectly, related to cognitive ability as measured in grade 12.

Panel B of Table 2 shows the results of probit models of risk-neutral and patient decision-making as a function of elementary-school achievement. We have standardized the independent variables so that a one-unit increase can be interpreted as an increase of one sample standard deviation. As the first column shows, a one-standard-deviation increase in elementary-school mathematics

[13] We follow Wolff, Schiefelbein and Schiefelbein (2002) in defining elementary school to consist of grades 1 through 6. Results are similar when we define elementary school as consisting of grades 1 through 5.

GPA is associated with an 8% greater probability of making risk-neutral choices. This effect is both statistically significant and economically large, although it is somewhat smaller than the estimated effect of cognitive ability estimated in Table 2. Again consistent with the prior evidence, column (2) includes measures of both mathematics and language GPA and suggests that the effects of mathematical achievement are greater than those of achievement in language. Columns (4) and (5) report similar results on the relationship between early-life achievement and impatience: Mathematical achievement has a statistically significant and economically large effect, and the effects of language achievement are smaller and statistically weaker.

Our findings raise the question of whether mathematical ability measured at the end of high school has an effect independent of its correlation with early-life achievement. Though the high degree of collinearity between elementary school grades and our measures of cognitive ability make separate identification difficult, it is possible to make some preliminary statements regarding the relative importance of elementary-school and post-elementary-school achievement. Columns (3) and (6) show estimates from probit models of risk-neutral and patient decision-making on measures of current ability and elementary-school performance. In general, controlling for elementary-school GPA, current ability measures have essentially no effect on decision-making. The point estimates of these effects are close to zero, and the standard errors are small enough to reject any economically meaningful effect. In contrast, elementary-school GPAs are estimated to have a large effect on decision-making in our study, even controlling for current ability levels. These effects are imprecisely estimated, however, and therefore cannot be statistically distinguished from zero. If we take the point estimates at face value, then early-life achievement accounts for essentially all of the effect of cognitive ability on economic decision-making. This may be less surprising in our particular context, since the participants in our study have had essentially identical post-elementary schooling.

Grades in elementary school strongly predict behaviors at the end of secondary school. Moreover, when both elementary-school achievement and current ability are included in a regression, the former carries essentially all of the explanatory power. Though not conclusive, this finding hints that differences in cognitive ability that are apparent early in life may have an important and lasting effect on economic decision-making.

3 Evidence from the field

The evidence presented above begs the question of whether cognitive ability leads to more normative decision-making in real-world contexts. In this section we therefore turn to survey evidence on real-world behavior and ask whether individuals with greater cognitive ability behave in a way more consistent with the predictions of standard economic models.

3.1 Data

The National Longitudinal Survey of Youth 1979 (NLSY) is compiled from repeated interviews of a nationally representative sample of 12,686 Americans. All respondents were between the ages of 14 and 22 in 1979, the first survey year. Interviews were conducted annually through 1994 and biennially thereafter. Although the survey is primarily focused on labor force information, it has also obtained information on respondents' household finances, consumption and savings, pension plans, and a host of other choices and characteristics.

In 1980, 94% of survey respondents were administered the Armed Services Vocational Aptitude Battery (ASVAB), which consists of 10 exams designed to measure different areas of knowledge and ability.[14] On the basis of each respondent's ASVAB results, data processors constructed an approximation to the respondent's percentile in the Armed Forces Qualifying Test (AFQT), a measure developed by the U.S. Department of Defense.[15] Each constructed score was then compared to the overall distribution of scores for respondents age 17 and over to yield a percentile score ranging from 0.01 to 0.99. This percentile score will serve as our primary measure of cognitive ability in this section.

The AFQT score has a correlation of over 0.94 with the first principal component of the scores on the ten sections of the ASVAB, which cognitive psychologists believe to be a good measure of "general cognitive ability." This principal component is in turn highly correlated with SAT scores (Frey and Detterman, 2004), which we use as a measure of cognitive ability in our own laboratory work (see Section 2). Moreover, SAT scores in turn have a very strong correlation with scores on

[14] These areas are: general science, arithmetic reasoning, word knowledge, paragraph comprehension, numerical operations, coding speed, auto and shop information, mathematics knowledge, mechanical comprehension and electronics information. See http://www.goarmy.com/army101/asvabtes.htm for sample questions.

[15] In particular, for each respondent a score was calculated by summing the raw scores from arithmetic reasoning, word knowledge, and paragraph comprehension, plus one-half of the score from the numerical operations exam.

a Raven's Matrices Task, another common tool for measuring general cognitive ability (Frey and Detterman, 2004).

As in our laboratory study, we will be interested in separating the effects of mathematical and verbal ability. We have therefore computed a mathematical ability score as the sum of performance on the arithmetic reasoning, numerical operations, and mathematical knowledge sections of the ASVAB and a verbal ability score as the sum of performance on the word knowledge and paragraph comprehension sections.

Our dependent variables are drawn from various measures of household financial decisions, and are chosen to be related to behaviors that economists have identified as anomalous (see Section 3.2 for details). In particular, we will study:

1. *Financial market participation.* In 1998, respondents were asked, "Not counting any individual retirement accounts (IRA or Keogh) 401K or pre-tax annuities...Do you [or your spouse] have any common stock, preferred stock, stock options, corporate or government bonds, or mutual funds?"

2. *Asset accumulation.* In 1990, 1992, 1993, 1994, 1996, and 1998, respondents were asked, "Suppose you [and your spouse] were to sell all of your major possessions (including your home), turn all of your investments and other assets into cash, and pay all of your debts. Would you have something left over, break even, or be in debt?"

3. *IRA participation.* Beginning in 1994, special efforts were made in the NLSY to track retirement savings. Respondents were asked in the 1994, 1996 and 1998 surveys, "Do you [or your spouse] have any money in individual retirement accounts (IRAs or Keogh)?"

4. *Knowledge about pension plan.* In 1994, respondents participating in an employer pension plan were asked, "We would like to ask about [this plan]. How much do you know about this...pension plan? A lot, something or very little?" In the case of respondents with multiple pension plans, we will use the first pension plan listed at the respondent's primary job.

Since cognitive ability is valuable in the labor market, respondents with higher AFQT scores will tend to have greater human capital wealth. Wealth may directly affect preferences over savings and asset allocation, so we might observe a correlation between some outcome and AFQT scores even absent any direct effects of cognitive ability. We will therefore present specifications that include controls for labor market income. In particular, we will control for the log of family income in every available survey year from 1979 to 1998 (18 years of data in all), with dummies proxying for missing data. Though such controls will not perfectly capture lifetime human capital wealth, the availability of so many years of data allows for much richer specifications than would be possible in purely cross-sectional data.

In some specifications we will also control for household net assets, defined as follows:

$$\begin{aligned} \text{net assets} \quad &= \quad \text{home equity + vehicle equity + liquid savings + CDs} \\ &\quad + \text{other assets - debts outstanding} \end{aligned} \tag{1}$$

We will argue that asset accumulation is itself a function of cognitive ability, but we will include this control to suggest that the other effects we identify exist, even holding constant total asset wealth.

Finally, to account for family background characteristics that may be correlated with cognitive ability, we will take advantage of the fact that many of the respondents in the NLSY are siblings (53.8% have at least one sibling who is also a respondent). We will therefore be able to estimate our models with "sibling group" fixed effects to difference out family-specific factors that might be correlated with both cognitive ability and financial sophistication.

3.2 Results

In this section we present results of probit specifications of the form

$$\Pr(Y_{it} = 1) = \Phi\left[\alpha\left(AFQT_i\right) + X_{it}\beta\right]$$

where Y_{it} is an outcome of interest, X_{it} is a vector of controls, Φ is the normal CDF, i indexes individuals and t indexes years. Since respondents were at different ages when taking the ASVAB, we include dummies for age in all specifications. We also include a dummy for gender in all models.[16]

[16] Controlling for gender does not meaningfully affect our results, since gender and AFQT score are statistically unrelated in our sample. When we split the sample by gender, in most cases the estimated effect of AFQT score is quite similar for male and female respondents.

For all outcomes Y_{it} measured in multiple years, we include dummies for survey year to control for time trends. We also adjust standard errors for within-individual correlation in the error structure whenever we have repeated measures for a given individual. In all cases, we report coefficients as marginal effects evaluated at the sample mean of the data.

Financial market participation. A large economic literature on the equity premium puzzle (Mehra and Prescott, 1985) documents that "even though stocks appear to be an attractive asset—they have high average returns and a low covariance with consumption growth—investors appear very unwilling to hold them" (Barberis and Thaler, 2003). A number of deviations from standard models—most notably myopic loss aversion (Benartzi and Thaler, 1995; Barberis, Huang and Thaler, 2003)—have been proposed to explain this behavior. Since narrow framing of risks is not an element of normative decision theory, we predict that cognitive ability will be associated with greater financial market participation.

Table 3 shows that this prediction is borne out in the data. With no income controls, column (1) shows that respondents at the top of the AFQT distribution are 45 percentage points more likely to own stocks, bonds, or mutual funds than respondents at the bottom of the distribution. As column (2) shows, controlling for 18 years of income data reduces this coefficient to 27 percentage points, but it remains strongly statistically significant. Figure 2 illustrates these findings visually: in general there is a clear and consistent relationship between AFQT decile and the propensity to participate in financial markets.

As column (3) shows, adding a control for net assets does not meaningfully change our estimate of the importance of cognitive ability. In column (4) we go a step further and include fixed effects for the respondent's sibling group. In this specification, all identification is coming from within-family variation in cognitive ability. Although the standard errors increase somewhat because we have taken out much of the variation in AFQT scores, column (4) shows that there is still a significant positive effect of cognitive ability on financial market participation. The estimated magnitude of the effect is comparable in size to the effect estimated without including family fixed effects. The fact that greater cognitive ability is associated with more sophisticated economic choices even among respondents from the same family implies that the relationship is not driven by differences in family background.

The estimates from our laboratory studies suggest that mathematical ability may play a more important role in economic decision-making than verbal ability. To check whether this pattern persists in the NLSY data, in column (5) we report the effect of both math and verbal ability on financial market participation. To make these scores comparable in magnitude, we have expressed them as percentiles of the sample distribution, so that each coefficient can be interpreted as the effect of moving from the bottom to the top of the measured ability distribution. Both mathematical and verbal ability are positively related to financial market participation, and both effects are statistically significant and economically large. Consistent with our laboratory evidence, we find that the effect of mathematical ability is greater than that of verbal ability, although we cannot distinguish the two coefficients statistically.

Asset accumulation. Bernheim (1991) argues that Americans have anomalously low levels of retirement savings. In Table 4 we investigate whether asset accumulation behavior is related to cognitive ability. We code a dummy equal to one if the respondent indicates that she would have money left over if she liquidated all of her assets and paid off her debts.[17] (Results are similar if we construct the answer to this question using our measures of asset positions rather than using the respondent's self-report.) Even controlling for income, there is a large and statistically strong relationship between AFQT score and the propensity to have positive net assets.

Figure 2 presents the relationship between asset accumulation and AFQT decile visually, showing a strong and consistent positive association between the two variables. As column (3) shows, this relationship persists (although it is somewhat weakened) even when we estimate only from variation among siblings. In column (4), we include separate measures of mathematical and verbal ability, and find that the effect of mathematical ability is large, positive and statistically significant, and the effect of verbal ability is small, negative, and statistically insignificant. We can reject that these two coefficients are equal ($p < 0.0001$), indicating that in this case the effect of mathematical ability is statistically more important than the effect of verbal ability.

IRA participation. Some researchers, such as Engen, Gale and Scholz (1994), have noted that actual levels of IRA participation fall short of the levels predicted by calibrated economic models. Table 5 shows the relationship between IRA participation and AFQT scores. The coefficients can

[17] Since one might question the wisdom of significant saving in early adulthood, we have checked that our results are robust to restricting attention to the 1998 survey. In this year respondents were ages 33 to 41.

be interpreted as the estimated marginal effect (evaluated at the sample mean) of moving from the bottom to the top of the distribution of AFQT scores. Column (1) of Table 5 shows that, with no income controls, an individual at the top of the score distribution is 39 percentage points more likely to have an IRA or Keogh account than an individual at the bottom of the score distribution. The effect is strongly statistically significant. This coefficient is reduced to about 20 percentage points when we include income controls in column (2). As column (3) shows, including a control for net assets does not significantly affect the coefficient of interest.

These results suggest that there is a robust positive relationship between AFQT scores and IRA participation. The effects are economically large: a movement of 10 percentile points in AFQT score is associated with an increase of about 2 percentage points in IRA participation, or about 10% of the baseline participation rate of 18%. As figure 2 demonstrates, this effect is present throughout the distribution of AFQT scores. Column (4) shows that the effect of cognitive ability arises within as well as between families, although the coefficient estimated with sibling group fixed effects is somewhat smaller than the estimate without these controls. Finally, column (5) shows that the estimated effect of mathematical ability is slightly larger than that of verbal ability. Although this difference is not statistically significant, it supports the view that mathematical skill is the more relevant component of cognitive ability for the economic choices we analyze.

Knowledge about pension. There is considerable evidence that employees are not well-informed about their retirement plans and that this misinformation can affect behavior in important ways. Mitchell (1988) matched employees' self-reported pension plan information with firm records using the 1983 Survey of Consumer Finances. She found widespread "missing and misinformation" among employees about the details of their plans. Gustman and Steinmeier (2001) used the Social Security earnings and employer pension records of Health and Retirement Survey participants to compare self-reported expectations to calculations based on administrative records. They found significant misinformation, as well as some evidence of a connection between poor information and retirement behavior. Chan and Stevens (2003) showed more direct evidence that behavior responds to mistaken perceptions of the financial incentives to retire. They also showed that more knowledgeable individuals are significantly more responsive to correctly-calculated incentives.

Table 6 addresses the question of whether higher cognitive ability is associated with less ignorance about pension plans. The dependent variable is a dummy equal to one when the respondent indicates that she knows “very little” about her primary pension plan (about one-third of respondents indicate very little knowledge). The results in the table demonstrate an economically and statistically significant relationship between pension knowledge and AFQT scores. As in the other contexts we examine, the result is robust to income and asset controls. Because of a small sample size, our estimate becomes imprecise and statistically insignificant when we control for family fixed effects, though the point estimate has the correct sign and a similar magnitude to our other estimates. Additionally, as with asset accumulation, only mathematical ability has a statistically significant effect on pension knowledge, and there is a statistically significant difference between the effects of mathematical and verbal ability ($p = 0.0015$). While a more objective determination of knowledge would be preferable to a subjective self-evaluation, these results seem to suggest that individuals with greater cognitive ability are better informed about at least some of their financial opportunities.

Overall, then, we find that a number of anomalous behaviors often attributed to deviations from standard decision theory are more significant for individuals with lower cognitive ability. In all of the cases we examine, the estimated effect of cognitive ability is economically large and statistically robust to including detailed controls for income over the entire survey period of 1979-1998. In general, these effects remain strong even when we take identification only from differences among siblings, indicating that family background does not drive the effects. Finally, we estimate that mathematical ability has a larger impact than verbal ability, and in two instances we can strongly reject the equality of these two coefficients. These results suggest that the patterns identified in the laboratory do indeed generalize to market contexts, and that individuals with greater cognitive ability are better able to approximate the behavioral prescriptions of economic models.

4 Can Human Capital Policy Affect Sophistication?

The results presented in the previous two sections demonstrate a strong and robust positive relationship between cognitive ability and sophistication in economic decision-making. Given recent interest in policy approaches to address psychological biases (Benjamin and Laibson, 2003; Jolls, Sunstein and Thaler, 1998; Camerer, Issacharoff, Loewenstein, O’Donoghue, and Rabin, 2003), it is natural to ask whether human capital policy—by contributing to the development of cognitive ability—may have the effect of improving economic decision-making.

By construction, our laboratory data are poorly suited to answer this question. In both studies, our populations have had similar educational experiences; indeed, the Chilean high school students who participated in the second study have had identical schooling for most of their lives. We will therefore take advantage of measures of schooling quality available through the NLSY.

NLSY surveyors contacted the last school attended by respondents and obtained a set of basic administrative data about the institution. Variables collected include information on the race and gender of students and teachers, teachers' qualifications and turnover rates, the availability of a set of vocational curricula at the school, the number of faculty members, the number of library books, and total enrollment.

Because we want to estimate the effects of human capital policy, we ignore demographic characteristics such as the race and gender of respondents' schoolmates and focus on the parameters most directly under policy control:

1. The number of library books per student.
2. The number of full-time-equivalent teachers per student.
3. The percent of teachers with Master's degrees or PhD's.
4. The percent of teachers who left the school during the 1978-79 school year for reasons other than death or retirement.
5. The number of occupational training programs offered by the school among the following seven categories: agricultural, business/office, distributive education, health, home economics, trade/industry, technical.

We will use these measures as instruments for cognitive ability (as measured by AFQT score) to determine whether the component of cognitive ability that is affected by human capital policy has an effect on financial sophistication. Measures of this sort have frequently been treated as exogenous in the study of wages and rates of return to schooling (see, e.g., Card and Krueger, 1992), but we wish to stress that we consider this exercise only a first step in exploring the effects of human capital policy on economic decision-making.

Appendix Table 3 presents the results of a regression of AFQT scores on these measures of school quality. In general we find effects with reasonable magnitudes in the expected direction. Statistically, the effects of library books per capita, the teacher-student ratio, and the share of teachers with advanced degrees are quite strong. The effect of teacher turnover is statistically weaker but still significant, and the effect of occupational programs is statistically indistinguishable from zero. The F-statistic on the null hypothesis that the instruments exert no effect is 29.15, large enough to reject the presence of a weak instruments problem with high confidence (Stock and Yogo, 2002).[18]

Table 7 reports the results of two-stage least squares (2SLS) estimation using the model in Appendix Table 3 as a first stage to predict AFQT scores. Panel (A) of the table presents estimates of the effect of school quality on financial market participation. As column (1) shows, with no controls there is a positive and statistically significant effect of AFQT score. As columns (2) and (3) show, the estimate remains positive and statistically significant when we control for income and net assets. The coefficient is also economically large: an increase of 10 percentile points in AFQT is estimated to increase participation by six percentage points. Column (4) shows, however, that when we include family fixed effects, our standard errors become too large to permit meaningful inferences. Given that siblings presumably attend the same or similar schools, it may not be too surprising that there is not enough within-family variation in our schooling quality measures to identify the effect of AFQT score on financial market participation.

Since we have several different measures of school quality, we can perform a test of overidentifying restrictions to check the validity of the instruments. In the case of financial market participation, our test, which uses Hansen's J statistic (Hansen, 1982; Hoxby and Paserman, 1998; Baum, Schaffer and Stillman, 2002), cannot reject the null hypothesis that the instruments are uncorrelated with the error term.

In Panels (B)-(D) of Table 7, we report the effects of school quality on the other financial

[18] As a second check on weak instruments concerns, we have re-estimated our models using Limited Information Maximum Likelihood (LIML) and constructed confidence intervals using the conditional likelihood ratio approach proposed by Moreira (2003). The results are quite similar.

decision-making outcomes. For asset accumulation and IRA participation, we again find large and statistically significant effects of AFQT score. There is some evidence of within-family effects, although the standard errors are quite large. Because pension knowledge is available for only a limited subset of respondents, our estimates of the effect of school quality on pension knowledge are statistically imprecise. In a number of specifications, our test of overidentifying restrictions rejects the null hypothesis, which suggests that caution is needed in interpreting the magnitudes of the estimated effects. Despite these limitations, however, the evidence in Table 7 seems to suggest that human capital policy may indeed be able to improve economic decision-making.

5 Conclusions

The findings presented in this paper suggest a substantial return to cognitive skill in economic decision-making. Evidence from a laboratory study indicates that higher cognitive ability is associated with less short-run discounting and lower levels of small-stakes risk-aversion. Results from the NLSY show that individuals with greater cognitive ability make consumption and personal financial decisions that more nearly approximate the normative implications of economic models, even among siblings. While each of the empirical methodologies we employ has its limitations, we believe that the confluence of evidence from a range of contexts and methodologies provides a very strong case for a robust positive relationship between cognitive ability and normative decision-making.[19] These findings are consistent with the idea that cognitive constraints prevent people from making decisions in a manner consistent with standard economic models.[20]

One speculative interpretation of our results builds on recent research that posits two distinct cognitive systems (e.g., Kahneman 2003; Loewenstein and O'Donoghue, 2004; Fudenberg and Levine, 2004; Stanovich and West, 2000; McClure, Laibson, Loewenstein, and Cohen, 2004). Ac-

[19] Interestingly, however, there appears to be evidence that the behavior of cognitively deficient individuals may in some contexts conform *more* closely with standard economic predictions. For example, patients with orbitofrontal cortical damage seem to be immune to ambiguity aversion (the Ellsberg paradox), and autistic children typically make very small offers in Ultimatum Games (Camerer, personal communication, 2004). Documenting cases like these may help to elucidate more precisely how people without such deficits make economic decisions and at what point in the process cognitive ability plays a role.

[20] We note, however, that even the most skilled individuals still exhibit substantial psychological biases. This fact suggests that institutions (private or public) designed to provide advice on economic decision-making cannot simply rely on the most cognitively able to determine optimal choices.

cording to this framework, an automatic system makes quick, intuitive judgments. With effort, a much slower deliberative system can reason and override the intuitive decisions.[21] On this interpretation of our evidence, "cognitive ability" measures the ease or frequency with which the deliberative system overrides the automatic system.[22]

Our results suggest the existence of substantial non-labor-market returns to cognitive ability.[23] For example, we calculate that, in a portfolio choice problem, a typical investor would be willing to give up about 5% of lifetime wealth in order to avoid having her investment decisions made in accordance with myopic loss-aversion.[24] Evidence presented in Appendix Table 2 suggests that an increase of one standard deviation in measured cognitive ability corresponds to a 10 percentage point decrease in the probability of loss aversion. Hence, we might conjecture that a one-standard-deviation increase in cognitive ability is worth about 0.5% of lifetime wealth due to improved portfolio allocation alone.[25] Since portfolio choice is only one of many important household decisions that are affected by cognitive ability, the total value of cognitive ability's effect on decision-making could be quite substantial. Such calibrations also suggest that our results may have potentially important policy implications. To the extent that education can increase cognitive ability (Cascio and Lewis, 2005), human capital policy may be an important tool for addressing biases in decision-making in a wide range of contexts.

[21] The fact that math ability most strongly predicts normative decision-making might suggest that the more cognitively skilled take attractive small-stakes gambles because they know how to compute expected values, and they participate in the stock market because they understand compound interest. In its simplest form, however, this hypothesis cannot account for the relationship between SAT scores and normative decision-making among Harvard undergraduates. Virtually all Harvard undergraduates, even the lowest SAT scorers, can calculate expected values in the very simple gambles we offered them. The dual-systems interpretation we discuss can be understood as a more nuanced version of this hypothesis, in which more cognitively skilled individuals are better able to *apply* rules like expected value and compound interest when facing decision problems.

[22] Ongoing work by Shane Frederick (personal communication, 2004) promises to shed light on which types of time-preference-related decisions are most strongly predicted by cognitive ability.

[23] See also Haveman and Wolfe (1984), who estimate non-labor market returns to years of schooling.

[24] We assume a constant coefficient of relative risk aversion $\rho = 5$, an exponential discount rate $\gamma = 0.08$, and log-normal portfolio returns in an infinite-horizon model. We use Campbell and Viceira's (2002, p. 104) estimates of equity and bond returns and Benartzi and Thaler's (1995) calculation that a loss-averse investor would hold around 40% equities. Details are available from the authors upon request.

[25] For comparison, Cawley, Heckman, and Vytlacil (2001) estimate that a one-standard-deviation increase in cognitive ability corresponds to an increase in wages of 10-16%.

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Table 1 *Preferences and Cognitive Ability: Harvard Undergraduates*

Panel A: Whole sample

Dependent variable	Risk neutral (dummy)			Patient (dummy)		
	(1)	(2)	(3)	(4)	(5)	(6)
Math SAT $\geq$ median	0.2352	0.2062	0.2776	0.1385	0.2260	0.1415
	(0.1166)	(0.1214)	(0.1223)	(0.1384)	(0.1580)	(0.1429)
Verbal SAT $\geq$ median			-0.1367			-0.0125
			(0.1290)			(0.1456)
Demographics?	No	Yes	No	No	Yes	No
Mean of dependent variable	0.2982	0.2982	0.2982	0.5294	0.5294	0.5294
Pseudo-R^2	0.0555	0.1012	0.0716	0.0140	0.0997	0.0141
N	57	57	57	51	51	51

Panel B: Sample accepting some delay in payment

Dependent variable	Patient (dummy)		
	(7)	(8)	(9)
Math SAT $\geq$ median	0.2527	0.3249	0.2469
	(0.1388)	(0.1492)	(0.1420)
Verbal SAT $\geq$ median			0.0314
			(0.1509)
Demographics?	No	Yes	No
Mean of dependent variable	0.5745	0.5745	0.5745
Pseudo-R^2	0.0482	0.1275	0.0489
N	47	47	47

Notes: Standard errors in parentheses. Data are from laboratory study of Harvard undergraduates. Risk neutral indicates that participant made six decisions of the following form in a way consistent with risk-neutral expected-value maximization:

Please circle either Choice A or Choice B.

(A) You get $0.50 for sure.

(B) If the die comes up 1, 2, or 3, you get X. If the die comes up 4, 5, or 6, you get nothing.

Patient indicates that participants made six decisions of the following form in a way consistent with the maximization of undiscounted wealth:

Please circle either Choice A or Choice B.

(A) You get $5.00 right now.

(B) You get X a week from now.

Data on SAT scores based on participants' self-reports. Median refers to sample median Demographics include dummies for gender and year in school.

Table 2 *Preferences and Cognitive Ability: Chilean High School Students*

Panel A: Current cognitive ability

Dependent variable	Risk neutral (dummy)			Patient (dummy)		
	(1)	(2)	(3)	(4)	(5)	(6)
Standardized math score	0.1358	0.1250	0.1237	0.1804	0.2643	0.1665
	(0.0489)	(0.0564)	(0.0535)	(0.0831)	(0.1027)	(0.0933)
Standardized verbal score			0.0326			0.0389
			(0.0710)			(0.1238)
Demographics?	No	Yes	No	No	Yes	No
Mean of dependent variable	0.1087	0.1087	0.1087	0.2826	0.2826	0.2826
Pseudo-R^2	0.1480	0.2006	0.1513	0.0444	0.1148	0.0453
N	92	92	92	92	92	92

Panel B: Early life achievement

Dependent variable	Risk neutral (dummy)			Patient (dummy)		
	(1)	(2)	(3)	(4)	(5)	(6)
Standardized math GPA	0.0854	0.0816	0.0540	0.1481	0.1371	0.1218
(elementary school)	(0.0326)	(0.0416)	(0.0423)	(0.0560)	(0.0711)	(0.0713)
Standardized language GPA		0.0053			0.0154	
(elementary school)		(0.0363)			(0.0621)	
Standardized math score			0.0006			0.0006
(current)			(0.0006)			(0.0011)
Mean of dependent variable	0.1059	0.1059	0.1059	0.2824	0.2824	0.2824
Pseudo-R^2	0.1171	0.1175	0.1391	0.0742	0.0748	0.0777
N	85	85	85	85	85	85

Notes: Standard errors in parentheses. Data are from laboratory study of Chilean high school students in grade 12. Demographics include dummies for gender, age, and municipality of residence. Risk neutral indicates that participant made six decisions of the following form in a way consistent with risk-neutral expected-value maximization:

Please circle either Choice A or Choice B.

(A) You get 250 pesos for sure.

(B) If the die comes up 1, 2, or 3, you get X. If the die comes up 4, 5, or 6, you get nothing.

Patient indicates that participants made six decisions of the following form in a way consistent with the maximization of undiscounted wealth:

Please circle either Choice A or Choice B.

(A) You get 500 pesos right now.

(B) You get X a week from now.

Test scores standardized by the population standard deviation. Average GPA in elementary school (grades 1-6) standardized by sample standard deviation.

Table 3 *Financial Market Participation and Cognitive Ability*

Dependent variable: Does respondent/spouse have any stocks, bonds, or mutual funds outside of retirement accounts?
(About 19% say yes)

	(1)	(2)	(3)	(4)	(5)
AFQT	0.4511 (0.0143)	0.2692 (0.0161)	0.2640 (0.0164)	0.2742 (0.0762)	
Math score percentile					0.1750 (0.0233)
Verbal score percentile					0.1269 (0.0227)
log(income) 1979-1998?	No	Yes	Yes	Yes	Yes
Net assets?	No	No	Yes	No	No
Sibling group?	No	No	No	Yes	No
Model	Probit	Probit	Probit	FE	Probit
Pseudo-R^2	0.1354	0.2036	0.2138	—	0.2061
N	7955	7955	7955	7955	7955

Notes: Data from NLSY. Standard errors in parentheses are heteroskedasticity-robust. Coefficients in probit models are marginal effects evaluated at the sample mean of the independent variables. AFQT score is a percentile ranging from 0.01 to 0.99. Math score is the sum of performance on the arithmetic reasoning, numerical operations, and mathematical knowledge sections of the ASVAB, expressed as a percentile in sample distribution. Verbal score is the sum of performance on the word knowledge and paragraph comprehension sections of the ASVAB, expressed as a percentile in sample distribution. All specifications include a dummy for gender and dummies for age in 1979. Question asked only in 1998. Income controls include controls for log of family income for all available years of data, 1979-1998, with dummies for missing values. Net asset control is a linear term for net assets in the current survey year, calculated as in equation (1), with a dummy for missing data.

Table 4 *Asset Accumulation and Cognitive Ability*

Dependent variable: Would respondent have something left over after liquidating all assets and paying off all debts?
(About 66% would have something left over.)

	(1)	(2)	(3)	(4)
AFQT	0.5494	0.2546	0.1639	
	(0.0138)	(0.0155)	(0.0276)	
Math score percentile				0.1782 (0.0375)
Verbal score percentile				-0.0030 (0.0032)
log(income) 1979-1998?	No	Yes	Yes	Yes
Sibling group?	No	No	Yes	No
Model	Probit	Probit	FE	Probit
Pseudo-R^2	0.0780	0.1418	—	0.1428
N	51575	51575	51575	51575

Notes: Data from NLSY. Standard errors in parentheses are clustered by individual whenever multiple observations per individual are used. Coefficients in probit models are marginal effects evaluated at the sample mean of the independent variables. AFQT score is a percentile ranging from 0.01 to 0.99. Math score is the sum of performance on the arithmetic reasoning, numerical operations, and mathematical knowledge sections of the ASVAB, expressed as a percentile in sample distribution. Verbal score is the sum of performance on the word knowledge and paragraph comprehension sections of the ASVAB, expressed as a percentile in sample distribution. All specifications include a dummy for gender, dummies for age in 1979, and dummies for survey year (whenever multiple years of data are pooled). Question asked in 1990, 1992, 1993, 1994, 1996, and 1998. Income controls include controls for log of family income for all available years of data, 1979-1998, with dummies for missing values.

Table 5 *IRA Participation and Cognitive Ability*

Dependent variable: Does respondent/spouse have an IRA or Keogh account?
(About 18% say yes)

	(1)	(2)	(3)	(4)	(5)
AFQT	0.3869 (0.0115)	0.1986 (0.0129)	0.1967 (0.0130)	0.1143 (0.0354)	
Math score percentile					0.1334 (0.0182)
Verbal score percentile					0.0926 (0.0178)
log(income) 1979-1998?	No	Yes	Yes	Yes	Yes
Net assets?	No	No	Yes	No	No
Sibling group?	No	No	No	Yes	No
Model	Probit	Probit	Probit	FE	Probit
Pseudo-R^2	0.1050	0.1711	0.1775	—	0.1733
N	24587	24587	24587	24587	24587

Notes: Data from NLSY. Standard errors in parentheses are clustered by individual. Coefficients in probit models are marginal effects evaluated at the sample mean of the independent variables. AFQT score is a percentile ranging from 0.01 to 0.99. Math score is the sum of performance on the arithmetic reasoning, numerical operations, and mathematical knowledge sections of the ASVAB, expressed as a percentile in sample distribution. Verbal score is the sum of performance on the word knowledge and paragraph comprehension sections of the ASVAB, expressed as a percentile in sample distribution. All specifications include a dummy for gender, dummies for age in 1979, and dummies for survey year. Question asked in 1994, 1996, and 1998. Income controls include controls for log of family income for all available years of data, 1979-1998, with dummies for missing values. Net asset control is a linear term for net assets in the current survey year, calculated as in equation (1), with a dummy for missing data.

Table 6 *Knowledge about Pension and Cognitive Ability*

Dependent variable: Respondent claims to know "very little" about pension plan.
(About 34% know very little)

	(1)	(2)	(3)	(4)	(5)
AFQT	-0.2443 (0.0331)	-0.1470 (0.0376)	-0.1472 (0.0377)	-0.1054 (0.2930)	
Math score percentile					-0.3285 (0.1077)
Verbal score percentile					0.0165 (0.0913)
log(income) 1979-1998?	No	Yes	Yes	Yes	Yes
Net assets?	No	No	Yes	No	No
Sibling group?	No	No	No	Yes	No
Model	Probit	Probit	Probit	FE	Probit
Pseudo-R^2	0.0307	0.0522	0.0522	—	0.0556
N	2755	2755	2755	2755	2755

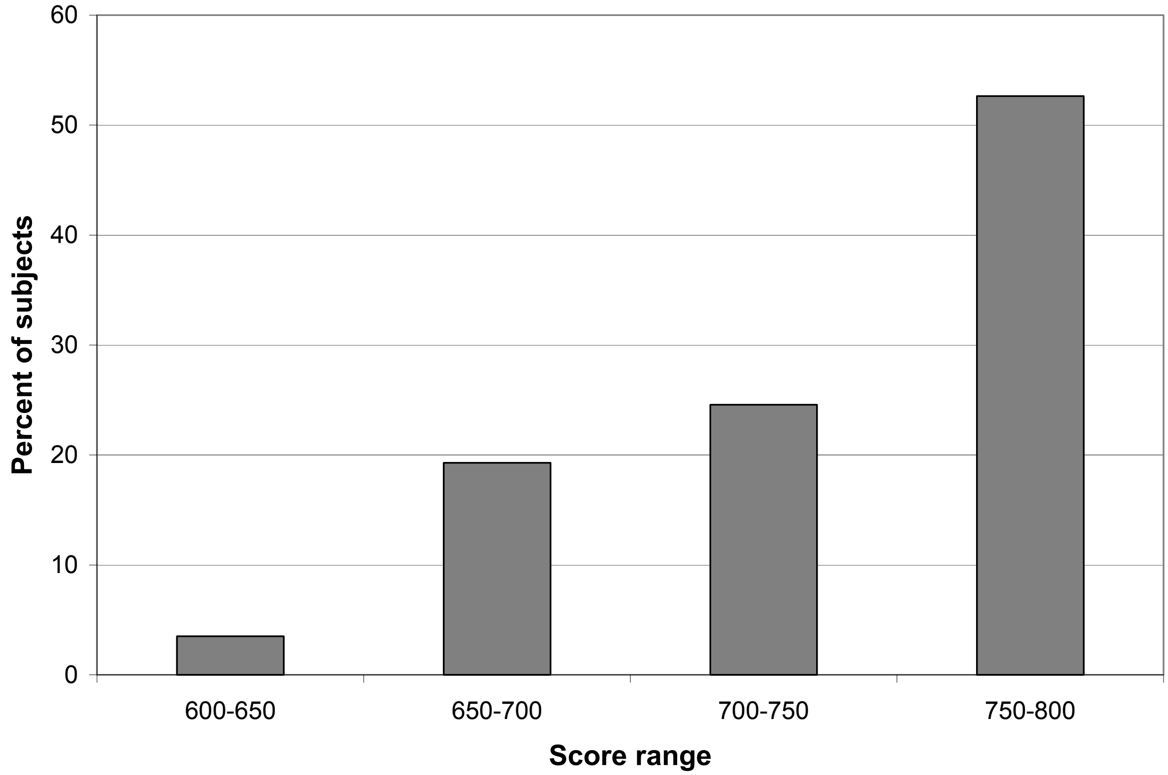
Notes: Data from NLSY. Standard errors in parentheses are heteroskedasticity-robust. Coefficients in probit models are marginal effects evaluated at the sample mean of the independent variables. AFQT score is a percentile ranging from 0.01 to 0.99. Math score is the sum of performance on the arithmetic reasoning, numerical operations, and mathematical knowledge sections of the ASVAB, expressed as a percentile in sample distribution. Verbal score is the sum of performance on the word knowledge and paragraph comprehension sections of the ASVAB, expressed as a percentile in sample distribution. All specifications include a dummy for gender and dummies for age in 1979. Question asked only in 1994. Income controls include controls for log of family income for all available years of data, 1979-1998, with dummies for missing values. Net asset control is a linear term for net assets in the current survey year, calculated as in equation (1), with a dummy for missing data.

Table 7 *Financial Sophistication and Human Capital Policy*

	(1)	(2)	(3)	(4)
(A) Does respondent/spouse have stocks, bonds, or mutual funds?				
AFQT	0.7970	0.6822	0.6278	2.9219
	(0.1172)	(0.1820)	(0.1833)	(1.8629)
Hansen's J-test (p-value)	0.2105	0.5453	0.7114	
N	5775	5775	5775	5775
(B) Does respondent have positive net assets?				
AFQT	0.7262	0.5300	—	1.0546
	(0.0973)	(0.1413)		(0.4103)
Hansen's J-test (p-value)	0.0008	0.0001		
N	37394	37394		37394
(C) Does respondent/spouse have an IRA or Keogh account?				
AFQT	0.4566	0.2774	0.2184	0.4175
	(0.0965)	(0.1461)	(0.1457)	(0.4565)
Hansen's J-test (p-value)	0.2540	0.0232	0.0178	
N	17807	17807	17807	17807
(D) Does respondent claim to know "very little" about pension plan?				
AFQT	-0.1643	0.0342	0.0332	0.8230
	(0.2774)	(0.4286)	(0.4279)	(1.6443)
Hansen's J-test (p-value)	0.9610	0.9748	0.9758	
N	2130	2130	2130	2130
log(income) 1979-1998?	No	Yes	Yes	Yes
Net assets?	No	No	Yes	No
Sibling group dummies?	No	No	No	Yes

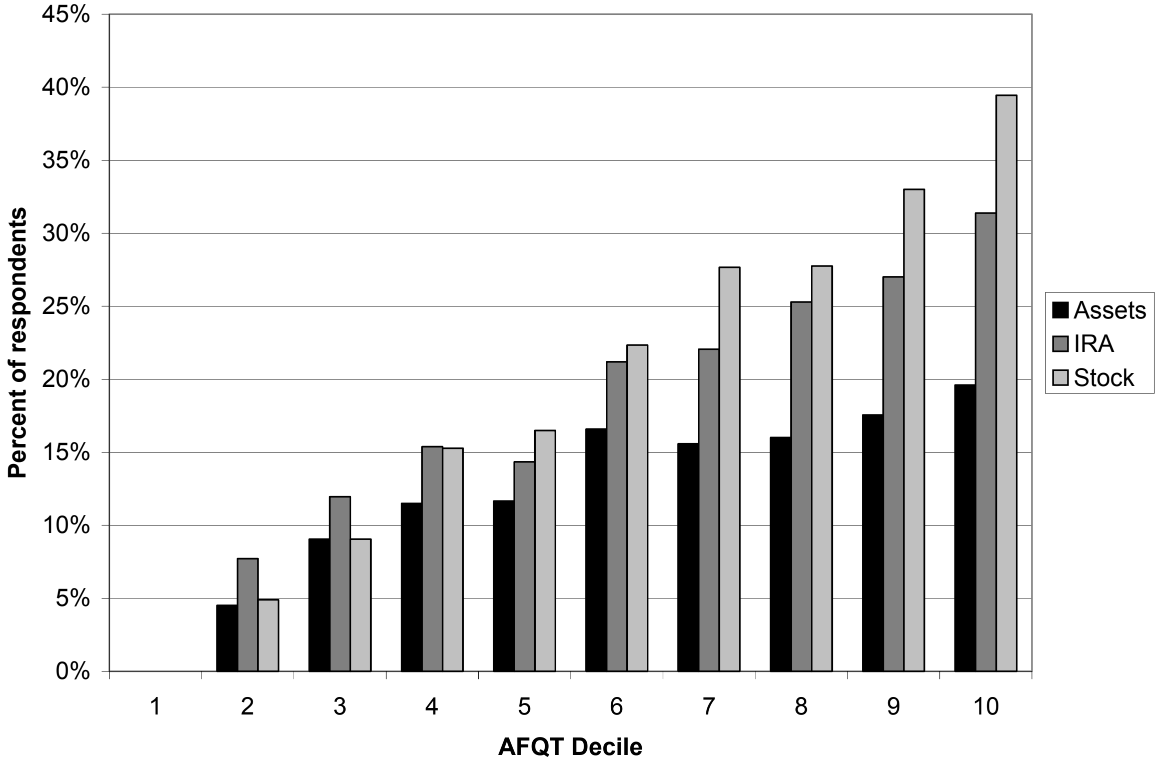
Notes: Data from NLSY. AFQT score is a percentile ranging from 0.01 to 0.99. Estimates formed using two-stage least squares. In columns (1)-(3), standard errors in parentheses are heteroskedasticity-robust and clustered by individual where multiple observations per individual have been used. Instruments for AFQT are log(number of library books per student), log(number of teachers per student), percent of teachers with advanced degrees, percent of teachers leaving in 1978-79, and number of occupational training programs offered, with dummies for missing values. Respondents with no valid data on instruments are excluded. All models include dummies for respondent age and gender. Income controls include controls for log of family income for all available years of data, 1979-1998, with dummies for missing values. Net asset control is a linear term for net assets in current survey year, calculated as in equation (1), with dummy for missing data.

Figure 1 *Distribution of Participants' Math SAT Scores*



Notes: Data are from laboratory study of Harvard undergraduates. Data on SAT-scores based on participants' self-reports.

Figure 2 *Financial Sophistication and Cognitive Ability*



Notes: Data from NLSY. Figure reflects marginal effects from a probit model, including dummies for age and gender and controls for the log of family income for all survey years 1979-1998, with dummies for missing values. Percentage for bottom AFQT decile is normalized to 0. Dependent variable for IRA bars is a dummy indicating that the respondent answered "yes" to the question "Not counting any individual retirement accounts (IRA or Keogh) 401K or pre-tax annuities...Do you [or your spouse] have any common stock, preferred stock, stock options, corporate or government bonds, or mutual funds?" (Question asked in 1998.) Dependent variable for Assets bars is a dummy indicating that the respondent answered "yes" to the question "Suppose you [and your spouse] were to sell all of your major possessions (including your home), turn all of your investments and other assets into cash, and pay all of your debts. Would you have something left over, break even, or be in debt?" (Question asked in 1990, 1992, 1993, 1994, 1996 and 1998.) Dependent variable for Stocks bars a dummy indicating that the respondent answered "yes" to the question "Do you [or your spouse] have any money in individual retirement accounts (IRAs or Keogh)?" (Question asked in 1994, 1996 and 1998.) In lowest decile, 41% of respondents have positive net assets, 3.4% own stocks, and 3.5% have an IRA or Keogh.

Appendix Table 1 *Preferences and Cognitive Ability: Ordered Probit Estimates*

Dependent variable	Number of risk-neutral choices			Number of patient choices		
	(1)	(2)	(3)	(4)	(5)	(6)
Parameter estimates:						
Standardized math score	0.5355	0.7059	0.4138	0.3720	0.6344	0.2577
	(0.2086)	(0.2533)	(0.2305)	(0.2008)	(0.2428)	(0.2260)
Standardized verbal score			0.3468			0.3030
			(0.2888)			(0.2855)
Average marginal effects:						
Standardized math score	0.5055	0.6461	0.3868	0.4594	0.7458	0.3157
Standardized verbal score			0.3242			0.3712
Demographics?	No	Yes	No	No	Yes	No
Mean of dependent variable	3.4565	3.4565	3.4565	4.5000	4.5000	4.5000
N	92	92	92	92	92	92

Notes: Data are from laboratory study of Chilean high school students in grade 12. Demographics include dummies for gender, age, and municipality of residence. Number of risk neutral choices counts the number of choices of the following form made in a way consistent with risk-neutral expected-value maximization:

Please circle either Choice A or Choice B.

(A) You get 250 pesos for sure.

(B) If the die comes up 1, 2, or 3, you get X. If the die comes up 4, 5, or 6, you get nothing.

Number of patient choices counts the number of decisions of the following form made in a way consistent with the maximization of undiscounted wealth:

Please circle either Choice A or Choice B.

(A) You get 500 pesos right now.

(B) You get X a week from now.

Test scores standardized by the population standard deviation.

Appendix Table 2 *Loss Aversion and Cognitive Ability*

Dependent variable	Loss neutral			Number of loss-neutral choices		
Model	Probit			Ordered probit		
	(1)	(2)	(3)	(4)	(5)	(6)
Parameter estimates:						
Standardized math score	0.1013	0.0907	0.1184	0.3821	0.4224	0.4218
	(0.0490)	(0.0587)	(0.0575)	(0.2001)	(0.2227)	(0.2248)
Standardized verbal score			-0.0505			-0.1123
			(0.0777)			(0.2888)
Average marginal effects:						
Standardized math score	—	—	—	0.2930	0.2779	0.3148
Standardized verbal score		—	—			-0.0594
Demographics?	No	Yes	No	No	Yes	No
Mean of dependent variable	0.1087	0.1087	0.1087	3.6087	3.6087	3.6087
N	92	92	92	92	92	92

Notes: Data are from laboratory study of Chilean high school students in grade 12. Demographics include dummies for gender, age, and municipality of residence. Number of loss neutral choices counts the number of choices of the following form made in a way consistent with risk-neutral expected-value maximization:

Please circle either Choice A or Choice B.

(A) You get 0 pesos for sure.

(B) If the die comes up 1, 2, or 3, you get X. If the die comes up 4, 5, or 6, you lose 250 pesos.

Loss neutrality indicates that all six choices were made in a way consistent with risk-neutral expected-value maximization. Test scores standardized by the population standard deviation.

Appendix Table 3 *School Quality and Cognitive Ability*

	AFQT
log(number of library books per student)	0.0584 (0.0065)
log(number of full-time equivalent teachers per student)	0.0597 (0.0126)
Share of teachers with advanced degrees	0.0652 (0.0143)
Share of teachers leaving school in 1978-79	-0.1075 (0.0388)
Number of occupational programs available	-0.0012 (0.0038)
R^2	0.0714
N	7911

Notes: Data from NLSY. AFQT score is a percentile ranging from 0.01 to 0.99. Regression includes dummies for missing values of all measures of school quality and for respondent age. Respondents with no valid data on school quality excluded from regression.