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Research Article

Causal Effects of Schooling on Memory at Older Ages in Six Low- and Middle-Income Countries: Nonparametric Evidence With Harmonized Datasets

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Abstract

Objectives: Higher schooling attainment is associated with better cognitive function at older ages, but it remains unclear whether the relationship is causal. We estimated causal effects of schooling on performances on the Consortium to Establish a Registry for Alzheimer's Disease (CERAD) word-recall (memory) test at older ages in China, Ghana, India, Mexico, Russia, and South Africa.

Methods: We used harmonized data (n = 30,896) on older adults (≥50 years) from the World Health Organization Study on Global Ageing and Adult Health. We applied an established nonparametric partial identification approach that bounds causal effects of increasing schooling attainment at different parts of the schooling distributions under relatively weak assumptions.

Results: An additional year of schooling increased word-recall scores by between 0.01 and 0.13 *SD*s in China, 0.01 and 0.06 *SD*s in Ghana, 0.02 and 0.09 *SD*s in India, 0.02 and 0.12 *SD*s in Mexico, and 0 and 0.07 *SD*s in South Africa when increasing schooling from never attended to primary. No results were obtained for Russia at this margin due to the low proportion of older adults with primary schooling or lower. At higher parts of the schooling distributions (e.g., high school or university completion), the bounds cannot statistically reject null effects.

Discussion: Our results indicate that increasing schooling from never attended to primary had long-lasting effects on memory decades later in life for older adults in 5 diverse low- and middle-income countries.

Keywords: Bounds, Causal inference, Cognition, Partial identification

Older age is the leading risk factor for Alzheimer's disease and related dementias, and aging populations are rapidly expanding globally. Thus, tackling dementia is a pressing challenge for initiatives focused on global population health. Policies to increase schooling attainment could be effective long-term primary prevention strategies or part of multipronged approaches that could help mitigate dementia risk. The 2020 Lancet Commission on Dementia Prevention, Intervention and Care (Livingston et al., 2020) estimated 7% of global dementia cases were attributable to low schooling attainment. Research in the United States (Hudomiet et al., 2022) has also found increases in schooling attainment are significantly associated with trends in lowering dementia prevalence.

One channel through which schooling might reduce dementia risk is by improving cognition at older ages. A burgeoning literature has shown that more schooling is associated with better performance on cognition tests at older ages. A meta-analysis of 53 observational studies (Opdebeeck et al., 2016) for older adults (age > 60 years) in high-income countries (HICs) estimated a correlation of 0.30 between schooling and global cognition. Studies have also found positive associations between schooling and later-life cognition in several low- and middle-income countries (LMICs, e.g., Gonçalves et al., 2023; Kobayashi et al., 2019, 2021; Kohler et al., 2023; Prynn et al., 2023; Trani et al., 2022; Westrick et al., 2024; Zhang et al., 2024). A meta-analysis of 92 studies by Seblova et al. (2020) found the associations between schooling and changes in cognitive performance, however, were small in magnitude and not statistically significant. Even if schooling does not protect against cognitive decline, higher

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baseline cognition due to more schooling is associated with later ages at dementia onset. This is because higher baseline cognition implies greater cognitive reserve which in turn protects against cognitive decline (Stern et al., 2023). The positive relationship between schooling and cognitive performance at older ages suggests lower likelihoods of dementia diagnosis among individuals with more schooling. Furthermore, even small delays in age at dementia onset can lead to large reductions in disease burdens (Brookmeyer et al., 1998).

Relationships between schooling and older-age cognition may be causal, with more schooling improving cognition by increasing cognitive reserve (Stern et al., 2023) and socioeconomic outcomes (Langa, 2018). Alternatively, these relationships may be noncausal, driven by other factors (e.g., genetics, early-life nutrition, family background) that are correlated with both schooling and older-age cognition.

To identify the causal relationship, several studies have used changes in compulsory schooling laws as natural experiments within the instrumental variables (IVs) framework and found positive effects of schooling on memory test scores. An extra year of schooling increased memory scores by 0.35 *SD*s in the United States (Glymour et al., 2008), 0.42–0.50 *SD*s in England (Banks & Mazzonna, 2012; Gorman, 2023), and 0.10–0.16 *SD*s in Europe (Crespo, 2014; Mazzonna, 2014; Schneeweis et al., 2014). Other studies have employed within-sibling comparisons with genetic (polygenic scores) controls for schooling and cognition. An extra year of schooling increased immediate memory and verbal fluency scores by 0.07 *SD*s in the United States (Herd & Sicinski, 2022) and fluid intelligence scores by 0.12 *SD*s in the United Kingdom (Fletcher et al., 2021). Another study (Clouston et al., 2012) using propensity score matching found that fluid intelligence scores of individuals with a university education were 0.17–0.55 *SD*s higher than individuals with secondary schooling in the United States and United Kingdom.

Overall, evidence on causal relationships between schooling and older-age cognition is limited, especially for LMICs. This is a major gap because LMICs are projected to have the largest increase in dementia cases in coming decades and have also rapidly increased schooling attainment. If schooling causally improves cognition, effects on lowering dementia risk could be particularly pronounced in more-recent birth cohorts with substantially higher schooling attainment.

Quasi-experimental estimates for HICs may not be policy-relevant for LMICs. Estimates using compulsory schooling laws in HICs only provide local average-treatment effects (LATEs) for the proportionately small subset of individuals who were affected by these laws, which mostly target high school completion. This may not be informative about the effects of completing primary schooling, which is arguably more relevant for LMICs where a substantial proportion of the aging adult population did not complete primary schooling.

Only two studies have estimated schooling effects on older-age cognition in LMICs. Huang and Zhou (2013) found an extra year of schooling was associated with increased memory scores of 0.11 *SD*s in China, using the 1959–1961 famine as a natural experiment. Nikolov and Yeh (2022) found effect sizes of 0.18–0.23 *SD*s for older adults in a South African province, using past schooling expenses as IVs.

The objective of this study is to provide new evidence on the causal relationship between schooling and older-age cognition in LMICs. Specifically, we estimated causal effects of schooling on memory (word-recall) test performance for older adults (born 1900–1959) in China, Ghana, India, Mexico, Russia, and South Africa using harmonized data from the World Health Organization (WHO) Study on Global Ageing and Adult Health (SAGE).

The key innovation of our paper is the application of a nonparametric partial identification approach (Manski & Pepper, 2000) that allows for arbitrary correlations between schooling and unobserved factors that can affect cognition and produces a range of possible values—or bounds—on the causal effects based on weak, credible assumptions. Importantly, this approach estimates bounds on population average-treatment effects (ATEs) of increasing schooling at different parts of schooling distributions, not just at schooling levels affected by compulsory schooling laws.

Amin et al. (2025) used this partial identification approach to estimate schooling effects on older-adult cognition using the U.S. Health & Retirement Study (HRS)-Harmonized Cognition Assessment Protocol (HCAP) data. But it has not been applied to estimating schooling effects on older-adult cognition in LMICs. Moreover, by using harmonized data and a common econometric framework, we were better able to compare schooling effects across countries with varying development, institutional oversight, and cultures.

Method

Partial Identification Framework

Our approach is rooted in the theory of causal inference based on counterfactuals (Imbens & Rubin, 2015). Let $Y_i(t_1)$ and $Y_i(t_2)$ be values of older-age cognition that individual i would obtain with no schooling (t_1) and primary schooling (t_2), respectively. We are interested in the population ATE of increasing schooling from none to primary on cognition:

$$\Delta(t_1, t_2) = E[Y(t_2)] - E[Y(t_1)] \tag{1}$$

ATE estimation is complicated because of missing counterfactuals. Letting S denote realized schooling, the expected potential outcome $E[Y(t_2)]$ is:

$$E[Y(t_2)] = E[Y(t_2) \mid S < t_2] * P(S < t_2) + E[Y(t_2) \mid S = t_2] * P(S = t_2) + E[Y(t_2) \mid S > t_2] * P(S > t_2) \tag{2}$$

where $P(S < t_2)$ is the probability that schooling is less than t_2, $P(S = t_2)$ is the probability that schooling is equal to t_2, and $P(S > t_2)$ is the probability that schooling is greater than t_2.

The data identify sample analogues of all right-side quantities except the counterfactuals $E[Y(t_2)|S < t_2]$ and $E[Y(t_2)|S > t_2]$, that is, the average cognition under primary schooling (t_2) for individuals with realized schooling (S), respectively, less and higher than primary schooling. A similar equation and the arguments below apply to $E[Y(t_1)]$. The bounding approach we employed makes assumptions to bound each of the counterfactuals in the expressions for $E[Y(t_2)]$ and $E[Y(t_1)]$ to then bound the ATE $\Delta(t_1,t_2)$. Assumptions are outlined below and illustrated in Figure 1 with further details in the Supplementary Methodology Section.

Bounded support

This assumption exploits that our outcome has minimum (Y_{min}) and maximum (Y_{max}) values, which are used in place of the counterfactuals (Figure 1A).

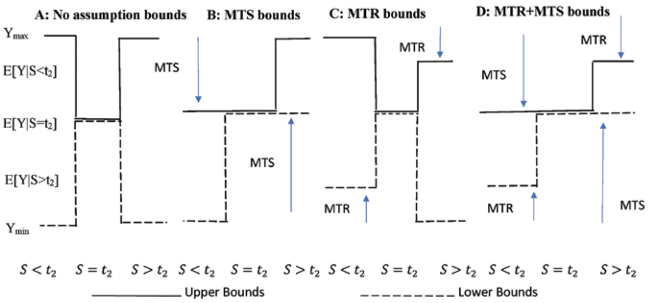

Figure 1. No assumption, MTS, and MTR bounds for $E[Y(t_2)]$. Each panel illustrates graphically the construction of nonparametric bounds on $E[Y(t_2)]$ under a given assumption. The x-axis represents levels of the realized schooling (S), while the y-axis represents values of the outcome (Y). Panel (**A**) illustrates the construction of the no-assumption bounds, where the maximum (Y_{max}) and minimum (Y_{min}) of the outcome are used to replace the missing counterfactuals for those with $S < t_2$ and $S > t_2$. In panel (**B**), the MTS assumption tightens the lower and upper bounds relative to the no-assumption bounds by using the observed mean outcomes for those receiving t_2 ($E[Y|S = t_2]$) to replace the missing counterfactuals for those with $S < t_2$ and $S > t_2$. In panel (**C**), the MTR assumption reduces the upper bound for those with $S > t_2$ by using the observed mean outcome for those with $S > t_2(E[Y|S > t_2])$. The MTR assumption increases the lower bound for those with $S < t_2$ by using the observed mean outcome for those with $S < t_2(E[Y|S < t_2])$. The combination of panel (**B**) and panel (**C**) would correspond to the case when the MTS and MTR assumptions are imposed together in panel (**D**). MTR = monotone treatment response; MTS = monotone treatment selection.

Monotone treatment selection

Assume that individuals with higher schooling attainment on *average* have weakly higher potential outcomes at every schooling level t. For example, monotone treatment selection (MTS) requires that the average potential older-age cognition at any schooling level t of individuals with primary schooling is greater than or equal to the average potential older-age cognition of individuals with no schooling. Under MTS, $E[Y(t_2)S = t_2]$ is used to replace the counterfactuals (Figure 1B).

Monotone treatment response

Assume that more schooling does not decrease older-age cognition for any individual: $Y_i(t_2) \geq Y_i(t_1)$ for all i. Monotone treatment response (MTR) compares potential cognition under primary schooling versus potential cognition under no schooling for the same individual. Under MTR, $E[Y(t_1)|S = t_1]$ is used to replace $E[Y(t_2)|S < t_2]$, and $E[Y(t_3)|S = t_3]$ replaces $E[Y(t_2)|S > t_2]$ (Figure 1C).

Monotone treatment selection ± monotone treatment response

MTS and MTR assumptions can be combined to provide tighter bounds (Figure 1D).

Monotone instrumental variable

In spite of its name, a monotone instrumental variable (MIV) is not a conventional IV. In the traditional IV framework to address endogeneity, a valid IV is a variable that is related to treatment and satisfies the exclusion restriction, which requires that the IV affects the outcome only through treatment. For example, one may consider height as a possible instrument, as it is a marker of early-life nutrition affecting cognitive development and is associated with schooling attainment. However, the exclusion restriction is unlikely to be satisfied as nutrition (the cause of height) can directly affect older-age cognition, and height can have direct effects through better health. Improvements in childhood nutrition have been linked to increases in height (Fogel, 1986), health (Fogel, 1994), and life expectancies (Fogel, 2012) in later-born generations, which are possible explanations for the Flynn effect (increase in cognitive performance over time).

In contrast, an MIV is a variable with a weakly monotone (increasing or decreasing) mean relationship with potential outcomes $Y(t)$, and is used to tighten bounds derived under other assumptions. An MIV *does not* need to be related to treatment or satisfy the exclusion restriction. Therefore, its requirement is fairly mild and considerably less restrictive than the assumptions required in the classic IV estimation. Specifically, taking the case of a weakly increasing MIV denoted by Z, it satisfies the following for all schooling levels t:

$$m_1 \leq m \leq m_2 \;\Rightarrow\; E\left[Y(t) \mid Z = m_1\right] \leq E\left[Y(t) \mid Z = m\right] \leq E\left[Y(t) \mid Z = m_2\right] \quad (3)$$

For height to be a valid MIV, Equation (3) requires that the average potential older-age cognition of individuals with greater height (e.g., m_2) is no less than that of individuals with lesser height (e.g., m_1). This assumption allows height to affect older-age cognition both directly and indirectly. The MIV assumption would be violated if the direction of the inequality in Equation (3) switched such that average potential cognition of taller individuals is *less than* average potential cognition of shorter individuals.

With an MIV, one then divides the sample into subsamples based on values of Z and obtains bounds for each subsample using the other assumptions. Bounds may be relatively tight for some subsamples and wider for others. The possibility of obtaining tighter bounds using an MIV comes from the weakly increasing relationship between the MIV (used to partition the sample) and the potential outcomes. Thus, an MIV lower (respectively, upper) bound on $E[Y(t_2)]$ is obtained by first taking the maximum (minimum) lower (upper) bound over different subsamples to obtain conditional-on-Z MIV bounds (Supplementary Figure 1). Then, MIV bounds on $E[Y(t_2)]$ are obtained by taking the weighted average of the conditional-on-Z MIV bounds over Z.

Bounds on the ATE $\Delta(t_1,t_2)$

Bounds on the ATE of increasing schooling from t_1 to t_2 for a given set of assumptions are calculated by:

$$LB\ E\left[Y(t_2)\right] - UB\ E\left[Y(t_1)\right] \leq \Delta\ (t_1,\ t_2) \leq UB\ E\left[Y(t_2)\right] - LBE\left[Y(t_1)\right] \quad (4)$$

where LB and UB denote lower and upper bounds, respectively. Under the MTR assumption, the lower bound on $\Delta(t_1,t_2)$ is never below zero, because MTR rules out that more schooling worsens cognition. Bounds for other schooling margins (e.g., increasing schooling from primary to high school) are computed analogously.

Assessment of identifying assumptions

We believe MTS is likely to hold. Human capital models posit that individuals with higher innate abilities have more schooling (Ben-Porath, 1967), and polygenic scores (that proxy for innate ability) for schooling and cognition predict cognition and memory at older ages (Ding et al., 2019; Fletcher et al., 2021), indicating that higher innate ability is also likely related to better older-age memory. Given that individuals with higher innate ability are more likely to have more schooling and better older-age cognition, it is plausible that, on average, individuals with higher schooling attainment have higher potential cognition at all schooling levels.

MTR is stronger than MTS as it is required to hold for each individual, rather than on average. MTR would be violated if more schooling leads to worse cognition for some individuals. More schooling may increase alcohol consumption for individuals who work in stressful jobs, for example, which could worsen cognition. MTR does not rule out such negative channels, but rather assumes that positive channels linking schooling to cognition dominate for each individual.

As noted earlier, the MIV assumption would be violated if average potential cognition of shorter individuals were *greater than* average potential cognition of taller individuals. This could be the case for *certain* individuals—for example, some tall individuals could be bullied in school, resulting in lower schooling and older-age cognition. However, the proportion of such individuals would have to be implausibly high to reverse the direction of the weakly monotonic *mean* relationship in Equation (3).

Estimation and inference

Bounds were estimated by plugging in sample analogs for expectations and probabilities in the corresponding expressions provided in the Supplementary Methodology Section. Confidence intervals (CIs) on the population ATE were constructed using methods developed by Chernozhukov et al. (2013), which deal with biases arising from taking the minimum or maximum (intersections) over several candidate bounds with the MIV assumption. We implemented the methods using the user-written STATA command *mpclr* (Germinario et al., 2021).

Data and Measures

We used the Wave 1 (2007–2010) WHO SAGE surveys conducted for China, Ghana, India, Mexico, Russia, and South Africa. Multistage cluster sampling strategies were employed in each country. Country-specific strata were typically defined by region/state/province/district and locality. Primary sampling units (enumeration areas, villages, or districts) were identified to generate a list of households, which were then classified as being a "50+ household" or "18–49 household." All persons aged 50+ were selected from "50+ households" and one person aged 18–49 years from "18–49 households." The number of individuals in "50+ households" was as follows: China: 13,368; Ghana: 4,732; India: 7,150; Mexico: 4,412, Russia: 4,511, and South Africa: 3,842.

Memory was assessed using the Consortium to Establish a Registry for Alzheimer's Disease (CERAD) memory test, in which interviewers read a set of 10 words (arm, bed, plane, dog, clock, bike, ear, hammer, chair, and cat) and asked respondents to recall as many as they can. Respondents were given three immediate trials to recall as many words as possible. The immediate memory score was the total correctly recalled words from the three attempts (range 0–30). The delayed memory score (range 0–10) was the total correctly recalled after delays of about 5–10 min. Our outcome was the sum of the immediate memory and delayed memory recall scores.

Respondents were asked to report their highest schooling level, grouped into no schooling, some primary, primary, secondary, high school, university, and postgraduate degree. In some HICs, secondary schooling and high school are synonyms. To avoid confusion, we labeled secondary schooling as middle school/junior high. We assigned 0 years of schooling for no schooling, 3 for some primary, 6 for primary, 10 for middle/junior high, 12 for high school, and 16 for university. Measured height was the MIV.

Sample Restrictions

Sample restrictions are shown in Supplementary Figure 2. We restricted the analysis to individuals with completed/partial interviews. We then excluded observations with missing data on word-recall, schooling, height, age, and sampling probability weights. For South Africa, we also dropped observations that were part of a single strata unit, since the bootstrapping procedure used for inference requires the number of strata be greater than one. Final sample sizes were 12,269 observations for China, 4,142 for Ghana, 6,360 for India, 1,787 for Mexico, 3,406 for Russia, and 2,932 for South Africa. The sample size for Mexico was much lower because of the 4,142 individuals in 50+ households, only 2,112 were interviewed. This low response rate was due to short field work time, which meant respondents who were not at home at the initial visit were not revisited (Kowal et al., 2012).

Results

Sample Characteristics

Table 1 presents descriptive statistics. Average age of respondents was lowest in India (61.29) and highest in Mexico (68.56). Birth cohorts were concentrated between 1930 and 1959. Russia had the highest proportion of female respondents (~61%). Ghana and India had the lowest average schooling attainments, with over 50% of respondents having never attended school. Despite similar schooling distributions, respondents in Ghana scored better on word-recall tests (22.04) than respondents in India (20.58). The word-recall distributions in Ghana and India were not equal (Kolmogorov–Smirnov test p-value < .01), and there was indication of ceiling effects for Ghana (Supplementary Figure 3). Respondents in Mexico performed the worst on average. This could be due to there being more older individuals and more women in the analytical sample. In terms of our MIV, the tallest individuals on average were found in Russia and the shortest in Mexico. Height distributions were similar across countries (Supplementary Figure 4). On average men had more schooling than women and performed better on the word-recall test (Supplementary Table 1). Average word-recall scores were weakly increasing in schooling, consistent with the implication of the MTS + MTR assumptions (Supplementary Table 2).

Effects of Completing Primary Schooling

Results are summarized in Figure 2 (presented in Supplementary Table 3), which shows OLS estimates, MTS + MTR + MIV bounds, and corresponding 95% CIs for increasing schooling from (i) none to some primary, (ii) some primary to primary, and (iii) none to primary. In Figure 2A and B, 95% CIs included zero, whereas they excluded zero for all countries except South Africa in Figure 2C. Ghana

Table 1. Summary Statistics

Variable	China	Ghana	India	Mexico	Russia	South Africa
Demographics						
Age (year), mean (*SD*)	62.43 (8.84)	64.26 (10.64)	61.29 (8.72)	68.56 (8.70)	63.45 (10.09)	61.44 (9.28)
10-year age category, %						
50–59 years	45	40	49	18	46	50
60–69 years	32	28	31	41	25	31
70–79 years	18	23	16	29	21	14
80–89 years	4	8	4	12	7	5
≥90 years	1	2	1	8	1	0
Gender, %						
Female	51	47	49	54	61	60
Schooling						
Years of schooling, mean (*SD*)	5.94 (4.63)	4.04 (5.19)	4.10 (1.61)	4.70 (4.41)	11.83 (2.72)	5.53 (1.47)
Schooling category, %						
No schooling	23	54	51	21	1	24
Some primary schooling	20	10	10	41	1	24
Primary schooling	22	11	15	20	5	23
Secondary schooling	20	4	10	7	20	15
High school	12	17	9	3	56	9
University	4	4	5	8	18	6
Memory						
CERAD score, mean (*SD*)	21.83 (6.65)	22.05 (5.84)	20.58 (5.52)	19.89 (5.78)	24.43 (6.87)	23.35 (6.44)
MIVs						
Height (cm), mean (*SD*)	159.19 (8.70)	161.81 (9.43)	156.80 (9.95)	156.25 (9.82)	164.15 (8.96)	159.38 (11.74)
Observations	12,269	4,142	6,360	1,787	3,406	2,932

Notes: CERAD = Consortium to Establish a Registry for Alzheimer's Disease; MIV = monotone instrumental variable; *SD* = standard deviation. Weighted summary statistics using country-specific sampling weights.

had the tightest bounds followed by India. Increasing schooling from none to primary increased word-recall scores by 0.07–0.37 *SD*s in Ghana and by 0.10–0.55 *SD*s in India. The widths of the bounds were similar for China (0.08–0.76 *SD*s) and Mexico (0.12–0.74 *SD*s). No results were presented for Russia due to the low proportion of individuals with primary schooling or lower.

Effects of Completing High School and University

Figure 3 shows results for increasing schooling from (i) middle school/junior high to high school, (ii) primary to high school, and (iii) high school to university. Figure 3A and B suggests benefits of high school completion are likely to be larger for individuals who only completed primary schooling. Bounds indicated increasing schooling from primary to high school improved word-recall scores by 0.16–0.48 *SD*s in China, 0.26–0.53 *SD*s in India, and 0.41–0.71 *SD*s in Mexico, with 95% CIs also excluding zero. In Figure 3C, bounds for India showed completing university increased word-recall scores between 0.21 and 0.46 *SD*s, but the 95% CI included zero. Bounds and 95% CIs included zero for all other countries. Bounds were not displayed for Mexico because they crossed, with the lower bound (0.00) being larger than the upper bound (−0.80). This was likely because only 3% (38 individuals) completed high school. The consequences of this small number of observations are compounded when breaking up the sample into three height categories to compute the MTS + MTR + MIV bounds.

Gender Differences

Gender-stratified results are presented in Supplementary Figures 5–10. Since bounds overlapped, we were unable to statistically determine differences by gender. There were two interesting observations for increasing schooling from none to primary (Supplementary Figure 7). First, in Ghana and India, 95% CIs excluded zero for women but not men. Second, China was the only country where 95% CIs excluded zero for both men and women, and bounds were narrower for women.

Discussion

We contributed new causal evidence for six LMICs with different development levels and schooling distributions using harmonized data and employing a common statistical framework that bounds causal effects under relatively weak assumptions. We found increasing schooling from none to primary increased word-recall scores by 0.08–0.76 *SD*s in China, 0.08–0.76 *SD*s in Ghana, 0.10–0.55 *SD*s in India, 0.12–0.74 *SD*s in Mexico, and 0–0.42 *SD*s in South Africa. The results imply an extra year of schooling increased word-recall scores by 0.01–0.13 *SD*s in China, 0.01–0.06 *SD*s in Ghana, 0.02–0.09 *SD*s in India, 0.02–0.12 *SD*s in Mexico, and 0–0.07 *SD*s in South Africa, assuming a difference of 6 years of schooling between individuals with no schooling and primary schooling. These effects are meaningful when compared to associations of an additional year of aging on cognition. OLS regressions of standardized word-recall scores on age showed an extra

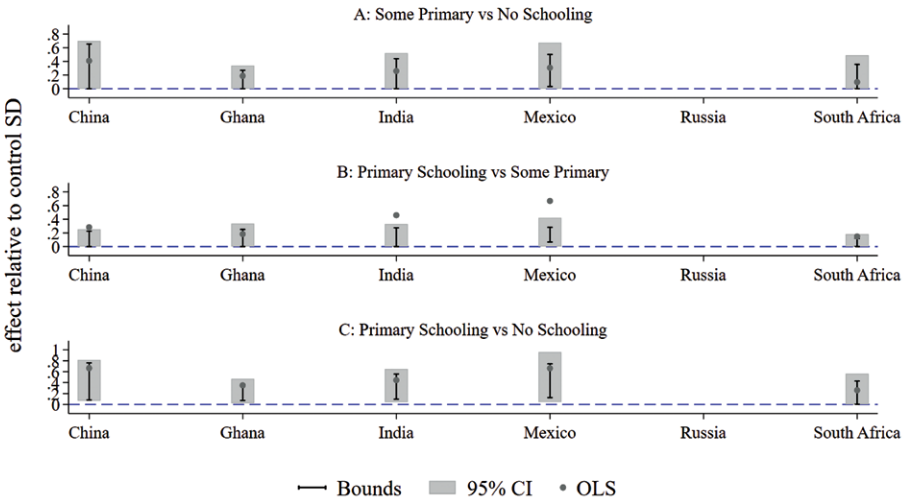

Figure 2. Effects of primary schooling under MTS + MTR + MIV assumptions. The y-axis indicates the size of the effects relative to the corresponding control *SD*. The control group in panels (**A**) and (**C**) are individuals with no schooling, while individuals with some primary represent the control group in panel (**B**). The bands indicate the estimated bounds, while the large dots represent the corresponding OLS estimate from regression of word-recall scores on dummy variables for schooling comparisons. The shaded boxes represent the valid 95% confidence intervals on the parameter of interest obtained using the Chernozhukov et al.'s (2013) method described in Supplementary Methodology Section. Height with three bins was used as the MIV. No results are shown for Russia due to the low proportion of individuals with primary schooling or lower. CI = confidence interval; MIV = monotone instrumental variable; MTR = monotone treatment response; MTS = monotone treatment selection; OLS = ordinary least squares; *SD* = standard deviation.

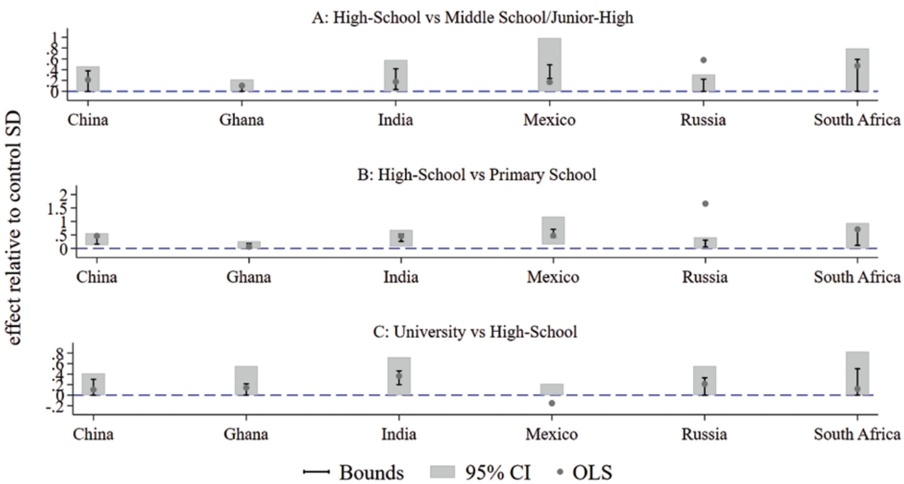

Figure 3. Effects of completing high school and university under MTS + MTR + MIV assumptions. The y-axis indicates the size of the effects relative to the corresponding control *SD*. The control group in panel (**A**) is individuals with middle school/junior high, while individuals with primary schooling represent the control group in panel (**B**). The control group in panel (**C**) is individuals with high school. The bands indicate the estimated bounds, while the large dots represent the corresponding OLS estimate from regressions of word-recall scores on dummy variables for schooling comparisons. The shaded boxes represent the valid 95% confidence intervals on the parameter of interest obtained using the Chernozhukov et al.'s (2013) method described in Supplementary Methodology Section. Height with three bins was used as the MIV. In panel (**C**), the bounds for Mexico cross and are thus not shown. Estimated bounds for Mexico are [0, −0.80]. CI = confidence interval; MIV = monotone instrumental variable; MTR = monotone treatment response; MTS = monotone treatment selection; OLS = ordinary least squares; *SD* = standard deviation.

year of aging was associated with lower word-recall scores of 0.02 *SD*s in Ghana and South Africa, 0.03 *SD*s in India, 0.04 *SD*s in China, and 0.05 *SD*s in Mexico. Thus, increased word-recall scores due to an additional year of schooling can offset 0.25–3 additional years of aging in China, 0.5–3 additional years of aging in Ghana, 0.67–4 additional years of aging in India, 0.4–2.4 additional years of aging in Mexico, and 0–2.5 additional years of aging in South Africa.

Our results additionally highlight the importance of country-specific contexts. Ghana and India had similar schooling distributions with over 50% of older adults never attending school. Yet the estimated largest possible effect of an additional year of schooling from completing primary school was higher in India (0.09 *SD*s) than Ghana (0.06 *SD*s). China and South Africa also had similar schooling distributions, yet the upper bound of an additional year of completing primary schooling was much larger in China (0.13 *SD*s) than in South Africa (0.07 *SD*s).

What are possible explanations for differences in estimated upper bounds of an additional year of schooling from completing primary schooling across countries that had similar schooling distributions? One possibility is school quality, which can affect cognitive stimulation and hence cognitive reserve. Though schooling levels are harmonized, we conjecture individuals who completed primary schooling had different school quality experiences. Many respondents were born during colonial eras for Ghana, India, and South Africa, and schooling quality likely was poor. Aboagye (2021) notes that when Ghana became self-governing in 1951, there were 3,198 schools of which 1,378 were unassisted and 52 were governmental schools. Furthermore "unassisted schools were started by local communities without reference to the Education Department, and were unregistered, ill-housed, lacked teaching and learning equipment and were mainly staffed with untrained indigenous teachers." Similarly, when the Union of South Africa was formed in 1910, schooling was under the control of Provincial Councils. There were separate curricula for Black and White individuals at primary schools. Jansen (1990) notes that "the syllabus for Black schools was inferior and conceived to ensure that the great majority of Black individuals were fitted for menial jobs." The Welsh committee found problems of mass absenteeism, overcrowding, poor facilities, and understaffing in Black schools in 1935–1936. Smaller upper bounds for Ghana and South Africa could be due to lower school qualities.

China and India have also historically emphasized different levels of schooling. After independence, the first Indian 5-year plan (1951–1956) prioritized university education over primary schooling. Between the first and fifth 5-year plans (1974–1978), spending on primary schooling as a percentage of total governmental schooling expenditure decreased from 56% to 36%, whereas spending on university education increased (Tilak, 1990). In contrast, China emphasized primary schooling in the 1950s. India's literacy rate increased from 28% in 1961 to 41% in 1981, whereas China's literacy rate increased from 43% to 68%. The difference between India and China can be seen as a consequence of India's relative focus on university education (Drez & Loh, 1995). Carron and Châu (1996) compared teaching and learning in China and India based on case studies by Cheng (1996) and Govinda and Varghese (1993). They reported rural Chinese schools had much better physical facilities and teaching resources compared to comparable Indian schools, and that teacher absenteeism was a bigger problem in India than in China. Thus, individuals in China and India also likely experienced widely different primary schooling quality.

Our results also differ from findings for HICs. Based on educational reforms that target high school completion, previous studies have found an additional year of schooling increased word-recall scores by 0.35 *SD*s in the United States (Glymour et al., 2008), 0.42–0.50 *SD*s in England (Banks & Mazzonna, 2012; Gorman, 2023), and 0.10–0.16 *SD*s in Europe (Crespo, 2014; Mazzonna, 2014; Schneeweis et al., 2014). We did not find similar effects on the comparable margin of increasing schooling from middle school/junior high to high school. Assuming a difference of 4 years of schooling between middle school/junior high to high school, bounds indicated an extra year of schooling increased word-recall scores by *at most* 0.09 *SD*s in China, 0.02 *SD*s in Ghana, 0.11 *SD*s in India, 0.25 *SD*s in Mexico, 0.22 *SD*s in Russia, and 0.15 *SD*s in South Africa, all smaller than findings for the United States and England. The bounds were also consistent with null effects. Using the same partial identification approach, Amin et al. (2025) found evidence of a causal effect when going from being a high school to university graduate, with an extra year of schooling increasing memory scores by 0.03–0.10 *SD*s at this margin for older adults in the U.S. HRS-HCAP, whereas we could not reject null effects at this margin.

Our smaller estimated effects of high school completion in LMICS compared to estimates based on compulsory schooling laws for HICs could be due to: (1) Differences in identified treatment effects: estimates identified from changes in compulsory schooling laws capture effects only for individuals who continued their schooling because of compulsory schooling laws. These individuals are likely to differ from the rest of the population, so estimates are likely not representative of the entire population. In contrast, we estimated bounds on the population ATE of high school completion. (2) Differences in schooling access, curriculum and quality: older adults in the United States and England likely had easier access to schooling. The period 1910–1940 is known as the "high school movement" in the United States due to the dramatic increase in secondary school enrollment and high school graduation rates, that was not seen anywhere else (Goldin, 1998). In contrast, older adults in LMICs likely had limited access to schooling, especially those in Ghana, India, and South Africa who were in school ages during colonization. As discussed earlier, schooling quality varied across the LMICs, which we hypothesize was higher in the United States and England. (3) Selective survival: life expectancies in the LMICs were much lower compared to the United States and England for the cohorts studied here. Hence, individuals who survived to the time of the WHO SAGE survey may have unmeasured characteristics that promote survival and cognitive health, which could mute the effects of schooling.

Our study has some limitations. First, our approach does not deal with possible biases arising from sample selection issues. For Mexico, almost half the study sample was excluded from the analysis due to low interview response rates. Since our estimated upper and lower bounds rely primarily on estimated averages for subpopulations defined by schooling status, biases on these estimated averages resulting from missing data will translate into biases of the estimated (lower or upper) bounds. If those biases are such that the bounds are wider, then the true parameter of interest is still covered by the estimated bounds, but if those biases lead to narrower bounds, then the true parameter could lie outside of the estimated bounds. Biases could also arise if individuals were excluded because of missing data on key variables (cognition, schooling, height) had low schooling attainment and low cognitive performance. We examined this concern in Supplementary Table 4. There were no statistically significant differences (with the exception of China) in years of schooling attainment between individuals with missing data on memory and nonmissing data. South Africa was the only country where a significant proportion (16%) of individuals had no schooling data, and these individuals had statistically significantly higher memory scores than individuals with schooling data. For all countries, individuals excluded due to missing data on height had statistically significantly lower memory scores than individuals with data on height. For Ghana and India, these excluded individuals also had statistically significantly fewer years of schooling. Second, we were unable to distinguish between age and birth cohort effects, which is needed to elucidate whether effects of schooling on memory are larger in later-born cohorts. This is important because schooling attainment and quality have likely increased over time. Third, we were unable to identify the mechanisms through which schooling affected old-age cognition nor the relevance of intercountry differences, such as school quality. Fourth, our study is not fully comprehensive as we did not have data on LMICs in every region (e.g., Caribbean, South America), and there is likely substantial heterogeneity across as well as within regions. Fifth, we only examined memory because the data did not include other measures of cognitive domains. Nevertheless, findings for memory are still informative about dementia risk as poor memory, as measured in our study, has been shown to be associated with an elevated risk of Alzheimer's dementia even independent of clinical thresholds in HICs (Schreiner et al., 2016; Waragai et al., 2017) and LMICs (Prince et al., 2003; Stewart et al., 2016).

Despite limitations, our study adds novel findings about causal relationships between schooling and older-age cognition in six LMICs. This is important because low schooling is a prominent risk factor for dementia and used as an input in dementia prediction models. Evidence has shown that not all dementia models developed for HICs can be extrapolated to LMICs (Stephan et al., 2020), and thus there is a need for country-specific estimates. We have demonstrated that partial identification can yield informative policy-relevant findings. While the literature has examined impacts of schooling attainment on older-age cognition, there is little evidence for school *quality*. This is an important gap because school quality can influence cognitive reserve. A promising avenue for future research is using partial identification to elucidate on the causal relationship between school quality and older-age cognition in LMICs.

Supplementary Material

Supplementary data are available at *The Journals of Gerontology, Series B: Psychological Sciences and Social Sciences* online.

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Conflict of Interest

None.

Data Availability

The data are publicly available through the Inter-university Consortium for Political and Social Research. Code to replicate the results is available from the corresponding author. The study was not preregistered.

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