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Teenage Driving, Mortality, and Risky Behaviors†

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We investigate the effect of teenage driving on mortality and risky behaviors in the United States using a regression discontinuity design. We estimate that total mortality rises by 5.84 deaths per 100,000 (15 percent) at the minimum legal driving age cutoff, driven by an increase in motor vehicle fatalities of 4.92 deaths per 100,000 (44 percent). We also find that poisoning deaths, which are caused primarily by drug overdoses, rise by 0.31 deaths per 100,000 (29 percent) at the cutoff and that this effect is concentrated among females. Our findings show that teenage driving contributes to sex differences in risky drug use behaviors. (*JEL* I12, J13, J16, R41)

Suicides and motor vehicle accidents are the two leading causes of death for teenagers in the United States. Both often involve substance abuse, which itself is also a leading cause of teenage death. Over 25 percent of all teenage hospitalizations are related to mental health or substance abuse disorders (Heslin and Elixhauser 2016). Yet, the determinants of risky behaviors among youth remain poorly understood, and what little we know about the drivers of risky behaviors among adults may not apply to teenagers. For example, drug overdose deaths among teenagers declined between 2007 and 2014, in stark contrast to the significant nationwide increase in adult "deaths of despair" that has received much attention from researchers (Case and Deaton 2015, 2017).

We investigate the effects of teenage driving on mortality and risky behaviors. While driving undoubtedly increases motor vehicle mortality risk, the magnitude of this risk is hard to quantify. Moreover, driving enables teenagers to participate in unsupervised risky behaviors away from home, which may in turn lead to changes in mental health or drug use that have additional effects on mortality risk. Investigating these links can shed light on the determinants of risky behaviors among youth. However, identifying the effects of driving on teenage outcomes is challenging. Individual-level data on driving behaviors are scarce, and comparing the behaviors

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of drivers to nondrivers is unlikely to yield causal estimates because the decision to obtain a license is voluntary. In addition, detecting changes in important but rare outcomes such as drug-related mortality requires large sample sizes.

We overcome these challenges by using a regression discontinuity (RD) approach to identify the causal effect of teenage driving on a number of outcomes.[1] Our research design exploits variation in driving eligibility caused by the minimum legal driving age (henceforth, "minimum driving age" or "MDA"), which creates large differences in the teenage driver population on either side of the MDA cutoff. We employ a confidential dataset that includes information about month and year of birth for over 500,000 teenage deaths during 1983–2014, which enables us to compare mortality rates for teenagers just above the MDA to mortality rates for teenagers just below the MDA. We estimate that driving eligibility increases teenage mortality by 5.84 deaths per 100,000 (15 percent) at the MDA cutoff. This effect is driven primarily by an increase in motor vehicle fatalities of 4.92 per 100,000 (44 percent).

We also estimate that teenage poisoning deaths rise by 0.314 deaths per 100,000 (29 percent) at the cutoff. This poisoning effect is driven by a stark rise in female drug overdose deaths of 0.646 per 100,000 (78 percent) and an accompanying rise in female gas poisoning deaths of 0.127 per 100,000 (82 percent).[2] These deaths reflect changes in both suicides and accidental deaths, although our analysis suggests that the increase in poisoning suicides reflects substitution away from other methods of suicide. By contrast, we find no effect on male poisoning deaths. This null effect is meaningfully precise: the 95 percent confidence interval rules out the corresponding female estimate.

The most common MDA during our sample period is 16, which raises the concern that our results might be driven by a "birthday effect" or some other regulation that takes effect at age 16, such as the federal minimum legal working age or a state's minimum school-leaving age. However, the discontinuities that we observe are long lasting, which is inconsistent with a temporary birthday effect, and we do not detect changes in the probability of working or leaving school at the MDA cutoff. Furthermore, our main results are similar when we limit our analysis to the subsample of states with MDAs other than 16. We therefore conclude that the MDA is the causal mechanism underlying our results.

Youth risky behaviors differ significantly by sex. During our sample period, 29 percent of young female deaths from suicides and accidents were caused by poisonings, compared to just 15 percent for young males. Prior studies have also noted that poisonings make up a significantly larger share of young female suicides than young male suicides (Ruch et al. 2019). Our results reveal that teenage driving contributes to these behavioral differences among male and female youth.

Much of the research on the causal determinants of drug abuse comes from studies of adults. For example, several studies find that drug overdose deaths increase after the first of the month and following the receipt of money, suggesting a causal relationship between a "full wallet" and substance abuse (Phillips, Christenfeld, and

[1] Our analysis controls for the family-wise error rate in order to address the multiple inference concern that arises when testing many hypotheses. See Section III for details.

[2] During our sample period, about 80 percent of teenage poisoning deaths are caused by drug overdoses, and about 20 percent are caused by gas poisonings. Most gas poisoning deaths are caused by carbon monoxide.

Ryan 1999; Riddell and Riddell 2006; Dobkin and Puller 2007; Evans and Moore 2012). Less is known about the causal determinants of drug abuse among adolescents.[3] Indeed, Gruber (2001) calls for economists to pay more attention to the risk-taking behavior of youth, noting that while economic incentives appear to matter, most of the variation remains unexplained. Anderson (2010), who finds no effect of a Montana anti-drug advertising campaign on methamphetamine use among high schoolers, emphasizes the need for further research on the determinants of illegal drug use among the young. Our study advances this literature by using a novel source of exogenous variation to uncover a strong, causal relationship between driving and drug overdose deaths among teenagers.

The remainder of our paper is organized as follows. Section I provides background information. Section II describes our data. Section III outlines our empirical strategy. Section IV describes our results, and Section V concludes.

I. Background

A. *Teenage Driving*

Every US state requires drivers to be licensed. Teenagers begin the licensing process by obtaining a learner's permit, which allows them to drive under adult supervision. Depending on the state, the adult must be at least 18–25 years of age and have up to five years of driving experience. The minimum legal age for obtaining a learner's permit ranges from 14 to 16 over our 1983–2014 sample period. With rare exception, teenagers must then complete a driver's education course and behind-the-wheel training to become eligible to take their state's driving test, which typically consists of two components: a written test and a behind-the-wheel test. The teenager receives her driver's license after passing both components.

Beginning in 1996, states began adopting Graduated Driver Licensing (GDL) programs. Teenagers who live in a state with a GDL program earn an "intermediate" driver's license after passing their state's driving test. An intermediate license imposes driving restrictions on newly licensed teenagers in an effort to encourage safer driving. For example, it typically prohibits unsupervised driving during certain nighttime hours and limits the number and age of passengers in the teenage driver's vehicle. These restrictions are lifted after a set period of time that ranges from three months up to two years, at which point the teenager earns a "full" license.

Our study focuses on the MDA, the minimum legal age at which teenagers become eligible to pass their state's driving test. In states with GDL programs in place, this eligibility threshold corresponds to the minimum legal intermediate licensing age. The MDA, which varies by state and over time, ranges from 14 to 18 years of age during our 1983–2014 sample period. On average, about 40 percent of 16-year-olds

[3] Most of the prior literature on youth drug use focuses on tobacco, alcohol, and marijuana (e.g., Glied 2002; Cawley, Markowitz, and Tauras 2004; Carpenter et al. 2019). Drug overdose deaths among teenagers, however, are mostly caused by opioids (both illegal and prescription) and sedatives. The teenage driving literature has investigated the effects of graduated driver licensing laws on motor vehicle fatalities (Dee, Grabowski, and Morrisey 2005; Morrisey et al. 2006; Karaca-Mandic and Ridgeway 2010; Gilpin 2019) and crime (Deza and Litwok 2016) and has investigated the effects of the MDA on crash risk (Foss et al. 2011; Chapman, Masten, and Browning 2014; Curry et al. 2015).

have a driver's license during this period, although this declines to less than 30 percent in more recent years (online Appendix B.1). A similar decline occurs for vehicle miles traveled.

B. *Teenage Mortality*

Accidents and suicides are the two leading causes of death among teenagers, respectively. Motor vehicle fatalities, which comprise over 60 percent of all accidental teenage deaths, are nearly as common as teenage suicides (CDC 2018). About 25 percent of teenage motor vehicle fatalities are alcohol related, and over 50 percent occur during the nighttime (Dee and Evans 2001). Motor vehicle fatality rates have declined significantly in recent decades, falling by over 50 percent during our 1983–2014 sample period. Prior studies have attributed this decline to a reduction in drunk driving, an increase in seat belt use, and the introduction of GDL programs (Dee and Evans 2001, Gilpin 2019). As of 1998, all states have implemented zero-tolerance drunk driving laws that set a limit of 0.02 percent blood alcohol concentration ($\sim$1 drink) or lower for drivers under age 21 (Carpenter 2004). Air bags were also introduced during this time period, but Dee and Evans (2001) argue that they played only a small role in reducing teenage motor vehicle fatalities.

Poisonings are the second leading cause of accidental death among teenagers (CDC 2018). Two-thirds of drug overdose deaths for youth ages 15–24 are related to heroin and illegal opioids, and one-third are related to sedatives and prescription opioids (National Institute on Drug Abuse 2019). Prior studies find that young females exceed males in their nonmedical use of sedatives and prescription opioids and are also more likely to overdose than males (Cotto et al. 2010, Lyons et al. 2019).

Poisonings also make up a significantly larger share of young female suicides than young male suicides (Ruch et al. 2019). Teenage females show higher rates of suicidal thinking and are more likely to attempt suicide than teenage males, but males die by suicide at higher rates (Swahn and Bossarte 2007). Section IIB documents sex disparities in poisoning deaths—both accidental and suicidal—over our sample period.

II. Data

A. *MDA Laws*

We obtained data on MDA laws from the Insurance Institute for Highway Safety for the years 1995–2014 (IIHS 1995–2014). Data for the years 1983–1994 were hand collected from databases of state session laws (HeinOnline 1983–1994). These data are reported in the online Appendix (Table B.1).

B. *Mortality*

We measure mortality using the National Vital Statistics (NCHS 1983–2014). This dataset is based on death certificate records and includes information on decedents' month and year of death, cause of death, race, and sex. The medical classification codes used to define the cause of death outcomes employed in our analysis are reported in the online Appendix (Table B.2). We obtained a restricted-use version of the National Vital Statistics for the years 1983–2014 that includes information on decedents' state of residence and month and year of birth. We use these data to calculate age in months at death for all decedents. We then aggregate to the age-in-months level and calculate age-specific deaths per 100,000 person-years by combining these count data with population estimates provided by the Surveillance, Epidemiology, and End Results (SEER) Program (NCI 1983–2014).[4] Our final dataset includes information on 501,193 teenage deaths.

The average death rate for teenagers ages 15–19 is 68.63 deaths per 100,000 during our sample period (online Appendix Table B.3). The majority of these deaths are attributable to external causes, and the leading external cause of teenage death is motor vehicle accidents.[5] The motor vehicle mortality rate for ages 15–19 is over five times larger than that for ages 10–14. This mortality rate increases further for ages 20–24 but then decreases for ages 25–29.

Suicides and accidents comprise 25 percent of all deaths for ages 15–19. Males in this age group are four times more likely than females to die by suicide or accident. The cause of these deaths varies by sex. For ages 15–19, 38 percent of male suicides and accidents are caused by firearms, compared to just 22 percent for females. By contrast, only 15 percent of male suicides and accidents are caused by poisonings, compared to 29 percent for females.

Deaths from all causes and deaths from motor vehicle accidents have both declined significantly for males and females during our sample period (online Appendix Figure B.3). However, although male deaths from suicides and accidents have declined steadily since 1994, female deaths from suicides and accidents increased beginning in 2003.

Male and female poisoning deaths evolved similarly from 1983 until 2000 (online Appendix Figure B.4). The trends then diverged: the male poisoning death rate climbs and falls significantly over the next 15 years, while the female poisoning death rate remains relatively constant after 2003. Neither group experiences sustained increases in poisoning deaths after 2007. By contrast, Case and Deaton (2015) document a steeply increasing trend in poisoning deaths among midlife whites for 2007–2013.

Our main estimates combine suicides and accidents into one category to minimize measurement error concerns (Cutler, Glaeser, and Norberg 2001; Alexander and Schnell 2019). When someone dies from a drug overdose, for example, it may not be clear whether the death should be classified as a suicide or an accident. We report estimates separately for suicides and accidents in the online Appendix.

C. *Driving Behaviors*

We measure driving behaviors using the National Longitudinal Study of Adolescent to Adult Health (hereafter, "Add Health"; Add Health 1995–1996). This

[4] These population data are available only for integer ages. When calculating age-specific death rates, we divide the count of deaths for a specific age in months by one-twelfth of the corresponding integer-age population.

[5] Deaths not caused by external factors are attributable to "internal causes." Cancer, circulatory system diseases, and nervous system diseases are the three largest subcategories within internal causes, and they make up more than half of all deaths due to internal causes at these ages.

nationally representative study began in 1994 with a classroom survey of about 20,000 students in grades 7–12. The study then followed up with a series of in-home interviews in 1995 and 1996. We obtained a restricted-use version of the in-home survey data that includes month and year of birth. After excluding observations with missing data, our sample includes 32,307 person-year observations (online Appendix B.4). This sample includes respondents ranging in age from 11 to 21; 97.9 percent of respondents are between the ages of 13 and 19. We aggregate these data to the age-in-months level using Add Health's cross-sectional weights.

The in-home survey asks respondents whether they have a driver's license and whether they drive 0, 1–50, 51–100, or "over 100" miles per week, which we use to measure vehicle miles driven. We assign values of 25 and 75 to respondents who selected the ranges 1–50 and 51–100, respectively. For "over 100," our baseline specification assigns a value of 150. By way of comparison, the typical adult driver drove 265 miles per week in 1996 (Federal Highway Administration 1997, 1992–2002). To account for uncertainty, we also report results from an alternate specification that instead assigns a value of 265 to the "over 100" response.

Add Health also asks questions about drug consumption and mental health. We do not present results for these outcomes because the low prevalence of the most important behaviors (suicide attempts and illegal drug consumption) and the survey's small sample size cause the analysis to be underpowered.

III. Empirical Strategy

We employ an RD design to identify the effect of driving eligibility on teenage driving behaviors and mortality. Eligibility depends on year, age, and state of residence. For analytical convenience, we recenter the age variable for decedents in our data by measuring it in months from the MDA law in force during the month of death.

Our main identifying assumption is that assignment to either side of the MDA threshold is as good as random. This assumption is very reasonable: age cannot be manipulated, and we do not suffer from sample selection bias because we observe the universe of deaths.

We estimate the following model:

$$Y_a = \alpha AGE_a + \beta POST_a + \gamma(POST_a \times AGE_a) + \delta D_a + \varepsilon_a. \tag{1}$$

The dependent variable, Y_a, is an outcome for the one-month age cell a. The running variable, AGE_a, is measured in months from the MDA, and $POST_a$ is an indicator equal to 1 if $AGE_a \geq 0$. This indicator suffers from measurement error at the age cutoff because we do not know whether a teenager who died in the month she reaches the MDA was over or under the MDA on the day of her death. We remove the bias associated with this measurement error by including the indicator variable D_a, which is equal to 1 when $AGE_a = 0$ and is 0 otherwise (Dong 2015).

All of our regressions employ a triangular kernel. Our preferred specification employs a mean-squared error (MSE) optimal bandwidth that is constant on each side of the cutoff but allowed to vary across outcomes. We report robust bias-corrected confidence intervals that account for the possibility that our estimating equation is misspecified (Calonico, Cattaneo, and Titiunik 2014).

We address multiple inference concerns by controlling for the family-wise error rate using the Sidak-Holm step-down correction.[6] In our setting, the family includes all 13 mortality outcomes reported in our main table. When estimating models separately for males and females, we include outcomes from both subgroups in the family, that is, we report *p*-values that adjust for testing 26 different hypotheses.

IV. Results

A. *Driving*

We begin by estimating the effect of driving eligibility on license status and vehicle miles driven. Figure 1, panel A shows that about 25 percent of teenagers obtain a license within their first two months of eligibility and that this proportion rises to over 50 percent after 12 months. The increase at the age cutoff (value 0 on the *x*-axis) is attenuated because of the measurement error discussed in Section III. Figure 1, panel B shows a corresponding rise in vehicle miles driven. Miles driven is positive prior to the MDA because teenagers with learner's permits can drive under parental supervision.

The first three rows of panel A in Table 1 report RD estimates of β from equation (1) for our driving outcomes. Column 2 reports that driving eligibility increases a teenager's probability of obtaining a driver's license by 18.6 percentage points. It also increases her annual driving amount by 375 miles using the baseline definition of average vehicle miles driven, or by 575 miles using the alternate definition (online Appendix Figure A.1). The increase in licensing rates at the cutoff is similar for males and females, but the increase in vehicle miles driven is larger for males.

B. *Mortality*

Figure 1, panels C and D show the effect of driving eligibility on total mortality and motor vehicle fatalities. Both of these mortality outcomes increase by about 5 deaths per 100,000 during the first two months of driving eligibility, for both males and females. There is little change in their overall trends: total mortality and motor vehicle fatalities both increase with age at about the same rate in the periods before and after the age cutoff.

Figure 2, panel A illustrates a stark increase in female poisoning deaths in the months immediately following the age cutoff. Figure 2, panels B and C show that female deaths from drug overdoses and gas poisonings—which together comprise all poisoning deaths—both rise significantly at the age cutoff. By contrast, Figure 3 shows that trends in male poisoning deaths appear continuous at the age cutoff.

Panel B in Table 1 reports estimates of β from equation (1) for 13 different causes of death.[7] Column 2 reports that total mortality increases by 5.84 deaths per 100,000

[6] This correction is conservative because it does not account for the significant collinearity among our outcomes (see Appendix C of Jones, Molitor, and Reif 2019). We are unaware of any resampling-based multiple testing corrections for RD designs with discrete running variables.

[7] Because bandwidths vary by outcome, estimates for specific causes of death do not add up to the estimate for total deaths. Estimates from a specification that uses a constant bandwidth of 24 months are available in the online Appendix (Table A.12). When using that specification, subcategory estimates do add up to the total estimate.

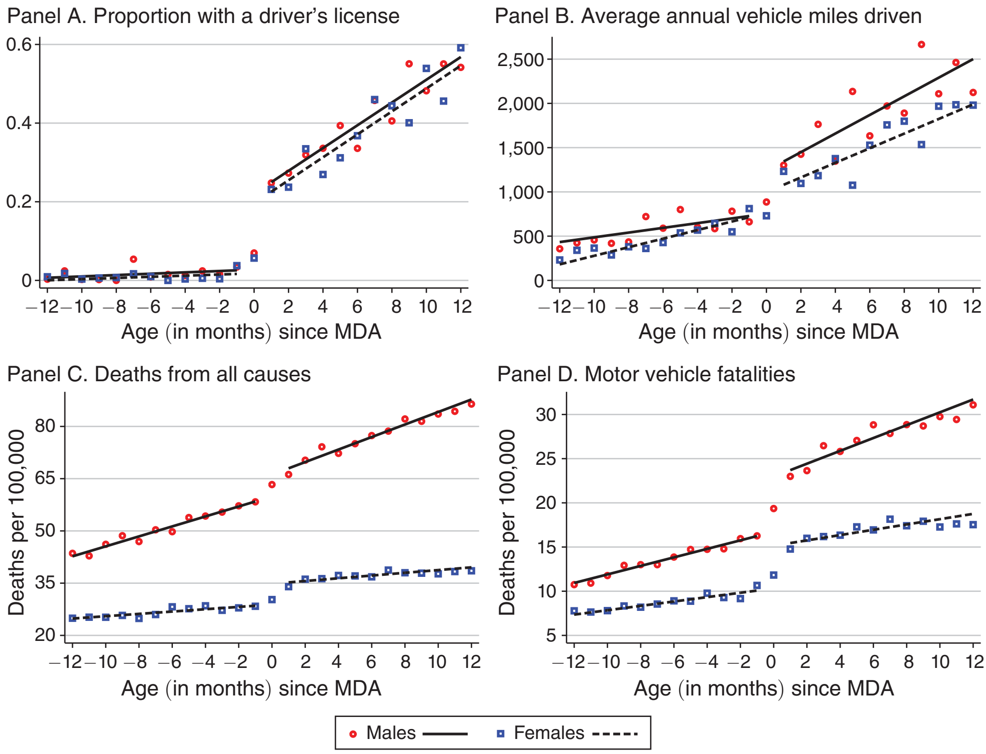

FIGURE 1. TEENAGE DRIVING AND MORTALITY

Notes: The figure shows the proportion of teenagers with a driver's license, average annual vehicle miles driven (baseline specification), all-cause death rates, and motor vehicle accident death rates by age relative to the MDA. Estimates in panels A and B are based on weighted responses to the 1995–1996 Add Health surveys. Estimates in panels C and D are based on death counts from the 1983–2014 National Vital Statistics and population data from the 1983–2014 SEER Program. The fitted lines are estimated using equation (1) with a bandwidth of 24 months. Table 1 provides RD estimates.

at the age cutoff, an increase of 15 percent relative to a mean of 38.9 deaths per 100,000. The estimated effect on deaths from internal causes is small and statistically insignificant. By contrast, the estimated effect on deaths from external causes is 5.20 deaths per 100,000 (19 percent) and remains marginally significant after conservatively accounting for multiple inference (family-wise $p = 0.0523$). Scaling the all-cause mortality estimate by the driver's license estimate reported in panel A implies that driving increases total mortality at the cutoff by 31.4 deaths per 100,000 (81 percent) among teenagers who receive a license.

Columns 4 and 6 of Table 1 report RD estimates separately for males and females. Comparing column 3 to column 5 shows that—for every cause of death—death rates are higher for males than females in the year prior to reaching the MDA. Thus, while the absolute increase in all-cause mortality at the age cutoff is about the same for both males and females, the relative increase for females (22 percent) is double the relative increase for males (11 percent).

Table 1 also decomposes deaths due to external causes into four main subcategories: motor vehicle accidents, suicides and accidents, homicides, and other. Column 2

Table 1—Effect of Driving Eligibility on Teenage Driving and Mortality

Outcome variable	Full sample		Male		Female	
	Mean (1)	RD (2)	Mean (3)	RD (4)	Mean (5)	RD (6)
Panel A. Driving						
Has driver's license	0.0130	0.186 [0.124, 0.231]	0.0163	0.193 [0.139, 0.231]	0.0101	0.179 [0.103, 0.232]
Miles driven (miles/yr) (baseline)	514	375 [159, 530]	569	486 [195, 734]	458	234 [−105, 479]
Miles driven (miles/yr) (alternate)	549	575 [231, 856]	613	753 [328, 1,194]	484	327 [−144, 676]
Panel B. Mortality						
All causes	38.9	5.84 [1.99, 9.36] {0.0252}	50.6	5.72 [−0.809, 11.3] {0.643}	26.7	5.76 [4.35, 7.53] {<0.0001}
Internal causes	12.2	0.406 [−0.120, 1.17] {0.560}	13.8	−0.0589 [−0.979, 1.03] {1.00}	10.5	0.820 [−0.0420, 2.00] {0.554}
External causes	26.7	5.20 [1.42, 8.47] {0.0523}	36.8	5.56 [0.0377, 10.3] {0.524}	16.1	4.82 [2.81, 6.66] {<0.0001}
Motor vehicle accident	11.2	4.92 [2.36, 7.07] {<0.001}	13.6	5.67 [2.76, 8.10] {<0.01}	8.75	4.46 [2.41, 6.14] {<0.001}
Suicide and accident	10.5	0.167 [−0.680, 0.924] {0.998}	15.6	−0.0506 [−1.63, 1.22] {1.00}	5.07	0.337 [−0.0259, 0.849] {0.555}
Firearm	3.64	0.0914 [−0.326, 0.474] {0.998}	5.87	0.529 [0.0108, 1.04] {0.524}	1.29	−0.333 [−0.715, -0.0560] {0.313}
Poisoning	1.08	0.314 [0.183, 0.522] {<0.001}	1.17	0.133 [−0.218, 0.458] {0.998}	0.984	0.747 [0.591, 1.07] {<0.0001}
Drug overdose	0.864	0.315 [0.233, 0.496] {<0.0001}	0.897	0.0447 [−0.242, 0.305] {1.00}	0.830	0.646 [0.476, 0.999] {<0.0001}
Carbon monoxide and other gases	0.214	0.103 [−0.0301, 0.215] {0.593}	0.270	0.0798 [−0.149, 0.258] {0.998}	0.154	0.127 [0.0333, 0.243] {0.163}
Drowning	1.53	−0.294 [−0.576, −0.0967] {0.0523}	2.64	−0.690 [−1.20, -0.352] {<0.01}	0.367	0.126 [−0.00258, 0.270] {0.544}
Other	4.23	0.105 [−0.316, 0.463] {0.998}	5.93	0.0406 [−0.511, 0.512] {1.00}	2.43	0.0749 [−0.519, 0.639] {1.00}
Homicide	4.80	−0.0423 [−0.623, 0.534] {0.998}	7.33	−0.0320 [−1.18, 1.10] {1.00}	2.14	−0.0779 [−0.335, 0.154] {0.998}
Other external	0.243	0.00608 [−0.148, 0.154] {0.998}	0.328	−0.0571 [−0.316, 0.154] {0.998}	0.154	0.143 [0.0872, 0.247] {<0.001}

Notes: The driving estimates in panel A are based on weighted responses to the 1995–1996 Add Health surveys. The mortality estimates in panel B, which are measured in deaths per 100,000 person-years, are based on death counts from the 1983–2014 National Vital Statistics and population data from the 1983−2014 SEER Program. Columns 1, 3, and 5 report means of the dependent variable one year before reaching the MDA. Columns 2, 4, and 6 report MSE-optimal estimates of β from equation (1). Robust, bias-corrected 95 percent confidence intervals are reported in brackets. Family-wise p-values, reported in braces, adjust for the number of outcome variables in each family and for the number of subgroups.

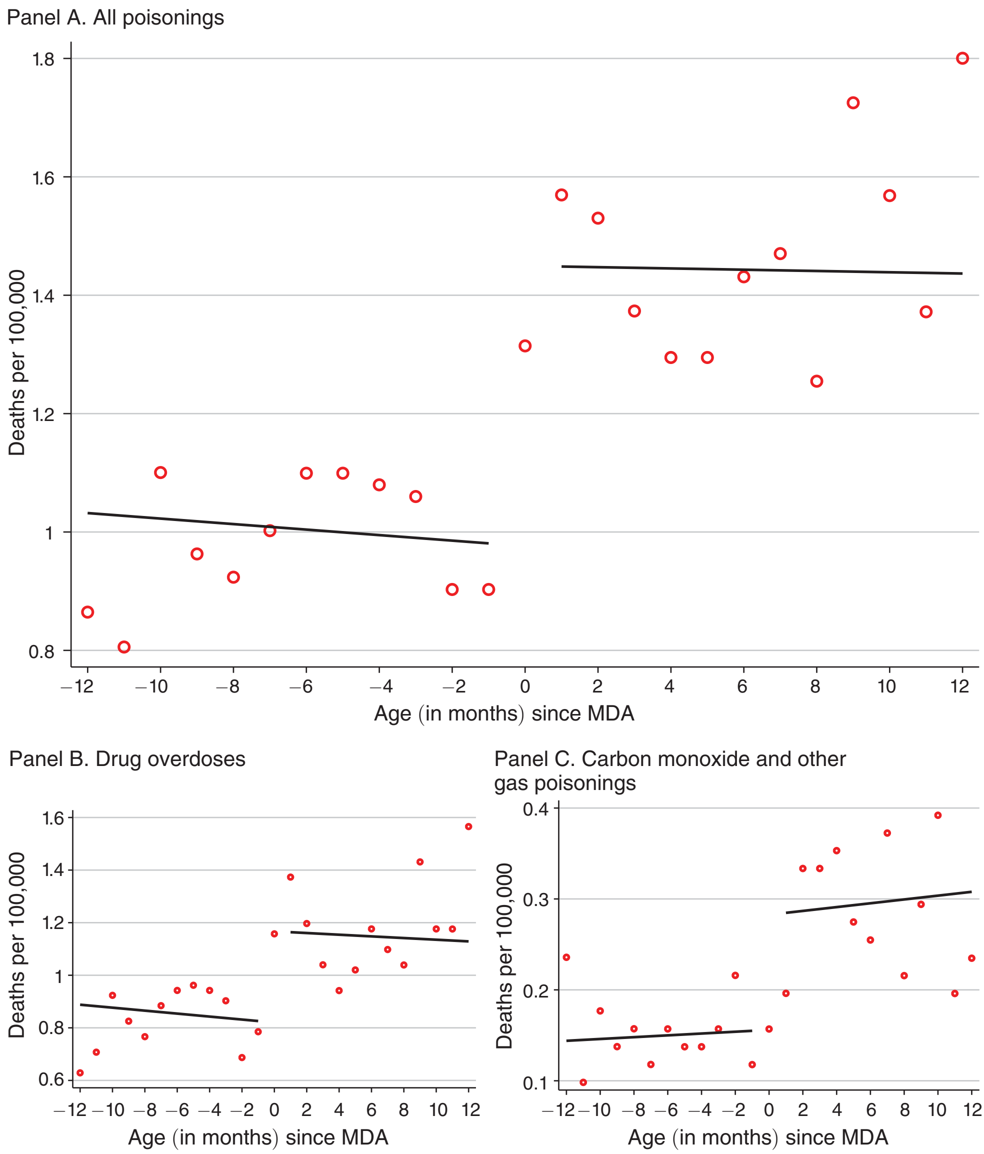

Figure 2. Female Poisoning Deaths, 1983–2014

Notes: The figure shows female death rates for different causes of death by age relative to the MDA. Deaths from all poisonings, shown in panel A, equal the sum of deaths from drug overdoses and from carbon monoxide and other gas poisonings. The fitted lines are estimated using equation (1) with a bandwidth of 24 months. Table 1 provides RD estimates.

reports a significant increase in motor vehicle fatalities of 4.92 deaths per 100,000 (44 percent) at the age cutoff. Columns 4 and 6 report that this increase is significant for both males and females (family-wise $p < 0.01$) and can explain the majority of the increase in total mortality for both subgroups. As with all-cause mortality, the absolute increase in motor vehicle fatalities at the cutoff is similar for males and females, but the relative increase is larger for females.

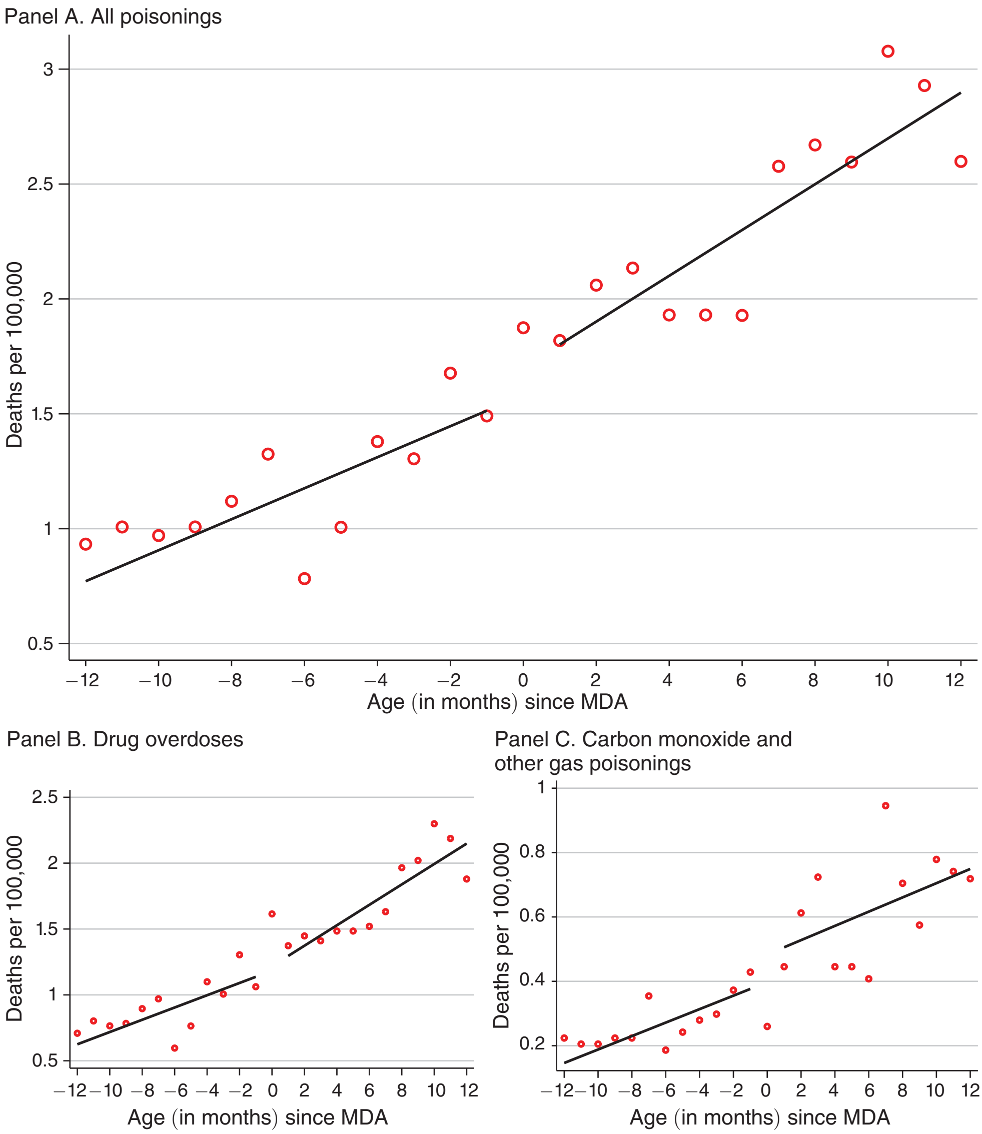

FIGURE 3. MALE POISONING DEATHS, 1983–2014

Notes: The figure shows male death rates for different causes of death by age relative to the MDA. Deaths from all poisonings, shown in panel A, equal the sum of deaths from drug overdoses and from carbon monoxide and other gas poisonings. The fitted lines are estimated using equation (1) with a bandwidth of 24 months. Table 1 provides RD estimates.

Column 2 of Table 1 also reports a significant increase in poisoning deaths of 0.314 per 100,000 (29 percent). This poisoning effect can be further decomposed into a 0.315 per 100,000 (36 percent) increase in drug overdose deaths and a 0.103 per 100,000 (48 percent) increase in gas poisoning deaths.[8] Comparing the

[8] Nearly 90 percent of all gas poisoning deaths are caused by carbon monoxide, an odorless gas produced by motor vehicle exhaust (based on authors' calculations using the 1983–1998 National Vital Statistics).

estimates in column 6 to those in column 4 reveals that this effect is driven by female poisoning deaths, which increase by 0.747 per 100,000 (76 percent) at the cutoff (family-wise $p < 0.0001$). This increase can be attributed primarily to a 0.646 per 100,000 (78 percent) increase in drug overdose deaths (family-wise $p < 0.0001$). Gas poisoning deaths rise by 0.127 per 100,000 (82 percent) at the cutoff, although this effect is not statistically significant after accounting for multiple hypothesis testing (family-wise $p = 0.163$).

Column 4 of Table 1 reports that the estimate for male poisoning deaths is small and statistically insignificant and that its 95 percent confidence interval rules out the corresponding female estimate. We do estimate a statistically significant decrease in male drownings of 0.690 per 100,000 (26 percent) (online Appendix Figure A.3a). Unlike the change in female poisoning deaths shown in Figure 2, panel A, this change in male drownings is short lived. Similarly, while column 6 of Table 1 reports a statistically significant increase in female deaths due to "other external" causes, the RD plot does not provide compelling evidence of an effect (online Appendix Figure A.4b).

Finally, we assess how our estimates change over time by estimating our model for different four-year bins. Figure 4, panels A and C reveal a steady decline in our estimated effects for male and female motor vehicle fatalities beginning in the mid-1990s. Likewise, Figure 4, panel D exhibits a declining trend in the estimates for female poisoning deaths. By contrast, Figure 4, panel B shows that the effects of driving eligibility on male poisoning deaths remain centered around zero for the majority of our sample period.

We cannot directly investigate the cause of the declines shown in Figure 4, panels A, C, and D because our driving behavior data are available only for 1995 and 1996. That said, these declines coincide with a decline in teenage driving documented in national statistics. For example, license counts published by the Federal Highway Administration show a steady decline in the proportion of US teenagers with a driver's license over our sample period (online Appendix Figure B.1), and McGuckin and Fucci (2018) report a decline in average vehicle miles traveled among licensed teenagers beginning in 1990 (online Appendix Figure B.2). In addition, prior studies of GDL programs—which were introduced beginning in 1996—suggest that these programs successfully reduced teenage motor vehicle fatalities by limiting, rather than improving, teenage driving (Karaca-Mandic and Ridgeway 2010, Gilpin 2019).

C. *Extensions*

Our RD estimates account neither for accidents with multiple fatalities nor for people killed by a teenage driver who survives the accident. According to data published by the Fatality Analysis Reporting System (FARS), on average an additional 0.24 people died for every car accident where a newly eligible teenage driver died at the wheel. In addition, among all fatal car accidents involving a newly eligible teenage driver at the wheel, the accidents where that teenage driver died accounted for only 45 percent of the total fatalities.[9] If we assume these statistics apply proportionally

[9]These statistics are calculated using Fatality Analysis Reporting System data published for the years 1983–2014. Those data report only integer ages.

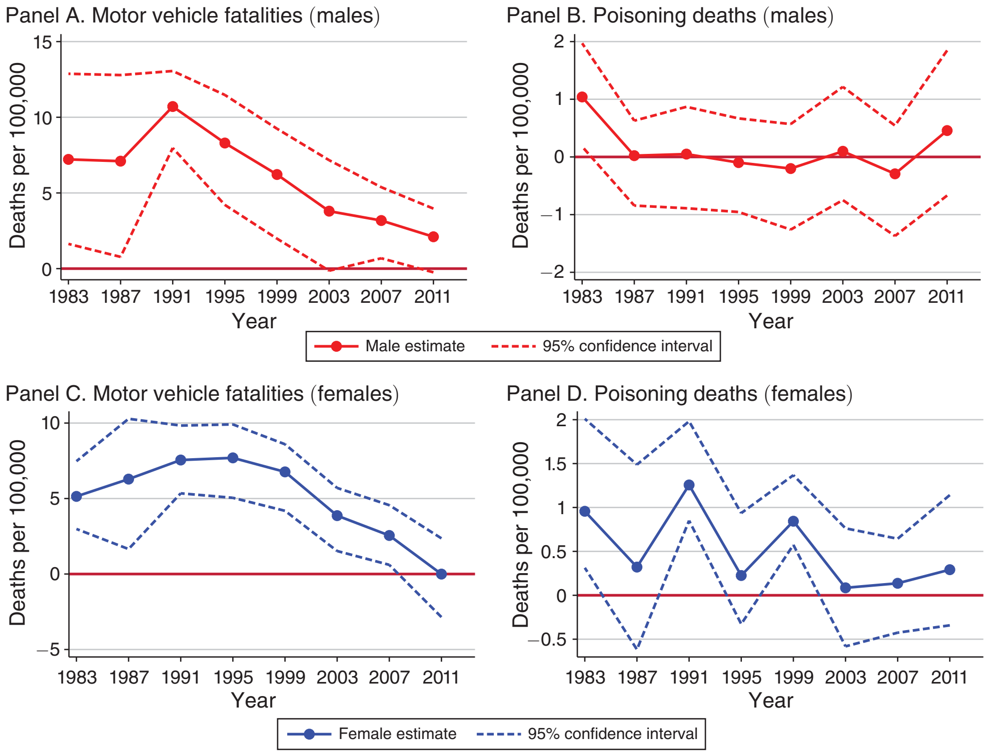

Figure 4. Trends in the Effect of Driving Eligibility on Motor Vehicle Fatalities and Poisoning Deaths

Notes: The figure plots MSE-optimal estimates of β from equation (1) separately for four-year bins. The dependent variable is deaths per 100,000 person-years. The dashed lines report robust bias-corrected 95 percent confidence intervals. Table 1 provides estimates for outcomes measured over the whole 1983–2014 sample period.

to our RD estimate, then they imply that total motor vehicle fatalities increase by 13.6 $(= 4.92 \times 1.24/0.45)$ deaths per 100,000 at the age cutoff.

In the online Appendix, we decompose the female poisoning death estimates by estimating our model separately for deaths classified as suicides versus accidents (Table A.1). A few results stand out in this exploratory analysis. First, the increase in female drug overdose deaths is caused by both accidents and suicides. Second, although most gas poisoning deaths among female teenagers are accidental in the aggregate, the increase in these deaths at the age cutoff is driven by suicides. Finally, the net effect of driving eligibility on total suicides is small and statistically insignificant because of an offsetting reduction in firearm suicides, suggesting that female teenagers who die by suicide substitute away from using firearms and toward using drugs and gas poisoning upon gaining access to a car.

We caution that the reliability of the suicide/accident classification is unclear. For example, some of the increase in accidental drug overdose deaths may in fact be suicides. We also lack statistical power to discern with confidence whether the increase in poisoning suicides reflects a net increase in suicide or substitution away from other methods of suicide, making it difficult to test different theories of youth suicide (Cutler, Glaeser, and Norberg 2001). Overall, we conclude that the rise in female poisoning deaths represents changes in both accidental deaths and suicides and that the effects on suicide might reflect a compositional shift in the method of suicide.

D. *Robustness*

The most common MDA in our sample is 16 years, which also happens to be the federal minimum legal working age as well as the minimum legal school-leaving age in many states. We do not believe these other laws confound our estimates, however. Analysis of the Add Health data shows small and statistically insignificant changes in working for pay or leaving school at the MDA cutoff (online Appendix Figure A.5). Moreover, our results are similar when we limit our analysis to states with MDAs that are not 16 years (online Appendix Tables A.2 and A.3).[10] Because the MDA differs from the minimum legal ages for both working and leaving school in this subsample (online Appendix B.5), we conclude that the MDA is the causal mechanism underlying our results.

Our analysis examines a large number of outcomes across two different subgroups. Although we adjust for multiple inference, our outcomes and subgroups were not specified prior to analysis. However, we emphasize that our most surprising result—the increase in female poisoning deaths illustrated in Figure 2—is far too large to be spurious. A multiple testing correction would need to adjust for many thousands of hypotheses to increase the unadjusted p-value ($p < 0.00001$) above the conventional significance level of 0.05.

Our online Appendix reports our main results separately by race, sex, and their pairwise combinations (Tables A.4–A.7). Whites are more likely than non-Whites to obtain a driver's license upon becoming eligible, consistent with prior studies (Shults and Williams 2013). This differential driving effect is reflected in the mortality estimates: motor vehicle fatalities increase the most at the cutoff for White males and White females, and poisoning deaths increase the most for White females. We also detect some modest seasonality in our estimate for motor vehicle fatalities, with effect sizes peaking during the summer months in both absolute and relative terms (Figure A.6).

Our estimates are not sensitive to using different bandwidth selection procedures or polynomial approximations or to imposing a uniform bandwidth of 24 months (online Appendix Tables A.10–A.12). Estimating our model using placebo cutoffs generates estimates that are centered around zero, and our motor vehicle fatality and female poisoning death estimates lie in the far tails of those placebo distributions (online Appendix Figure A.8).

V. Conclusion

Employing a regression discontinuity design, we estimate that driving eligibility increases teenage mortality at the MDA cutoff by 5.84 deaths per 100,000 over the

[10] For example, female poisoning deaths rise by 0.509 deaths per 100,000 at the cutoff in states where the MDA is not 16. This estimate is statistically significant, and its 95 percent confidence interval includes the full sample estimate of 0.747 deaths per 100,000 (Table 1).

1983–2014 time period. The relationship between this threshold estimate and changes in the MDA is ambiguous. For example, raising the MDA could reduce the effect of eligibility on mortality if older teenagers are more careful drivers or increase the effect if older teenagers are more likely to drink and drive. If we assume the effect remains constant, then our estimate implies that a one-year increase in the MDA would have saved 228 teenage lives annually during our sample period.[11] This estimate, which does not account for the additional people killed in motor vehicle accidents involving newly eligible teenage drivers, is similar in magnitude to the estimated benefits of raising the minimum legal drinking age (Carpenter and Dobkin 2009).

We also estimate that female poisoning deaths increase by 76 percent following driving eligibility, but we find no effect on male poisoning deaths. These findings reveal that teenage driving is an important causal factor behind sex disparities in teenage poisoning deaths. While the increase in female poisoning deaths reflects changes in both accidents and suicides, the specific behavioral mechanisms underlying our results remain unclear. Driving may enable female teenagers to purchase or consume drugs more easily and may affect mental health by altering social environments. We encourage future researchers to investigate these different possibilities.

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[11] There was an average of 3.9 million 16-year-olds alive in the United States during 1983–2014.

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