Alcohol-Related Relative Risk of Driver Fatalities and Driver Involvement in Fatal Crashes in Relation to Driver Age and Gender: An Update Using 1996 Data*

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ABSTRACT. *Objective:* To re-examine and refine estimates for alcohol-related relative risk of driver involvement in fatal crashes by age and gender as a function of blood alcohol concentration (BAC) using recent data. *Method:* Logistic regression was used to estimate age/gender specific relative risk of fatal crash involvement as a function of the BAC for drivers involved in a fatal crash and for drivers fatally injured in a crash, by combining crash data from the Fatality Analysis Reporting System with exposure data from the 1996 National Roadside Survey of Drivers. *Results:* In general, the relative risk of involvement in a fatal vehicle crash increased steadily with increasing driver BAC in every age/gender group among both fatally injured and surviving drivers. Among 16-20 year old male drivers, a BAC increase of 0.02% was estimated to more than double the relative risk of fatal single-vehicle crash injury. At the midpoint of the 0.08% - 0.10% BAC range, the relative risk of a fatal single-vehicle crash injury varied between 11.4 (drivers 35 and older) and 51.9 (male drivers, 16-20). With only very few exceptions, older drivers had lower risk of being fatally injured in a single-vehicle crash than younger drivers, as did women compared with men in the same age range. When comparable, results largely confirmed existing prior estimates. *Conclusions:* This is the first study that systematically estimated relative risk for drink-drivers with BACs between 0.08% and 0.10% (these relative risk estimates apply to BAC range midpoints at 0.09%.) The results clearly show that drivers with a BAC under 0.10% pose highly elevated risk both to themselves and to other road users. (*J. Stud. Alcohol* **61:** 387-395, 2000)

BASED ON extensive research over several decades, we now have overwhelming evidence showing that even blood alcohol concentration (BAC) levels as low as 0.02% impair driving-related skills. One such line of evidence grows out of laboratory research with dosed subjects (Moskowitz and Robinson, 1987; see also National Institute on Alcohol Abuse and Alcoholism, 1997, chapter 7). Confirming evidence also comes from field research that compares the BACs of crash-involved with noncrash-involved drivers to determine the relative risk of crash involvement (for a review, see Perrine et al., 1989; Zador, 1991).

According to National Highway Traffic Safety Administration (NHTSA) information, as of September 1999, 31 states defined driving with a BAC above 0.10% as a crime per se, while another 17 states plus the District of Columbia set their per se limit at 0.08%. (Under a per se law it is a crime to drive with a BAC at or above the proscribed level; two states, Maryland and South Carolina, do not have a per se law but a presumptive limit.) Due to a combination of legal measures, enforcement actions and changes in voluntary behavior patterns, alcohol-related fatalities have been declining for nearly 2 decades, both in absolute numbers and as a proportion of all fatalities. Nonetheless, there were still 15,936 alcohol-related traffic fatalities in the United States that accounted for nearly 38% of total traffic fatalities in 1998 (NHTSA, 1999), indicating that much more needs to be done.

The objective of the present research is to re-examine and refine relative fatal crash risk estimates, in a systematic fashion using more recent data. It extends similar prior work by the first author, in three important ways. First, we estimate relative risk for the policy-relevant BAC range of 0.08% to 0.10%. Second, we estimate relative risk for six driver groups: (1) driver fatalities in single-vehicle crashes, (2) driver involvements in single-vehicle fatal crashes, (3) driver fatalities in two-vehicle crashes, (4) driver involvements in two-vehicle fatal crashes, (5) driver fatalities in all crashes and (6) driver involvement in all fatal crashes. Third, we employ statistical methods to estimate both the effect of sampling roadside exposure and the effect of multiple imputation of missing BACs on the uncertainty of relative risk estimates.

Method

Data sources

Driver exposure data: the 1996 Roadside Survey. The 1996 National Roadside Survey (96NRS) of weekend nighttime drivers in the 48 contiguous states followed the same principles as its two predecessors (in 1973 and 1986). A

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sample of noncommercial operators of four-wheel motor vehicles was interviewed and breath-tested during a roughly 1-month period in the fall of 1996. Counties with a population of less than 20,000 were not sampled, and in counties with larger populations, roadways with average daily traffic below 2,000 were excluded from the surveys (for details, see Lestina et al., 1999). Using a geographically stratified multistage cluster sample, drivers were selected for interviews and breath tests. This survey was designed based on the National Automotive Sampling System/Crashworthiness Data System (NASS/CDS) (NHTSA, 1995). The first stage of the design comprised 24 primary sampling units (PSUs) employed by NASS/CDS, six each in the Northeast, South, West and Midwest regions. The second stage comprised a total of 46 police jurisdictions: 11-12 per region. At the third stage, square grids with sides roughly equal to 1 mile were superimposed on the sampled jurisdictions and then randomly sampled to obtain the requisite number of squares (this procedure was modified for areas with low road density). Once a square was chosen, the survey was conducted at the first safe area found in it by the survey team leader. Driver selection represented the final stage: the first driver who approached the site after an interviewer became available was stopped for the next interview. Field operations were conducted on Friday and Saturday nights during two 2-hour periods at separate sites: at one site between 10 pm and midnight, and at the other between 1 am and 3 am. Data from the 96NRS is only representative of locations and periods when drinking and driving is most prevalent (i.e., not of all times or roadways in the 48 contiguous states).

We adjusted driver sampling weights from the 96NRS for nonrespondents, and used the adjusted weights to approximate the statistical distribution of drivers on weekend nights (exposure), by gender, age (16-20, 21-34 and ≥ 35) and % BAC (0.000, 0.001-0.019, 0.020-0.049, 0.050-0.079, 0.080-0.099, 0.100-0.149 and 0.150+). For details on nonresponse adjustments, see Lestina et al. (1999) and Zador et al. (in press).

Data on drivers in fatal crashes. The Fatality Analysis Reporting System (FARS) is a census of all fatal motor vehicle crashes that occur on public trafficways in the United States and result in a fatality within 30 days. Although FARS is maintained by NHTSA of the U.S. Department of Transportation, the data in FARS are obtained through cooperative agreements with agencies in each state's government and are managed by regional contracting officer's technical representatives located in the 10 NHTSA regional offices. For basic data elements associated with a fatal vehicle crash, reporting is usually of very high quality with relatively few missing values; however, there is one exception: even in recent years, BACs were not available for many drivers involved in fatal crashes. To deal with this problem, NHTSA has employed a statistical method since the early 1980s for imputing missing BACs (Klein, 1986). More recently, the method of multiple imputation (Rubin, 1987) was adopted to handle the problem of missing BACs on FARS (Rubin et al., 1999). Under multiple imputation, each missing value is replaced by a small number of imputed values (10, in the present case) that are generated by a statistical procedure designed to reflect the statistical properties of the missing driver BACs. We used the 10 complete-data versions of FARS in our statistical analyses. Note that, although the data files for the multiple imputation method are available, NHTSA is not yet using the multiple imputation method for its published alcohol estimates. The same method used in previous years is to be used for the 1998 FARS estimates.

We selected drivers of four-wheel passenger vehicles who were 16 years of age and over and were involved in fatal crashes during 1995 or 1996 in 1 of the 48 contiguous states (NHTSA, 1995-96). The crash had to have occurred on a weekend night in a county with a 1990 population of at least 20,000; outside of special jurisdictions; and on a paved road that was not classed as an interstate, other urban freeway or expressway. There were only two notable differences between the exposure and the crash screening criteria, and both were disregarded to increase the sample size for drivers retained for the analyses. First, we accepted crashes that occurred between midnight and 1 AM, since those crashes were excluded from the exposure sample only to permit the survey team to change location, and not because BAC distribution between midnight and 1 AM was thought to be different. Second, we did not restrict crashes to the weekend nights during which the surveys were conducted. Including weekend nights for the whole year increased sample sizes almost 12-fold and introduced no substantial difference in the distribution of driver BACs since driver BACs varied little between the survey period and the rest of the year. We classified the drivers meeting these selection criteria by the number of crash-involved vehicles (one, two, and any number of vehicles) and by whether the driver was just involved in the crash or was also fatally injured in the crash. We thus defined six driver groups for analysis: drivers fatally injured in single-vehicle crashes, drivers involved in fatal single-vehicle crashes, drivers fatally injured in two-vehicle crashes, drivers involved in fatal two-vehicle crashes, drivers fatally injured in a motor-vehicle crash and drivers involved in a fatal motor-vehicle crash. We classified the six groups of driver fatalities and involvements by gender, age group and BAC, in the same way we classified the exposure sample.

Statistical methods

Using odds ratios and logistic regression to estimate relative risk. Following Zador (1991), we base our methods on the intuitive notion that comparisons between the frequency distribution of fatal-crash involvement by gender, age and BAC, and the frequency distribution of roadside exposure by gender, age and BAC, can provide a good yardstick for measuring the effect of these factors on the relative likelihood of fatal-crash involvement per unit of driving exposure. Since the 96NRS did not provide a national estimate for total miles

TABLE 1. Fatality and weighted survey counts for two populations

Population	Fatality count	Exposure count
Group 1	F1	E1
Group 2	F2	E2

driven on weekend nights, it was not possible to scale fatal involvement and exposure count ratios to the corresponding involvement rates per miles driven. However, since our involvement (or fatality) count per exposure count ratios are proportional to national involvement (or fatality) rates, dividing two such ratios (e.g., at differing BACs) gives the corresponding ratio of involvement (or fatality) rates. We have used data on driver fatalities or driver involvement from FARS and driving exposure from 96NRS to estimate driver involvement (or fatalities) per exposure, then divided such ratios to approximate the relative risk of driver involvement and driver fatality (for a general discussion of relative risk, see Schlesselman, 1982).

More specifically, consider a two-way table formed of fatality and (weighted) survey counts for two populations (Group 1 and Group 2) (Table 1).

The odds ratio (OR) compares the fatality/exposure ratio between Groups 1 and 2: OR = (F1/E1)/(F2/E2) (1). Taking Group 2 as the baseline, this odds ratio compares fatality odds in Group 1 to fatality odds in Group 2. Since odds ratios are scale invariant, we can substitute $cE1$ and $cE2$ in (1) for exposure counts $E1$ and $E2$, where c is the scaling constant, without affecting the numeric value of the OR. Now, for a large value of c, the odds in the numerator and the denominator of (1) have approximately the same value as the corresponding crash rates: F1/(F1+cE1) and F2/(F2+cE2). The unknown value of c that would scale up survey-based exposure counts to the national total of miles driven is extremely large relative to observed involvement/exposure ratios. Therefore, it is legitimate to use the odds ratios (F1/E1)/(F2/E2) to estimate relative risk, F1/(F1+cE1)/F2/(F2+cE2). Given this discussion, and following the common practice in epidemiology (Schlesselman, 1982), we used odds ratios to estimate relative risk, and henceforth we will refer to estimates of relative risk, rather than to estimates of odds ratios.

We used logistic regression, implementing the SAS procedure (SAS Institute, Inc., 1996), to model involvement (or fatality) counts relative to exposure counts in two sets of analyses for each of our six driver groups, and then estimated relative risk and its upper and lower 95% confidence intervals (CI) from model parameters. In this article, because of space limitations, we present results from only two of the six driver groups (for complete results, see Zador et al., in press). In one set of analyses, we estimated risk among sober drivers (BAC = 0%) by age and gender relative to the risk of male drivers between the ages of 21 and 34 (see Figure 1). In the other, to estimate risk as a function of BAC relative to sober drivers, we modeled involvement (or fatality) to exposure count ratios in terms of variables for age, gender and interactions of age, gender and BAC (see Table 2). Note that any nonsignificant interaction terms were not retained in the final model. We then estimated the relative risk (RR), adjusted for all covariates in the model, from the regression coefficient, say b, of the BAC variable in the form RR(BAC) = exp(bBAC). (For convenience, positive BACs were rescaled by a factor of 1,000 so that 0.1% would be entered in formulas as 100.) To illustrate: The relative risk of receiving a fatal injury in a single-vehicle crash by a driver age 21-34 whose BAC is 0.13% was estimated at RR(0.13) = exp(0.029 × 130) = 43.4, where 130 = 0.13 × 1,000 and b = 0.029 is the regression coefficient from Table 2 for the parameter "BAC, age 21-34" (i.e., among drivers age 21-34, a BAC of 0.13% increased the chance of being killed in a single-vehicle crash by a factor of about 43).

For the purpose of fitting models, we represented BAC class intervals (by gender and age) by average driver BAC. However, to facilitate comparisons across driver populations, we present relative risk estimates at constant BACs (0.0%, 0.01%, 0.035%, 0.065%, 0.090%, 0.125%) that correspond to class interval midpoints for the first five BAC categories, and 0.220% for the last BAC category, corresponding to the average BAC of those with BAC values greater than 0.15%.

Model performance

We assessed model performance using four statistics: (1) the heterogeneity factor estimated as the Pearson chi-square statistic divided by its degrees of freedom, (2) the maximum of rescaled R^2, (3) the Hosmer-Lemeshow goodness-of-fit test p value and (4) the Shapiro-Wilk statistic p value for testing the normality of standardized Pearson residuals. For data that are conditionally binomial, the dispersion parameter is approximately equal to 1. A heterogeneity factor substantially larger than 1 is indicative of overdispersion (i.e., more variation than would be expected under the assumption that conditionally, on the sum of exposure and fatality counts, the fatalities were binomially distributed). We adjusted all variance and confidence bounds for the presence of overdispersion. The maximum rescaled R^2 (Nagelkerke, 1991) can be used to assess model quality somewhat in the manner of the customary R^2 statistic for linear regressions. The Hosmer-Lemeshow test statistic is a direct measure for a logistic regression model's ability to predict outcome probabilities (Hosmer and Lemeshow, 1989). We have applied the Shapiro-Wilk test to determine whether standardized residuals followed an approximate normal distribution. In the ideal case, heterogeneity factor and rescaled R^2 are near 1, and the p values are between 0.05 and 0.95.

Variance estimation

As described earlier, the 1996 National Roadside Survey had a complex, multistage design (Voas et al., 1997). If no special steps were taken, standard statistical packages would tend to underestimate true variability for data collected under a complex design. We used Fay's method of balanced repeated replications (BRR) to obtain design-consistent variance estimates for important model parameters. This involved three steps: First, we used Westat's implementation of Fay's method (SPSS, 1998) to create replicate weights from the adjusted full-sample weight. Second, we repeatedly estimated our models using the full-sample weight and each set of replicate weights. Finally, we combined the resulting estimates using a simple formula to obtain design-unbiased variance estimates (for details, see Zador et al., in press).

As part of data preparation, NHTSA had replaced missing driver BACs on FARS by 10 imputed BAC values (Rubin and Schaffer, 1998). On the one hand, having imputed values on FARS made it possible to employ complete-data methods in our analyses. On the other hand, unless special care is exercised, using imputed values as if they were actual values will result in underestimating variability (Rubin, 1987). We followed Rubin's two-step method to eliminate this bias. As a first step, we re-estimated our models using each of the 10 imputed BAC values, and averaged the results. Next, we combined the 10 sets of estimates into a single unbiased variance estimate. Since there were 12 replicate weights for exposure and 10 multiply-imputed completed sets of driver counts, we estimated model parameters for each of the six driver populations a total of 130 times in order to compute imputation-adjusted and design-consistent parameter variances (for details, see Zador et al., in press).

Results

Figure 1 displays fatality risk in single vehicle crashes and involvement risk in all fatal crashes for sober drivers by age and gender, relative to the comparable risk for sober male drivers ages 21-34 (for whom, in these comparisons, the baseline risk was taken to be 1.00). With few exceptions, relative risk decreased with age, was lower among women than men, and was higher for fatalities in single-vehicle crashes than for involvements in fatal crashes. Some of the baseline risk differences were striking, especially for fatalities in single-vehicle crashes by age. Among sober male drivers, the relative risk of a fatal driver injury in a single-vehicle crash was 1.75 for drivers 16-20 and 0.71 for drivers over age 34. For women, it was 1.18 in the youngest group and 0.28 in the oldest group. In general, relative risk estimates differed significantly between drivers of the same gender from adjacent age groups (except for the two highest age-group women) and between men and women of the same age group (except for the youngest age group).

Figure 1 also shows that the pattern of results for involve-

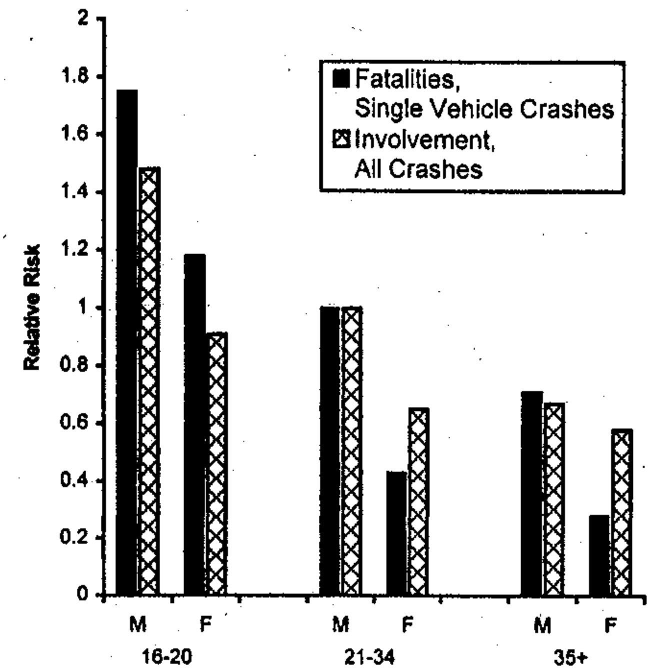

FIGURE 1. Risk of driver fatalities in single-vehicle crashes and driver involvement in fatal crashes at BAC = 0, by age and gender, relative to men 21-34

ment in fatal crashes and for fatalities in single-vehicle crashes were somewhat different, especially for women. First, although relative risk still decreased with increasing age among both men and women, the comparable decreases by age were less pronounced for driver involvement in all crashes than for fatalities in single-vehicle crashes. Moreover, for women in the age groups 21-34 and 35 and over, the relative risk of being involved in a fatal crash exceeded the relative risk of a fatal injury in a single-vehicle crash. For analogous results for the other four driver groups, and for the distribution by age and gender of exposure, involvement and fatality counts, see Zador et al. (in press).

Tables 2, 3 and 4 present logistic regression model parameter estimates and standard errors, model-based estimates of proportionate increase in relative involvement risk and fatal injury risk associated with a 0.02% increase in BAC, and model-based relative risk estimates with their confidence intervals, respectively.

Models in Table 2 contain nine parameters: indicator variables ($n = 3$) for age groups (16-20, 21-34 and 35+); indicator variable for being female; indicator variable for interaction term involving low positive BAC (.001 - .019%) among drivers >20; continuous variables (interaction terms; $n = 3$) for the effect of BAC by age group (16-20, 21-34 and 35+); and continuous variable (interaction term) for the effect of BAC for women ages 16-20.

As evidenced by the results of the Hosmer-Lemeshow goodness-of-fit test and the Shapiro-Wilk test for the normal distribution of standardized Pearson residuals, relative risk

TABLE 2. Logistic regression coefficients in models for risk of driver fatalities and driver involvement in single-vehicle crashes, in two-vehicle crashes and in all crashes as a function of variables for age, gender and interactions of age, gender and BAC. Data from the 96NRS and the 1995-96 FARS.

		Single-vehicle crashes		Two-vehicle crashes		All crashes	
Variable	Parameter	Fatalities	Involvements	Fatalities	Involvements	Fatalities	Involvements
Age 16-20	Coefficient[b]	−1.547	−0.572	−2.184	−0.873	−1.077	0.085
	SE	0.072	0.063	0.060	0.057	0.065	0.057
Age 21-34	Coefficient	−2.352	−1.205	−2.643	−1.187	−1.654	−0.331
	SE	0.042	0.028	0.051	0.034	0.036	0.025
Age 35+	Coefficient	−2.540	−1.656	−2.425	−1.291	−1.672	−0.591
	SE	0.043	0.039	0.037	0.036	0.036	0.039
Female	Coefficient	−0.580	−0.509	−0.065	−0.265	−0.351	−0.356
	SE	0.069	0.053	0.054	0.043	0.053	0.042
BAC < .019, age 21+	Coefficient	−2.861	−1.889	−1.593	−2.004	−2.031	−1.925
	SE	0.375	0.126	0.121	0.134	0.137	0.106
BAC[a], age 16-20	Coefficient	0.044	0.039	0.032	0.031	0.041	0.035
	SE	0.007	0.006	0.005	0.005	0.006	0.005
BAC, age 16-20, female	Coefficient	−0.014	−0.015	−0.006	−0.015	−0.016	−0.016
	SE	0.006	0.005	0.006	0.005	0.006	0.005
BAC, age 21-34	Coefficient	0.029	0.024	0.023	0.019	0.026	0.020
	SE	0.001	0.001	0.001	0.001	0.001	0.001
BAC, age 35+	Coefficient	0.027	0.024	0.020	0.018	0.023	0.020
	SE	0.001	0.001	0.001	0.001	0.001	0.001
Model diagnostic							
Heterogeneity factor		1.6979	1.7774	1.8783	3.3159	2.0918	3.7070
Max-rescaled R^2		0.6844	0.4935	0.6524	0.3142	0.5297	0.3171
H-L goodness-of-fit, p		0.1998	0.6806	0.0317	0.0001	0.4008	0.0002
Normality of residuals, p		0.2813	0.0606	0.5701	0.4175	0.2189	0.0165

[a]BAC represents driver BAC as a continuous variable.
[b]A positive (negative) parameter indicates that variable and risk change in the same (opposite) directions.

was adequately represented by the models in Table 2 for three of the driver groups: drivers involved in a fatal single-vehicle crash, drivers killed in a single-vehicle crash and drivers killed in any vehicle fatal crashes. While the Hosmer-Lemeshow test statistic ($p = .032$) rejected the hypothesis of model fit for fatally injured drivers in two-vehicle crashes, the regression model explained 65% of all explainable relative risk variation, and the standardized Pearson residuals were normally distributed. Overall, we deem model fit acceptable for driver fatalities in two-vehicle crashes. In contrast, the models performed poorly for the two remaining driver groups—drivers in fatal crashes involving two vehicles or drivers in fatal crashes involving any number of vehicles.

We explored, in considerable detail, the way our models broke down for fatal two-vehicle crashes. We examined model fit statistics for the models in Table 2 and for several other model specifications, including specifications obtained by stepwise regression (for a summary of results for a few of the dozens of models that were examined, see the Appendix). The results showed clearly that sober driver involvement in two-vehicle crashes is not closely related to driver involvement at positive BACs, and we discovered that only the inclusion of indicator variables representing overall sober driver risk, and sober driver risk by age and gender, would produce acceptable model fit. This result was, in fact, not too surprising—for two reasons. First, in crashes involving more than a single vehicle, some drivers may be innocent (and probably sober) victims whose vehicles were struck by a high BAC at-fault driver. Second, in multivehicle crashes, crash configuration and vehicle occupancy become important determinants of relative risk. However, we decided not to use regression models that included sober driver risk variables (e.g., main effect for zero BAC, zero BAC by age interaction, etc.; see Appendix) because it was not clear how these models can be used to estimate relative risk with BAC = 0 as the baseline. Therefore, relatively poor model fit notwithstanding, we believe that the relative risk estimates presented from the model parameter estimates in Table 2 provide reasonable, albeit conservative, approximations of the true relative risk, even for driver involvement in multi-vehicle fatal crashes. Additional research will be needed to improve model fit for these driver groups.

Table 3 shows model-based estimates for factor of proportionate increase in relative risk associated with an increase of 0.02% in BAC level for each driver group, by age and gender. Of noteworthy mention, it was estimated that each 0.02 percentage point increase in the BAC of a driver with a nonzero BAC more than doubled the risk of receiving a fatal injury in a single-vehicle crash among male drivers aged 16-20, and nearly doubled the comparable risk among the other driver groups. Proportionality factors were estimated from age-specific regression coefficients of BAC in Table 2, except that for female drivers aged 16-20 the estimates were adjusted for the effect of being female. For the relative risk estimates in subsequent tables, relative risk was

TABLE 3. Model-based estimates for factor of proportionate increases in the relative risk of driver fatalities and driver involvement associated with a 0.02% increase in BAC. Estimates for single-vehicle crashes, two-vehicle crashes and all crashes by age and gender. Data from the 96NRS and the 1995-96 FARS.

	Single-vehicle crashes		Two-vehicle crashes		All crashes	
	Fatalities	Involvements	Fatalities	Involvements	Fatalities	Involvements
Male						
Age 16-20	2.41	2.17	1.94	1.84	2.29	2.01
Age 21-34	1.78	1.62	1.56	1.45	1.66	1.51
Age 35+	1.73	1.62	1.49	1.44	1.61	1.50
Female						
Age 16-20	1.80	1.63	1.71	1.39	1.65	1.47
Age 21-34	1.78	1.62	1.56	1.45	1.66	1.51
Age 35+	1.73	1.62	1.49	1.44	1.61	1.50

also adjusted for the effect of low BAC (0.001 - 0.019%) for drivers 21 or older.

The relative risk of receiving a fatal injury in a single-vehicle crash increases steadily with increasing driver BAC for both men and women in every age group with one exception (see Table 4). Among all male and female drivers, except those in the 16-20 age group, the relative risk of receiving a fatal injury is lower for drivers with a positive BAC under 0.02% than for drivers with 0.0% BAC. Remarkably, however, for the 16-20 age group, the comparable relative risk was substantially increased even at this low positive BAC, by 55% among men and by 35% among women. Looking at relative risk across the six age and gender groups, we find that at a BAC of 0.035%, it was elevated by a factor between 2.6 and 4.6; at a BAC of 0.065%, by a factor between 5.8 and 17.3; at a BAC of 0.09%, by a factor between 11.4 and 52; at a BAC of 0.125%, by a factor between 29.3 and 240.9; and at a BAC of 0.220%, by a factor between 382 and 15,560. The relative risk increased fastest for men aged 16-20 and slowest for drivers of either gender aged 35 and over. In general, controlling for age, relative risk increased faster for men than for women, and controlling for gender, it increased faster for drivers aged 16-20 and slowest for drivers 35 and over. Table 4 shows that this pattern of results is quite similar for the relative risk of driver involvement in fatal crashes. In addition, in every comparison, relative risk increased faster with increasing BAC for fatally injured drivers than for driver involvement in fatal crashes.

TABLE 4. Model-based risk of driver fatalities in single-vehicle (SV) crashes and driver involvement in all fatal crashes (All) as a function of driver BAC by gender and age, relative to sober drivers of the same age and gender (data from the 96NRS and the 1995-96 FARS[a])

	Crash type	000	BAC 010-019	020-049	050-079	080-099	100-149	150+
Male								
Age 16-20	SV	1.00	1.55	4.64	17.32	51.87	240.89	15,559.85
		–	(1.36-1.76)[b]	(2.97-7.26)	(7.56-39.70)	(16.45-163.57)	(48.87-1,187.33)	(939.22-257,777.67)
	All	1.00	1.42	3.44	9.94	24.03	82.73	2,371.74
		–	(1.28-1.58)	(2.37-4.99)	(4.98-19.82)	(9.23-62.53)	(21.91-312.31)	(228.91-24,574.14)
Age 21-34	SV	1.00	0.08	2.75	6.53	13.43	36.89	572.55
		–	(0.04-0.16)	(2.53-2.98)	(5.61-7.60)	(10.89-16.57)	(27.57-49.36)	(342.99-955.76)
	All	1.00	0.18	2.04	3.76	6.25	12.74	88.13
		–	(0.14-0.22)	(1.90-2.19)	(3.28-4.30)	(5.18-7.54)	(9.81-16.54)	(55.68-139.51)
Age 35+	SV	1.00	0.07	2.57	5.79	11.38	29.30	381.68
		–	(0.04-0.16)	(2.34-2.84)	(4.84-6.93)	(8.87-14.60)	(20.73-41.42)	(207.56-701.86)
	All	1.00	0.18	2.02	3.70	6.13	12.41	84.13
		–	(0.14-0.22)	(1.83-2.24)	(3.06-4.48)	(4.71-7.98)	(8.61-17.89)	(44.19-160.17)
Female								
Age 16-20	SV	1.00	1.35	2.86	7.04	14.91	42.63	738.36
		–	(1.21-1.50)	(1.96-4.16)	(3.50-14.14)	(5.68-39.15)	(11.15-163.01)	(69.67-7,824.89)
	All	1.00	1.22	1.98	3.56	5.80	11.50	73.62
		–	(1.10-1.34)	(1.40-2.80)	(1.88-6.76)	(2.39-14.10)	(3.35-39.47)	(8.41-644.68)
Age 21-34	SV	1.00	0.08	2.75	6.53	13.43	36.89	572.55
		–	(0.04-0.16)	(2.53-2.98)	(5.61-7.60)	(10.89-16.57)	(27.57-49.36)	(342.99-955.76)
	All	1.00	0.18	2.04	3.76	6.25	12.74	88.13
		–	(0.14-1.22)	(1.90-2.19)	(3.28-4.30)	(5.18-7.54)	(9.81-16.54)	(55.68-139.51)
Age 35+	SV	1.00	0.07	2.57	5.79	11.38	29.30	381.68
		–	(0.04-0.16)	(2.34-2.84)	(4.84-6.93)	(8.87-14.60)	(20.73-41.42)	(207.56-701.86)
	All	1.00	0.18	2.02	3.70	6.13	12.41	84.13
		–	(0.14-0.22)	(1.83-2.24)	(3.06-4.48)	(4.71-7.98)	(8.61-17.89)	(44.19-160.17)

[a]Although relative risk estimates for the male and female 21-34 (and 35+) age groups are the same, the groups were depicted separately in the tables.
[b]95% confidence interval for relative risk.

In general, the pattern of results for the other driver groups was quite similar to the pattern described above (see Zador et al., in press). There are two major differences among the other driver groups: (1) For fatally injured drivers, relative risk increased more slowly with increasing BAC in two-vehicle than in single-vehicle crashes. As indicated earlier, this was to be expected since in multivehicle fatal crashes some involved drivers were likely to be no more than marginally at-fault. (2) Since most fatally injured drivers were killed in a single-vehicle or in a two-vehicle crash, the overall rate of increase in relative risk was bracketed by the rates of increase for single-vehicle and two-vehicle crashes.

Discussion

Confirmatory findings

This study generally confirmed that the relative risks of fatal injury and fatal crash involvement increase steadily with increasing driver BAC within each of the six driver age and gender groups studied. The only exception was that among drivers 21 and over, relative risk was lower at near-zero positive BAC than at zero BAC. The classic Grand Rapids study by Borkenstein et al. (1974) found a similar "dip" in the risk curve. Hurst (1973) showed that controlling self-reported drinking frequency eliminates the Grand Rapids dip. The customary interpretation of these results is that the anomalous dip probably results from differing alcohol tolerance between crash-involved and noncrash-involved drivers. Since drinking frequency data were not available in our study, we were unable to estimate risk curves by drinking frequency. With few exceptions, relative risk was found to decrease with increasing driver age at every BAC level, for both men and women—a finding that extends similar age trends reported for more moderate BACs by Zador (1991).

The current study also confirms the substantially higher relative risk for involvement in a single-vehicle crash of young drivers at a zero BAC as previously reported by Mayhew et al. (1986). In addition, female drivers exhibited substantially lower relative risk than male drivers of the same age. To a somewhat lesser extent, both sets of findings were also true for most of the other five driver groups studied.

In this study, lower and upper 95% confidence bound estimates for relative risk as a function of driver BAC take into account both the sampling variation of the roadside driver exposure sample and the effect of multiple BAC imputations performed by Rubin and Schaffer (1998) for NHTSA. Not surprisingly, relative risk confidence intervals are wide (e.g., lower and upper confidence bounds were 16.5 and 164 for male drivers ages 16-20 killed in single-vehicle crashes with a BAC between 0.08% and 0.10%; these relative risk estimates apply to BAC range midpoints at 0.09%). We note that the width of 95% confidence intervals increases with increasing BACs for mathematical reasons (both relative risk and its confidence bounds depend exponentially on the corresponding logistic regression parameters). We also note that, allowing for comparable variation in prior estimates, the relative risk estimates presented here are largely in line with estimates published elsewhere. (Relative risk estimates presented in this article differ in several ways from similar estimates in Zador [1991]. In the earlier study, the baseline BAC group was defined to include drivers at or below a BAC of 0.01%, age groups and BAC groups were defined differently, driver fatalities were included from only 29 states with low rates of missing BACs, missing BACs were not imputed, and the numeric BAC values were not used in analyses except to classify drivers.)

New findings

This is the first study that estimated relative risk from compatible data sources using the same methods for six groups of drivers involved in fatal crashes that were defined by the number of crash-involved vehicles and by whether the driver was only involved or also fatally injured in the crash. Drivers killed in single-vehicle crashes are of particular interest for assessing the *pure* effect of drink-driving because in single-vehicle crashes: (1) driver fault is not shared, (2) crash configuration is less of a factor, (3) vehicle occupancy is not relevant and (4) the seating position of the fatally injured occupant is fixed. In two-vehicle crashes, the possibility that fault may be split between two drivers, one or both of whom may have a (possibly different) positive BAC, would seem to make it difficult to estimate the pure effect of BAC on crash risk. It was all the more gratifying to find that the relative risk of a fatal driver injury depends on driver BAC in almost the same way for single-vehicle crashes and two-vehicle crashes, provided that the relative risk model of two-vehicle crashes statistically accounted for the possible roles of not at-fault sober drivers (see Appendix). In this study, we focused on the general effect on relative risk of a positive driver BAC, rather than on its pure effect. Our main statistical model for estimating relative risk did not, therefore, adjust relative risk estimates for the overrepresentation of sober (probably not-at-fault) drivers. Consequently, the model we used in this study appears to have generally underestimated the pure effect of positive driver BAC on relative risk, except for drivers in single-vehicle crashes.

As noted earlier, this study confirmed that relative risk and driver age are inversely related at every BAC. However, somewhat surprisingly and in part contrary to Zador (1991), we also found that for the 16-20 age group, women had lower relative risk than men at every BAC. For BACs of 0.02% and over, this lower relative risk was roughly comparable to relative risk among adult drivers aged 21 to 34—an important finding because of the increasing nighttime presence of young female drinking drivers observed in the 96NRS. That most recent survey found more, although not significantly more, female than male drinking drivers in the 16-20 age group. Perhaps the lower relative risk could be attributed to women driving more cautiously than their male age counterparts.

Finally, this is the first study that systematically estimated relative risk for drink-drivers with BACs between 0.08% and 0.10% (these estimates apply to BAC range midpoints at 0.09%), and the relative risk estimates obtained here provide clear evidence that drink-driving at BACs under 0.10% is very dangerous. For driver fatalities in single-vehicle crashes with a BAC in this range, relative risk estimates ranged from a low of 11.4 for drivers age 35 and over, to a high of 51.9 for male drivers under 21, and even the lowest among the six lower confidence bounds indicated a nearly six-fold rise in fatality risk. We found a similar pattern overall (i.e., for all driver groups studied), although with smaller magnitudes. Drivers age 35 and over had the lowest relative risk of involvement at about 6.1, followed by those in the 21-34 age group at 6.3. The youngest male age group had a relative risk of about 24—four times that of the other age groups (Table 4). Of course, relative risk was considerably higher for drivers with a BAC between 0.10% and 0.15% and ranged, for driver fatalities in single-vehicle crashes, from 29 for drivers age 35 and over, to 241 for male drivers under 21. Relative risk for drink-drivers with a BAC at or above 0.15% ranged from 382 for drivers age 35 and over, to 15,560 for male drivers under 21.

Policy implication

There is considerable evidence that lowering state BAC limits to 0.08% from 0.10% reduced fatal vehicle crashes (e.g., Hingson et al., 1996; Voas and Tippetts, 1999) and, according to recent NHTSA information, 17 states plus the District of Columbia have defined a driver BAC of 0.08% as illegal per se. New findings from this study lend support to lowering the illegal per se limit by showing that driving at BACs under 0.10% is indeed very dangerous. In addition, an ongoing laboratory investigation at the Southern California Research Institute, with participation by the first author of this article, has provided strong further evidence that driving-related performance is seriously impaired at even very low BAC levels, even among experienced drinkers.

Obviously, baseline differences are important for comparing driver groups in absolute terms since overall crash risk is affected both by baseline risk differences among sober drivers and by age- and gender-related differences in the effects of drink-driving. When considering policy options for young drivers, for example, it is important to bear in mind overall risk, not just sober-driving or drink-driving risks. Since young male drivers in the 16-20 age group start from an already high baseline risk level in all driver groups (two driver groups are shown in Figure 1), even at the slightly elevated BACs in the 0.02% - 0.05% range, this group experienced fatal driver injuries in single-vehicle crashes more than eight times as often as did sober male drivers age 21-34 at comparable BACs. Policy measures designed to reduce drink-driving and alcohol-related crashes in the youngest age group include the enforcement of minimum drinking-age laws that prohibit the purchase of alcoholic beverages by persons under age 21, and the establishing and enforcing of near-zero BAC limits (zero tolerance) for drivers under 21. Complementary strategies designed to reduce both sober-driving and drink-driving crashes among the youngest drivers include graduated licensure and nighttime driving restrictions (NHTSA, 1998).

With the new findings of this study, and parallel results currently being observed in laboratory studies at the Southern California Research Institute, one can state with confidence that driving at BAC levels below 0.10% is very dangerous. Thus, reducing BAC limits from 0.10% to 0.08% is an effective method of saving lives. Moreover, these results show that with such elevated relative risks, reducing drink-driving at any BAC is likely to further reduce alcohol-related motor vehicle fatalities in the United States.

Appendix

Alternative logistic regression models for involvement/exposure ratios

We used stepwise logistic regression to model involvement/exposure ratios in terms of optimally selected combinations of 18 main effects and interactions for age; age by gender; zero BAC (BAC0 = 1 if BAC = 0, BAC0 = 0 if BAC > 0); zero BAC by age; BAC under 0.02% (BAC02 = 1 if BAC < 0.02, BAC02 = 0 if BAC ≥ 0.02); and BAC under 0.02% by age. The model selected by stepwise regression for driver involvement in fatal two-vehicle crashes brought all diagnostic statistics into the acceptable range. Specifically, the final model was judged adequate (based on $p = 0.57$, 6 df) for the Hosmer-Lemeshow statistic of model fit. Unfortunately, this improvement was achieved by including two terms for sober drivers that showed sober drivers, especially those under age 35, to be overrepresented relative to their expected involvement frequency computed from models in which the effect of BAC on involvement is linear on the Logit scale. However, this model was not used for generating our estimates because it was not clear how it could be used to estimate relative risk with BAC = 0 as the baseline.

Following are the regression coefficients of BAC by age in three models for involvement/exposure ratios: Driver involvement, two-vehicle crashes, zero BAC term included: 16-20, 0.041; 21-34, 0.024; 35+, 0.024; Fatally injured drivers, single-vehicle crashes: 16-20, 0.044; 21-34, 0.029; 35+, 0.027; Driver involvement, two-vehicle crashes: 16-20, 0.031; 21-34, 0.019; 35+, 0.018. Two of these models, that did not include zero-BAC terms (one for fatally injured drivers in single-vehicle crashes and the other for driver involvement in two-vehicle fatal crashes) are also included in Table 2. The third model, which includes a zero-BAC term, is for driver involvement in two-vehicle fatal crashes. It is not surprising that, in every age group, the regression coefficients of BAC for driver involvement in fatal two-vehicle crashes are substantially higher in the model that incorporates a zero-BAC term than in the corresponding model that does not (this finding is actually a mathematical consequence of the fact that zero-BAC coefficients are always positive). It is surprising, however, that in every age group the regression coefficients of BAC in the model for driver involvement in fatal two-vehicle crashes that incorporates a zero-BAC term are only slightly smaller than similar age-group regression coefficients for fatally injured drivers in single-vehicle crashes. This suggests that positive BAC affects single-vehicle fatalities and two-vehicle crash involvement to roughly the same extent, provided that not-at-fault sober drivers are suitably accounted for. However, until confirmed by additional research, this finding must be considered more as a hypothesis than a definitive conclusion. Note, however, that similar suggestions were also made in Zador (1991).

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