

**Evaluation of California's
Graduated Licensing System**

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ABSTRACT

Objective: Evaluations of several state graduated licensing programs have indicated reductions in injury and fatal crashes involving 16-year-old drivers. However, results of evaluations of California's July 1, 1998 graduated licensing law have been mixed, with one study showing no overall reduction in crashes. The present study attempted to clarify the effects of California's law.

Methods: Auto-Regressive Integrated Moving Average time series analysis was used to account for pre-existing downward trends and seasonality in monthly 16-year-old driver involvements in injury or fatal crashes per 10,000 population, based on 1995-2003 data. Monthly 24-55-year-old driver involvements in injury/fatal crashes per 10,000 population were used as an additional check for statewide trends.

Results: Implementation of the graduated licensing law resulted in an estimated 23 percent reduction in 16-year-old driver injury and fatal crash involvements. An estimated 8,052 16-year-old driver involvements in injury and fatal crashes were prevented in the 66 months following the law's implementation. Sixteen-year-old driver involvements in injury and fatal crashes with injured teenage passengers declined by an estimated 38 percent, or 3,953 fewer crash involvements. Nighttime and daytime 16-year-old driver involvements in injury and fatal crashes were reduced by similar amounts.

Conclusions: California's graduated licensing law has reduced 16-year-old driver involvements in crashes. The results differ from those of Masten and Hagge (2003) who found no overall effects for 16 year-olds. Their modeling process was overly restrictive and did not adequately account for seasonal and other systematic nonlinear periodic trends in injury and fatal crashes.

INTRODUCTION

Since the mid-1990s all US states and most Canadian provinces have adopted some form of graduated licensing. Graduated licensing phases in young beginners to full-privilege driving, protecting them while they learn. A basic system starts with an extended learner stage, allowing the accumulation of low-risk practice driving over time. Once licensed, restrictions are placed on late-night driving and transporting young passengers, which are especially high-risk activities for young drivers (Chen et al., 2000; Williams and Preusser, 1997).

There is considerable variation among jurisdictions in their licensing laws. Some do not have all the stages of a full graduated system, and others have weak versions of requirements, for example, night restrictions that begin as late as 1 a.m. or passenger restrictions that allow as many as three young companions. Nevertheless, state and provincial evaluations typically have found reductions in crashes resulting from graduated licensing. There is substantial variation in reported effects, but most are in the 20-30 percent range for crashes among young beginners. Positive effects were reported in Connecticut, Florida, Georgia, Kentucky, Michigan, North Carolina, Ohio, Oregon, Wisconsin and in the provinces of

British Columbia, Nova Scotia, Ontario, and Quebec (Agent et al., 2001; Boase and Tasca, 1998; Bouchard et al., 2000; Foss et al., 2001; Fuhr et al., 2005; Mayhew et al., 2003; Ohio Department of Public Safety, 2001; Raymond et al., 2004; Rios et al., 2006; Shope and Molnar, 2003; Ulmer et al., 2000; Ulmer et al., 2001; Wiggins, 2005).

California was one of the early adopters of graduated licensing. Its system became effective July 1, 1998. The main features included an increase in the learner period from 1 to 6 months, a requirement that parents or guardians certify at least 50 hours of practice accumulated during this period, a 12-month restriction after licensure on driving unsupervised between midnight and 5 a.m., and a 6-month restriction on driving unsupervised with passengers younger than 20.

Statewide evaluations of California's graduated system have yielded conflicting results. A study by Rice et al. (2004) compared 1 year of pre-law and 2 years of post-law data, with 25-34 year-olds as a control. The authors reported that graduated licensing reduced the crash rate for 16-17 year-olds by 28 percent, comparing 1997 with 2000, and by 17 percent comparing 1997 with 2001. A study by Cooper et al. (2004) reported a 17 percent reduction in 16-year-old licensed driver crashes, based on data from 21 pre-law and 21 post-law months. However a third, more rigorous, study of the law change by Masten and Hagge (2003) found no overall effects on 16-year-old driver injury and fatal crash rates per 1,000 16 year-olds. All three studies found evidence of positive effects of California's passenger restriction, and two studies, including the one finding no overall effects, reported reductions in crashes during restricted nighttime hours.

The three studies used the same crash data source but differed in analytic techniques, outcome measures, and the pre- and post-law periods. Because 16 year-olds' crashes were trending downward in California since the early 1990s, methods have to be used to avoid interpreting mere continuations of the downward trend as indicating positive effects of the law. The study by Masten and Hagge (2003) had a relatively long baseline period, beginning in 1994, and used time series analysis, which can take into account pre-existing data trends, seasonality, and serial dependence among successive observations (Wilson, 2001). The authors found no effect of the law change and noted that the lack of effect was "not surprising," primarily because of survey findings that prior to the law many teenage drivers already were complying with the requirements that subsequently were legislated and that post-law compliance was incomplete. However, it is likely that those conditions have existed in every jurisdiction where graduated licensing was enacted, including jurisdictions where crash reductions were found.

Given the conflicting results of existing studies, further analyses are needed to clarify the effects of California's graduated licensing law. Prior studies were limited to post-law periods through 2001. Two more years of post-law data available at the time of the present study were used to consider longer-term effects of the law.

METHOD

Masten and Hagge (2003) used a common analytical technique for evaluating the law change, relying on month-to-month differencing to cancel out a pre-existing linear trend. However, that procedure does not remove seasonal and other periodic, inherent nonlinear trends. An alternative is the use of autoregressive terms in the modeling process. This amounts to describing and adjusting for inherent linear and nonlinear trends.

Monthly statewide injury and fatal crash data from January 1995 through December 2003 were analyzed to determine the effects of the July 1, 1998 graduated licensing law changes. The starting date for the pre-law period was chosen to minimize any potential effects of California's zero tolerance law that went into effect January 1, 1994, tightening the restrictions on drinking and driving for drivers younger than 21. It is possible that the zero tolerance law had an effect on teenage driver rates of involvement in injury and fatal crashes as enforcement and awareness levels of that law grew throughout 1994. Crash data for years prior to 1994 are not available, so any effects of the zero tolerance law cannot be determined. Therefore, the starting date for the present study was January 1, 1995 to minimize any potential confounding effects of the zero tolerance law.

Crash data were obtained from the California Highway Patrol's Statewide Integrated Traffic Records System (SWITRS). Analyses were based on the number of drivers of the age of interest involved in injury and fatal crashes each month per 10,000 population of that age. Population estimates came from the US Census Bureau for each year. To create monthly population totals from the annual estimates, any population increase or decrease in successive years was interpolated equally across months of the year.

Analyses

Analyses of trends in driver involvements in injury and fatal crashes per 10,000 population of that driver age used an interrupted time series design. The Auto-Regressive Integrated Moving Average (ARIMA) method was used to determine if there was a change in the rate of driver crash involvements coincident with the law change. The analyses used the rate of 24-55-year-old driver involvements in injury and fatal crashes as a covariate to account for fluctuations in driving conditions and unknown factors influencing the overall crash rates.

The intervention for each analysis was modeled as a gradual permanent change beginning in July 1998 and increasing linearly each month to full implementation by July 1999. The linear gradual change in the law's implementation during the first year was designed to mirror the steadily increasing number of licensed 16 year-olds subject to the law each month until July 1999, when all licensed 16 year-olds would be subject to the graduated licensing law changes. For analyses involving 16 year-olds, an additional gradual temporary intervention parameter was added beginning June 1, 1998, increasing to a peak on September 1, 1998, and decreasing linearly until December 31, 1998. This parameter was designed to

account for the fact that many more teenagers than usual obtained licenses in June, just prior to the July 1, 1998 law change, to avoid being subject to the new graduated licensing requirements. The large influx of newly licensed teenage drivers in June 1998 was accompanied by an acute temporary increase in injury and fatal crashes involving 16-year-old drivers. An alternative early licensure parameter increasing in May and June, peaking in July and decreasing linearly through October also was modeled but was not used because it did not reach statistical significance in the ARIMA model. The alternative parameter was based on past research indicating that the first month of teenage licensure holds the greatest crash risk after which it decreases steadily (McCartt et al., 2003). Results for graduated licensing were not different.

A combination of parameters was entered into each analysis to account for periodic linear and nonlinear trends in the data. In addition, the crash series for the 24-55-year-old age group, a temporary early licensure intervention variable term, and the law change intervention variable term were included in the model. Analyses were conducted using the trends module of the SPSS 11.5 software package.

Estimated changes in the rate of driver crash involvements per 10,000 population of that age due to the law change were used to calculate estimated changes in the number of driver involvements in injury and fatal crashes. The ARIMA-based monthly estimate was the average monthly change in the rate of driver crash involvements per 10,000 population following the law change. To derive the estimated change in the number of driver crash involvements during the 66 months following the law change, the average monthly estimated rate of crash involvements was multiplied by the sum of the population estimates after the law change and then divided by 10,000 to derive the total number of avoided driver involvements in crashes. Crash involvements prevented during the first 12 months of the law change were calculated slightly differently because the full law change was not estimated to be in effect for all 16 year-olds until July 1999. For those 12 months, an increasing proportion of the monthly ARIMA estimate was used to calculate crash involvements that were prevented. The totals for the first 12 months and the following 54 months then were summed. To obtain the percentage change, the estimated total change in crash involvements was divided by the difference between actual crash involvements for the period and the estimated total without the law change.

Outcome Measures

Sixteen year-olds are the primary target of the California law. The following measures (per 10,000 population) were used to assess the law's effects, with crashes involving 24-55-year-old drivers as a control:

- 16-year-old drivers involved in injury and fatal crashes
- 16-year-old drivers involved in injury and fatal crashes with injured 13-19-year-old passengers (data for injured or killed passengers was used because of concerns about a change in how passengers were reported that began in 2002)
- 16-year-old drivers involved in injury and fatal crashes between midnight and 4:59 a.m.

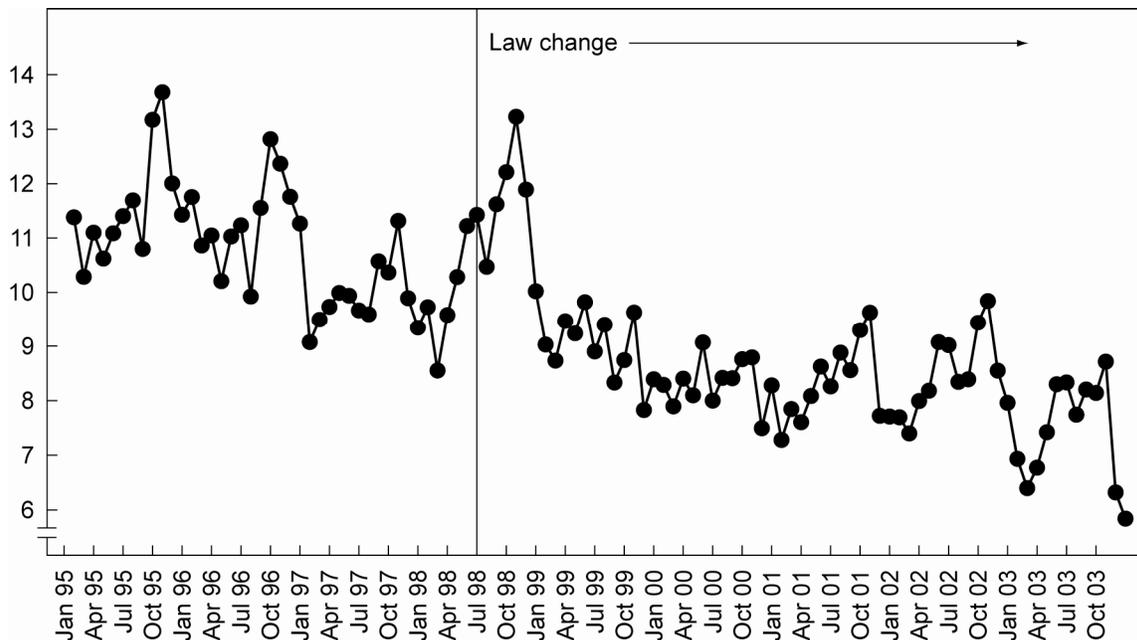
Masten and Hagge (2003) estimated the law's effects on 16 year-olds by examining crashes involving 16-year-old drivers, excluding crashes involving 24-55-year-old drivers, and counting crashes involving two or more 16-year-old drivers only once. The authors used crashes involving 24-55-year-old drivers, excluding all those involving 16-year-old drivers, as the control. Using the Masten and Hagge approach, the present study obtained the same results as those obtained when 16-year-old driver involvements in injury and fatal crashes were used, with 24-55-year-old driver involvements in injury and fatal crashes as the control group.

To determine whether the law simply shifted the ways in which teenagers were killed or injured, the rates of 16 year-olds involved in any capacity in injury or fatal crashes (e.g., driver, passenger, pedestrian) and the rates of 16-year-old drivers involved in injury/fatal crashes with no injured or killed passengers of any age also were analyzed to determine if these measure increased after the law change (the latter is possible if more teenagers drive alone after the law change). To determine whether the law shifted the crash problem to older teenagers, before-after changes in the rates of 17- and 18-19-year-old drivers involved in crashes were examined; the gradual temporary term accounting for the licensure increase for 16 year-olds was not included for these analyses.

RESULTS

Figure 1 shows the monthly rates of 16-year-old drivers involved in injury and fatal crashes in California per 10,000 16 year-olds from January 1995 through December 2003. These rates exhibited a

Figure 1
Monthly Rates of 16-Year-Old Drivers Involved in Injury and Fatal Crashes per 10,000 Population, 1995-2003



downward trend both prior to and following the July 1998 law change. There also was a clear seasonal effect, with rates higher during the summer. The corresponding monthly rates of 24-55-year-old drivers involved in injury and fatal crashes, as shown in Figure 2, did not exhibit any obvious trend, and there was no discernable change after the graduated licensing law went into effect.

Figure 2
Monthly Rates of 24-55-Year-Old Drivers Involved in Injury and Fatal Crashes per 10,000 Population, 1995-2003

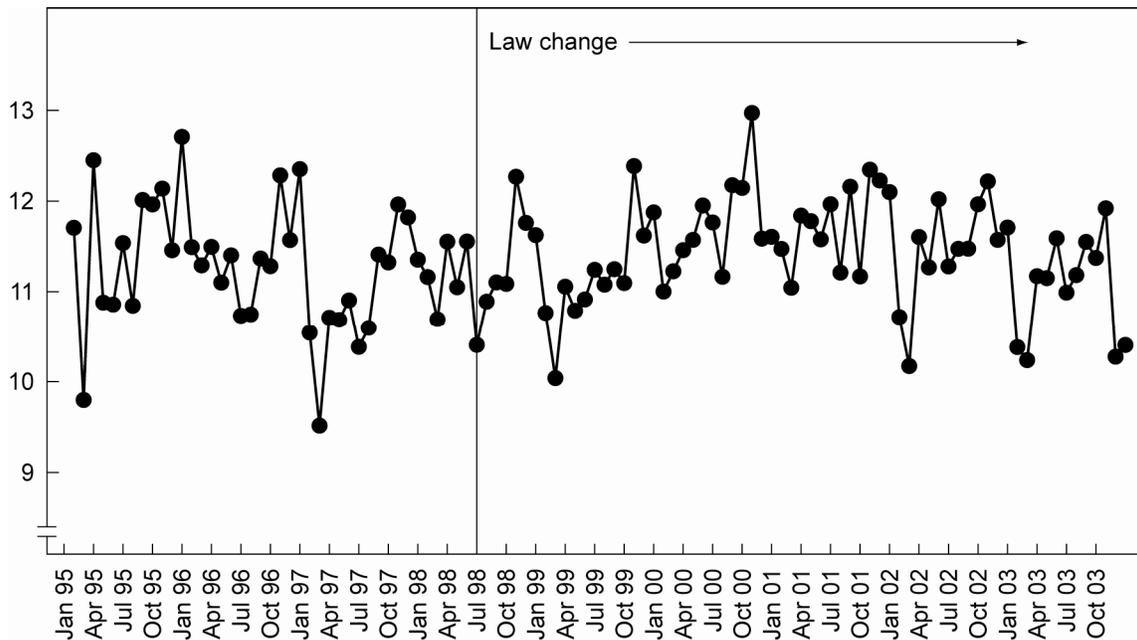


Figure 1 suggests a downward trend in 16-year-old driver crash involvement rates even before the law change, with seasonal and other periodic fluctuations. Therefore, time series models are appropriate to account for the pre-existing downward trend and the seasonal and other periodic linear and nonlinear trends so that any effect of the law change can be isolated. For ease of comparison, ARIMA models were of a single type that adequately explained the pre-existing trend as well as seasonal and other systematic trends. The model related the crash rate for any month to the rate for the previous month (i.e., autoregressive term of order one), the intervention variables and covariate for the given month and the previous month (i.e., moving average term of order one), and the values for the same two months in the previous year (i.e., integrated seasonal autoregressive and moving average of order one). The common shorthand notation for such a model is (101) (101).

According to the ARIMA analysis (Table 1), the rates of 16-year-old drivers involved in injury/fatal crashes were significantly lower after introduction of the law change ($p < 0.001$). The law change resulted in an estimated average monthly reduction of 2.8 16-year-old drivers involved in injury and fatal crashes per 10,000 population when all 16 year-olds were subject to the law. There would have

Table 1
Involvements of 16-Year-Old Drivers in Injury and Fatal Crashes per 10,000 Population

		Estimates	Standard error	t	Significance
Nonseasonal lags	AR1	8.38	0.87	9.593	0.000
	MA1	4.13	1.42	2.903	0.005
Seasonal lags	Seasonal AR1	9.88	0.46	21.276	0.000
	Seasonal MA1	8.86	2.23	3.968	0.000
Regression coefficients	Full law average estimated monthly per 10,000 change	-2.75	0.41	-6.665	0.000
	Early licensure	1.59	0.63	2.524	0.013
	Covariate*	4.34	1.26	3.446	0.001
Constant		5.93	1.52	3.898	0.000

*Covariate = crashes involving 24-55-year-old drivers

been an estimated 35,395 16-year-old drivers involved in crashes during these 66 months without the law change, but there were only 27,343. This translates into 8,052 fewer driver involvements in injury and fatal crashes than would have been expected without the law change (122 per month, on average, for the entire 66-month period), an overall 23 percent decline.

Figure 3 shows the actual monthly rates of 16-year-old driver crash involvements and the estimated monthly rates that would have occurred without the law change. The estimated crash involvement rates without the law represent the total number of involvements in injury and fatal crashes each month combined with the number of estimated involvements avoided by the law change, divided by the population and then multiplied by 10,000.

Figure 3
Monthly Rates of 16-Year-Old Drivers Involved in Injury and Fatal Crashes per 10,000 Population Compared with Estimates of the Rates without the Law Change, 1995-2003

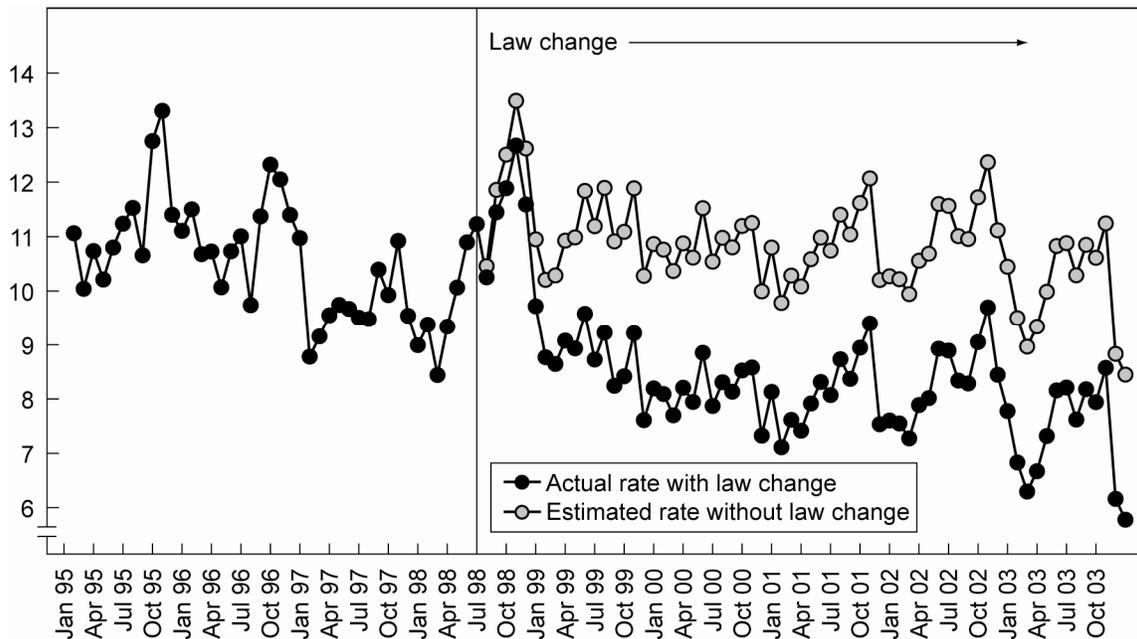


Figure 4 shows the monthly rates of 16-year-old driver involvements in injury/fatal crashes with injured passengers 13-19 years old, per 10,000 population. Again, there was a downward trend both prior to and following the July 1998 law change. However, the rates after the law change appear lower than would have been expected based on the pre-law rates. Table 2 shows the results of the ARIMA analysis. The rate of 24-55-year-old drivers involved in crashes in which teenage passengers were injured or killed was used as a covariate. The law change resulted in an estimated average monthly reduction of 1.4 16-year-old drivers involved in injury and fatal crashes with injured or killed teenage passengers per 10,000 population when all 16 year-olds were subject to the law ($p < 0.001$), an overall 38 percent decline resulting in an estimated 3,953 fewer 16-year-old drivers involved in such crashes.

Figure 4
Monthly Rates of 16-Year-Old Drivers Involved in Injury and Fatal Crashes with Injured 13-19-Year-Old Passengers per 10,000 Population, 1995-2003

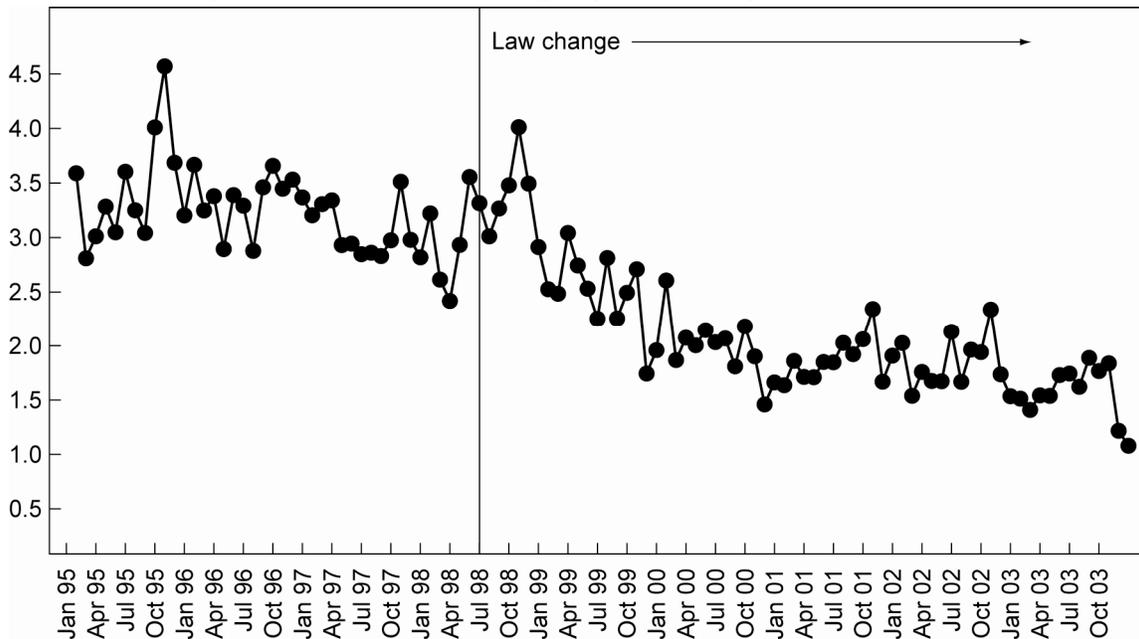


Table 2
Involvements of 16-Year-Old Drivers in Injury and Fatal Crashes with Injured or Killed 13-19-Year-Old Passengers per 10,000 Population

		Estimates	Standard error	t	Significance
Nonseasonal lags	AR1	7.51	1.19	6.31	0.000
	MA1	3.50	1.74	2.01	0.047
Seasonal lags	Seasonal AR1	9.80	0.71	13.83	0.000
	Seasonal MA1	8.80	2.22	3.95	0.000
Regression coefficients	Full law average estimated monthly per 10,000 change	-1.35	0.15	-8.93	0.000
	Early licensure	0.48	0.30	1.62	0.109
	Covariate*	23.28	12.87	1.81	0.073
Constant		2.58	0.39	6.64	0.000

*Covariate = crashes involving 24-55-year-old drivers carrying teenage passengers

Figure 5 highlights the reduction in the monthly rates of 16-year-old driver involvements in nighttime crashes (midnight to 4:59 a.m.) per 10,000 population. The analyses indicate that crash rates decreased significantly after the law ($p < 0.001$). As shown in Table 3, the law change resulted in an estimated average monthly reduction of 0.1 16-year-old driver involvements in injury and fatal crashes per 10,000 population when the law covered all 16 year-olds, an overall 27 percent decline. Table 4 shows that during the unrestricted hours, when most crashes occur, there was an estimated average monthly rate reduction of 2.6 16-year-old drivers involved in injury and fatal crashes per 10,000 population when the law covered all 16-year olds, an overall 23 percent decline ($p < 0.001$).

Figure 5
Monthly Rates of 16-Year-Old Drivers Involved in Injury and Fatal Crashes during Nighttime Restricted Hours (Midnight to 4:59 a.m.) per 10,000 Population, 1995-2003

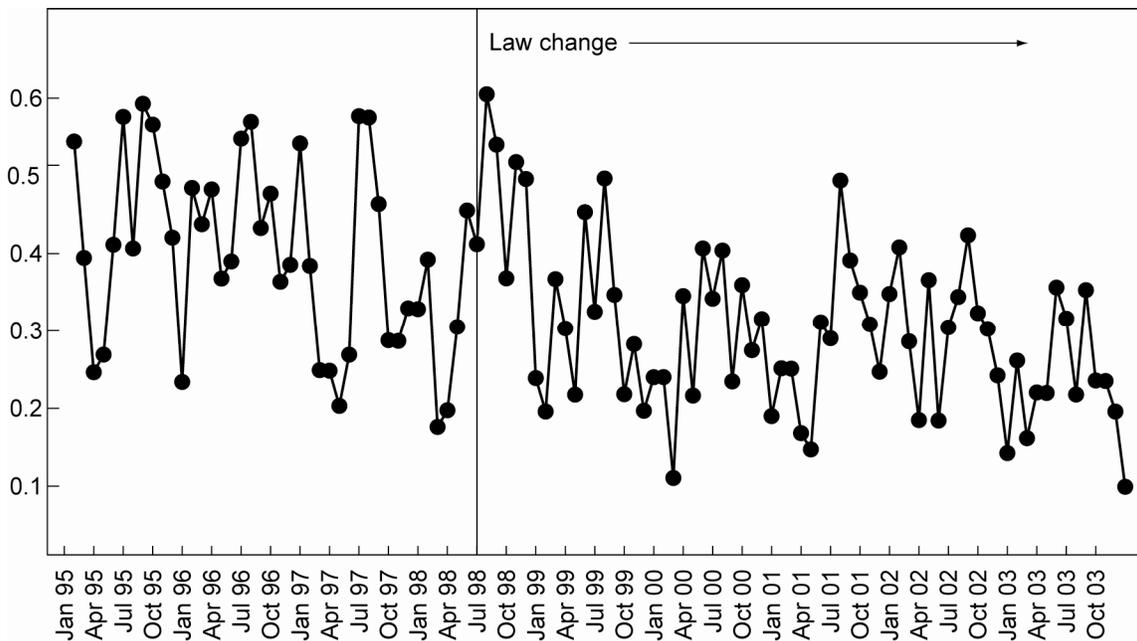


Table 3
Involvements of 16-Year-Old Drivers in Injury and Fatal Crashes during Nighttime Restricted Hours (Midnight to 4:59 a.m.) per 10,000 Population

		Estimates	Standard error	t	Significance
Nonseasonal lags	AR1	0.78	6.88	0.11	0.910
	MA1	-0.67	6.87	-0.10	0.922
Seasonal lags	Seasonal AR1	9.61	1.11	8.68	0.000
	Seasonal MA1	8.47	2.42	3.50	0.001
Regression coefficients	Full law average estimated monthly per 10,000 change	-0.12	0.02	-5.33	0.000
	Early licensure	0.12	0.07	1.86	0.066
	Covariate*	3.34	2.20	1.52	0.132
Constant		0.24	0.11	2.23	0.028

*Covariate = crashes involving 24-55-year-old drivers at night

Table 4
Involvements of 16-Year-Old Drivers in Injury and Fatal Crashes during
Nonrestricted Hours (5 a.m. to 11:59 p.m.) per 10,000 Population

		Estimates	Standard error	t	Significance
Nonseasonal lags	AR1	8.28	0.89	9.27	0.000
	MA1	4.19	1.44	2.91	0.004
Seasonal lags	Seasonal AR1	9.95	0.56	17.92	0.000
	Seasonal MA1	9.34	3.68	2.54	0.013
Regression coefficients	Full law average estimated monthly per 10,000 change	-2.56	0.35	-7.33	0.000
	Early licensure	1.46	0.57	2.56	0.012
	Covariate*	7.37	1.84	4.02	0.000
Constant		4.75	1.41	3.38	0.001

*Covariate = crashes involving 24-55-year-old drivers during nonrestricted hours

Table 5 shows the results of the ARIMA analysis of all 16 year-olds involved in any capacity (as drivers, passengers, pedalcyclists, or passengers) in injury and fatal crashes per 10,000 population. Introduction of the law was coincident with a significant decrease in this more general crash involvement rate ($p < 0.001$). The law change resulted in an estimated average monthly reduction of 2.4 16 year-olds involved in any capacity in injury/fatal crashes per 10,000 population when the law covered all 16 year-olds, an overall 17 percent decline. As shown in Table 6, there was an estimated average monthly rate

Table 5
Involvements of All 16 Year-Olds in Injury and Fatal Crashes per 10,000 Population

		Estimates	Standard error	t	Significance
Nonseasonal lags	AR1	8.94	0.70	12.78	0.000
	MA1	5.20	1.25	4.15	0.000
Seasonal lags	Seasonal AR1	9.90	0.62	16.09	0.000
	Seasonal MA1	9.06	2.87	3.16	0.002
Regression coefficients	Full law average estimated monthly per 10,000 change	-2.43	0.54	-4.46	0.000
	Early licensure	1.03	0.67	1.53	0.129
	Covariate*	10.73	2.02	5.31	0.000
Constant		4.12	1.86	2.22	0.029

*Covariate = all 24-55 year-olds in crashes

Table 6
Involvements of 16 Year-Old Drivers in Fatal and Injury Crashes with
No Injured or Killed Passengers per 10,000 Population

		Estimates	Standard error	t	Significance
Nonseasonal lags	AR1	5.92	5.25	1.13	0.262
	MA1	4.82	5.71	0.84	0.400
Seasonal lags	Seasonal AR1	9.51	1.52	6.26	0.000
	Seasonal MA1	8.53	2.73	3.12	0.002
Regression coefficients	Full law average estimated monthly per 10,000 change	-0.30	0.07	-4.37	0.000
	Early licensure	0.30	0.19	1.57	0.119
	Covariate*	4.98	1.16	4.28	0.000
Constant		0.63	0.53	1.19	0.239

*Covariate = crashes involving 24-55 year-old drivers with no injured or killed passengers

reduction of 0.3 16-year-old drivers involved in injury/fatal crashes per 10,000 population with no injured or killed passengers of any age, an overall 12 percent decrease ($p < 0.001$).

Finally, the ARIMA model was used for analyses of 17-year-old driver involvements in injury and fatal crashes (Table 7) and 18-19-year-old driver involvements (Table 8). The rate of 17-year-old driver involvements decreased by 8 percent ($p = 0.011$), whereas the rate of 18-19-year-old driver involvements increased by 4 percent ($p = 0.067$) after the licensing law change. The results from all analyses are summarized in Table 9.

Table 7
Involvements of 17-Year-Old Drivers in Injury and Fatal Crashes per 10,000 Population

		Estimates	Standard error	t	Significance
Nonseasonal lags	AR1	9.42	0.43	21.67	0.000
	MA1	6.95	1.00	6.96	0.000
Seasonal lags	Seasonal AR1	9.44	0.56	16.92	0.000
	Seasonal MA1	7.29	1.54	4.74	0.000
Regression coefficients	Full law average estimated monthly per 10,000 change	-1.38	0.54	-2.58	0.011
	Covariate*	9.91	1.33	7.45	0.000
Constant		3.83	1.70	2.26	0.026

*Covariate = crashes involving 24-55-year-old drivers

Table 8
Involvements of 18-19-Year-Old Drivers in Injury and Fatal Crashes per 10,000 Population

		Estimates	Standard error	t	Significance
Nonseasonal lags	AR1	8.99	0.80	11.29	0.000
	MA1	6.35	1.33	4.76	0.000
Seasonal lags	Seasonal AR1	9.08	0.70	12.92	0.000
	Seasonal MA1	5.92	1.49	3.97	0.000
Regression coefficients	Full law average estimated monthly per 10,000 change	0.94	0.51	1.85	0.067
	Covariate*	15.16	1.39	10.94	0.000
Constant		2.49	1.68	1.48	0.142

*Covariate = crashes involving 24-55-year-old drivers

Table 9
Summary of Results

Measure	Estimated change*	Percent change
16-year-old drivers involved in crashes	-2.75	-23
16-year-old drivers involved in crashes with teenage passengers	-1.35	-38
16-year-old drivers involved in crashes at night	-0.12	-27
16-year-old drivers involved in crashes during nonrestricted hours	-2.56	-23
All 16 year-olds involved in crashes	-2.43	-17
16-year-old drivers involved in crashes without passengers	-0.30	-12
17-year-old drivers involved in crashes	-1.38	-8
18-19-year-old drivers involved in crashes	+0.94	+4

*Estimated changes are based on the average monthly estimate per 10,000 16 year-olds when all 16 year-olds were subject to the law

DISCUSSION

Results of the present study indicate that California's 1998 graduated licensing law reduced the rate of 16-year-old drivers involved in injury and fatal crashes per 10,000 population, particularly those involving injured or killed teenage passengers. There is a concern that restricting teenage passengers may lead to more crashes as more 16 year-olds may drive themselves. Because the presence of teenage passengers in a vehicle was not recorded unless they were injured, it is not possible to determine with certainty when a driver was alone. However, there was a reduction in 16-year-old driver crash involvements in which no passengers were recorded. The effectiveness of the nighttime restriction on crash involvement was not much greater than that experienced during unrestricted hours. This may be because the restriction did not begin until midnight and, according to a survey of teenagers before the law change, many did not drive after midnight during the first 6 months of licensure with no restriction (Williams et al., 2002). The recent adoption of an 11 p.m. start time may lead to greater reductions during restricted hours.

There often is a concern when delaying full-privilege licensure that reductions in crash involvements among newly licensed drivers may be offset by increases among older teenagers who are less experienced than their predecessors. However, there was no correspondingly significant increase in the rates of involvements for 17- or 18-19-year-old drivers in injury and fatal crashes.

The results of the present study differ from those of Masten and Hagge (2003). A careful examination of the time series analysis of 16-year-old driver injury/fatal crash involvements per 10,000 population in the present study revealed significant seasonal and nonseasonal trends in the data, even after month-to-month differencing was applied to remove the linear downward trend. The differencing used by Masten and Hagge canceled out the pre-existing downward trend in the time series, but it did not remove or adjust for inherent seasonal and other nonlinear systematic trends. The present study used an alternative approach to differencing by replacing the differencing term with integrated seasonal and nonseasonal autoregressive terms and seasonal and nonseasonal moving average terms, which adjusted for inherent nonlinear month-to-month and seasonal trends rather than trying to eliminate them. In addition, the present analysis allowed for changes in the form of the model after the covariate and intervention terms had been added. It is possible that Masten and Hagge removed any effects of the intervention by including these terms after finding a model resulting in a stable series rather than during the modeling process.

The present study included 2 years of post-law data not available in the previous studies. However, using this model the results essentially are the same with or without the additional years. For example, using data for 1995-2001 the reduction in 16-year-old driver crash involvements was 19 percent compared with 23 percent for 1995-2003. The present study also differed from the Masten and Hagge

study by starting in 1995 rather than 1994 because of the potential confounding due to any effects of the zero tolerance law that took effect January 1, 1994. Starting in 1994 changes the present results slightly, indicating an 18 percent reduction in 16-year-old driver crash involvements for the 1994-2001 data (the Masten and Hagge time period) and a 26 percent reduction based on the 1994-2003 period.

Limiting the evaluation of the passenger restriction to crashes in which a teenager was injured may mean the results do not completely reflect the impact of the passenger restriction on all 16-year-old driver crash involvements with teenage passengers.

It is possible that the effects of California's law have increased since 2003. On January 1, 2004 the law was amended to raise the starting permit age from 15 years to 15, 6 months. Learners are required to remain in this stage for a minimum of 6 months, so the increase in minimum age for obtaining a learner's permit is likely to result in a slight delay in licensing among 16 year-olds and a subsequent reduction in their crash rate per 10,000 population. In 2005 the night restriction was expanded to include the period 11 p.m. to midnight, and the duration of the passenger restriction was increased from 6 months to 1 year. These changes make California's graduated licensing law one of the strongest in terms of recommended features (Williams and Mayhew, 2004). Many states other than California also have strengthened their original legislation, an important trend that should increase the effects of graduated licensing.

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