Analyses of the Impact of the Ohio Smoke-Free Workplace Act

September 1

2011

On Nov. 7, 2006, 58 percent of voters approved the Smoke-Free Workplace Act, and Ohio became the first Midwestern state to institute a workplace smoking ban.

Introduction

On Nov. 7, 2006, 58 percent of voters approved the Smoke-Free Workplace Act, and Ohio became the first Midwestern state and first tobacco-growing state to institute an indoor smoking ban.

The law impacts approximately 280,000 "public places" and "places of employment" in Ohio. These workplaces must prohibit smoking, remove ashtrays and post no-smoking signs with the toll-free enforcement number, 1-866-559-OHIO (6446). The Ohio Department of Health (ODH) was charged with writing enforcement rules and worked with stakeholders to draft rules for the enforcement, which began on May 3, 2007.

In 2011, ODH and its public health partners looked at five sources of data to determine the impacts of the law. ODH also conducted an analysis of attitudes and behaviors of Ohio adults related to the law. It is important to note that these studies only represent initial findings on the impact of the law as additional studies are currently underway.

Table of Contents

Summary of the Economic Impact of Ohio's Smoke-Free Workplace Act

Prepared by Elizabeth G. Klein, PhD, MPH & Nancy Hood, MPH
August 2011

To evaluate the economic effects on bars and restaurants associated with the Ohio Smoke free Workplace Act, an interrupted time-series analysis was conducted for the state of Ohio. County-level taxable sales liability from bars and restaurants were summarized for the state as the primary outcome, evaluated separately to investigate whether the comprehensive Ohio Smoke free Workplace Act policy influenced either business type differently.

Methods
Data used in these analyses were collected and provided by the Ohio Department of Taxation (ODT). Ohio businesses are required by law to report collected sales tax on a regular basis; large businesses report every month and small businesses (defined as less than $1,200 in state sales tax liability in a six month period) report every six months. ODT aggregated the number of reporting entities and the dollar amount of local (county) permissive sales tax liability for each month and county. All counties were summarized in order to describe the business taxable sales liability for the state of Ohio.

Entities were further classified by industry. Businesses most likely to be licensed to sell alcohol were selected using the North American Industry Classification System (NAICS) industry codes: (1) full-service restaurants (NAICS code 7221) and (2) drinking places (NAICS code 7224). Data were also provided for (3) all hospitality businesses (NAICS code 72). For confidentiality, data were suppressed when there were fewer than five reporting entities for a given county and month. Also, businesses that reported every six months were excluded from analyses because data were not available monthly. Data were provided for June 2003 through May 2010 (84 months).

The primary predictor variable was the enforcement of the Ohio Smoke free Workplace Act. An indicator variable was created by assigning a "0" to the months prior to policy enforcement (June 2003-April 2007), and assigning a "1" to the months when the policy was in effect and enforced (May 2007-May 2010).

Two outcome variables were created:
Restaurant Sales Ratio: ratio of restaurant sales to hospitality sales (minus restaurant sales). In terms of the measures described above, Restaurant Sales Ratio = (1)/[(3) – (1)]
Bar Sales Ratio: ratio of bar sales to hospitality sales (minus bar sales). In terms of the measures described above, Bar Sales Ratio = (2)/[(3) – (2)].

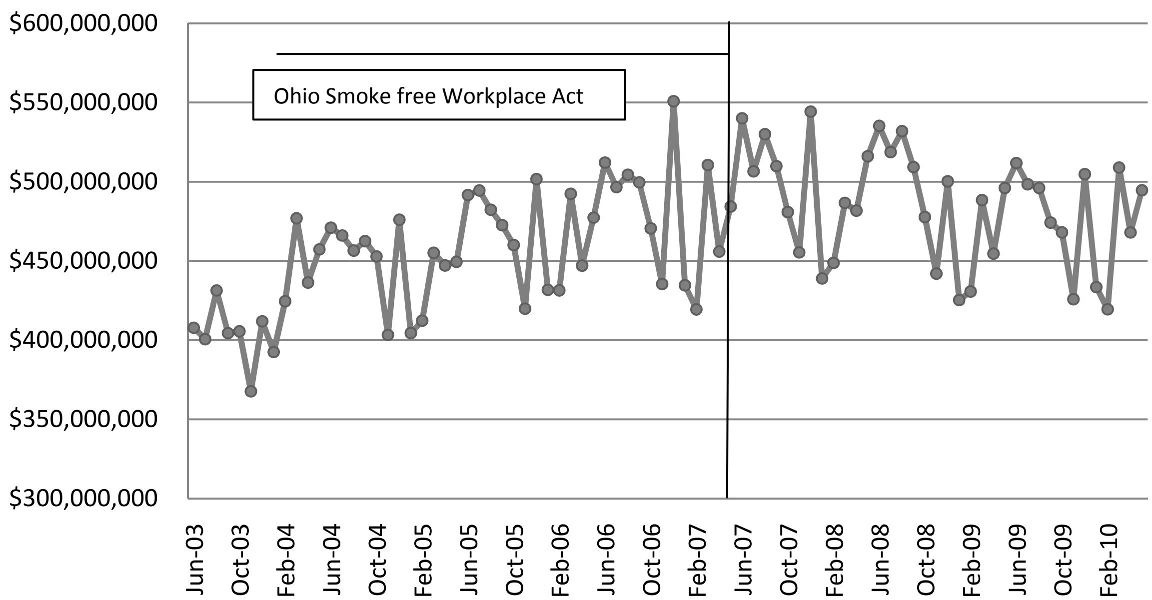
To account for the underlying economic trends, seasonality, and other factors unrelated to the enactment of the Ohio Smoke free Workplace Act, an interrupted time-series design was used. After the model was identified for each time series, three potential shapes to a change in taxable sales were tested: 1) gradual permanent change, 2) abrupt permanent change, or 3) abrupt temporary change. The final model was chosen by model fit statistics.

The interrupted time-series model includes an estimate of the initial change in taxable sales liability due to the effect of the Ohio Smoke free Workplace Act (ω), as well as the rate of effect (δ) which represents the amount of time at which any initial changes occurred. Statistical significance was determined by a *p*-value of 0.05 or lower.

Results

Graphs 1 and 2 show the unadjusted taxable sales liability for restaurants and bars, respectively, over the period of study. Table 1 presents the ARIMA model results for analyses of restaurant and bar sales in Ohio. After accounting for unemployment and seasons of the year, there were non-significant increases in taxable sales for restaurants and bars (ω=0.02 and ω=0.001, respectively). Since these were not statistically significant changes in taxable sales for either restaurants or bars, the conclusion of this study was that there were no significant changes in taxable sales associated with the Ohio Smoke free Workplace Act.

Graph 1: Unadjusted taxable sales for Ohio restaurants: June 2003 – May 2010



Graph 2: Unadjusted taxable sales for Ohio bars: June 2003 – May 2010

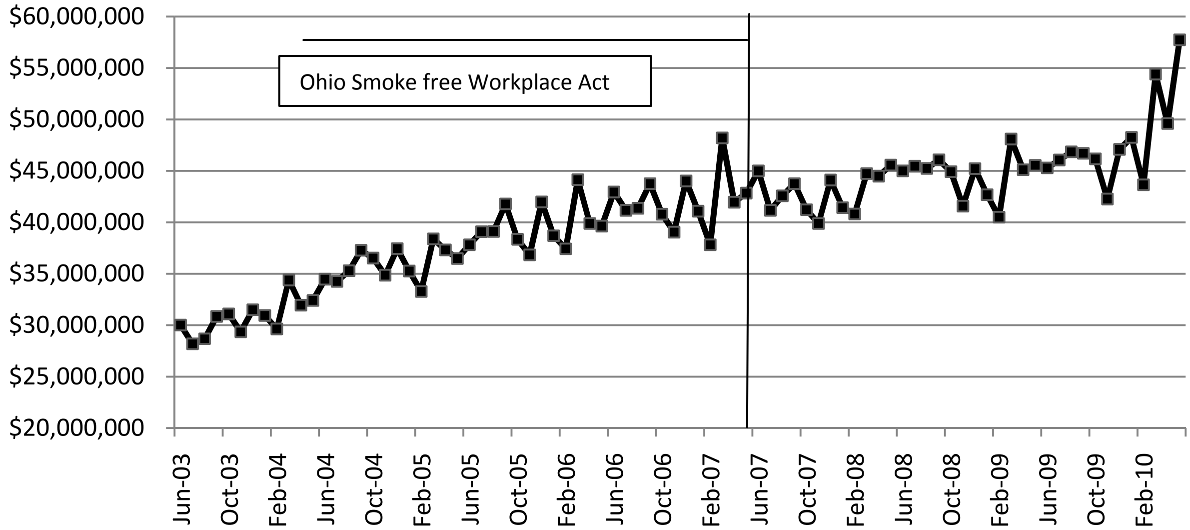


Table 1.

	Ratio of restaurant to total hospitality sales		Ratio of bar to total hospitality sales	
Model	Gradual permanent, 0 months delay		Gradual permanent, 0 months delay	
	Estimate	*p-value*	*Estimate*	*p-value*
Initial effects (ω)	0.02	0.67	0.001	0.53
Rate of effects (δ)	0.46	0.79	-0.40	0.75
Unemployment rate	0.0003	0.93	0.0002	0.25
Noise model	ARIMA(0,2,0)(0,1,0)$_{12}$		ARIMA(0,1,0)(0,1,0)$_{12}$	
AIC	-232.6		-657.5	
SBC	-225.9		-650.8	

Using Chief Complaint Data to Evaluate the Effectiveness of a Statewide Smoking Ban

William E. Storm, MPH, Brandi L. Bennett, BA, Brian E. Fowler, MPH
Bureau of Infectious Disease Epidemiology & Surveillance, Center for Public Health Statistics & Informatics (CPHSI), Ohio Department of Health

Objective

Preliminary analysis was completed to evaluate whether or not the smoke-free law in Ohio has made a positive change in reducing the effects of secondhand smoke (SHS) exposure by comparing syndromic surveillance data (trends for emergency department (ED) and urgent care (UC) chief complaint visits)] related to heart attack and/or acute myocardial infarction (AMI) before and after the smoking ban.

Background

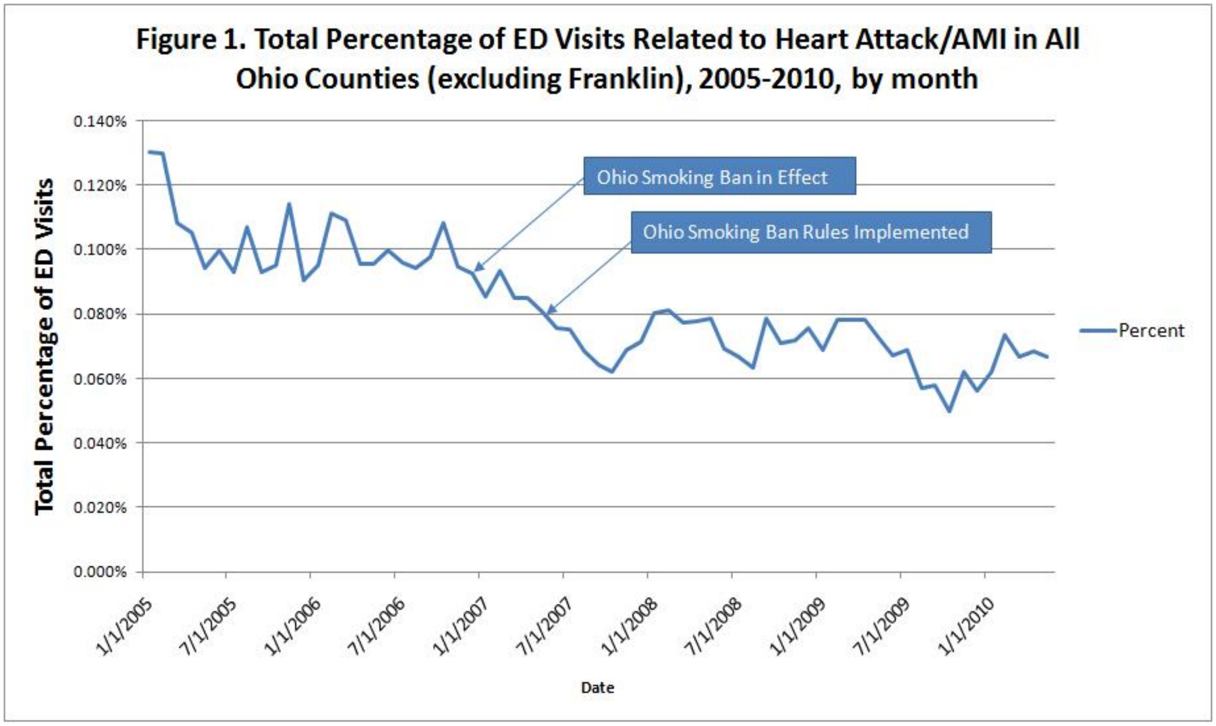
In November 2006, Ohioans supported a statute which set into law a requirement that all public places and places of employment in Ohio prohibit smoking (Ohio Administrative Code: Chapter 3701-52) [1]. The law took effect in December 2006; however, the rules for implementation were not finalized until June 2007. The primary purpose of the law was to protect employees in all workplaces from exposure to environmental tobacco smoke. When determining how best to evaluate the health impact of a smoke-free law as it relates to SHS exposure, most studies have reviewed the incidence of heart attacks or AMIs. In the 2006 Surgeon General's Report, "The Health Consequences of Involuntary Exposure to Tobacco Smoke," [2] SHS exposure is causally associated with cardiovascular events, including AMI. The Institute of Medicine also released a report in 2009 from a meta-analysis, "Secondhand Smoke Exposure and Cardiovascular Effects: Making Sense of the Evidence," [3] of eleven epidemiologic studies reviewing the incidence of acute coronary events following the passing of a smoke-free law. Each of the eleven studies showed a decrease in heart attack rates after implementation of smoke-free laws. The purpose of this study was to evaluate this relationship in Ohio.

Methods

Syndromic surveillance data from hospital ED and UC chief complaints were collected and analyzed from Ohio's EpiCenter system, for 2005-2010. Although these data types are pre-diagnostic in nature, they are more readily accessible than discharge data. Heart attack and AMI were defined rather specifically in the analysis (chief complaints must have included a reference to heart attack/pain/problems or AMI and excluded common visits solely for cardiac conduction or volume concerns or general respiratory problems). These data were combined and analyzed as a total percentage of visits by month, using SAS v 9.2. Data analyses were performed in 87 of Ohio's 88 counties. Franklin County was excluded from analyses as Columbus, Ohio (located within this county) passed its own smoke-free ban prior to the state ban.

RESULTS

Figure 1 below shows the trends of total percentage of ED and UC visits related to heart attack/AMI from 2005-2010 for all Ohio counties, excluding Franklin County. When comparing the means pre- and post-smoking ban, the data showed almost a 30% reduction in mean total percentage of visits for heart attack/AMI post-smoking ban.



CONCLUSIONS

Based on these results, the data suggest since the smoke-free law in Ohio has been passed, a reduction in the harmful effects of SHS exposure has also been observed by reducing heart attack and AMI, as defined by pre-diagnostic chief complaint data; however, no causal assumptions can be made. Additional analyses should be completed to further evaluate this relationship and to control for age and gender of the patients. Further, collection of patient diagnosis from the healthcare facilities would provide strength in validating the observed results.

REFERENCES

[1] Ohio Administrative Code: Chapter 3701-52. **Clean Indoor Air (Smoking Regulations).** http://www.odh.ohio.gov/rules/final/f3701-52.aspx

[2] A Report of the Surgeon General, 2006. "The Health Consequences of Involuntary Exposure to Tobacco Smoke." http://www.surgeongeneral.gov/library/secondhandsmoke/report

[3]IOM (Institute of Medicine). 2010. Secondhand Smoke Exposure and Cardiovascular Effects: Making Sense of the Evidence. Washington, DC: The National Academies Press.

Significant change in statewide rates of hospital discharge data for myocardial infarction due to the enactment of Ohio's Smoke-Free Work Place Law

David Bruckman, MS, MT(ASCP), Chief Systems Analyst, Cleveland Department of Public Health, and Core Faculty, Prevention Research Center for Healthy Neighborhoods, Case Western Reserve University, on behalf of the Ohio Department of Health

Brandi Bennett, BA, Epidemiologist, Chronic Disease and Behavioral Epidemiology, Center for Public Health Statistics and Informatics, Ohio Department of Health.

Last revised July 26, 2011 (V5)

Background:

Tobacco smoking and secondhand exposure to tobacco smoke is strongly implicated to an increased incidence of acute myocardial infarctions. Studies of hospital data on acute myocardial infarctions (AMIs) and acute coronary syndromes in states and localities that have instituted indoor smoking bans have shown a decrease in these occurrences after enactment (Juster et al, APHA, 2007. Sargent et al, BMJ, 2004, Bartecchi et al 2005, Gupta et al, 2011). Ohio voters approved the Ohio Smoke Free Workplace Act in November 2006. While the law went into effect in December 2006, the rules were not enacted until May 2007. These findings were reported in an environment of decreasing rates of AMI-associated hospital discharges nationally.

Objective:

We wanted to determine if the rate of discharges for AMIs from hospitals in Ohio decreased after the enactment of the state's smoke-free law. Our primary research hypothesis was that there would be a significant change in age- and age-and sex-adjusted rates of hospital discharges due to AMIs after the May 2007 enactment of the rules for Ohio's Smoke-Free Workplace law. We would test for this change on age-adjusted discharge rates separately for males and females. As a secondary hypothesis, we considered whether the May 2007 enactment date of the rules would be a significant factor in any change in rates. We hypothesized May 2007 would act as a statewide intervention and, as an indicator variable, would signal a significant change in slope. To model this intervention, we hypothesized that there may be some diffusion of an effect leading up to the intervention date as businesses began to adapt at different times due to a lag in the time between passage of the law and enactment of the rules. Lastly, we wanted to use more advanced statistical methods that benefitted from the monthly time-series data obtained that would test for the time of the law's enactment.

Method:

The Ohio Department of Health obtained de-identified data from the Ohio Hospital Association. The data reflected AMIs among Ohio inpatient discharges from 2004-2009 with a principal diagnosis of ICD-9 410. Age-adjusted rates were calculated for all persons and for males and females separately using SAS 9.2 (SAS Institute, Inc., Cary NC). The rates were reported per 1,000 persons.

Rates were calculated for each month in the five year span. The final data set had 60 records. The rates were age-adjusted to the 2000 census population for Ohio. No information on the total number of discharges for MI was obtained from OHA. OHA data was assumed to reflect over 90% of hospitals across Ohio.

We used mixed linear models with a varying covariance structures to determine if age-adjusted rates decreased each year, assuming a linear fixed effect for year, but also testing dummy variables for year using 2009 as the reference year. We made no assumptions to the covariance structure of the data but were interested in the size of the covariance parameter.

To determine if there was a significant change in slope, we fit simple spline polynomial functions, using optimized knot selection to determine the best inflection point in the monthly rate data (Ruppert, Wand, and Carroll, 2003). We used SAS/GLIMMIX to select optimal inflection points (including a central knot) using a generalized linear model with a radial smoothing option (SAS Institute, Inc., 2010). SAS fits the minimal amount of inflections, or knots, using a k-d tree method to partition the data (space). Based on the optimal inflections, we fit a simple cubic spline function to predict age-adjusted discharge rates for males and females, separately, using the midpoint of each month as the random variable in the radial smoothing method. No assumptions were made on the covariance structure, using the default variance components (i.e. covariance and residual parameters). We plotted the functions, checked for fit, zero mean and covariance for the estimated random effect (months), and checked residuals for normality and zero mean.

To test our secondary hypothesis, we used SAS 9.2 Analyst Time Series Forecasting System application (SAS Institute, Inc., Cary NC) to perform initial modeling, then to use the editing and interactive features to customize models. (Brocklebank and Dickey, 2003. Woodfield, 2000) Our intention was to develop a family of simple models to characterize the underlying time series trends, minimizing root mean square errors for selection ranking. Next, we introduced an intervention (interruption) model with the May

2007 enactment of the rules as an indicator point (i.e. equal to 0 times an additive slope indicator before the date, and 1 times an additive slope indicator after the date). Also, we tested the indicator date as a step and ramp function. We tested various lags, 3 to 5 month periods prior to the May 2007 intervention to simulate the effect of early adoption of smoke free policies between December 2006 and April 2007.

Model evaluation considered minimizing root mean squared error (RMSE), R-squared, measurements of maximized deviance with respect to parameters used (Schwartz Bayesian Coefficient or Akaike Information Coefficient), and visual examination of the model against data points. All ARIMA and IMA (autoregressive and non-autoregressive integrated moving average) models used Schwartz Bayesian criterion for model optimizing and Bayesian information criterion to estimate autocorrelation function parameter estimates. The software was permitted to find the best model parameters in the ARIMA and IMA modeling.

The May 2007 rules enactment date was introduced into the second phase of modeling to determine if its introduction into the model improved model criteria (reduced root mean square error, minimizing information criterion measures). The intervention model was introduced as a ramp function, assuming that there was a gradual adaption to the law prior to May 2007. (Hamilton, 1994)

We did not use a hold-out sample for validation nor intend to forecast future rates. The entire data set was used.

Results:

Mean age-adjusted rates are presented by year for females, males and all persons in Tables 1, 2, and 3, respectively. Each table provides the mean age-adjusted rates, rates from the linear model (with a single regression coefficient reflecting the annual change in rates), and a dummy variable model that uses the rate for 2009 as a reference. For each group, mean rates decreased consistently through the 5 year period. For all discharges due to AMI, the five year change in discharge rates due to AMI dropped from 1.9775 per 1,000 (1,977 discharges per one million Ohio residents) in 2005 to 1.680 per 1,000 (1,680 discharges per one million Ohio residents) in 2009.

For males, females and all persons, the predicted mean age-adjusted discharge rates significantly decreased through across the five years for each group; all models were significant, $p<0.0001$. Linear (single term) models were just as efficient as the dummy variable models, with very low covariance residual (variation explained by month-to-month correlation), and lower AIC measures.

Figure 1 reveals the age- and sex-adjusted rates for discharge due to AMI. The plotted line will be discussed later. Note the plunging drop in rate in mid-2007, from a 2007 high of 1.96 discharges per 1,000 in March 2007 (followed by 1.91 per 1,000 in May, 2007) to a low of 1.63 per 1,000 in September. The plunge from May to September in 2007 is visually striking.

Discharge rates for AMI decreased the most among males, an average decrease of 0.08542 per 1,000 males per year (about 85 fewer discharges per million Ohio males Ohio each year) for AMI. For females, an average of 0.05275 per 1,000 females per month (about 53 per million Ohio females annually) was predicted. **Lastly, an average decrease of 0.06867 per 1,000 persons (about 69 fewer discharges per million Ohioans) annually for acute myocardial infarction was predicted from 2005 to 2009.** Each linear change was statistically significant.

Using general linear mixed modeling with a radial smoother, we determined that June 2007 was the optimal inflection point (knot) between non-parametric spline functions for rates in both males and females, separately. The fitted cubic spline function for discharge rates for males is illustrated in Figure 1. Both models for discharge rates converged quickly (in 4 iterations), requiring no additional penalties (data restrictions) to refit the model, and fit the data well. Model residuals were normal and were near zero (estimate 0.0203, SD=0.0039, and 0.0051, SD=0.00097, respectively).

In successive time series models of age-adjusted discharge rates for males, females and all persons, performed separately, a seasonal effect was observed. In general, lowest rates were observed in July of each year, with the highest rates in November through March. Rates dropped consistently for February, perhaps due to being the shortest month of the year. The model for the age-adjusted rates for males and females together (total) is presented in Figure 2.

Among data for males, models improved (smaller RMSE) when May 2007 was introduced in the form of an interruption model, either as a ramp or step function. Automatic model filtering resolved a seasonal exponential smoothing model with 12-month periodicity. The model with the lowest root mean square error (0.06382) used a combination of twelve seasonal dummy variables, a linear trend term, and series of parameters describing a ramped effect to the May 2007 indicator date. While standard mixed effect linear models explained about 49% of the model variance (R-squared in Table 2), and combined seasonal exponential/ARIMA models 76.8%, this combined time-series model resolved 86.1% of total model variance, the best performance found of those tested. Addition of the May 2007 parameters for the April and May effects were significant ($p<0.0050$ and $p<0.0012$, respectively.)

Modeling for females also resolved the 12-month periodicity where a log linear trend with seasonal dummy indicators provided the lowest root mean square error (-0.153), RMSE (0.0484) and explained 79.7% of the model variance. Model performance did not improve significantly when adding any of the forms of a May 2007 interruption indicator, either with varying forms (ramp or step interruption) or with varying lags (0-5 months).

The model for all persons judged as best by both RMSE and Akaike Information Criterion was an indicator model similar to that used for males: a seasonally adjusted model (with 12 parameters) and a 5-month (lag) ramped increase to the May 2007 month indicator. Reflecting a diminished effect of the May 2007 indicator, the parameter for April 2007 was no longer significant ($p<0.1079$) while the May

2007 indicator continued to be significant ($p<0.0469$) as in the model for males. Figure 2 illustrates the model. Mean percent error (-0.0707), RMSE was minimized (0.0486) against IMA (2,2) and ARIMA (0,1,0) moving average models, while maximizing the amount of variance explained, 84.0%.

Discussion:

This study utilizes statewide hospital discharge data of acute myocardial infarction to show a significant decrease in discharge rates before and after a statewide indoor tobacco smoke exposure ban went into effect. Because AMI discharge trends have been decreasing across much of the US, we purposely chose two different advanced methods (time-series interruption models and cubic spline regression) other than the methods used by Juster et al 2007 using New York State discharge data to reveal the significant statewide effect associated with the date of enactment of Ohio's law.

Our modeling efforts reveal that age- and sex-adjusted discharge rates per 1,000 population for acute myocardial infarction in Ohio hospitals decreased significantly from 2005 through 2009, and that an indicator for the May 2007 enactment was a significant factor in modeling this trend using time-series analysis. Moreover, June 2007 was independently shown to be a separation point (knot or inflection point) within a cubic spline polynomial curve reflecting a non-linear change before and after enactment. Finally, time series modeling was able to explain a remarkable amount of the variation in the data, signaling a robust model of only adjusted rates of discharge without other stratification (smoking status, comorbidities, etc.)

As in our second research hypothesis**, decreases in discharge rates for males can be described by a one-month lagged effect prior to May 2007, where a gradual adoption of businesses to Ohio's Smoke Free Work Place law is believed to have occurred. For males and females together, the lagged effect was significant only for May 2007, reflecting a significant influence on decreasing discharge rates thereafter.** In this final model, 84% of the variation in the data was explained in the final prediction model. This use of an interrupted time series modeling confirms the findings in the cubic spline polynomial modeling that demonstrated a significant change in age-adjusted discharge rates for MI at the time of enactment.

Some of the more statistically robust analyses have been done on using state level data. Based on the work by Heckman and Hotz 1989, Juster et al., we also examined New York State principal admissions data for AMI (and stroke) also choosing an interrupted time series model adding the having with the flexibility of adjusting for interactions between time and changing levels of restrictive legislation across counties. In contrast, we chose a priori not to develop predicted rates of AMI discharge but to use time series interrupted models to examine the direct effect of Ohio's legislation on existing data across over 90% of Ohio's hospitals. Our findings support theirs of an accelerated decline in AMI discharges after introduction of more comprehensive anti-tobacco exposure laws.

More recent work using quarterly hospital discharge data across Delaware by Moraros et al 2010 showed reduced incident rates of AMI and asthma discharges in Delaware residents, where a statewide smoking ordinance was enacted in 2002, but not in non-Delaware residents. They used a Poisson model testing for seasonal and linear trends. Similar to their work, we found a strong linear trend in decreasing AMI discharge rates using an entirely different modeling method.

In a meta-analysis, Lightwood and Glantz (2009) combined community-level rates of heart attacks to determine changes in rates of admission due to myocardial infarctions after adoption of a smoke free law. This study supports their findings that the ban provides this continuing public health benefit over time. In fact, 69 fewer heart attack admissions to Ohio hospitals per year is a direct system savings of well over $1.1 million dollars assuming each case has an average cost of $16,200 (2004 figures), a very conservative estimate based on analysis of Healthcare Cost and Utilization Project (HCUP) data (Russo, 2007).

Comparing the estimated total cost of AMI hospitalizations over 2005-2006 ($7,206,490)versus 2008-2009 ($6,468,708), Ohio residents have conservatively saved at least $737,782 in hospital stay costs over the two year periods before and after the Smoke-Free law went into effect. Since these costs do not include professional (physician) fees, the amount should be considerably greater.

Limitations of this analysis are that we only have the age-adjusted rates on a monthly basis. Seasonal variation in the models may dampen the effect of the intervention indicator, spread over five months. In addition, we have no quantitative evidence of business practices or of other person-level explanatory rates, or separate rates for smokers and non-smokers as lucidly performed by Gupta et al. Data only reflect discharges for acute myocardial infarctions and do not reflect other co-morbidities that may be associated with secondhand or primary tobacco smoke exposure (e.g. asthma, COPD, etc.) Unlike Gupta and other analysis methods, we found that time series analysis of monthly data provided a stronger model. Monthly rates analyzed iteratively resolved seasonal patterns and a significant effect of the enactment of the Smoke Free Law across an entire state more clearly and robustly than using Poisson regression.

Additional investigation is needed to examine for any change in discharge rates for these other co-morbidities sensitive to tobacco exposure.

References:

Bartecchi B, Alsever N, Nevin-Woods C, et al. A city-wide ordinance reduces the incidence of acute myocardial infarction. Paper presented at: Scientific Sessions of the American Heart Association; November 14, 2005; Dallas, Tex.

Brocklebank JC, Dickey DA. SAS for Forecasting Time Series. 2003. SAS Institute, Inc: Cary, NC.

Fuller WA. Introduction to Statistical Time Series. 1976. Wiley Publishing: New York, NY.

Gupta R, Luo J, Anderson RH, Ray A. Clean indoor air regulation and incidence of hospital admissions for acute coronary syndrome in Kanawha County, West Virginia. Prev Chronic Dis 2011;8(4):A77. http://www.cdc.gov/pcd/issues/2011/jul/10_0200.htm. Accessed June 17, 2011.

Hamilton JD. Time Series Analysis. 1994. Princeton University Press: Princeton, NJ.

Juster HR, et al. Declines in Hospital Admissions for Acute Myocardial Infarction in New York State After Implementation of a Comprehensive Smoking Ban. AJPH, 2007;97:2035-2039.

Khuder SA, Milz S, Jordan T, Price J, Silvestri K, Butler P. The impact of a smoking ban on hospital admissions for coronary heart disease. Prev Med. 2007 Jul;45(1):3-8. Epub 2007 Apr 4.

Lightwood JM, Glantz SA. Declines in acute myocardial infarction after smoke-free laws and individual risk attributable to secondhand smoke. Circulation. 2009 Oct 6;120(14):1373-9. Epub 2009 Sep 21.

Moraros J, Bird Y, Chen S, Buckingham R, Meltzer RS, Prapasiri S, Solis LH. The impact of the 2002 Delaware smoking ordinance on heart attack and asthma. Int J Environ Res Public Health. 2010 Dec;7(12):4169-78. Epub 2010 Dec 2.

Ruppert, D., Wand, M. P., and Carroll, R. J. Semiparametric Regression. 2003. Cambridge:Cambridge University Press.

Russo, C. A. (Thomson Medstat), Ho, K. (AHRQ), and Elixhauser, A.(AHRQ). Hospital Stays for Circulatory Diseases, 2004. HCUP Statistical Brief #26. February 2007. Agency for Healthcare Research and Quality, Rockville, MD. http://www.hcup-us.ahrq.gov/reports/statbriefs/sb26.pdf

Sargent RP, Shepard RM, Glantz SA. Reduced incidence of admissions for myocardial infarction associated with public smoking ban: before and after study. BMJ. 2004;328:977–980.

SAS Institute, Inc. SAS/STAT® 9.22 User's Guide. 2010. Cary, NC:SAS Institute, Inc.

Woodfield TJ. Course Notes: Instructor-based Training. Introduction to Time Series Forecasting Using SAS/ETS Software. 2000. SAS Institute, Inc: Cary, NC.

Table 1. Mean (SD) age-adjusted discharge rates for MI among females, and estimated mean rates results from linear mixed modeling using year as a linear effect (row 2) and year represented as dummy variables (row 3)

MODELING FOR RATES AMONG FEMALES	Intercept	Linear effect	Estimated rate per year (p-value of corresponding model coefficient)					Covariance residual	DF	AIC
			2005	2006	2007	2008	2009			
Mean rates (SD)	n/a	n/a	1.4325 0.0942	1.3442 0.0940	1.3150 0.0994	1.2833 0.0841	1.1992 0.0480	n/a	59	n/a
Linear model	1.1566 p<0.0001	0.05275 p<0.0001	1.4204	1.3676	1.3149	1.2621	1.2094	0.006145	58	-119.9
Dummy variable model	1.1992 p<0.0001	n/a	1.4325 p<0.0001	1.3442 p<0.0001	1.3150 p=0.0007	1.2834 p=0.0114	1.1992 p<0.0001	0.006207	55	-109.0

Table 2. Mean (SD) age-adjusted discharge rates for MI among males, and estimated mean rates results from linear mixed modeling using year as a linear effect (row 2) and year represented as dummy variables (row 3)

MODELING FOR RATES AMONG MALES	Intercept	Linear effect	Estimated rate per year (p-value of corresponding model coefficient)					Covariance residual	DF	AIC
			2005	2006	2007	2008	2009			
Mean rates	n/a	n/a	2.6333 0.1612	2.5325 0.1435	2.4492 0.1749	2.4317 0.1626	2.2567 0.1241	n/a	59	n/a
Linear model	2.2044 p<0.0001	0.085 p<0.0001	2.6315	2.54608	2.46066	2.37524	2.28982	0.024	58	-41.9
Dummy variable model	2.2567 p<0.0001	n/a	2.6334 p<0.0001	2.5325 p<0.0001	2.4492 p=0.0035	2.4317 p=0.0075	2.2567 p<0.0001	0.024	55	-35.1

Table 3. Mean (SD) age- and sex-adjusted rates and estimated mean rates from linear mixed modeling using year as a linear effect (row 2) and year represented as dummy variables (row3)

MODELING FOR RATES AMONG ALL DISCHARGES	Intercept	Linear effect	Estimated rate per year (p-value of corresponding model coefficient)					Covariance residual	DF	AIC
			2005	2006	2007	2008	2009			
Mean rates (SD)	n/a	n/a	1.9775 0.1167	1.8850 0.1109	1.8333 0.0994	1.7933 0.1114	1.6800 0.0780	n/a	59	n/a
Linear model	1.6278 p<0.0001	0.069 p<0.0001	1.9712	1.9025	1.8338	1.7651	1.6965	0.011	58	-88.3
Dummy variable model	1.6800 p<0.0001	n/a	1.9775 p<0.0001	1.8850 p<0.0001	1.8333 0.0007	1.7933 0.0101	1.6800 p<0.0001	0.011	55	-78.2

Figure 1. Cubic spline predicted rate of age-adjusted discharge rates per 1,000 males for myocardial infarction from Ohio hospitals and emergency departments, 2005-2009. The plotted line reflects a best-fit prediction line using a cubic spline model that automatically generated a knot or inflection on June 2007, one month after the May 2007 enactment of the Smoke Free Work Place law. Average annual age-adjusted rates are shown in the histogram.

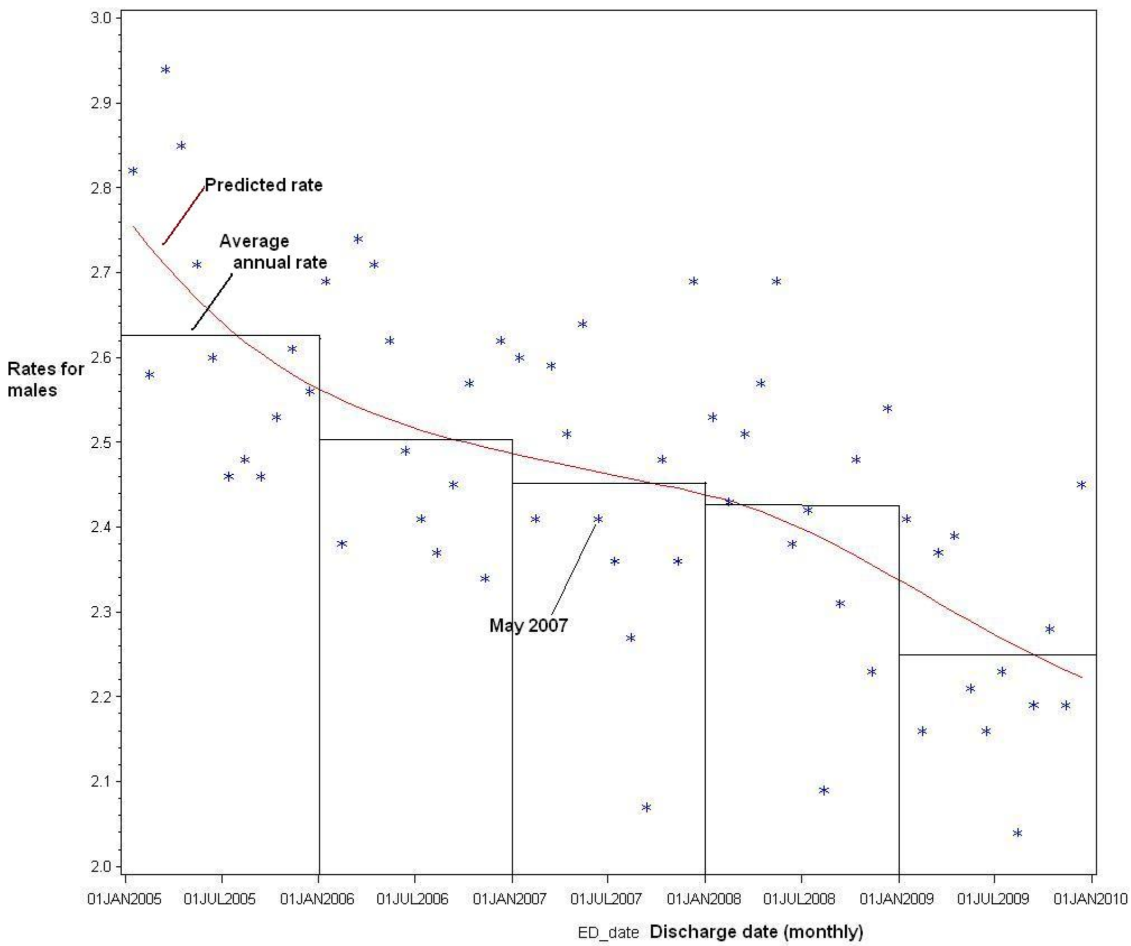
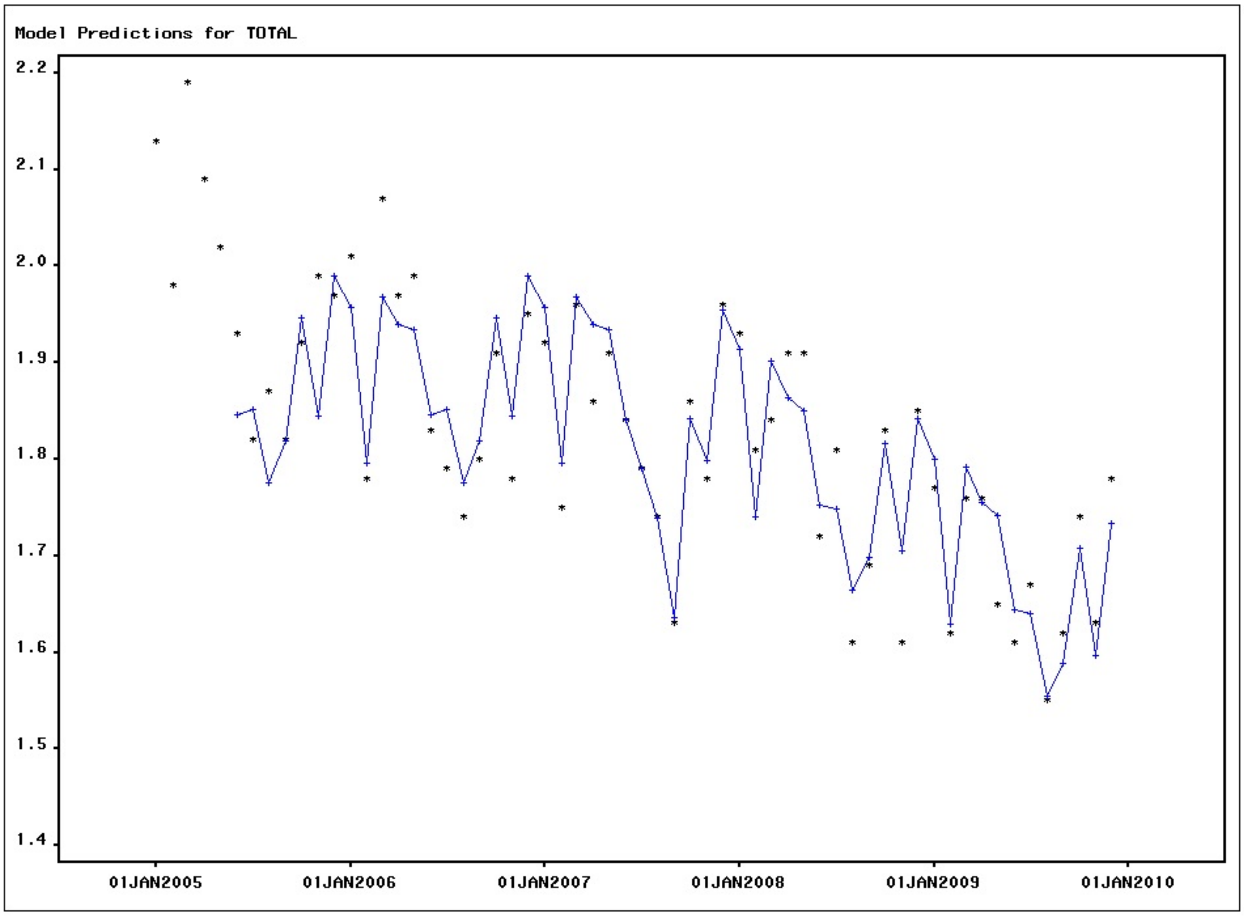


Figure 2. Age- and sex-adjusted discharge rates per 1,000 for myocardial infarction from Ohio hospitals and emergency departments, 2005-2009. The plotted line reflects a best-fit prediction line using a model with 12 seasonal dummy variables and an intervention indicator for the May 2007 enactment of the Smoke Free Work Place law. The indicator for the May 2007 enactment is lagged 5 months using a ramp function.



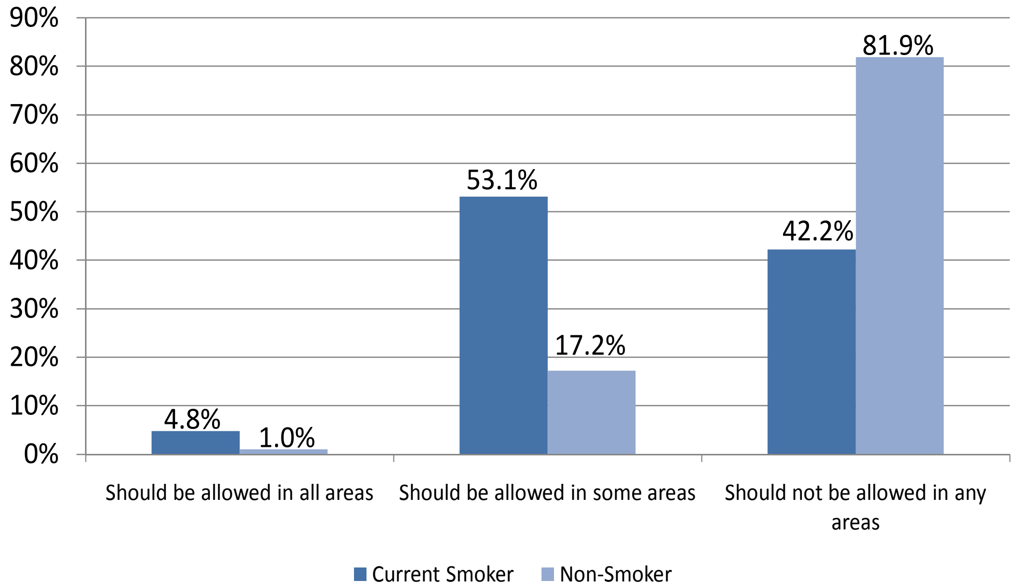
Attitudes and Behaviors Related to Smoke-Free Policies among Ohio Adults, 2009

Brandi Bennett, Epidemiologist

The Ohio Behavioral Risk Factor Surveillance System (BRFSS), conducted since 1984 at the Ohio Department of Health (ODH), is the primary source of Ohio-specific information regarding diseases, injuries, and health risk behaviors among Ohio residents 18 years and older. It is a telephone-based survey. Randomly selected telephone numbers among households are contacted and surveyed to determine the prevalence of selected diseases, injuries, and associated health behaviors among Ohio's diverse population. The data are used to identify current and emerging health problems, monitor trends, and develop, manage, and evaluate public health programs and policies. These data are provided to and used by internal ODH programs, local health departments, community health organizations, and other stakeholders concerned with disease and injury prevention and control. In 2007, the Adult Tobacco Survey (ATS) was combined with the BRFSS. It asks questions specific to tobacco use, behaviors, attitudes, knowledge, and beliefs. In 2009, the sample size of the ATS was more than 5,000 persons. The following analyses of questions related to smoke-free policies were calculated using 2009 data from the BRFSS/ATS.

The majority of adults in Ohio do not believe smoking should be allowed in indoor work areas or in the indoor dining areas of restaurants. Nearly 74 percent of respondents do not believe smoking should be allowed at all in indoor work areas, and approximately 75 percent of respondents do not believe smoking should be allowed at all in the indoor dining area of restaurants. Responses to these questions varied by smoking status, as shown in Figures 1 and 2.

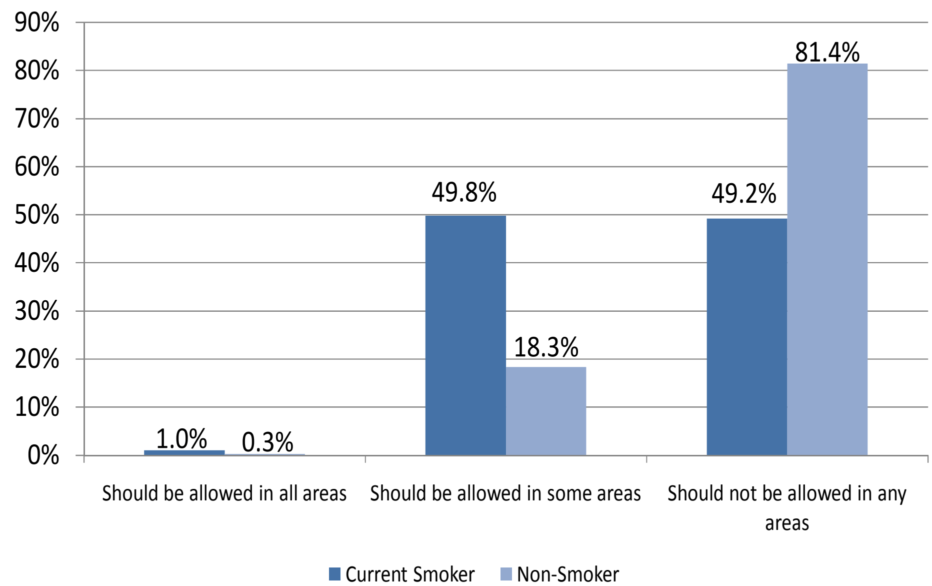
Figure 1: Estimated Prevalence of Opinions about Smoking in Indoor Work Areas among Adult Ohioans, by Smoking Status, 2009[1]



[1] Source: 2009 Ohio Behavioral Risk Factor Surveillance System and Adult Tobacco Survey, Chronic Disease and Behavioral Epidemiology, Center for Public Health Statistics and Informatics, Ohio Department of Health, 2010.

In indoor work areas, current smokers were more likely to state that smoking should be allowed in some areas (53 percent) than never smokers (17 percent) (Figure 1). In the indoor dining areas of restaurants, current smokers were more likely to state that smoking should be allowed in some areas (50 percent) than never smokers (18 percent) (Figure 2).

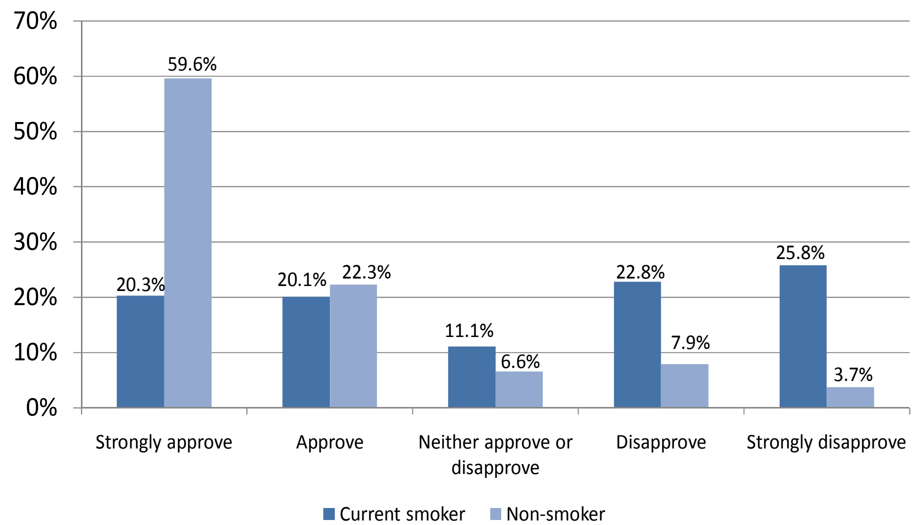
Figure 2: Estimated Prevalence of Opinions about Smoking in the Indoor Dining Area of Restaurants among Adult Ohioans, by Smoking Status, 2009[1]



[1] Source: 2009 Ohio Behavioral Risk Factor Surveillance System and Adult Tobacco Survey, Chronic Disease and Behavioral Epidemiology, Center for Public Health Statistics and Informatics, Ohio Department of Health, 2010.

The ATS includes questions about support for Ohio's smoke-free law. Around 73 percent of adult Ohioans either strongly approve of Ohio's smoke-free law (51 percent) or approve of the law (22 percent). Eight percent of Ohioans do not approve or disapprove of the law. Eleven percent disapprove, and 8 percent strongly disapprove. Analyses were conducted by smoking status (Figure 3).

Figure 3: Estimated Prevalence of Opinions about Ohio's Smoke-Free Law among Adult Ohioans, by Smoking Status, 2009[1]

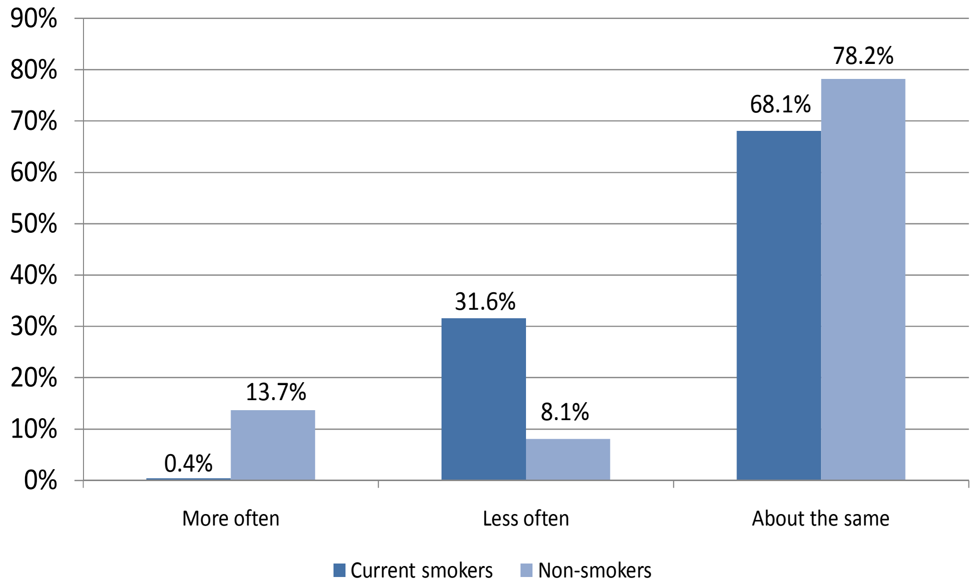


[1] Source: 2009 Ohio Behavioral Risk Factor Surveillance System and Adult Tobacco Survey, Chronic Disease and Behavioral Epidemiology, Center for Public Health Statistics and Informatics, Ohio Department of Health, 2010.

Among current smokers, 20 percent strongly approve of the law, and 20 percent approve of the law. This is compared with 60 percent of never smokers who strongly approve of the law, and 22 percent who approve of the law. Nearly 49 percent of current smokers disapprove or strongly disapprove of the law, compared with 12 percent of never smokers.

According to the ATS, approximately 76 percent of respondents have gone out to restaurants about the same as they did before Ohio's smoke-free law went into effect. About 68 percent of current smokers and 78 percent of never smokers have gone out to restaurants about the same as before the law went into effect (Figure 4).

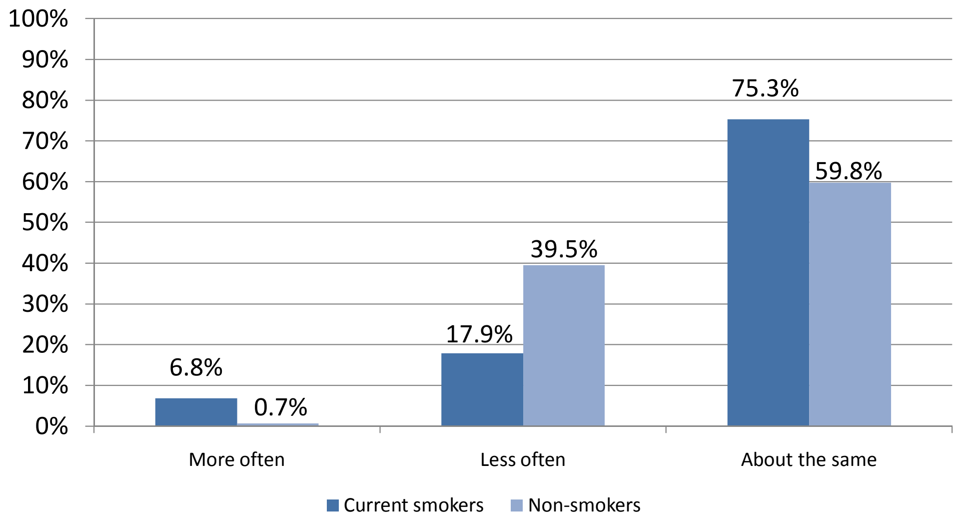
Figure 4: Estimated Prevalence of Frequency of Visits to Restaurants Since Passage of the Smoke-Free Law, by Smoking Status, among Adult Ohioans, 2009[1]



[1] Source: 2009 Ohio Behavioral Risk Factor Surveillance System and Adult Tobacco Survey, Chronic Disease and Behavioral Epidemiology, Center for Public Health Statistics and Informatics, Ohio Department of Health, 2010.

Nearly 72 percent of respondents have gone out to bars about the same as they did before the smoke-free law went into effect. About 75 percent of current smokers and 60 percent of never smokers have gone out to bars about the same as before passage of the smoke-free law (Figure 5).

Figure 5: Estimated Prevalence of Frequency of Visits to Bars Since Passage of the Smoke-Free Law, by Smoking Status, among Adult Ohioans, 2009[1]



[1] Source: 2009 Ohio Behavioral Risk Factor Surveillance System, Chronic Disease and Behavioral Epidemiology, CPHSI, Ohio Department of Health, 2010.

In conclusion, review of these analyses suggests a high level of support for smoke-free policies. The majority of respondents to the ATS favor smoke-free indoor work areas and smoke-free indoor eating areas in restaurants. There is, overall, a high level of support for Ohio's smoke-free law. Approximately three out of four respondents stated they visit restaurants and bars with about the same frequency as they did before the smoke-free law went into effect. There are differences in level of support for the law based upon smoking status, with non-smokers having a higher level of support compared to smokers.

Analysis of the Association between Birth Outcomes and the Ohio Tobacco Ban

Prepared by: Erin Hade, MS
Center for Biostatistics, the Ohio State University

Monthly rates of LBW and preterm birth (gestation less than 37 weeks) were calculated from Ohio's certificates of live birth from the period January 2006 through December of 2009. Statistical modeling was then employed to assess whether rates of these outcomes were lower after enforcement of the smoking ban began in May 2007. Furthermore, individual births were examined using another statistical modeling approach to compare the odds of occurrence of LBW (and preterm birth) before and after the smoking ban, while controlling for other factors.

Across the entire period, 8.6 percent (50,185) babies were born with birth weight classified as low (less than 2,500 grams). A large amount of variability was observed in monthly LBW rates across time. Controlling for a tendency of rates to vacillate between lows and highs approximately every 4 months, the rate of LBW was found to be similar before and after the ban. Thus, no association was found between the tobacco ban and rates of LBW. Similar findings were observed for rates of preterm births. Furthermore, no associations were observed between the smoking ban and rates of poor outcomes among subgroups, such as women with Medicaid as the principal payment for delivery, those with private insurance, and those with other payment sources. Finally, the model to explore the association between LBW and the tobacco ban at the subject level, found no statistically significant relationship between the adjusted odds of LBW before and after the smoking ban enforcement date.

Statistical modeling of data from Ohio vital birth records produced no evidence of a change in these outcomes attributable to the smoking ban. These results are limited by the fact that the analyses were based on a relatively small number of years of observation, both before and after the ban.

Summary

There is little evidence in the vital statistics data compiled between January 2006 and December 2009, that the enforcement of the tobacco ban had a substantial effect on poor birth outcomes. These trends may be more apparent with increased pre and post ban data.

Background

In November 2006, voters approved a state wide tobacco ban, which subsequently went into effect in December of that year. Formal enforcement of the ban began in May 2007. The current study seeks to explore the potential impact of this ban on birth outcomes. Using vital statistics/birth record data from January 2006 through December 2009, we explored how the rate of babies with low birthweight (LBW) varied before and after the ban and how individual level factors may have influenced a woman's probability of having a baby with low birthweight.

Results

Between January 2006 and December 2009 583,530 births we identified as taking place in Ohio to Ohio residents. Of these 8.6% (50,185) babies were born with birthweight below 2500 grams.

	2006	**2007**	**2008**	**2009**
2500 grams +	134,896 (91.3)	135,041 (91.3)	133,459 (91.5)	129,949(91.5)
< 2500 grams	12,842 (8.7)	12,838 (8.7)	12,456 (8.5)	12,049 (8.5)
	147,738	147,879	145,915	141,998

To model how the rate of LBW varied over time, we calculated the rate of LBW for each month during this timeframe. Exploration of changes in these monthly rates (on the log scale) before and after the ban was conducted through descriptive and modeling methods. Figure 1 illustrates the time series of these rates. We notice quite a large amount of variability, which is consistently seen before and after the ban's enforcement (May 2007). Moreover, we also observe a strong (negative) autocorrelation of these rates every four months (see appendix for PACF and ACF plots). That is to say, that when the rate is observed to be high in one month, four months later it is likely to be low. Accounting for this autocorrelation at a 4 month lag, we find that rates of LBW are similar before and after the ban. We estimated the risk of LBW to be slightly increased by **1% (RR=1.01, 95% CI: 0.98-1.04%),** but not different substantially different from 0. Interestingly, we find that as time progresses, the rate of LBW declines approximately **1.4% per year (RR=0.986 95% CI: 0.985-0.987**).

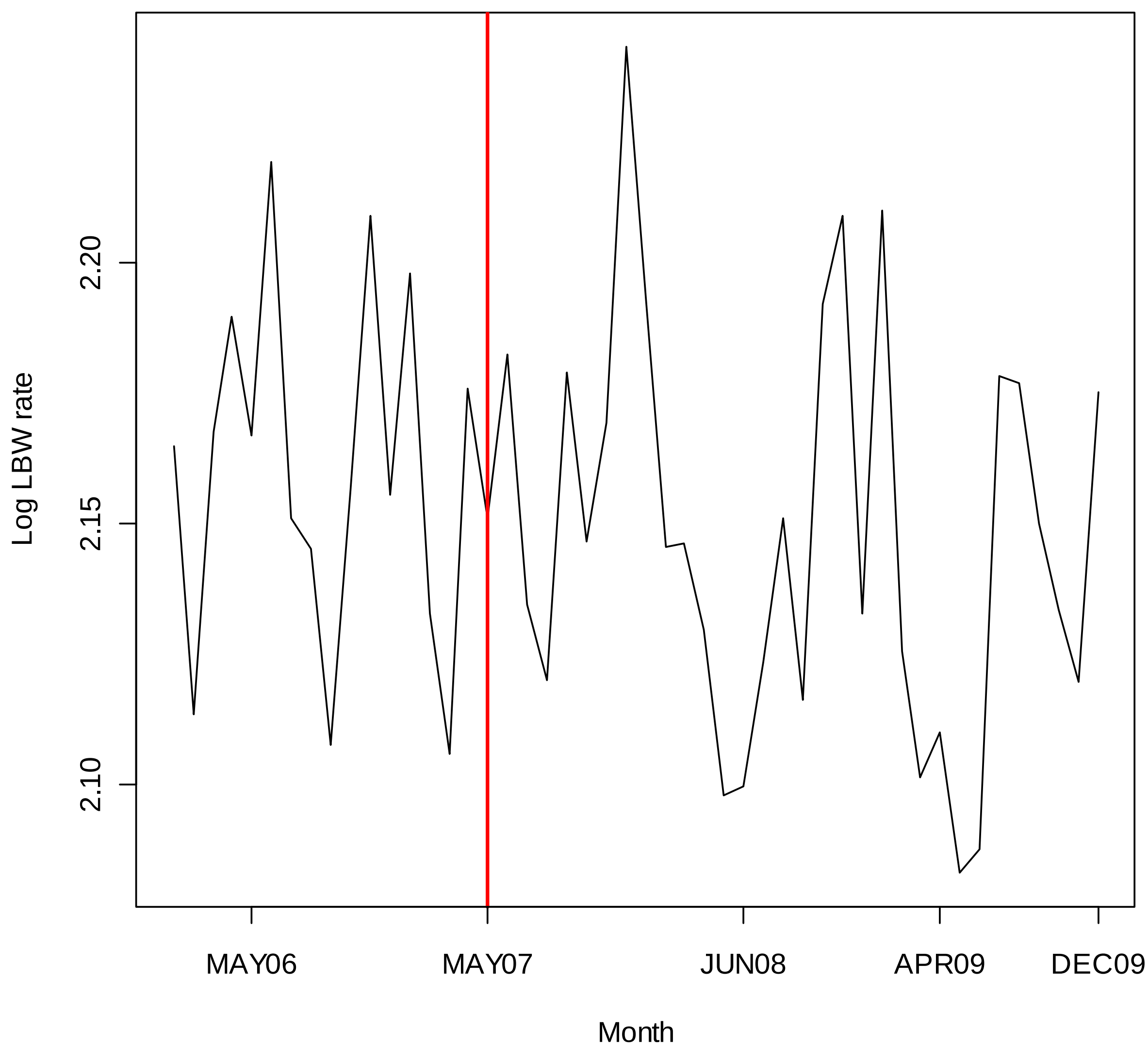


Figure 1:. Log LBW over time. The red line indicates the month of the tobacco ban enforcement.

Further we find no association between the tobacco ban and LBW for various subgroups, including women who indicate that the principal source of payment for this delivery was Medicaid, private insurance or some other payer source. These trends may have been limited by the short timeframe available before the tobacco ban and maybe strengthened by extending this work to further before and after the ban.

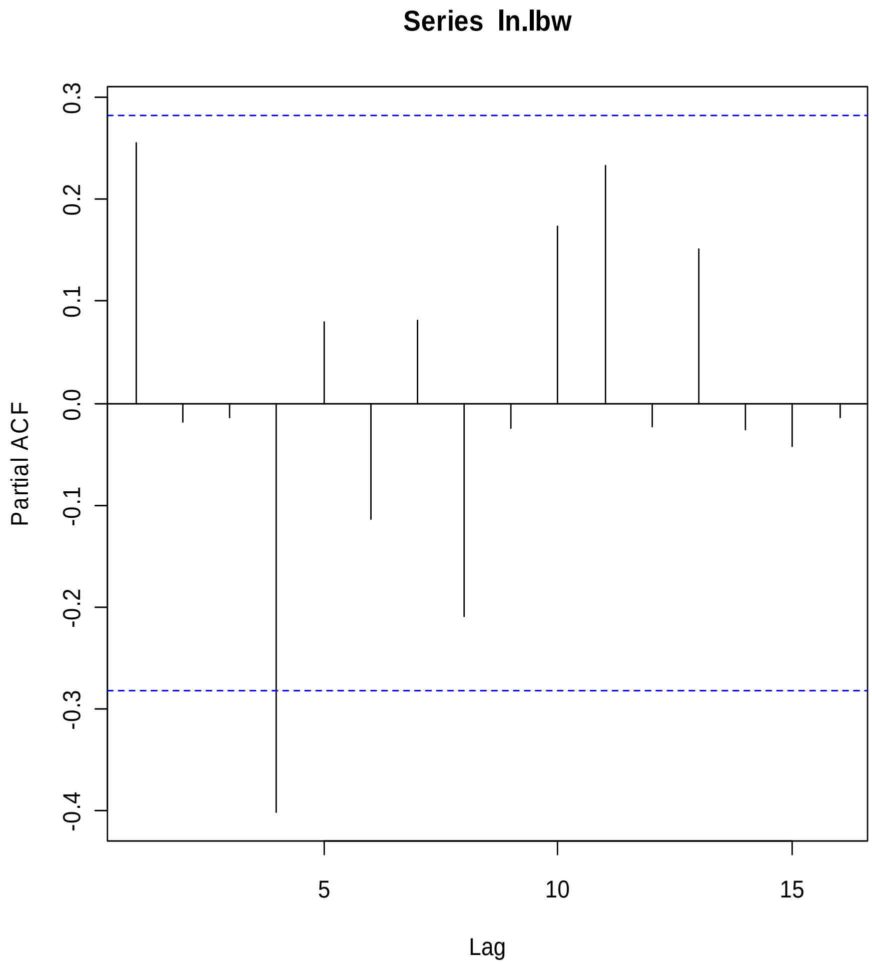
To explore the relationship between LBW and the tobacco ban in an alternate way, we considered modeling the probability of LBW before and after the ban with the subject level data. Again we found very little association between the tobacco ban and LBW status. We estimated that the adjusted odds of LBW were slightly increased after the ban **(OR: 1.02, 95% CI: 0.99-1.06),** but this was not significantly different from zero. These models adjusted for mother's race, ethnicity, marital status, insurance status and age. There was further little evidence that the effect of the tobacco ban varied/was modified by various demographic factors including race, insurance status, WIC status and marital status.

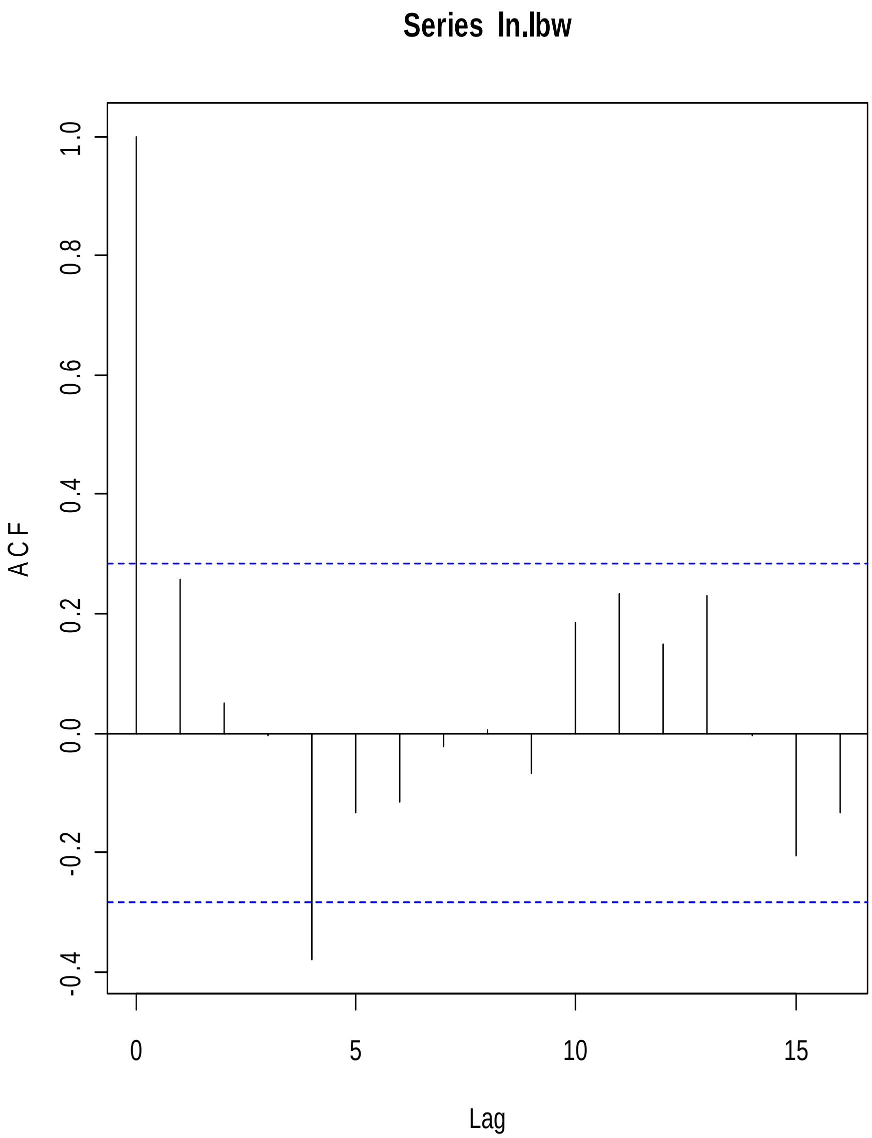
Finally, similar models were explored for pre-term birth (birth before 37 weeks) and again no association was found between the tobacco ban and the probability of pre-term birth.

<u>Methods</u>

To explore monthly rates over time, we utilized both time series and generalized linear models. Initial modeling included using ARIMA models to explore the effect of the tobacco ban and various autocorrelation structures. Generalized linear models for the monthly rate of LBW assumed a Poisson distribution and included terms with a four-month lag to account for the autocorrelation detected. Utilizing all of the individual data, the probability of LBW was modeled via logistic regression. Analyses were performed in SAS 9.2 and R version 2.11.1.

Appendix




Series ln.lbw
ACF
Lag
1.0
0.8
0.6
0.4
0.2
0.0
-0.2
-0.4
0
5
10
15

Ohio Pregnancy Risk Assessment Monitoring System (PRAMS) Data

By Connie Geidenberger, Ph.D

Executive Summary

Studies have shown that pregnant women exposed to environmental tobacco smoke (ETS) have 20% higher odds of giving birth to a low birth-weight (LBW) infant when compared to women without this exposure. Primary sources of significant ETS exposure to pregnant women are the home and workplace. According to the U.S. Census Bureau, two thirds of women who had their first child between 2001 and 2003 worked during pregnancy. This percentage has increased in the United States since the 1960's. The percentage of employed U.S. pregnant women who worked up through the month prior to their child's birth has similarly increased. Since work environments conferring secondhand smoke exposure to pregnant women may significantly contribute to poor pregnancy outcomes, a reduction in ETS exposure in the workplace and other public places could result in reduced rates of low birth weight and other poor outcomes. A recent study comparing preterm birth rates in Colorado cities with and without smoking bans found lower preterm birth rates in the city with the ban. Likewise, preterm birth rates declined in Ireland one year after introduction of a comprehensive Irish workplace smoking ban. The Ohio Smokefree Workplace Act was enacted in late 2006 with a goal of protecting the public from ETS exposure in the workplace. We therefore examined whether odds of LBW and preterm birth were reduced in Ohio after the Act's enforcement began in May 2007.

The Ohio Pregnancy Risk Assessment Monitoring System (PRAMS) is a representative survey of Ohio resident mothers of live born infants and is part of a cooperative state/federal effort to carry out ongoing surveillance designed to better understand and ultimately prevent poor birth outcomes. Data from Ohio PRAMS, covering birth years 2005-2009, were used to compare the odds of having a LBW infant among Ohio resident mothers of singleton live-born infants conceived before May 1, 2007 to that of those conceived on or after that date. Statistical modeling was employed to assess the relationship of LBW with pre/post smoking ban enforcement, while controlling for other factors, including maternal smoking, age, education, income, stressful events during pregnancy, race/ethnicity, WIC program participation, and body mass index. Similar methods were used to examine the odds of preterm birth and LBW among term infants before and after the ban.

Without controlling for other factors, the odds of LBW did not differ between infants conceived pre- and post- smoking ban enforcement. Controlling for possible confounding factors did not change these results. Thus, there is no evidence from this assessment to support a relationship between timing of

smoking ban enforcement and odds of LBW. Similar results were observed for preterm births and LBW among term infants.

These results are limited by the facts that, 1) information was self-reported (which sometimes leads to underreporting of undesirable behaviors) and 2) no information was available on ETS exposure or employment of the mother during pregnancy. It is possible that different results may have been observed among subgroups of women employed during pregnancy in industries most likely impacted by the ban (such as in restaurants or bars). It is unknown what proportion of pregnant women work in such jobs in Ohio. Thus, it is also unknown whether improvements in outcomes of these women could have been obscured by their incorporation into a larger group of women who were unexposed to ETS irrespective of the existence of a smoking ban. Nevertheless, the Ohio PRAMS survey was expressly developed to focus on examination of risk factors for LBW. Thus, an adequate sample size of women with LBW infants was available from this data source for statistical assessment. Furthermore, relationships observed in the data between LBW and known risk factors (e.g., smoking during pregnancy, education, race/ethnicity) were consistent with the published literature, lending support for the validity of the information from PRAMS.

In conclusion, no evidence of statewide improvement in odds of poor birth outcomes was observed from this assessment. Given that no information was available to assess industry of employment (or any employment) during pregnancy, a relationship between the ban and birth outcomes could not be examined among subgroups of pregnant women most likely to have been impacted by the ban, thereby limiting interpretation of these results.

Background

Birth weight and gestational age at birth are considered important predictors of infant survival and health, since low birth weight (<2500 g) and preterm birth (gestation less than 37 weeks) place infants at substantial increased risk of subsequent mortality and ongoing health and developmental problems. At the population level, the rates of low birth weight (LBW) and preterm birth are important indicators of a society's overall health. Maternal smoking during pregnancy is a well-established risk factor for both LBW and preterm birth[1]. Furthermore, pregnant women exposed to environmental tobacco smoke (ETS) have 20% higher odds of giving birth to a LBW infant when compared to women without this exposure[2]. The relationship between maternal ETS exposure and preterm birth is less clear, although a number of published studies have concluded that ETS exposure of pregnant women in the workplace is a hazard for the developing fetus[3].

According to the U.S. Census Bureau, two thirds of women who had their first child between 2001 and 2003 worked during pregnancy. This percentage has increased in the United States since the 1960's. The percentage of employed U.S. pregnant women who worked up through the month prior to their child's birth has similarly increased[4]. While the number of pregnant women exposed to ETS in the workplace is unknown, the fact that many women work throughout their pregnancies, combined with the known hazards of ETS exposure to the developing fetus, makes workplace ETS exposure during pregnancy an important public health concern. To address the broader issue of known health hazards from ETS exposure, legislation banning smoking in the workplace and other public places has become more common in recent years, and improved health effects and lower smoking rates due to these measures have been demonstrated[5,6,7,8,9]. With regard to birth outcomes, a recent study comparing preterm birth rates in Colorado cities with and without smoking bans found lower preterm birth and maternal smoking rates in the city that had enacted a ban[8]. Likewise, preterm birth rates and maternal smoking rates declined in Ireland one year after introduction of a comprehensive Irish workplace smoking ban, although the rate of LBW increased [9].

The Ohio Smokefree Workplace Act was enacted in late 2006 with a goal of protecting the public from ETS exposure in the workplace. Consequently, it was of interest to explore whether changes in birth outcomes occurred in Ohio subsequent to the statewide smoking ban. We therefore examined whether the odds of LBW and odds of preterm birth were reduced in Ohio residents after the Act's enforcement began in May 2007.

Methods

Data from the Ohio Pregnancy Risk Assessment Monitoring System (PRAMS), covering Ohio resident women with singleton births from 2005 through 2009, were used for this assessment. Ohio PRAMS is a representative survey of Ohio resident mothers of live born infants and is part of a cooperative state/federal effort to carry out ongoing surveillance designed to better understand and ultimately prevent poor birth outcomes.

To establish pregnancies occurring before and after the ban, women were classified by timing of their last menstrual period (LMP), where women with LMP before May 1, 2007 were considered to have pre-smoke ban exposures while those with LMP on or after that date were considered to have post-smoke ban experiences. LBW was classified as less than 2,500 g. at birth while preterm birth was defined as gestational age less than 37 weeks.

Bivariate logistic regression methods for weighted survey data were used to calculate odds ratios of post to pre smoking ban exposures for outcomes of LBW, preterm birth, and LBW/term birth. Possible confounding factors were explored by assessing associations of covariates with outcome variables via bivariate logistic regression methods, and by examination of associations between potential confounders with the pre-post ban exposure variable. Maternal covariates that were examined included smoking during pregnancy, age, education, income, stressful events during pregnancy, race/ethnicity, WIC program participation, and pre-pregnancy body mass index. Multivariable models were then developed to control for possible confounding of the relationship between pre/post smoke ban exposure and birth outcomes.

Results

Results of bivariate logistic regression analyses for LBW are found in Table 1. Without controlling for other factors, the odds of LBW did not differ between infants conceived pre- and post- smoking ban enforcement. Statistically significant associations with LBW were observed for all of the potential confounding variables examined. Table 2 presents results of analyses comparing distributions of potential confounders with pre/post ban timing. Only maternal education was significantly related to smoke ban timing, with a slightly greater percentage of more hightly educated women found in the post ban period. Finally, results of the multivariable regression are found in Table 3. While most covariates remained statistically significantly associated with LBW, their presence in the model did not alter the observed relationship between smoke ban and LBW. Thus, there is no evidence from this assessment of a relationship between timing of smoking ban enforcement and odds of LBW. Similar results were observed for preterm births and LBW among term infants (not shown).

Table 1: Factors examined for an association with low birth weight (<2,500 g), unweighted frequencies, weighted percentages, and crude weighted odds ratios, singleton births, Ohio PRAMS, 2005-2009.

unweighted n=6,872

Variable	Low Birth Weight* # (%)	Crude Odds Ratio (95% CI)
Last Menstrual Period with respect to Smoking Ban Enforcement		
Before May 1, 2007	1,633 (6.6)	Ref
On or After May 1, 2007	1,083 (6.9)	1.1 (1.0 , 1.1)
Maternal Age (years)		
<18	156 (11.6)	1.9 (1.5 , 2.5)
18-35	2,186 (6.5)	Ref
>35	268 (6.9)	1.1 (0.9 , 1.3)
Maternal Race/Ethnicity		
White, non-Hispanic	1,462 (5.6)	Ref
Black, non-Hispanic	1,097 (12.2)	2.3 (2.2 , 2.5)
Hispanic	50 (5.0)	0.9 (0.6 , 1.2)
Other, non-Hispanic	109 (7.9)	1.4 (1.1 , 1.9)
Maternal Education		
< =High School	1,474 (8.8)	1.8 (1.6 , 2.0)
>High School	1,244 (5.1)	Ref
Maternal Smoking During Pregnancy		
Yes	680 (10.8)	2.0 (1.7 , 2.3)
No	1,851 (5.7)	Ref
Family Income		
Near or below 100% of poverty	1,199 (9.4)	1.9 (1.7 , 2.1)
Above 100% of poverty	1,278 (5.2)	Ref
Pre-pregnancy Body Mass Index		
<18.5	294 (11.6)	1.9 (1.6 , 2.3)
18.5-24.9	1,224 (6.5)	Ref
25.0-29.9	573 (5.8)	0.9 (0.8 , 1.0)
>30.0	600 (6.7)	1.0 (0.9 , 1.2)
Stressful Events in Pregnancy		
None	544 (5.5)	Ref
One or More	2,118 (7.1)	1.3 (1.2, 1.5)
WIC during Pregnancy		
No	1,231 (5.4)	Ref
Yes	1,429 (8.4)	1.6 (1.4, 1.8)

*unweighted frequencies and weighted percentages

Table 2: Distribution of risk factors for LBW, by timing of last menstrual period (pre and post smoking ban), unweighted frequencies, weighted percentages, and chi squared p value, singleton births, Ohio PRAMS, 2005-2009.

unweighted n=6,872

Variable	LMP Before May 1, 2007 # (%)	LMP After May 1, 2007 # (%)	Chi Squared P Value
Maternal Age (years)			
<18	193 (3.2)	119 (3.8)	0.53
18-35	3,468 (87.6)	2,197 (86.4)	
>35	383 (9.2)	265 (9.9)	
Maternal Race/Ethnicity			
White, non-Hispanic	2,590 (77.5)	1,468 (76.1)	0.40
Black, non-Hispanic	1,366 (15.4)	1,048 (16.0)	
Hispanic	94 (3.1)	65 (3.8)	
Other, non-Hispanic	147 (4.0)	94 (4.1)	
Maternal Education			
< =High School	1,999 (46.9)	1,209 (42.7)	**0.01**
>High School	2,198 (53.1)	1,466 (57.3))	
Maternal Smoking During Pregnancy			
Yes	818 (18.7)	473 (20.3)	0.39
No	3,254 (81.3)	1,885 (79.7)	
Family Income			
Near or below 100% of poverty	1,537 (34.7)	1,060 (35.3)	0.73
Above 100% of poverty	2,313 (65.3)	1,410 (64.7)	
Pre-pregnancy Body Mass Index			
<18.5	343 (7.2)	222 (7.2)	0.20
18.5-24.9	1,955 (49.6)	1,181 (46.7)	
25.0-29.9	954 (22.9)	596 (23.2)	
>30.0	902 (20.3)	634 (22.9)	
Stressful Events in Pregnancy			
None	954 (25.2)	595 (24.9)	0.06
One or More	3,150 (74.8)	2,025 (75.1)	
WIC during Pregnancy			
No	2,145 (57.5)	1,316 (56.4)	0.50
Yes	1,971 (42.5)	1,312 (43.6)	

*unweighted frequencies and weighted percentages

Table 3: Final multivariable logistic regression model of putative association between last menstrual period occurring pre/post the Ohio statewide smoking ban and low birth weight, while controlling for other factors, singleton births, Ohio PRAMS, 2005-09.
n=5,681

Variable	β	β Standard Error	Adjusted Odds Ratio	95% Confidence Intervals for Odds Ratio
Intercept	-3.21	0.07	0.04	0.04 , 0.05
LMP on or after May 1, 2007	**0.03**	**0.04**	**1.03**	**0.95 , 1.11**
Less than 18 years old	0.50	0.17	1.64	1.17 , 2.31
35 years or older	0.28	0.10	1.33	1.09 ,1.61
High school or less	0.30	0.07	1.35	1.18 , 1.55
Smoke during pregnancy	0.62	0.08	1.86	1.59 , 2.16
Underweight	0.34	0.12	1.41	1.12 , 1.78
Overweight	-0.17	0.08	0.85	0.73 , 0.98
Obese	-0.06	0.08	0.94	0.81 , 1.09
Black, nonHispanic	0.76	0.06	2.13	1.89 , 2.39
Hispanic	-0.21	0.21	0.81	0.54 , 1.21
Other, nonHispanic	0.45	0.16	1.57	1.15 , 2.15
Family income at or below 100% poverty	0.21	0.08	1.24	1.06 , 1.45
WIC during pregnancy	-0.08	0.08	0.92	0.79 , 1.08

Conclusion

These results are limited by the facts that, 1) information was self-reported (which sometimes leads to underreporting of undesirable behaviors) and 2) no information was available on ETS exposure or employment of the mother during pregnancy. It is possible that different results may have been observed among subgroups of women employed during pregnancy in industries most likely impacted by the ban (such as in restaurants or bars). It is unknown what proportion of pregnant women worked in such jobs in Ohio. Thus, it is also unknown whether improvements in outcomes of these women could have been obscured by their incorporation into a larger group of women who were unexposed to ETS irrespective of the existence of a smoking ban.

Nevertheless, the Ohio PRAMS survey was expressly developed to focus on examination of risk factors for LBW. Thus, the sample size of women with LBW infants from this data source had adequate statistical power to permit assessment these relationships. Furthermore, the size and direction of associations observed in the data between LBW and other known risk factors (e.g., smoking during pregnancy, education, race/ethnicity) were consistent with the published literature, lending support for the validity of the information from PRAMS.

In conclusion, no evidence of statewide improvement in odds of poor birth outcomes was observed from this analysis. Given that no information was available to assess industry of employment (or any employment) during pregnancy, a relationship between the ban and birth outcomes could not be examined among subgroups of pregnant women most likely to have been impacted by the ban, thereby limiting interpretation of these results.

References

1. Hofhuis W, de Jongste JC, Merkus PJ. Adverse health effects of prenatal and postnatal tobacco smoke exposure on children. Arch Dis Child 2003; 88:1086–90.
2. The Health Consequences of Involuntary Exposure to Tobacco Smoke: 2006 Surgeon General's Report. http://www.cdc.gov/tobacco/data_statistics/sgr/2006/index.htm
3. Misra DP and Nguyen R. Environmental Tobacco Smoke and Low Birth Weight: A Hazard in the Workplace? Environmental Health Perspectives 1999; 107, Supplement 6:897-904.
4. U.S. Census Bureau News Release. Women more likely to work during pregnancy. 2008. http://www.census.gov/newsroom/releases/archives/employment_occupations/cb08-33.html
5. Centers for Disease Control and Prevention (CDC) (January 2009). Reduced hospitalizations for acute myocardial infarction after implementation of a smoke-free ordinance—City of Pueblo, Colorado, 2002–2006.*MMWR Morb. Mortal. Wkly. Rep.* **57** (51): 1373–7.
6. Naiman A, Glazier RH, Moineddin R. Association of anti-smoking legislation with rates of hospital admission for cardiovascular and respiratory conditions. *CMAJ* 2010; **182** (8): 761–7.
7. Hopkins DP, Razi S, Leeks KD, Priya Kalra G, Chattopadhyay SK, Soler RE. Smokefree policies to reduce tobacco use. A systematic review. *Am J Prev Med.* 2010; **38** (2 Suppl): S275–89.
8. American Public Health Association News Release. Citywide smoking ban contributes to significant decrease in maternal smoking, preterm births. 2010. http://www.apha.org/about/news/pressreleases/2010/maternal+smoking+prevalence.htm
9. Kabir Z, Clarke V, Conroy R, et.al. Low birthweight and preterm birth rates 1 year before and after the Irish workplace smoking ban. BJOG. 2009; 116(13):1782-7.