Web Appendix 1: County names

Names of the 387 counties that satisfied the criteria for inclusion in the study are listed below:

Web Appendix 2: Sensitivity analyses

A collection of sensitivity analyses were performed by adjusting different aspects of the model reported in the main manuscript. Descriptions of the 6 separate variations are provided below. Models 1 and 2 explore the impact of analyzing a shorter or longer time window following the ban. Models 3 and 4 explore the impact of eliminating random effects from the model presented in the main manuscript. Model 5 explores county level models, and Model 6 explores the use of a time-ban interaction.

1) Extending the time span to include all post-ban months through 2008. The model is identical to that used in the primary manuscript, except the time range has been expanded to use all data from 1999—2008, inclusively. This approach would help identify long-term drops in acute myocardial infarction (AMI) following comprehensive smoking bans.

2) Limiting the time span to include just 2 post-ban months. The model is identical to that used in the primary run, except the time range has been shortened to only 2 months of post-ban data rather than the 12 months used in the primary manuscript. This modeling approach would help identify immediate reductions in AMI following comprehensive smoking bans.

3) Omitting the random effect for county-level linear trend (random slope and intercept). This analysis provides an estimate of the state-specific ban effect under the assumption that the admissions rate is the same across all counties in a state. The model may be written as

\[ Y_{c,a,g} \sim \text{Poisson}(\mu_{c,a,g}) \]

\[
\log(\mu_{c,a,g}) = \log N_{c,a,g} + \beta_0 + \beta_1 \times t + ns(t, df) + \sum_{j=2}^{12} \beta_j x_j \\
+ \beta_{13} 1_{(75<a<85)} + \beta_{14} 1_{(84<a)} + \beta_{15} 1_{(g=\text{male})} + (\beta_{16} + \gamma_{16}) B_t^c
\]
4) Omitting the random effect for the county-level ban effect. This analysis provides an estimate of the state-specific ban effect under the assumption that the ban effect is the same across all counties in a state. The model may be written as

\[ Y_{t,a,g}^c \sim \text{Poisson}(\mu_{t,a,g}^c) \]

\[
\log(\mu_{t,a,g}^c) = \log N_{t,a,g}^c + (\beta_0 + \gamma_0^c) + (\beta_1 + \gamma_1^c)t + ns(t, df) + \sum_{j=2}^{12} \beta_j x_j \\
+ \beta_{13} 1_{(75<a<85)} + \beta_{14} 1_{(84<a)} + \beta_{15} 1_{(g=male)} + \beta_{16} B_t^c
\]

5) The state-level Poisson model was adapted to allow for differences in secular trend and other effects at the county-level. The following model was fit for each county:

\[ Y_{t,a,g}^c \sim \text{Poisson}(\mu_{t,a,g}^c) \]

\[
\log(\mu_{t,a,g}^c) = \log N_{t,a,g}^c + \beta_0^c + \beta_1^c t + ns^c(t, df) + \sum_{j=2}^{12} \beta_j^c x_j \\
+ \beta_{13}^c 1_{(75<a<85)} + \beta_{14}^c 1_{(84<a)} + \beta_{15}^c 1_{(g=male)} + \beta_{16}^c B_t^c
\]

The random effects have been eliminated, and all state-level coefficients have been replaced by county-level estimates. County-level ban effects were pooled to estimate state-level ban effects using the methods described by DerSimonian and Laird (1986), then state-level ban effects were pooled using the same technique.

6) A fixed-effect term for a time-ban interaction was added to the model presented in the main manuscript, allowing for more flexible modeling of the ban effect.

\[ Y_{t,a,g}^c \sim \text{Poisson}(\mu_{t,a,g}^c) \]

\[
\log(\mu_{t,a,g}^c) = \log N_{t,a,g}^c + (\beta_0 + \gamma_0^c) + (\beta_1 + \gamma_1^c)t + ns(t, df) + \sum_{j=2}^{12} \beta_j x_j \\
+ \beta_{13} 1_{(75<a<85)} + \beta_{14} 1_{(84<a)} + \beta_{15} 1_{(g=male)} + (\beta_{16} + \gamma_{16}^c) B_t^c + \beta_{17} t B_t^c
\]
The new interaction term is represented by $\beta_{17} t B_{t}$. All other terms are equivalent to those described in the main manuscript. Thus, this new model includes a ban effect with both a step and a change in slope, allowing for the ban effect to grow (or shrink) over time. For a single reported value, the net ban effect was estimated at 12 months following the ban as $\beta_{16} + 12 \beta_{17}$. The standard error of this linear combination of terms was computed using the model covariance matrix relating $\beta_{16}$ and $\beta_{17}$ and assuming a normal model for the sample estimates.

Results from all sensitivity analyses are summarized in Figure 3. Estimates and 95% confidence intervals are shown, grouped by model degrees of freedom devoted to the nonlinear secular trend fit in each state (or in the case of sensitivity analysis #5, in each county). Results from all sensitivity analyses were in substantive agreement with results reported in the main manuscript: none of the models showed statistically significant ban effects after allowing for nonlinearity in the secular trend. The 6th sensitivity analysis was particularly interesting, because it explored potentially more realistic forms for the behavior of a possible ban effect for a 12-month period after the ban. A notable increase in the standard error associated with the ban effect for Model 6 as the degrees of freedom in the secular trend is increased is evident. We believe this increase in the standard error reflects increasing uncertainty in the slope of the secular trend, especially near the boundaries, as the degrees of freedom become large. This uncertainty would influence the standard error of the time-ban interaction coefficient. The results were substantively similar (to the model in the main manuscript). Therefore, we focused on the model (presented in the main manuscript) that allowed only for a shift in intercept. This also facilitates comparison to the large literature using the intercept-only approach for ban effect (Sargent et al. (2004), Barone-Adesi et al. (2006), Bartecchi et al. (2006), Juster et al. (2007), Cesaroni et al. (2008), Pell et al. (2008), Lemstra et al. (2008), Gasparrini (2009), Lightwood et al. (2009), Sims et al. (2010), Shetty et al. (2010), Bruinjtes et al. (2011)).
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Table 1: First row - mean number of Medicare enrollees (in millions) each year aggregating across the 387 counties studied. Second row - mean monthly hospitalization rate (per 100,000 Medicare enrollees) each year across the 387 US counties studied. (AMI, acute myocardial infarction)
Figure 1: The number of US counties that implemented any ban between January 1999 and December 2008 and had no previous smoking ban. The dashed line represents the 387 counties included in the study.
Figure 2: Map of US counties with comprehensive smoking bans and which satisfied other inclusion criteria (see Section 2), 1999—2008. Red represents older bans (2002—2004), orange represents bans in the mid-term (2005—2006), and yellow represents newer bans (2007—2008).
Figure 3: Estimated ban effect for the models reported in the main manuscript (labeled “reported”) and in each of the 6 sensitivity analyses (labeled 1-6 according to list in Web Appendix 2), US states, 1999—2008. The sensitivity analyses are grouped by degrees of freedom (df) in the secular trend. As with the reported results, a statistically significant effect near -5% is estimated for the linear secular trend in each sensitivity analysis. However, no estimates in any of the sensitivity analyses are statistically significant after relaxing the linearity assumption.
Figure 4: Point estimates and 95% confidence intervals of the percent decrease in hospital admission rates for acute myocardial infarction associated with comprehensive smoking bans comparing 4 models of secular trend in each US state, 1999—2008. States are labeled with the number of months before the ban and the number of case counties for the state. Results for each state are grouped by degrees of freedom (df) in the secular trend. In addition to secular trend, these effects are also adjusted for seasonality, gender and age group. The last 4 sets of estimates
show pooled results all counties and age groups, then 3 additional sets when a separate model is fit for each age group.