The recent commentary by Greenland et al. (1) criticizes the “floating absolute risk” (FAR) method for presenting uncertainty in estimates of polytomous exposures. However, it then goes on to propose “floating trends,” which are exactly the same, to within a constant!

The criticism that Greenland et al. raise of the FAR confidence intervals is that they are not, directly, confidence intervals for the relative risk estimates for each category relative to the baseline category. This is obviously true—they are not meant to be. The whole point of the FAR method is to allow computation of confidence limits for relative risks for any pair of categories. The variance of the log relative risk is obtained by adding the “floating variances” for the two categories, in exactly the same way as one does in, for example, an unpaired t-test.

Standard confidence intervals do not provide the same information. This is clear in Greenland et al.’s example on birth weight and breast cancer. For example, one can combine the risk estimates for the first two and the last three categories in table 1 (1, p. 1078), by taking an inverse (FAR) variance weighted average of the corresponding log relative risks, and can deduce that the breast cancer risks associated with birth weights lower than 3.5 kg are significantly lower than those at or above 3.5 kg (relative risk = 0.77; 95 percent confidence interval: 0.63, 0.95). The “correct” confidence limits are uninformative for this comparison. Greenland et al. suggest that the comparisons of other relative risk estimates would be possible if the average of the covariances among the log relative risk estimates were provided. This has the same rationale as the FAR method, providing information about the uncertainty associated with the baseline category which is lost in the standard analysis. Both methods are exact if there is no confounding due to other risk factors and approximate otherwise. The FAR method is aesthetically and perhaps formally preferable, because it always gives estimates that are invariant to the choice of the baseline category, and because it minimizes the average covariance (and hence the discrepancy between the confidence limits derived and the “exact” confidence limits that would be obtained from the full covariance matrix). The main reason for preferring the FAR confidence limits, however, is that they do provide a valid visual impression of the uncertainty associated with all of the estimates.

In the case of unmatched case-control (or cohort) analyses, an effect similar to that of the FAR method can also be achieved in logistic regression by fitting a separate parameter for each category, dispensing with the “intercept” parameter. Greenland et al. suggest reporting these results in terms of “case-control ratios” and describe the resulting curves as “floating trends.” As their name suggests, they are exactly the same in concept as the floating absolute risks except for an arbitrary multiplicative factor. This factor, which describes the overall ratio of cases and controls, is of course a reflection of the study design and contains no additional information on relative risks. The FAR method is invariant to multiplying all risks by an arbitrary constant. However, from a presentational point of view, it seems simpler to assign one group a risk of 1.0, since then the point estimates are the same as the standard relative risk estimates relative to that category, which one would normally wish to present anyway. The method suggested by Greenland et al. requires an additional column of estimates of no intrinsic value.

While the two methods are essentially identical in the simplest case with no confounding factors, there could be differences if other factors are entered into the model. The approach of Greenland et al. requires that any confounding factors be defined to have a (weighted) mean of zero, which will approximately remove the covariance between the estimates in each category due to confounding variables. The same effect is achieved by the FAR method automatically, without any reparameterization.

The other major advantage of the FAR method is that it can be applied directly to conditional logistic regression analyses of matched case-control studies and to Cox regression, where there is no intercept term and therefore no analogy of the method suggested by Greenland et al.

Thus, the only apparently substantial criticism of the FAR confidence intervals by Greenland et al. is that they “will never cover any parameter with 95 percent frequency” (1, p. 1083) and might therefore be misinterpreted. We chose the name “floating absolute risk” to describe the parameter for which they are correct confidence intervals, to draw attention to the fact that this is a new statistical concept. The FARs for a categorical variable are the logarithms of the actual values of the underlying rates in each group with the same unknown constant added to each. This explains the apparent paradox that this arbitrary constant can be chosen to make the estimate of the FAR in the reference group (as distinct from its actual value) equal to zero. It had not escaped our notice that it seems odd to assign a variance to zero; but that does not mean it is incorrect. We believe that the introduction of FAR unifies the mathematical structure as well as the analysis of data from case-control and cohort studies, and is thus conceptually as well as practically useful.

REFERENCE


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FOUR OF THE AUTHORS REPLY

We thank Easton and Peto (1) for their comments. Since their introduction (2), floating absolute risk (FAR) confidence limits have become controversial (3, 4), with one authority asserting that “the way most epidemiologists now calculate confidence intervals for relative risks is not appropriate” and that standard confidence limits “should generally be replaced” by FAR limits (3). As a result, some papers have presented FAR limits as if they were limits for relative risks (5–7). Easton and Peto (1) agree that such usage is a mistake, but they maintain that FAR intervals are correct for “a new statistical concept.” As we showed (8), however, this assertion is incorrect: FAR intervals are not valid confidence intervals for anything. Furthermore, it is logically and statistically incorrect to assign a nonzero variance to zero, as the original FAR method requires: Like any fixed number, zero has zero variance. Because FAR intervals are narrower than standard intervals, placing FAR intervals around relative risk estimates (as in the paper by Easton et al. (3) and later papers (3, 5–7)) misleads readers into thinking that the data contain more information about the relative risks than is present.

To avoid these problems, we proposed modification of the FAR method by anchoring the “floated” standard errors to the underlying rates or risks in a cohort study or to the underlying case-control ratios in a case-control study (8). Easton and Peto (1) incorrectly label as “arbitrary” the multiplicative constant used in this anchoring. The constant is instead just the one needed to create valid intervals from the FAR standard errors. In a cohort study, the resulting additional column of estimates has intrinsic value because it displays the trend on an absolute incidence scale, with incidence calculated at the average covariate values (a procedure that roughly approximates standardization to the total sample) (9, p. 1081).

The absolute scale is essential for evaluating the public health importance of the factor. In a case-control study, only relative values are estimable without external information. Nonetheless, plotting case-control ratios achieves the primary goal of giving a valid impression of trend uncertainty; furthermore, one can convert these ratios into absolute-incidence estimates if one divides them by the ratio of the case and control sampling fractions (9). Conversion to incidence can also be done if one knows the overall incidence in the source population (10), as in a study that takes its cases from a population-based registry.

Easton and Peto state that the FAR method does not require an additional column. The absence of that extra column is exactly why the FAR method gives the incorrect impression that the intervals it generates apply to the relative risks. It is also no advantage that the FAR method can be applied directly to conditional logistic and Cox regression, because the intervals it generates in those applications are just as invalid as the FAR intervals in other settings. Nor does the FAR method have any computational advantage: It requires rather awkward covariance manipulations (2), whereas our modification requires only subtracting the covariate means from the covariates.

In summary, although we agree with Easton et al. (2) regarding the unsatisfactory properties of conventional confidence intervals, we maintain that their original FAR intervals (2, 3) are invalid and must be modified if they are not to mislead readers. We have suggested some computationally simpler ways to accomplish the same goals as the FAR method while avoiding its problems (8). Our “floated trend” approach can be viewed as a direct correction of the FAR method. We believe that, without this correction, conventional intervals would be preferable to FAR intervals. Finally, we again caution that none of the methods under consideration (conventional, FAR, and our floated-trend approach) provide a satisfactory summary of trends compatible with the data under the analysis model (8, p. 1085).

REFERENCES


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We read with interest the article by Holmen et al. (1) regarding the increased frequency of somatic and psychological health problems among Norwegian adolescents who smoke daily. Their study adds epidemiologic support to previous observations, mainly from the United States and northern European countries. This association has not been assessed in other developed countries where the tobacco epidemic presents certain differential characteristics (2). The prevalence of smoking among young men and women in Catalonia, Spain, has increased in the last decade (3, 4). Smoking initiation rates were very low for women prior to the 1970s, and they have only recently begun to match those of men (5). We assessed the association between daily smoking among adolescents and self-reported health, use of medication, and use of health services, using data from the Catalan Health Interview Survey.

The Catalan Health Interview Survey was conducted in 1994 in a randomly selected sample (n = 15,000) of the noninstitutionalized population of Catalonia. A two-step random sample strategy was used (5, 6). The survey included sociodemographic data and information on self-perceived health and morbidity, lifestyle habits (including daily, occasional, and past smoking), utilization of health services, and prevalence of chronic conditions. Face-to-face interviews were conducted in respondents' homes (January–December 1994). For the present study, we included data from 564 boys and 524 girls aged 15–18 years who responded directly. Younger children (aged ≤14 years) were excluded from analysis, since data on their tobacco consumption was not obtained (7). The mean age at which smoking began, the mean number of cigarettes smoked per day, and the mean years of daily smoking, as well as their standard errors (SE), were calculated. The proportions of daily smokers and nonsmokers and the corresponding odds ratios (OR) (and 95 percent confidence intervals) among daily smokers versus never smokers for selected health conditions and use of health services were also derived.

In these data, 24.6 percent of boys and 20.8 percent of girls were daily smokers (≥1 cigarette/day). The mean age at starting smoking was 15.0 years (SE 0.1) for boys and 14.6 years (SE 0.1) for girls (t = 2.58, p = 0.01). No differences were found in the mean number of cigarettes smoked per day (11.3 (SE 0.6) in boys and 10.4 (SE 0.6) in girls) or in the mean number of years of smoking (1.9 years (SE 0.1) and 2.1 years (SE 0.1), respectively).

Boys and girls who smoked daily more frequently rated their health poor (OR = 2.5 in boys and OR = 4.3 in girls; p < 0.05 (table 1)). Asthma and bronchitis were significantly associated with daily smoking in boys but not in girls. Information on the use of selected medications was elicited for the 2 days before the interview; hence, the corresponding frequencies of use were low and are hardly comparable with those in the Norwegian study (table 1). In boys, there were no differences between the proportions of smokers and nonsmokers who visited a general practitioner, whereas daily smokers visited a specialist (OR = 1.6; p < 0.05) and were hospitalized (OR = 3.0; p < 0.05) more often. Girls who smoked daily were more likely to have visited a general practitioner, to have seen a specialist, and to have been hospitalized, although the associations were not statistically significant (table 1).

The almost twofold higher prevalence of daily smoking among adolescents in Catalonia as compared with the corresponding figures in Norway (12.3 percent in boys and 15.8 percent in girls, computed from Holmen et al.'s table 2 (1, p. 150)) is noteworthy. The patterns of association between daily smoking and the health variables examined were similar and were consistent with those in the Norwegian study. Smoking in adolescence is associated with health problems in adolescence itself, in addition to the relations of adolescent smoking with heavy smoking, lower cessation rates, longer duration of smoking, and higher nicotine dependence in adulthood (8).

The main limitation of the present study is its cross-sectional design, which limits the causal interpretation of the associations observed. This means that it is not possible to discern whether poor health status and somatic symptoms are caused by smoking or vice versa—that is, whether poor health promotes smoking or maintains smoking. Our sample was smaller than that of the Norwegian study, and hence the confidence intervals were wider. Nevertheless, the patterns of association were similar.

In summary, in terms of the association between daily smoking and health problems in adolescents, the results found in Catalonia mirror those reported from the United States and northern European countries (1, 9–12), despite differences in the patterns of smoking initiation and tobacco consumption in southern European countries. These short term tobacco-related conditions in adolescent smokers require prompt attention and should be considered alongside the well known long term health consequences of smoking.

References

Letters to the Editor


Editor’s note: Holmen and colleagues were asked if they wished to respond to this letter but chose not to do so.

TABLE 1. Association between daily smoking by teenagers aged 15–18 years and selected health conditions and use of health services, Catalan Health Interview Survey, Catalonia, Spain, 1994

<table>
<thead>
<tr>
<th></th>
<th>Boys</th>
<th>Girls</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Daily smokers (n = 139)</td>
<td>Non-smokers (n = 409)</td>
</tr>
<tr>
<td>Self-reported poor health*</td>
<td>12.6 (10.8, 13.0)</td>
<td>6.5 (2.5, 2.8)</td>
</tr>
<tr>
<td>Asthma</td>
<td>8.6 (7.9, 9.3)</td>
<td>4.9 (7.9, 6.6)</td>
</tr>
<tr>
<td>Bronchitis</td>
<td>15.8 (15.1, 16.5)</td>
<td>10.9 (10.2, 11.5)</td>
</tr>
<tr>
<td>Visited general practitioner‡</td>
<td>11.0 (10.0, 12.0)</td>
<td>6.5 (6.0, 7.0)</td>
</tr>
<tr>
<td>Visited specialist‡</td>
<td>9.6 (8.7, 10.5)</td>
<td>4.9 (4.4, 5.4)</td>
</tr>
<tr>
<td>Hospitalized‡</td>
<td>24.8 (23.9, 25.7)</td>
<td>12.7 (12.0, 13.4)</td>
</tr>
<tr>
<td>Medication use†</td>
<td>1.0 (1.0, 1.1)</td>
<td>0.7 (0.7, 0.8)</td>
</tr>
<tr>
<td>Pain medication†</td>
<td>0.6 (0.5, 0.7)</td>
<td>0.5 (0.4, 0.5)</td>
</tr>
<tr>
<td>Tranquilizers†</td>
<td>1.7 (1.6, 1.8)</td>
<td>1.1 (1.0, 1.2)</td>
</tr>
<tr>
<td>Visited primary care‡</td>
<td>26.6 (26.0, 27.2)</td>
<td>14.2 (13.6, 14.7)</td>
</tr>
<tr>
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<td>10.1 (9.6, 10.7)</td>
<td>6.5 (6.0, 7.0)</td>
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<td>1.1 (1.0, 1.2)</td>
</tr>
</tbody>
</table>

* Bad/fair health vs. excellent/very good/good health.† During the past 2 days.‡ During the past 12 months.