Hansen SN, Schendel DE, Parner ET. Explaining the increase in the prevalence of autism spectrum disorders: the proportion attributable to changes in reporting practices. *JAMA Pediatr.* Published online November 3, 2014.

eAppendix. The analytic model

This supplementary material has been provided by the authors to give readers additional information about their work.
The analytic model

Let \( T_1(z) \) and \( T_2(z) \) denote the ages at which the diagnostic change and inclusion of outpatients occurred for an individual born at calendar time \( z \). For each subcohort, \( t = 1, 2, \ldots, 6 \), we assume for the diagnostic rates that

\[
\lambda_t(t, z; \varphi_t, \alpha_t, \beta_t) = \lambda_{0t}(t) \exp(\varphi_t z) \exp(\alpha_t 1_{t > T_1(z)}) \exp(\beta_t 1_{t > T_2(z)}), \quad t > 0,
\]

where \( \lambda_{0t} \) is the baseline diagnostic rate in the \( i \)th subcohort. The parameter \( \varphi_t \) describes the change in diagnostic rates due to calendar time, the parameter \( \alpha_t \) describes the change in diagnostic rate due to the diagnostic change and the parameter \( \beta_t \) describes the change in diagnostic rates due to the inclusion of outpatient data. In terms of the undiagnosed risk function (survivor function) for a child born in the \( i \)th subcohort we have

\[
S_i(t, z; \varphi_t, \alpha_t, \beta_t) = \begin{cases} 
S_{0i}(t)^{\exp(\varphi_t z)} & , t \leq T_1 \\
S_{0i}(T_1)^{\exp(\varphi_t z(1-\exp(\alpha_t)))} S_{0i}(t)^{\exp(\varphi_t z + \alpha_t)} & , T_{1j} < t \leq T_2 \\
S_{0i}(T_1)^{\exp(\varphi_t z(1-\exp(\alpha_t)))} S_{0i}(T_2)^{\exp(\varphi_t z + \alpha_t)(1-\exp(\beta_t))} S_{0i}(t)^{\exp(\varphi_t z + \alpha_t + \beta_t)} & , t > T_2 
\end{cases}
\]

(1)

The parameters \( \exp(\alpha_t) \) and \( \exp(\beta_t) \) are the hazard ratios corresponding to the change in diagnostic criteria and inclusion of outpatient data in the \( i \)th subcohort respectively. The model can be fitted by a Cox regression model with 2 time-dependent covariates and date of birth entering as a linear covariate.

When estimating the overall effects \( \alpha_o, \beta_o \) of the changes in reporting practices we fit a stratified Cox proportional hazards model over all birth years assuming that the 2 effects of changing reporting practices are the same in all cohorts but allowing for different baseline rates and calendar effect.

The average expected prevalence curves of Figure 3 are obtained by averaging the expected prevalence curves obtained by inserting parameter estimates in (1) over all individuals in the study cohort. More specifically, the “no effect” curve (scenario 1) is given by

\[
\text{CIP}_{\text{noeffect}}(t) = 1 - \frac{1}{N} \sum_{j=1}^{N} S_i(t, z_j; 0, 0, 0) = 1 - S_{0i}(t)
\]

where \( N = 677,915 \) is number of individuals in the study. This prevalence curve reflects the baseline prevalence of the study cohort, i.e. the prevalence of an average individual in our study that is not under influence of calendar effect or changes in reporting practices. The “calendar effect” curve (scenario 2) is given by

\[
\text{CIP}_{\text{calendar}}(t) = 1 - \frac{1}{N} \sum_{j=1}^{N} S_i(t, z_j; \varphi_{i_j}, 0, 0),
\]

where \( i_j \) is the subcohort of the \( j \)th individual. The 3 effect curves “diagnostic change”, “outpatient” and “combined” (scenarios 3-5) are all estimated by inserting the respective overall parameter estimates as well as the calendar effect estimate:

\[
\text{CIP}_{\text{diagnostic change}}(t) = 1 - \frac{1}{N} \sum_{j=1}^{N} S_i(t, z_j; \varphi_{i_j}, \alpha_o, 0)
\]

\[
\text{CIP}_{\text{outpatient}}(t) = 1 - \frac{1}{N} \sum_{j=1}^{N} S_i(t, z_j; \varphi_{i_j}, 0, \beta_o)
\]

\[
\text{CIP}_{\text{combined}}(t) = 1 - \frac{1}{N} \sum_{j=1}^{N} S_i(t, z_j; \varphi_{i_j}, \alpha_o, \beta_o)
\]