Suicide mortality in Switzerland: Gender-specific differences and the impact of family integration

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Suicide has become one of the leading causes of death of Swiss males aged between 15 and 44 years, whose age-standardised rates are about three times higher than those for females. We compared suicide mortality of Swiss men and women aged 15–79 and investigated gender-specific differences from 1950–2007. Furthermore, we explored whether measures of family integration can explain temporal suicide trends. The use of multivariate age-period-cohort models avoids age aggregation and allows the exploration of heterogeneous time trends across age, period and birth cohort. In addition, explanatory variables can be included. We found strong gender-specific differences in suicide mortality. While the same risk factors may act on age and overdispersion, there was no significant correlation between gender-specific cohort effects. Family integration had an impact on Swiss suicide risk, but only partially explained the underlying trends.

Keywords: Age-Period-Cohort model; Bayesian analysis; Family integration; Suicide; Switzerland.

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1 Background and objectives

Age and gender are well established risk factors of suicide. In many countries especially elderly have a higher risk to commit suicide (Granizo *et al.*, 1996; Shah and De, 1998). Suicide rates of males are generally higher than those of women. Besides age- and gender-specific factors, environmental factors occurring at a specific date in time may influence suicide risk. For example, Lester (1990) examined suicide rates in Switzerland during the period when domestic gas was detoxified. This study did not only indicate a decline of suicide by means of domestic gas, but also a decline of the overall suicide rate, indicating that people did not switch to alternative methods. Thus, if one method of suicide is made less available, such as detoxification of domestic gas or strict gun control rules, a reduction in the overall suicide risk might be observed. Experiences common to a particular birth cohort, e.g. war, might also influence the suicide risk. Hence a better understanding of age, period and cohort effects might help to target effective preventive strategies. Over the last years a number of age-period-cohort (APC) analyses of suicide rates have been published, see for example Granizo *et al.* (1996); Etzersdorfer *et al.* (1996); Snowdon and Hunt (2002); Stockard and O'Brien (2002); Gunnell *et al.* (2003); Ajdacic-Gross *et al.* (2006).

Although Switzerland is an affluent country the suicide rates observed are quite high compared to other countries (Levi *et al.*, 2003). To gain more information on gender-specific differences in Switzerland Ajdacic-Gross *et al.* (2006) performed univariate APC analyses on suicide data over the last century. They found similar age and period effects, but stronger cohort effects for males than for females. However, since males and females might be subject to similar risk factors it seems justified to model them jointly suggesting the possibility of common time (age, period, cohort) effects.

In this paper we will apply multivariate APC models to capture trends in the sex ratio. Such models borrow strength from both genders for estimating common sets of time effects, for example the age effects, while the remaining parameter sets can be different across gender. The well-known identifiability problem of APC models is avoided since differences of gender-specific time effects are identifiable and can be interpreted as log relative risk (Riebler and Held, 2010). Since social aspects are strongly associated with suicide risk, we will further investigate whether changes in variables related to family integration can explain suicide trends. Following the theory of Durkheim (Durkheim, 1897) numerous papers were published investigating the relationship between marital status and suicide mainly based on time-series analyses (Breault and Barkey, 1982; Lester, 1986; Stack, 1987, 1990a,b, 1992; Surault, 1992; Rossow, 1993; Lester, 1994). Populations with high divorce rates, for example, were found to have high suicide rates even when controlling for confounding variables such as socio-economic status (Stack, 1990a). Nowadays only few publications investigate a correlation between family integration and suicide. Stack (1990b) assumed that the association between being divorced and committing suicide has changed over the years. For example, divorce and unmarried couples are more common and more accepted. We will apply univariate APC models separately to males and females replacing the period effects by an explanatory variable on family integration assuming either a parametric or non-parametric effect. Replacing one set of effects by an explanatory variable is also a valid solution to the non-identifiability problem of APC models (Brown and Kessler, 1988). However, this is only the case if the covariate-effect does not depend on the time scale for which it is used. Otherwise a linear dependence between the three time scales remains, see the introduction of Riebler (2010, page 4) for a discussion of this topic. Further, we apply a multivariate age-cohort (AC) model replacing the period effects by a correlated non-parametric covariate effect of family integration. The rate of unemployment is included in the model formulation to account for confounding. However, as the Swiss unemployment rate is so low, variation may be quite small to affect the overall suicide trends (Stack, 1989).

2 Data and Methods

Annual age- and gender-specific suicide mortality counts and mid-year population data, 1950–2007, were obtained from the Swiss Federal Statistical Office (statistics of causes of death). Age groups are stratified by five-year intervals: 15–19, …, 75–79, resulting in 13 age groups and

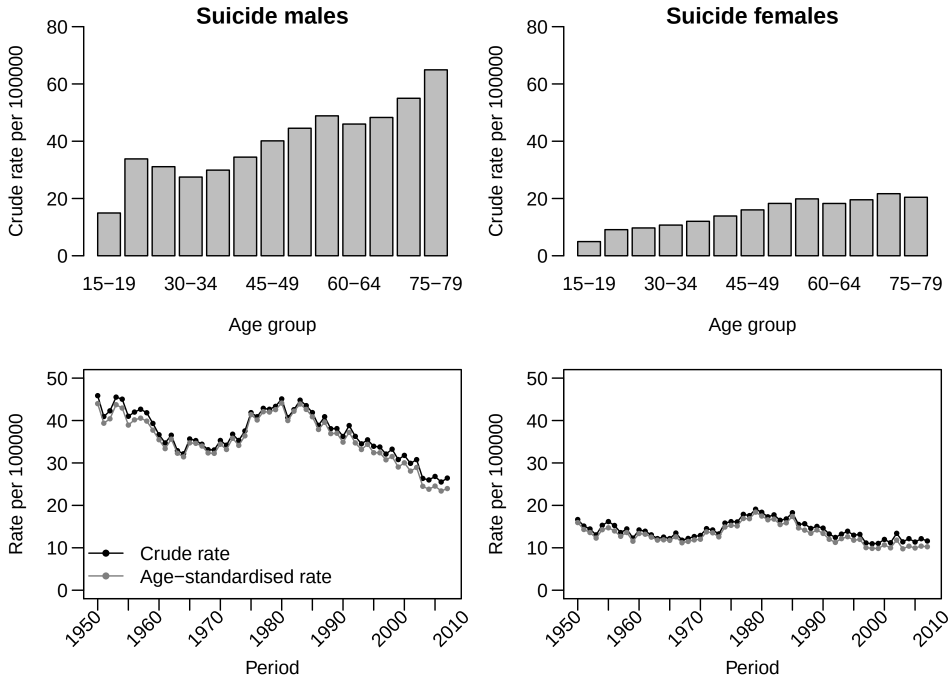

Figure 1: Suicide rates from 1950-2007. Top: Age-specific crude rates for males and females. Bottom: Crude and age-standardised rates of all periods for males and females.

58 one-year periods. We omitted data for children under the age of 15, when suicide is rare, and for adults over the age of 79 because for elderly assisted suicides have become frequent in Switzerland. Figure 1 shows age-specific crude rates and both crude and age-standardised rates per 100 000 persons of all periods for males and females. Age-standardisation was performed using the WHO world standard population (Ahmad *et al.*, 2001).

First, the data were analysed using Bayesian APC models for multiple outcomes (Riebler and Held, 2010). We assume that the number of suicides y_{ijg} of age group $i = 1, \dots, 13$, period $j = 1, \dots, 58$ and sex g is Poisson distributed with rate $n_{ijg} \times \lambda_{ijg}$. Here n_{ijg} is the number of persons at risk and $\log(\lambda_{ijg})$ denotes the linear predictor. Under the assumption of joint age effects, for example, differences of gender-specific period and birth cohort estimates are identifiable. The linear predictor is

$$\log(\lambda_{ijg}) = \mu_g + \theta_i + \varphi_{jg} + \psi_{kg} \tag{2.1}$$

where μ_g represents the gender-specific mean (intercept), θ_i the common age effect for age group i, φ_{jg} the effect of period j for sex g and ψ_{kg} the effect of the kth cohort for sex g. Note that the cohort index $k = 1, \dots, K$ depends on the age index i and period index j, but also on the width of age group and period intervals. We use the definition of Heuer (1997) which results in $K = 118$ birth cohorts. For identifiability of the gender-specific intercepts, we constrain all sets of time effects to sum to zero, i.e. here $\sum_{i=1}^{I} \theta_i = 0$, $\sum_{j=1}^{J} \varphi_{jg} = 0$ and $\sum_{k=1}^{K} \psi_{kg} = 0$ for both males ($g = 1$) and females ($g = 2$). If suitable, the linear predictor (2.1) can be modified to allow for separate period but common cohort effects or vice versa. Similarly, age and cohort effects may vary across gender but the period effects may be common, for example. However, keep in mind that differences are not identifiable if all sets of effects are allowed to vary (Riebler and Held, 2010).

Since we are in a Bayesian setting we treat all parameters as random and assign prior distributions. We follow Riebler and Held (2010) and use independent flat priors for the gender-specific intercepts. For the age, period and cohort effects we expect similarities between effects adjacent in time. Thus, we choose a Gaussian prior distribution based on independent second differences for all time effects, also known as random walk of second order (Besag *et al.*, 1995). This is a natural choice since second differences of time effects are identifiable (Clayton and Schifflers, 1987). Note that variance parameters are not gender-specific. Thus, there is one variance parameter for each time scale resulting in three variance parameters in total. The variances are treated as random and suitable prior distributions are assigned.

As is well known when working with registry data there are often inconsistencies in reporting systems or changes in reporting behaviour (Brillinger, 1986). For example the coding of causes of death might change over time. To adjust for such overdispersion, i.e. unobserved heterogeneity, we introduce further gender-specific variables z_{ijg} with mean zero and unknown variance into the linear predictor (2.1) (Besag *et al.*, 1995).

To validate and compare the "joint age effects" model (2.1) with other models we use the well-known deviance information criterion (DIC) (Spiegelhalter *et al.*, 2002). However, this criterion has recently been criticised for models with many random effects, e.g. (2.1), because complex models tend to be under-penalised (Plummer, 2008). For this reason we additionally calculate cross-validated proper scoring rules (Gneiting and Raftery, 2007), such as the mean Dawid-Sebastiani score (mean DSS), the mean ranked probability score (mean RPS) and the log score. Both DIC and proper scoring rules are negatively oriented such that smaller values are better. Having found the best model the question arises whether it is necessary to allow for correlation between gender-specific time effects and/or gender-specific overdispersion parameters. A model with correlated overdispersion parameters is similar in spirit to a seemingly unrelated regression models (Zellner, 1962). In this class of models, there are several regression equations which are assumed to be correlated via their error terms. Regarding the time effects (age, period and cohort effects) the inclusion of a correlation would actually represent a balance between separate and joint effects. More precise relative risk estimates may be obtained. For comparing models with and without correlation structure we additionally use the log marginal likelihood. Although improper priors are used, the use of the marginal likelihood is valid because the candidate models only differ by the inclusion of correlation between the priors, for more details see Riebler *et al.* (2010).

All models are estimated using both Markov chain Monte Carlo (MCMC) techniques as described in Riebler and Held (2010) and Riebler *et al.* (2010), and integrated nested Laplace approximations (INLA) (Rue *et al.*, 2009). INLA represents a deterministic alternative to MCMC. It computes directly very accurate approximations to the posterior marginal distributions and thus avoids time-consuming sampling. The INLA-program is freely available under `www.r-inla.org` and runs on Windows, Mac and Linux. Here we use the INLA version built on 16.05.2010. DIC can be calculated within both settings. The log score and marginal likelihood are calculated using INLA, in contrast mean DSS and the mean RPS are calculated with MCMC. For comparing correlated multivariate APC models the multivariate analogues of the mean RPS and DSS score are used to capture the potential correlation present between the male and female suicide rates (Riebler *et al.*, 2010).

For the MCMC analyses without inclusion of correlation we use a run of 120 000 iterations, discarding the first 20000 iterations and storing every 20th sample thereafter, leading to a total of 5000 samples. When including correlation we use a run of 520 000 iterations omitting the first 20000 iterations and storing every 200th sample thereafter, resulting in 2500 samples. This is necessary because of the high autocorrelations when accounting for correlations between time-effects and/or overdispersion parameters.

Explanatory variables on family integration are available for all observation years (1950–2007) and obtained for the 1970 to 2007 period from The Swiss Federal Statistical Office (2009) and for the period before 1970 from Calot (1998). Since social integration is difficult to measure, there are several proposals and it is debatable which one to choose as an indicator. Breault and Barkey (1982) proposed to measure family integration by the marriage rates minus divorce rates divided by marriage rates plus divorce rates, since married people are supposed to be better integrated than single persons. A high value is related to a better integration. We will call this indicator F-index throughout the paper. The F-index is preferable to a crude measure, such as divorce rates alone as there is a high correlation between divorce rates and marriage rates. Alternatively the number of divorces per 100 marriages could be used.

To calculate the F-index we first adjusted the divorce rate for the introduction of a new divorce law in 2000 in Switzerland. The new law resulted in an extreme increase of divorces in 1999 and a strong decrease in 2000. First there were more divorce proceedings terminated in 1999 to have more time to adjust to the new legal situation in 2000. Then the introduction of the new divorce law in 2000 caused a prolongation of proceedings and thus less decisions. Thus, we substituted the divorce rates in the years 1999 and 2000 by the average of this two years. Alternatively a moving average of first or second order could be applied to the whole time-series to smooth random variations over the whole period.

As an alternative measure for family integration we consider the total marriage rate. This index represents the mean percentage of unmarried persons aged below fifty who would marry in the course of time, if they showed the same age-specific marriage behaviour as in the observation year.

For analysing the association of family integration and suicide we start with the calculation of pairwise correlations and perform standard time-series analyses, as proposed in Stack (1989, 1990a). Then, we use INLA to estimate separate univariate APC models for males and females substituting the period effect block by a time-constant regression variable related to the rate of unemployment and an explanatory variable on family integration assuming either a parametric or non-parametric effect. Finally, we use the multivariate APC model classified as the best model without covariates and replace the period effects by a non-parametric covariate effect of the F-index assuming a correlation between the effect on males and females. Figure 2 shows the F-index, the total marriage rates for males and females, and the unemployment rate from 1950–2007. Both the F-index and also the total marriage rates strongly decrease from 1950 to 2007. Especially around the middle of the 1970s there is a big drop. Around the early 1990s a local maximum is visible while afterwards all markers are decreasing again. The unemployment rates show the typical economical patterns indicated within the figure. However, note that the rates are always below five per cent.

3 Results

We will first present the results of the multivariate APC analysis for detecting trends in the sex ratio in suicide mortality of males and females. Then we will explore whether the inclusion of explanatory variables related to family integration can explain changes in suicide rates.

3.1 Analysis of gender-specific differences

The model diagnostics for all models obtained by MCMC and INLA are shown in Table 1. The "joint period effects" model is classified by all model choice criteria as the best model, so that we assume gender-specific age and cohort effects. Including a correlation for overdispersion is clearly preferred to the model without correlation. The model with correlated overdispersion and the model with both correlated gender-specific effects and overdispersion are very similarly

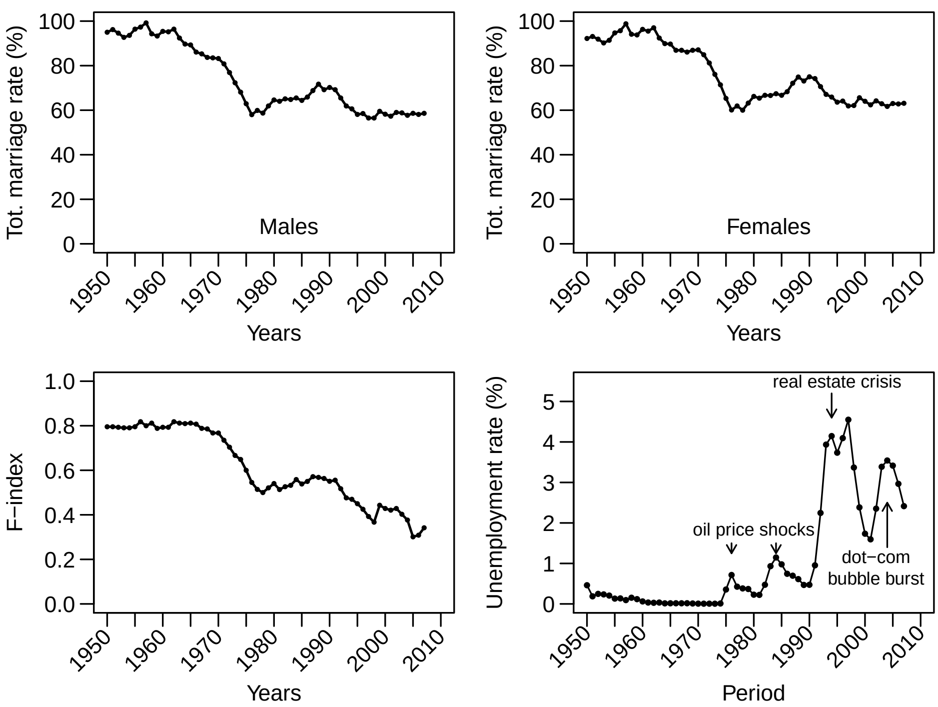

Figure 2: Covariate information from 1950–2007. Top (left to right): Total marriage rate for males and females given in per cent. Bottom (left to right): F-index measured as (marriage rate - divorce rate)/(marriage + divorce) and unemployment rate measured as unemployed/(unemployed + employee).

classified. The log-marginal likelihood slightly prefers the model with a correlation for both time scales and overdispersion, see Table 2. In this model, the correlation between overdispersion parameters was estimated as 0.56 with 95% credible interval $(0.18, 0.84)$. For the gender-specific age-effects the correlation was estimated as 0.82 $(0.50, 0.95)$ and for the cohort effects to 0.20 $(-0.61, 0.79)$. Thus, except for the cohort effects, the posterior distributions of all correlation estimates are clearly greater than zero, indicating that partly the same risk factors act on age and overdispersion for males and females.

Figure 3 shows the relative risks of suicide for males compared to females for the "joint period effects" model with both correlated age and cohort effects, and correlated overdispersion. Men have an about three times higher risk to commit suicide than women. Especially men between 16 and 25 have a high risk compared to the corresponding age group of females. While the risk

Table 1: Model choice criteria obtained from MCMC and INLA. For both approaches DIC estimates are given. In addition the mean Dawid-Sebastiani score $\overline{\text{DSS}}$ and the mean ranked probability score $\overline{\text{RPS}}$ are shown for MCMC and the log score for INLA. The column names indicate which effects (**A**ge, **P**eriod, **C**ohort) are assumed to be the same for males and females. The remaining effects are assumed to be gender-specific. The best value for each criterion is indicated in bold.

	Joint effects used for						
	A,P,C	P,C	A,C	A,P	C	P	A
MCMC model diagnostics							
Mean DSS	5.10	4.99	5.06	4.96	4.95	**4.88**	4.92
Mean RSS	4.62	4.37	4.53	4.34	4.28	**4.15**	4.21
DIC	2064.87	1971.86	2037.82	1942.87	1936.84	**1864.43**	1899.70
INLA model diagnostics							
Log Score	3.48	3.44	3.47	3.43	3.42	**3.39**	3.40
DIC	2064.68	1971.39	2037.47	1942.88	1936.74	**1864.16**	1899.29

Table 2: Model choice criteria obtained from MCMC and INLA for the "joint period effects" model without correlation, correlation between overdispersion parameters, between gender-specific age and cohort effects, between both overdispersion and gender-specific age and cohort effects. The best value for each criterion is indicated in bold.

	Correlation included for			
	-	overdispersion	time effects	both
MCMC model diagnostics				
Mean multivariate DSS	9.758	**9.752**	9.763	9.756
Mean multivariate RPS	6.653	**6.636**	6.647	6.640
DIC	1864.372	1854.804	1864.270	**1854.788**
INLA model diagnostics				
Log marginal likelihood	-5215.39	-5214.86	-5215.78	**−5214.72**
Log Score	3.390	**3.386**	3.390	3.386
DIC	1864.16	**1854.56**	1864.07	1854.60

for males between 30 and 74 is almost three times as high as in women it is increasing again for elderly men.

For cohorts born around 1870 the risk for males is about four times as high compared to females.

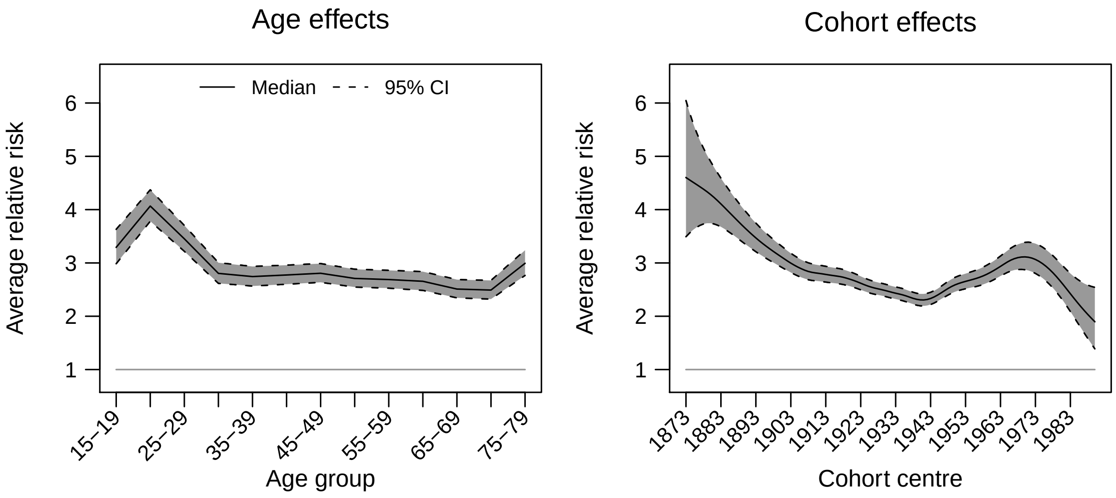

Figure 3: Relative risk of suicide for males compared to females for the "joint period effects" model assuming a correlation between males and females for the age and cohort effects, and also the overdispersion parameters.

However, with successive cohorts the average relative risk falls to being about twice as high for cohorts born around 1940. Then it increases again reaching a local maximum for cohorts born around 1970. For the youngest cohorts a decreasing trend is visible.

Posterior marginal distributions of MCMC and INLA were virtually identical for all variance parameters. For the gender-specific intercepts we found a small shift between MCMC and INLA which is probably related to the fact that the time effects (especially for the cohorts) of INLA do not exactly fulfil the sum-to-zero constraints. This may be solved by changing one of the default INLA specifications. Inspecting the identifiable linear predictor MCMC and INLA coincide perfectly.

3.2 Analysis of the impact of family integration

Following the Durkheim tradition, one may ask whether explanatory variables on social integration can explain changes in suicide rates. The usual approach to study this is to look at the age-standardised male and female suicide rates. The matrix of pairwise correlations is shown

Table 3: Pair-wise Spearman correlations.

	Y1	Y2	X1	X2	X3
Male suicide rate (Y1)					
Female suicide rate (Y2)	0.86				
F-index (X1)	0.40	0.29			
Total male marriage rate (X2)	0.39	0.26	0.94		
Total female marriage rate (X3)	0.27	0.14	0.92	0.98	
Unemployment (X4)	-0.25	-0.27	-0.81	-0.75	-0.70

in Table 3. Note, that neither the male suicide rates nor the female suicide rates in Switzerland are negatively correlated to the F-index. The total gender-specific marriage rates are positively correlated to both rates. Also the correlation between the unemployment rate and suicide mortality has the contrary sign of what we expected and indicates an inverse effect. Note that some of the explanatory variables are highly correlated, for example unemployment and the F-index have a correlation of -0.81 ($p < 0.05$). However, since correlations alone might not be meaningful we continued fitting a linear model using ordinary least squares. We used separate models for male and female mortality rates including three regressors, namely an intercept, the rate of unemployment and one covariate on family integration. To detect autocorrelation of first order between two successive residuals we used the Durbin-Watson test (Harvey, 1990). For each analysis the Durbin-Watson statistic was smaller than the lower bound of 1.50 for 58 observations and three regressors, indicating positive autocorrelation. Hence, we included a one-year lagged dependent variable as a regressor to eliminate autocorrelation. In autoregressive models the Durbin-Watson test statistics tends to underestimate autocorrelation, so that we used Durbin's h statistic (Harvey, 1990). Durbin's h is a normally distributed variable, so that the h-statistics should be within -1.96 and 1.96, which was the case for all models.

Table 4 shows the regression estimates for all models. The first two columns show the results for the models including the F-index as marker for family integration. In contrast the second two columns show the results when including total gender-specific marriage rates. The lagged suicide rate is positively related to the dependent variable for both gender in all regressions.

Table 4: The effect of family integration on suicide for men and women in Switzerland from 1950 to 2007. Standard errors are given in parentheses.

	Suicide			
	Males	Females	Males	Females
Suicide rate lag	$0.880_{(0.061)}$	$0.769_{(0.078)}$	$0.875_{(0.061)}$	$0.719_{(0.079)}$
F-index	$-2.280_{(2.744)}$	$-2.518_{(1.502)}$	-	-
Tot. marriage rate (males)	-	-	$-0.037_{(0.025)}$	-
Tot. marriage rate (females)	-	-	-	$-0.041_{(0.015)}$
Unemployment	$-0.665_{(0.351)}$	$-0.461_{(0.199)}$	$-0.728_{(0.301)}$	$-0.505_{(0.164)}$
Constant	$6.065_{(3.080)}$	$4.972_{(1.787)}$	$7.659_{(3.192)}$	$7.231_{(2.048)}$
Adjusted R-squared	0.858	0.771	0.862	0.786
Durbin's h statistic	-1.49	-1.74	-1.73	-1.94

Inspecting the results of the models including the F-index we see that neither the male nor the female suicide rate is related to the F-index. Unemployment is significantly negatively related to the female suicide rate.

Turning to the regressions that include the total gender-specific marriage rate the unemployment rate is again significantly negatively related to male and female suicide rates. For males the coefficient of the total male marriage rate is -1.48 $(-0.037/0.025)$ times its standard error. For females the coefficient of the corresponding female rate is -2.73 $(-0.041/0.015)$ times its standard error. Hence for higher values of the total marriage rate the female suicide mortality decreases.

To be able to keep the age-specific structure of the data, so that all information can be used, we integrate the explanatory variables into the APC model and estimate the models using INLA. In the following we assume a time-constant effect of the rate of unemployment. Thus, we replace the period parameters with one linear effect of the unemployment rate and either a linear, quadratic or non-parametric effect of the F-index or the total gender-specific marriage rates. Modelling the covariate in a non-parametric fashion we assume a random walk of second order as non-parametric prior for the covariate effect. This is a natural prior as it models deviations from

Table 5: Log score and DIC of univariate APC models including either no covariate, or a time-constant effect of unemployment rate and either a time-constant (linear), quadratic or non-parametric effect of an explanatory variable on family integration.

	Females		Males	
	Log score	DIC	Log score	DIC
No covariate	**3.090**	**849.23**	**3.677**	**1001.25**
Time-constant linear effect				
F-index	3.158	930.03	3.849	1135.35
Tot. marriage rate	3.151	922.39	3.814	1113.61
Time-constant quadratic effect				
F-index	3.116	881.53	3.721	1042.10
Tot. marriage rate	3.152	923.71	3.811	1113.00
Random walk of second order on covariate				
F-index	3.107	870.07	3.697	1021.06
Tot. marriage rate	3.119	885.49	3.730	1049.77

a linear trend but reduces to the linear model if its variance goes to zero (Natario and Knorr-Held, 2003). In Natario and Knorr-Held (2003) an inverse gamma distribution with shape equal to 1 and scale equal to 0.00005 is proposed as prior for non-equally spaced covariates with an average distance equal to one. We used this prior and scaled each covariate on family integration appropriately. In addition, a sum-to-zero constraint is applied on the covariate effects to ensure identifiability of the intercept.

Table 5 shows the model choice criteria for all models. For both sexes the model without covariates is clearly classified as the best model, which indicates that the covariates we proposed cannot fully replace the period effects. However, note that assuming a non-parametric effect of the F-index is much better than the corresponding parametric formulations assuming a linear or quadratic effect. In addition, this model is also not so far away from the standard APC model for both sexes. Table 6 shows the parameter estimates of all models with parametric covariate effects. The unemployment rate has a slight negative effect on the log female suicide rate in all models and on the log male suicide rate in models with a linear effect of F-index or total marriage

Table 6: Parameter estimates (posterior median, 2.5% and 97.5% quantile) for models with parametric covariate effects.

	Females			Males		
	2.5% qu.	Median	97.5% qu.	2.5% qu.	Median	97.5% qu.
Time-constant F-index						
(Intercept)	-8.05	-7.85	-7.66	-7.27	-7.11	-6.95
F-index	-1.96	-1.65	-1.34	-1.42	-1.17	-0.92
Unemployment	-0.09	-0.07	-0.06	-0.07	-0.05	-0.04
Time-constant total marriage rate						
(Intercept)	-8.03	-7.85	-7.66	-7.10	-6.95	-6.80
Tot. marriage rate	-0.02	-0.01	-0.01	-0.01	-0.01	-0.01
Unemployment	-0.06	-0.04	-0.02	-0.04	-0.03	-0.02
Time-constant quadratic F-index						
(Intercept)	-9.52	-9.15	-8.78	-9.11	-8.84	-8.57
F-index	1.58	2.69	3.79	3.82	4.65	5.47
F-index^2	-4.31	-3.45	-2.61	-5.27	-4.63	-4.00
Unemployment	-0.06	-0.04	-0.02	-0.02	-0.01	0.01
Time-constant quadratic total marriage rate						
(Intercept)	-9.23	-8.18	-7.13	-8.46	-7.78	-7.10
Tot. marriage rate	-0.03	-0.00	0.02	-0.01	0.01	0.03
Tot. marriage rate^2	-0.00	-0.00	0.00	-0.00	-0.00	-0.00
Unemployment	-0.06	-0.04	-0.02	-0.03	-0.02	-0.00

rate. The F-index has a negative linear effect for both males and females. Thus, increasing the F-index by 0.1 units reduces the female suicide risk by 15% ($= 100 \cdot (1 - \exp(-0.165))$). Modelling the F-index in a quadratic way $\beta_0 \cdot$F-index $+ \beta_1 \cdot$F-index^2, the estimate of β_0 is positive while the estimate of β_1 is negative for both males and females. Figure 4 shows the parametric and non-parametric estimates for the F-index and the total male and female marriage rate. Assuming a quadratic and non-parametric effect for the F-index results in very similar estimates. In contrast for the total marriage rate the quadratic and linear effect are very similar. Regarding the model choice criteria presented in Table 5 the inclusion of the F-index is preferred compared to the inclusion of the total marriage rate.

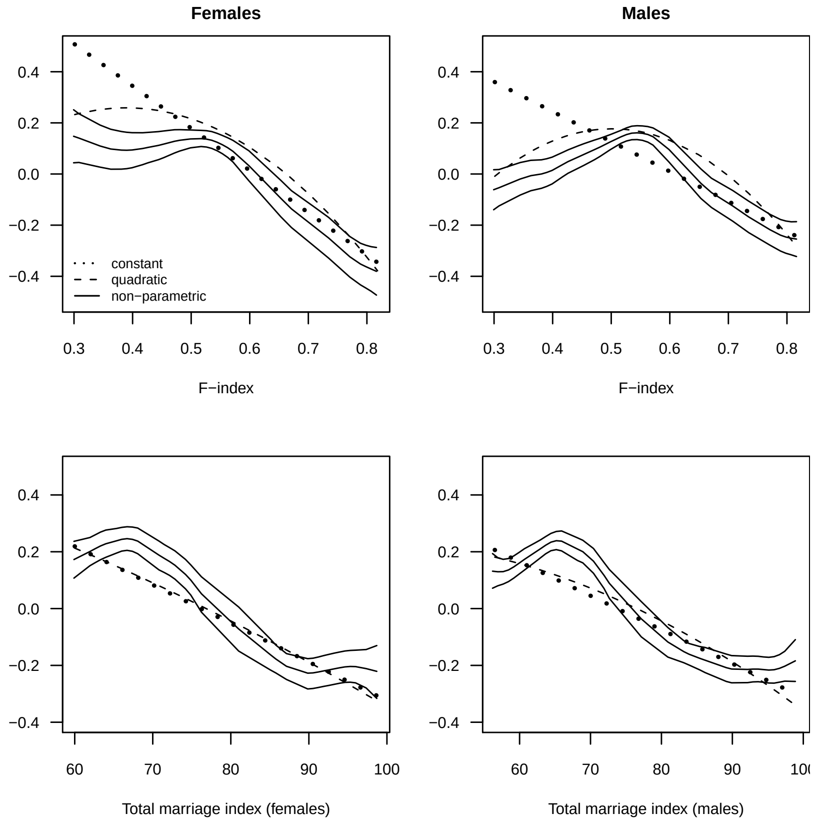

Figure 4: Estimated parametric (linear and quadratic) and non-parametric effects of covariates related to family integration. For the non-parametric trend 95% pointwise credible bands are shown.

Figure 4 shows that the effect of the F-index is very similar for males and females which suggests a joint analysis with correlated non-parametric F-index effects. We use the best-classified correlated multivariate APC model of Section 3.1, namely the "joint period effects" model with correlated age and cohort effects and correlated overdispersion parameters, and replace the period effects by a non-parametric effect of the F-index which is assumed to be correlated between males and females. Figure 5 shows the resulting age effects, cohort effects and covariate effects. The estimated time variables exhibit a similar pattern for males and females. Table 7 shows that the correlation estimates are, except for the cohort effects, very high and significantly different from zero. Thus, similar risk factors may act on age and overdispersion. In addition,

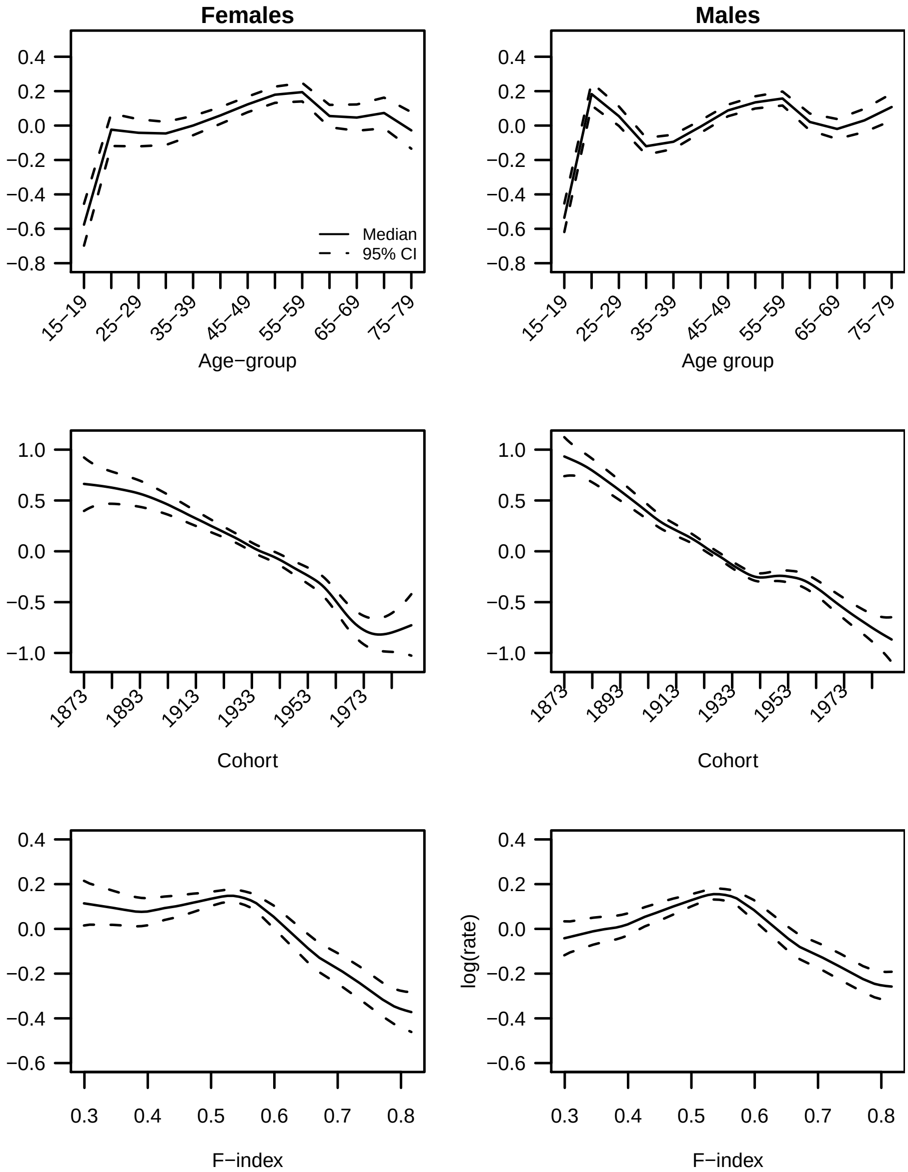

Figure 5: Estimated age effects, cohort effects and non-parametric effects of the F-index obtained from a multivariate AC model assuming correlation between each pair of gender-specific effects: age effects, covariate effects, cohort effects and overdispersion.

the F-index has a similar effect on males and females indicated by the high estimated correlation. Figure 5 shows that the age effects strongly increase from the youngest ages to the 20–24 years old persons. Then, for females, the effects slightly increase until the age of around 60 when

Table 7: Model choice criteria and correlation estimates obtained from INLA for the AC model with correlated **A**ge and **C**ohort effects, correlated **O**verdispersion and a correlated non-parametric effect of the F-index (third column). For comparison the corresponding model without correlation (second column) and the correlated model without covariate but joint period effects (first column) is given. The triple notation $_LP_U$ for the correlation parameters denotes the posterior median P with 2.5 per cent quantile L and 97.5 per cent quantile U.

		Period effects replaced by F-index	
	A,C,O	no correlation	A,F-index,C,O
Model choice			
Log Score	**3.386**	3.402	3.393
DIC	**1854.60**	1890.05	1870.26
Correlation coefficients			
Correlation age	$_{0.50}0.82_{0.95}$	-	$_{0.50}0.82_{0.95}$
Correlation F-index	-	-	$_{0.50}0.94_{1.00}$
Correlation cohort	$_{-0.61}{+0.20}_{+0.79}$	-	$_{-0.61}{+0.23}_{+0.81}$
Correlation overdispersion	$_{0.18}0.56_{0.84}$	-	$_{0.26}0.63_{0.85}$

they start to decrease again. For males the age effects clearly drop from the 20–24 age group to the 25–29 years old. Similar to the females the effects slightly increase until the age of 60, then start decreasing. For the oldest ones the age effects increase again. The cohort effects show a negative slope for both sexes from the oldest to the youngest cohorts. In contrast to males, the cohort effects in women stop decreasing for those born at the early 1970s. Finally, the effect of the F-index on females stays almost constant for values between 0.3 and 0.55. For larger values, representing a higher degree of social integration, the effect decreases. The effect on males increases slightly from a value of 0.3 to a maximum of about 0.55 and decreases for higher values as well. The joint time-constant effect of the unemployment rate is estimated as -0.00 (95% CI: $-0.02, 0.01$). Thus the effect of the unemployment rate is not significantly different from zero.

Figure 6 shows the estimated relative risks of suicide of males compared to females. The estimates are very similar to those obtained in Figure 3, but slightly lower and with larger credible bands.

In Table 7 the log score and DIC estimate compared to the corresponding model without assuming correlation for any of the effects and compared to the model regarded best in Section 3.1 are given. The correlated version is clearly preferred. However, the standard correlated multivariate "joint period effects" model is classified as the best model.

4 Discussion

The results of the present multivariate age-period-cohort analysis confirm previous findings of strong gender-specific differences in suicide rates. Nevertheless the correlation estimates between gender-specific effects were, except for the cohort effects, very high. So it seems that similar risk factors act especially on age and overdispersion, which results in similar effect curves but on different overall levels. For all years and all age groups men have, compared to women, a three-fold risk to commit suicide. Elderly men and those between 16 and 25 show a especially high relative risk compared to their female peers in the same age group. This is quite surprising considering the higher number of suicide attempts of females compared to males. An explanation might be that males use more lethal methods like hanging or firearms, while the most frequent method of females is poisoning (Ajdacic-Gross *et al.*, 2008).

We further found a pronounced elevation for males born around 1970. Ajdacic-Gross *et al.* (2006) found in univariate APC models an inflexion point in cohort effects around 1970 for males as well, but not for females. Explaining this observed pattern is difficult. There should be risk factors related to cohorts born around 1970, but not to those born before or after 1970. Since the effect seems to be present especially in males and not females, also reflected in an insignificant correlation estimate, we suppose that the causal factor must occur in later life. This is plausible in the context of suicide. Relevant risk factors for suicide are complex, so a better understanding is necessary to explain the observed pattern. Stockard and O'Brien (2002) suggested that cohorts experiencing less social integration and having a large relative cohort size have higher suicide rates. However, in Switzerland the persons around 1964 are the largest cohort and therefore do not explain the peak for cohorts born after 1970.

We further explored whether family integration could explain gender-specific differences in suicide rates. We started with standard time-series analysis based on age-standardised rates. A significant influence was only found for the total female marriage rate on female suicide rate, indicating that with a higher total female marriage rate suicide rates of women decrease. Exploiting the age-specific structure of the data we applied univariate age-period cohort models and replaced the period effects by parametric and non-parametric effects of variables related to family integration. To adjust for confounding induced by the rate of unemployment we additionally included a linear effect of the unemployment rate. The inclusion of the so-called F-index measured as (marriage rate - divorce rate)/ (marriage rate + divorce rate) was preferred compared to the gender-specific total marriage rate. Higher values, corresponding to better integration, have a decreasing effect on suicide risk. Of note, the effect of the F-index on males and females is very similar. Thus, we used a multivariate APC model and replaced the period effects by a correlated non-parametric effect of the F-index. The estimated correlation was very high. A difference we found for women and men was, that median values of the F-index, interpretable as normally integrated, have a higher effect on male suicide than being worse integrated. As the computation of the F-index is only based on marriage and divorce rates, this marker might be not suitable to measure bad integration. Many males might be not married, but nevertheless well integrated and happy.

By means of model choice criteria the standard APC model without covariates was preferred. This indicates that the measures of social integration we used cannot fully replace the period effects. If we had analysed only data from 1950 to 1980, we probably would have found a strong linear dependence between the F-index and the suicide rates, because from 1950 to 1970 the F-index value was almost constantly at 0.8. This indicates a high degree of social integration. In the same time-period suicide rates were strongly decreasing for both sexes. From 1970 to 1980 the F-index dropped and almost inversely the suicide rates increased. However after 1980 this linear relationship has vanished.

A problem of the present analysis is that it is very difficult to decide how to measure social or family integration. There are several proposals and all of them are controversially discussed. In addition, it is not completely clear how to properly include information on explanatory variables in the APC model (Knorr-Held and Rainer, 2001). Since social integration might not be the main risk factor for committing suicide, it could be necessary to adjust for further confounding variables, for example mental disease or religiousness. Note that it is also possible to include more covariates in a non-parametric fashion. In our analysis the two markers on social integration are strongly correlated, so that in this cases it is not meaningful to include both. Alternatively, the period effects could be kept in the model, but then the identifiability problem well known for APC models remains.

It could also be that the covariates on social integration exert a lagged influence on suicide mortality. For example, Wasserman (1984) determined a lag of nine months of the influence of divorce on suicide reported in the United States. Including a lagged covariate into a Bayesian APC model is particularly attractive, because then projections of suicide rates can be generated for future periods without any parametric assumptions. More research would be necessary to explore whether such a lagged effect is also present for Swiss suicide data. However, it is difficult to determine an exact time-lag because the separation of couples starts much earlier than the divorce proceeding is completed. Hence, the exact time-point which is relevant for suicide behaviour is difficult to determine.

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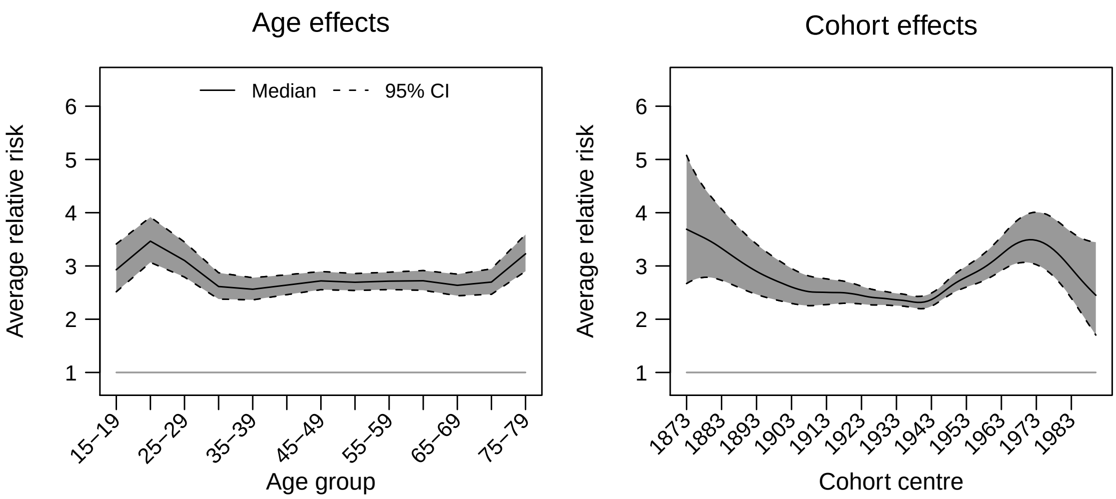

Figure 6: Estimated relative risk of suicide for males compared to females obtained from a multivariate model assuming correlation between each pair of gender-specific effects: age effects, covariate effects, cohort effects and overdispersion.

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