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Evidence for the “Suicide by Firearm” Proxy for Gun Ownership from Austria

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Evidence for the "Suicide by Firearm" Proxy for Gun Ownership from Austria[a]

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Abstract

When attempting to measure gun ownership in the United States, the problem of missing administrative data arises, making it necessary to find a valid proxy. Several such proxies are employed in economic studies, one of which is the fraction of "suicides by firearm" of "all suicides" (*FSS*). My work validates this proxy from out-of-sample data, namely, Austrian administrative data on firearm licenses. I also reevaluate, with appropriate statistical methods, a result on firearms and suicide from the medical that is often used for public policy advocacy. This result is, unfortunately, heavily biased due to ignoring a well-known fallacy and thus can be only partially confirmed.

JEL Classifications: C15; C51; I18
Keywords: Gun Ownership; Suicide; Ratio Fallacy; Spurious Correlation

[a]Thanks to Matthew Lang for sharing his data on suicide and firearms in the United States with me, which allowed me to perfectly replicate his findings.

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1 Introduction

The economic literature contains ample investigation into the relation between guns and crime. Seeing that in the United States there were 11,078 deadly assaults by firearm and 19,392 suicides by firearm in 2010,[1] a closer investigation of a possible association between firearms and suicide seems warranted. Two studies of Ireland (Kennelly, 2007; Yang and Lester, 2007) – with remarkably different outcomes – elaborate on the economic dimension of suicide in terms of cost. Furthermore, firearm suicide as a fraction of all suicides is believed to be a good proxy, at least in the cross-section, for gun ownership density (Azrael, Cook and Miller, 2004; Kleck, 2004). This association is exploited by Cook and Ludwig (2006) in a very detailed, albeit flawed (Westphal, 2013), analysis of the association between firearms and crime. Lang (2013) analyzes the association between firearms and suicide using U.S. National Instant Background Check data and confirms the validity of the FSS proxy.It seems valuable to investigate the interaction between firearms and suicide with high quality data from other than a U.S. sample, taking care to avoid methodological fallacies.

After World War II, European countries enacted much tighter firearms regulation than what exists in the United States. Therefore, much better administrative data are available. Austria has relatively low restrictions on the acquisition of firearms but has become increasingly concerned with monitoring legally purchased firearms. Austrian data on concealed carry licenses are available from 1982 to the present for all Austrian counties. This provides a reasonable, albeit imperfect, nationwide proxy for gun ownership taken directly from administrative data on firearm permits. These data have been used to compute correlations between firearm ownership rates and suicide rates in the medical literature (Etzersdorfer, Kapusta and Sonneck, 2006), and provide an intriguing starting point for possibly confirming, or not, the validity of the FSS proxy and at the same time further investigating the relationship between suicide and firearms.

Two questions are addressed in this paper: (1) Can the FSS proxy for gun ownership be confirmed from Austrian data on gun licenses? – and (2) What can be said about the relationship between firearms and suicide in Austria after a careful review of the methods used for analysis in former work? Answering these questions results in two main findings. First, I confirm the validity of the FSS proxy. An association between firearms and firearm suicides is persistent across all methods of analysis used and a variety of model specifications. If one prefers clustered standard errors over Driscoll-Kraay standard errors – a preference I do not advocate in my setting – a substitution between suicide methods

[1] ICD-10 codes X93, X94, and X95 used for "assault by firearm"; X72, X73, and X74 for "suicide by firearm." Values taken from United States Department of Health and Human Services (2010).

shows in the main model. Second, it is clear that, earlier correlation results in Etzersdorfer, Kapusta and Sonneck (2006) on the association between firearms and suicides are greatly overstated due to ignoring Pearson's (1896) finding on spurious correlations between ratio variables. Thus, the contributions of this paper include validation of earlier approaches to measuring gun ownership,[2] and a warning as to the hazards of using spurious results in public policy debate.

My paper is organised as follows. I revisit the literature on guns and suicide in Austria in Section 2. In Section 3.1, the results from Etzersdorfer, Kapusta and Sonneck (2006) are repeated. Sections 3.2 and 3.3 point out the statistical fallacy in Etzersdorfer, Kapusta and Sonneck (2006) and adjust for the problem using two approaches that both lead to numerically very close and qualitatively identical results. Section 4 motivates and estimates a fixed effects panel model based on a theoretical model from the economic literature. The main finding for the FSS proxy is found to be robust to several robustness checks in Section 4.2.

2 Former Analysis of Firearms and Suicides in Autria

Etzersdorfer, Kapusta and Sonneck (2006) (EKS hereafter) analyze correlations between *suicide rates* and *rates of firearm ownership*, proxied by the *rate of concealed carry licenses*, in all nine Austrian counties over the period from 1990 to 2000. Their results from a repeated cross-sectional analysis are *strong rank correlations between the firearms measure and firearm suicides, low-to-no rank correlations between firearms and other suicides, and weakly positive rank correlations between firearms and all suicides*. Based on these findings, their conclusion is to assume that overall suicides increase with more firearms, as depicted in Figure 1(a), as opposed to a substitution between suicide methods as shown in Figure 1(b). In EKS's (p. 468) opinion their findings "emphasise the need for political support" for stricter regulation on gun ownership in the interest of preventing suicide. Their finding is now propagated through the literature; for example, "it is a scientific fact ... that reducing the availability of guns ... will reduce deaths" (Leenaars, 2006, 439). There many references to similar studies[3] can be found.

[2]This measure is not without problems itself as can be seen in Westphal (2013).

[3]Notably similar to EKS of those mentioned are Markush and Bartolucci (1984), Killias (1993) and Leenaars et al. (2003).

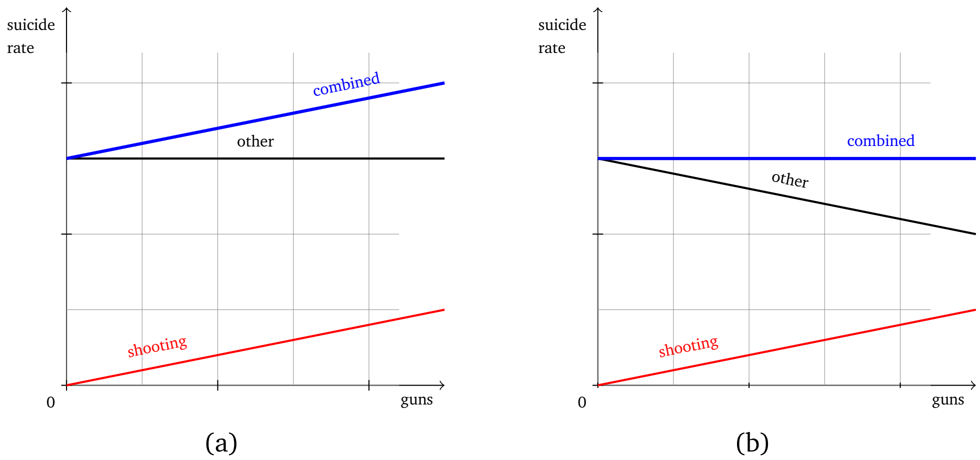

Figure 1: Competing models

3 Revisiting EKS's Results

3.1 Replication

In a first step, I replicate the results from EKS. EKS use four variables: *population size* (pop), *gun carry licenses*[4] (CCL), *suicides with firearms* ($E955$), and *all suicides* ($E95$). The latter two are based on their ICD-9[5] codes of the same name. The number of carry permits is used as a proxy for the number of gun owners. I obtained the data from their primary sources. Data on carry permits were obtained from the Austrian Interior Ministry, Department III/3. Statistik Austria provided population and suicide figures. Data for all variables were provided for the years $t = \{1982, 1985, 1987, 1990, 1992, 1994, 1995, \ldots, 2011\}$[6] and all $K = 9$ Austrian counties $k = 1, 2, \ldots, K$. I.e. $pop_{k,t}$ is population size in year t in county k. An overview of the variables is given in Table 1. These also are the variable names used in my program code and the data made available with this paper. Descriptives are set out in EKS (p. 464–465).

EKS compute rank correlations between *gun ownership rates* and *suicide rates*. Table 2 row I sets out my results for Spearman's rank correlation coefficient as

[4]"Waffenpaesse." CCL is the U.S. acronym for "concealed carry license." A CCL and a "Waffenpass" are not legally exactly identical, but are very similar and so I therefore use the acronym CCL based on its international recognition value.

[5]E955 includes suicides by explosives which cannot be distinguished from firearm suicides. However, for later years ICD-10 codes are available, which do differentiate between firearm and explosives suicides. The numbers indicate that there are very few suicides by explosives.

[6]Upon request, I could not be supplied with data for 1991 and 1993. Remarkable, EKS state results for these years.

Underlying	Absolute	Meaning	Rate	How computed?
Firearms	CCL	Number of carry permits	$CCLR$	CP/pop
Firearm suicides	$E955$	Number of fire-arm suicides	FSR	$E955/pop$
All suicides	$E95$	Number of *all* suicides	SR	$E95/pop$
Suicides *not* with firearms	$NE955$	Number of suicides *not* with firearms	OSR	$NE955/pop$
Population	pop	Number of inhabitants	NA	NA

Table 1: Variables and their meaning

well as Pearson's correlation coefficient, averaged over all years t.[7] Detailed result for individual years can be found in Table 4 in the Appendix. Neither small sample size nor non-normality of data, as claimed by EKS (p. 465), contradict the computation of Pearson's correlation coefficient. I therefore included these values in Table 2: values do not deviate much from the rank correlations. Tables 2 and 4 reveal the numerical and qualitative results from EKS are robust to the inclusion of years prior to and after their original period as well as to using either Pearson's or Spearman's method.

3.2 Accounting for Spurious Correlation between Ratios

Unfortunately, EKS fail to acknowledge Pearson's finding on correlations between ratios (Pearson, 1896): *using ratios for correlation analysis may lead to spurious results.*

Table 2, Row I shows there is little difference between rank correlations and Pearson's correlation for the data. Because of this and because of the availability of a theoretical result from Kim (1999), I now first use Pearson's correlation coefficient for illustration and examination of the ratio fallacy problem in EKS's results. A simulation study conducted in Section 3.3 shows that my findings do not change when using rank correlations.

Let there be three independent random variables X, Y, Z with known expected values and variances. To illustrate the problem at hand, let $X_{k,t}$ be the number of CCLs in county k in year t. $Y_{k,t}$ represents the corresponding number of suicides,

[7] As opposed to the *rank correlation of the average rates over time* (what is the meaning of that value aside from it being larger than the individual correlations in this setting?) as in EKS (Table 2).

Row	Firearm suicides Pearson	Firearm suicides Spearman	Other suicides Pearson	Other suicides Spearman	All suicides Pearson	All suicides Spearman
I	*Correlations between ratios...*					
	0.647	0.640	−0.053	0.044	0.157	0.222
IIa	*... rescaled by estimated reference points*					
	0.459	NA	−0.317	NA	−0.152	NA
IIb	*Reference points estimated by Equation* (1)					
	0.357	NA	0.397	NA	0.394	NA
IIIa	*... rescaled by simulated reference points*					
	0.433	0.444	−0.302	−0.190	−0.175	−0.089
IIIb	*Simulated reference points*					
	0.390	0.364	0.370	0.309	0.429	0.393

Table 2: Correlations between gun ownership rate and suicide rates and rescaling points. Both averaged over time

firearm suicides, or non-firearm suicides. $Z_{k,t}$ is the county's population in that year. In this setting, the coefficient of correlation $r_{X/Z,Y/Z}$ between X/Z and Y/Z in year t will not usually be zero, even if all three variables were truly uncorrelated. This *spurious correlation* is driven by the identical denominator common to both ratios.[8] The theoretical reference point for no correlation in this case is given in Kim (1999, Eq. (2.2)) as

$$r^0_{X/Z,Y/Z} = \frac{V^2_{1/Z}}{\sqrt{[V^2_X(1+V^2_{1/Z})+V^2_{1/Z}][V^2_Y(1+V^2_{1/Z})+V^2_{1/Z}]}} \tag{1}$$

for positive expected values of X, Y and when V_A is the coefficient of variation, i.e., $V_A = \sqrt{\mathscr{V}(A)}/\mathscr{E}(A)$. Positive expectation and finite variance is clearly fulfilled for population, suicides, and CCLs for all t. Therefore the empirical moments of population, carry permits and suicides will be used in Equation (1) to estimate the reference points for each year, as suggested by Kim (1999, 386). Using these yearly estimates to rescale the correlations based on Equations (2) and (3) gives us the rescaled correlations shown in Table 2, Row IIa with the estimated rescaling points in Row IIb. Detailed results for individual years are given in Table 5 in

[8] See Pearson (1896); Kronmal (1993); and Kim (1999).

the Appendix.

$$r^*_{X/Z,Y/Z} = \frac{r_{X/Z,Y/Z} - \hat{r}^0_{X/Z,Y/Z}}{1 + \hat{r}^0_{X/Z,Y/Z}} \quad \forall r_{X/Z,Y/Z} \leq r^0_{X/Z,Y/Z} \tag{2}$$

$$r^*_{X/Z,Y/Z} = \frac{r_{X/Z,Y/Z} - \hat{r}^0_{X/Z,Y/Z}}{1 - \hat{r}^0_{X/Z,Y/Z}} \quad \forall r_{X/Z,Y/Z} > r^0_{X/Z,Y/Z} \tag{3}$$

We obtain an average rescaled correlation of 0.46 between *firearms* and *firearm suicides*. This is neither a very strong nor a very weak correlation. Thus association between these two measures appears to persist after rescaling, albeit far more weakly than stated by EKS. Between *firearms* and *non-firearm suicides*, the average rescaled correlation over time takes a negative value of −0.32, which is rather weak. What is remarkable is the change in sign compared to the spurious results reported by EKS. Last, for *all suicides*, there is an average rescaled correlation of −0.15. This is hard to interpret without testing for significance, a problem addressed in Section 4. Without testing for significance, we have a not very strong, but clearly present, positive correlation between the measure for firearms and firearm suicides, a rather weak negative correlation between the measure for firearms and other suicides, and a negative correlation between the measure for firearms and all suicides too weak to base any findings on. However, it is still clear that rejecting Model (b) of Figure 1 in favor of Model (a), as done by EKS, is not advisable based on this empirical foundation.

3.3 Simulation Study

The results from Tables 2 (Rows IIa and IIb) and 5 while theoretically well founded are surprising given how much they change the initial results. Also my results report rescaled correlations for Pearson's method and not for Spearman's rank correlation: ranks are not ratios. So, do the results hold for ranked ratios? Because I could find no theoretical result for spurious correlation reference points for ranks of ratios, I conducted a simulation study.

I used a hotdeck simulation. For each year I repeatedly (10,000 times), randomly, and independently redistributed the observed numerators ($E95$, $E955$, CCL) across the counties, thus ensuring that, on average, there is no correlation between the numerators. Fortunately, $\max\{E95_{k,t}, E955_{k,t}, CCL_{k,t}\} < \min\{pop_{k,t}\}$ $\forall t$ so no ratios > 1 could occur. I next, for each repetition, computed the same ratios and ranks of ratios as done for the analysis in Sections 3.1 and 3.2. The random rank correlations of the numerators generated in this manner appear to be distributed around 0. (See Figure 2 in Appendix B for selected years.) The ratios' rank correlation distributions, on the other hand, are clearly shifted to the right and obviously skewed (Appendix B, Figure 3). The situation is persistent for all years and Pearson's correlations. The numerical results are well in accordance with the correction derived from Kim's (1999) theoretical result and Pearson's (1896) initial estimates of the problem size.

Year-wise simulated reference points can be calculated by computing the mean of the simulated (rank) correlations between the ratios. These values then can be used to rescale the results from the biased correlation analysis from Table 2, Row I, resulting in Rows IIIa, IIIb, and the detailed Tables 6 and 7 for rank correlations and Pearson's correlations found in Appendix B.

Results from Table 2, Row IIIa and Tables 6 and 7 are intriguing: after rescaling there remains, on average, a *negative* correlation between concealed carry licenses and all suicides. Interpreting such low correlation coefficients in favor of any side of any debate, however, is a tricky business. This question of valid inference is addressed in Section 4 of this paper.

4 Panel Regression

4.1 Model and Results

A more sophisticated method for analyzing the data seems appropriate. Panel regression, in contrast to EKS's repeated cross-sectional analysis, is capable of accounting for the contemporal and intertemporal structure of the data. Panel regression must be employed carefully as (a) either using ratios or (b) ignoring time series effects may cause spurious results. (a) can be dealt with by considering a risk model, as outlined below; (b) is solved by estimating the model on first differences.

The risk model is theoretically based on Duggan (2003).[9] A binary choice model is set up for individual i's suicide decision in the form of a linear probability model:

$$\Pr(Suicide_i) = \alpha + X_i\theta + \gamma Gun_i + \lambda_i + \varepsilon_i. \tag{4}$$

X_i represents individual characteristics, Gun_i is a dummy indicating if i owns a firearm, and the individual propensity λ_i tells us how strongly i is inclined to commit suicide. As Duggan notes, if the unobservable λ_i and Gun_i are not independent, we face a sample selection problem. Taking Duggan (2003, Eq. 2)

$$\lambda_i = \mu + \sigma Gun_i + \zeta_i \tag{5}$$

we see that unless $\sigma = 0$ in Equation (5), omitting λ_i introduces a bias into estimation of Equation (4)'s parameters. This problem can be overcome by using the data from above and imposing a risk model on the aggregate values. Let the

[9] I deviate slightly from Duggan's notation without changing the model to keep my formulas simple.

number of suicides (all, firearm, or non-firearm) have a conditional expectation of

$$\mathcal{E}[Y_{k,t}|\overline{X}_{k,t}, \overline{Gun}_{k,t}, \overline{\lambda}_{k,t}, pop_{k,t}] = pop_{k,t}(\alpha + \overline{X}_{k,t}\theta + \gamma\overline{Gun}_{k,t} + \overline{\lambda}_{k,t}) \tag{6}$$

where $\overline{X}_{k,t}, \overline{\lambda}_{k,t}$ are the averages of those values in county k in year t and $\overline{Gun}_{k,t}$ is the percentage of gun owners in k's population in year t. When we assume the average propensity and the average characteristics to be invariant over time, the (averages of) controls in $X_{k,t}$ and the propensities can be fully captured in a fixed effect model's county dummy. For those X_i that are individual characteristics, $pop\overline{X}$ become the count of persons with these characteristics.

Slightly relaxing the restriction of time-invariant unobserved variables, identical intertemporal changes in X and λ across all Austria can be captured in a time dummy. Then we arrive at a two-way fixed effects model of

$$y_{k,t} = \beta_0 pop_{k,t} + \beta_1 CCL_{k,t} + d_k + d_t + \varepsilon_{k,t}. \tag{7}$$

Here the number of concealed carry licenses is used as a proxy[10] for the number of gun owners. Then β_0 can be interpreted identically to α from Equations (4) and (6) as the baseline risk of an individual committing suicide. β_1 will be related to γ by the relation between concealed carry licenses and gun owners. A gun owner's relative risk of committing suicide from Equation (4), ignoring propensity for illustrative purposes, will be related to the ratio of coefficients from Equation (7):

$$\frac{\alpha + X_i\theta + \gamma}{\alpha + X_i\theta} \sim \frac{\beta_0 + \beta_1}{\beta_0}. \tag{8}$$

Given the nature of the data, i.e., a panel with time series, to rule out spurious results from time series effects (which may be numerous), Equation (7) is estimated on the first differences,[11] i.e.,

$$\Delta y_{k,t} = \beta_0 \Delta pop_{k,t} + \beta_1 \Delta CCL_{k,t} + \delta_t + \nu_{k,t}, \tag{9}$$

which is a common technique for circumventing many time series problems. The null hypothesis of poolability of the data, conducting a Chow test for poolability across periods, is rejected for firearm suicides and all suicides as the dependent variable with p-values[12] < 0.05. Results are shown in Table 3. The association

[10]This will be a somewhat noisy proxy, of course, so we expect coefficient estimates to be biased downward (Baltagi, 2008, Section 10.1).

[11]Including individual fixed growth parameters, i.e. a county dummy in the first differenced model, leads to highly insignificant county fixed effects and to no qualitative change to the results set out in Tables 3 and 8.

[12]All standard errors are computed according to Driscoll and Kraay (1998); available for Stata via `xtscc` from Hoechle (2007). Employing clustered robust standard errors does not qualitatively change the results and – for all results reported to be significant on any level – uniformly yields smaller p-values.

between firearms and firearm suicides, well known from the extant literature, is confirmed at a reasonable level of significance. No association is found for firearms and overall suicides. There are three obvious explanations of this result.

(a) Too much noise may be added to the model by including other suicides, so that significance can no longer be attained.

(b) Variables that *cannot* be captured by the dummies are missing from the model, somehow causing the coefficient to be biased toward zero.

(c) The negative coefficient for non-firearm suicides shows very weak significance:[13] using one sided testing, the p-value is 0.08, which may imply substitution between methods in accordance with the findings of Klieve, Barnes and De Leo (2009) and Leenaars et al. (2003). Given that the opposed effect estimates are of nearly identical size, this argument is intriguing. However, the extremely weak significance of the negative coefficient must be considered, meaning that this finding should be viewed with caution. It is not replicated in the alternative model specifications in Section 4.2.[14] Repeating my analysis with more data from different countries would be interesting.

exogenous	Dependent variable, differenced		
variable	all suicides	firearm suicides	other suicides
Δpop	-0.000094	-0.000094	0.0000004
	(0.000173)	(0.000153)	(0.000159)
ΔCCL	0.000351	0.003517^{***}	$-0.0031670^{\dagger}$
	(0.002080)	(0.001317)	(0.002235)
R^2_{within}	0.0008	0.0584	0.0097

Table 3: Estimation results for Model (9), standard errors in parentheses, $^{\dagger}/^{*}/^{**}/^{***}$ indicating two sided significance on 20/10/5/1% levels, robust standard errors according to Driscoll and Kraay (1998) computed with `vcovSCC` from Croissant and Millo (2008), $N = 198$ observations, R^2 adjusted

4.2 Robustness

The results from Section 4.1 do not exhibit strong significance, thus raising the question of their robustness to slight modifications[15] of the estimating equation. One possible modification is to standardise across counties by computing log

[13] With clustered robust standard errors, the coefficient becomes highly significant.

[14] All of those model specifications exhibit a worse fit. Thus they will all likely be further off the true underlying conditional expectation function.

[15] I use $\varepsilon_{k,t}$ for the error term multiple times in this section; however, I do not assume it to be identically distributed for all models. I also reuse β for coefficients with different interpretations. These are not identical across models.

growth rates on both sides of the estimating equation. This results in

$$\ln\Big(\frac{y_{k,t}}{y_{k,t-1}}\Big) = \alpha + \beta_0 \ln\Big(\frac{pop_{k,t}}{pop_{k,t-1}}\Big) + \beta_1 \ln\Big(\frac{CCL_{k,t}}{CCL_{k,t-1}}\Big) + \varepsilon_{k,t} \tag{10}$$

where once again $y_{k,t}$ may be either the number of all suicides, firearm suicides, or other suicides. The hypothesis of poolability across periods for Equation (10) is not rejected for any of the dependent variables. Thus the estimation is run without time dummies.[16] Results are found in Table 8 in the Appendix, in the row labeled "Log growth rates (pooled)." A moderately significant and positive estimate for β_1 is found when firearm suicides is the dependent variable.

Another feasible model specification is estimation directly on the ratios. Kronmal's (1993: 390) advice of including the inverse of the common denominator as an explanatory variable must be taken, however, or this model would fall prey to the same spuriousness found in EKS's results. Heterogenous time trends in ratios are addressed by taking first differences. The estimating equation becomes

$$\Delta y_{k,t} = \beta_0 pop_{k,t}^{-1} + \beta_1 \Delta CCLR_{k,t} + \delta_t + \varepsilon_{k,t} \tag{11}$$

where $\Delta y_{k,t}$ for all suicides is computed as $SR_{k,t} - SR_{k,t-1}$; for the other suicides, the respective ratios are used. There is no "common denominator" per se for the left- and right-hand sides. Constructing a common denominator would result in population from t and $t-1$ also appearing in the numerator of both sides. Therefore, instead of the true denominator, I follow Kronmal's advice by assuming population to be constant from $t-1$ to t. This allows controlling for the denominator by using $pop_{t,k}^{-1}$ as an additional variable on the right-hand side. Testing rejects poolability across time, at least for "all suicides"; therefore, the results given in Table 8 of the Appendix include time fixed effects. Again, the estimate of Equataion (11)'s β_1 when firearm suicides are the dependent variable is positive and moderately significant. The size of the estimate is notably similar to the result in Table 3 achieved by estimating Equation (9).

Following Duggan (2001) and Cook and Ludwig (2006), who run similar regressions for crime rates, we can also look for elasticity in suicide rates with respect to gun ownership rates by taking logarithms of the variables:

$$\Delta \ln y_{k,t} = \alpha + \beta_0 \Delta \ln pop_{k,t} + \beta_1 \Delta \ln CCLR_{k,t} + \varepsilon_{k,t}. \tag{12}$$

This model, in contrast to the models used by Duggan (2001) and Cook and Ludwig (2006), does account for spurious correlations between ratios by including the common denominator on the right-hand side. Note that this specification is

[16] Including time dummies does not change the results qualitatively; significance on the gun measure weakens, as is to be expected for an overspecified model.

very similar to Equation (10) model-wise, and as was the case for Equation (10), here, again, poolability is not rejected. Results for "firearm suicides" and "other suicides" in Table 8 in the Appendix are also very similar between these two models.[17] The estimate for β_1 when firearm suicides is the dependent variable is positive and moderately significant once again.

Thus, the results from the initial model in Section 4.1 hold up quite well to several modifications, which is what we would expect given an underlying – but, of course, unknown – conditional expectation function monotonous in the variables. From a goodness-of-fit point of view, Model (9) seems to be the best choice.

In light of this interesting result, a qualitative argument for causality can be made. Austrian firearm laws allow the purchase of firearms without need for a CCL. Therefore, persons intending to commit suicide by firearm do not need to acquire a CCL. This means the number of CCLs should not be driven by the number of firearm suicides. Firearm suicides, however, may very well be driven by the number of CCLs, given that those represent an underlying number of firearms owned by individuals.

5 Conclusion

I conclude this paper with two main findings.

(1) The "Fraction of Suicides by Firearms" (FSS) does indeed appear to be a valid proxy for gun ownership density. My results are in accordance with recent findings from U.S. data (Lang, 2013).[18]

(2) The correlation results from Etzersdorfer, Kapusta and Sonneck (2006) are greatly overstated because the authors fail to acknowledge spurious correlations between ratios.

Note, however, that finding (1), does not mean FSS should be used indiscriminately in regression analysis of models that contain firearms as an explanatory variable. FSS has its own problems, detailed in Westphal (2013), and may produce spurious results itself.

Finding (2) indicates that EKS's results cannot be used for public policy advocacy. Given that the journal that published Etzersdorfer, Kapusta and Sonneck (2006) is unwilling to acknowledge the partial spuriousness of the results,[19] EKS's article should be viewed with caution by the scientific and political community.

[17] An explanation for this can be found in Westphal (2013, Section 5.2).

[18] Lang's results are not affected by the ratio fallacy. The author unhesitatingly shared his data with me; running the usual specifications to check for spurious results due to ratio variables did not lead to different findings.

[19] The *Wiener Klinische Wochenschrift* rejected a short note on this problem for reasons unrelated to the scientific finding.

In conclusion, this paper demonstrates once again[20] that correlation and regression studies involving ratios need to take a close and careful look at the nature of the data and their possible implications. Recently, the ratio fallacy was demonstrated to occur in a prominent study on the association between guns and crime (Westphal, 2013), and it very well may be a problem in more analyses on that topic.

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[20] See Kronmal (1993) for more examples.

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A Correlation Tables, Repeated Cross-Section

	Rates					
	Firearm suicides		Other suicides		All suicides	
Year	Pearson	Spearman	Pearson	Spearman	Pearson	Spearman
1982	0.265	0.25	0.375	0.317	0.363	0.317
1985	0.734	0.667	0.131	0.233	0.331	0.200
1987	0.568	0.617	0.507	0.400	0.622	0.533
1990	0.870	0.867	0.014	−0.083	0.345	0.300
1992	0.658	0.517	0.006	0.200	0.209	0.183
1994	0.756	0.733	−0.111	0.083	0.144	0.283
1995	0.651	0.700	−0.038	−0.117	0.235	0.267
1996	0.381	0.300	−0.118	−0.067	0.044	0.200
1997	0.768	0.617	−0.181	−0.033	0.144	0.250
1998	0.512	0.417	0.215	0.300	0.353	0.400
1999	0.762	0.867	0.090	0.183	0.421	0.567
2000	0.745	0.817	−0.072	−0.083	0.155	0.250
2001	0.643	0.667	0.184	0.500	0.358	0.717
2002	0.527	0.583	0.049	0.167	0.196	0.133
2003	0.727	0.667	−0.156	−0.333	0.086	−0.267
2004	0.816	0.983	−0.152	−0.083	0.046	0.250
2005	0.213	0.367	−0.513	−0.383	−0.410	−0.267
2006	0.669	0.733	0.003	0.233	0.203	0.283
2007	0.463	0.300	−0.619	−0.400	−0.549	−0.317
2008	0.817	0.767	0.231	0.417	0.586	0.583
2009	0.767	0.700	−0.298	−0.200	−0.072	−0.033
2010	0.856	0.850	−0.252	−0.317	0.086	0.100
2011	0.712	0.733	−0.521	0.083	−0.296	0.167
Average over time	0.647	0.640	−0.053	0.044	0.157	0.222

Table 4: Correlations between gun ownership rate and suicide rates

	firearms suicides		other suicides		all suicides	
	estimated	rescaled	estimated	rescaled	estimated	rescaled
Year	r^0	r^*	r^0	r^*	r^0	r^*
1982	0.366	−0.074	0.378	−0.002	0.377	−0.010
1985	0.380	0.571	0.390	−0.186	0.391	−0.043
1987	0.358	0.327	0.378	0.207	0.377	0.393
1990	0.357	0.797	0.399	−0.275	0.392	−0.034
1992	0.364	0.462	0.388	−0.275	0.387	−0.128
1994	0.359	0.619	0.387	−0.359	0.384	−0.173
1995	0.353	0.460	0.404	−0.315	0.396	−0.115
1996	0.336	0.068	0.392	−0.366	0.383	−0.245
1997	0.375	0.629	0.393	−0.412	0.391	−0.177
1998	0.324	0.278	0.394	−0.128	0.382	−0.021
1999	0.338	0.641	0.401	−0.222	0.393	0.047
2000	0.384	0.586	0.414	−0.344	0.414	−0.183
2001	0.364	0.439	0.418	−0.165	0.412	−0.038
2002	0.383	0.234	0.394	−0.247	0.399	−0.145
2003	0.403	0.543	0.393	−0.394	0.396	−0.222
2004	0.391	0.698	0.407	−0.397	0.408	−0.257
2005	0.333	−0.090	0.412	−0.655	0.399	−0.578
2006	0.344	0.495	0.400	−0.284	0.391	−0.135
2007	0.359	0.163	0.402	−0.729	0.403	−0.678
2008	0.331	0.726	0.410	−0.127	0.406	0.303
2009	0.366	0.633	0.392	−0.496	0.395	−0.334
2010	0.340	0.782	0.420	−0.473	0.412	−0.231
2011	0.324	0.575	0.372	−0.651	0.367	−0.485
Average over time	0.357	0.459	0.397	−0.317	0.394	−0.152

Table 5: Correlations rescaled (r^*) with reference points of no correlation (r^0) between carry permit rates and suicide rates; reference points estimated by Equation (1)

B Simulation Results

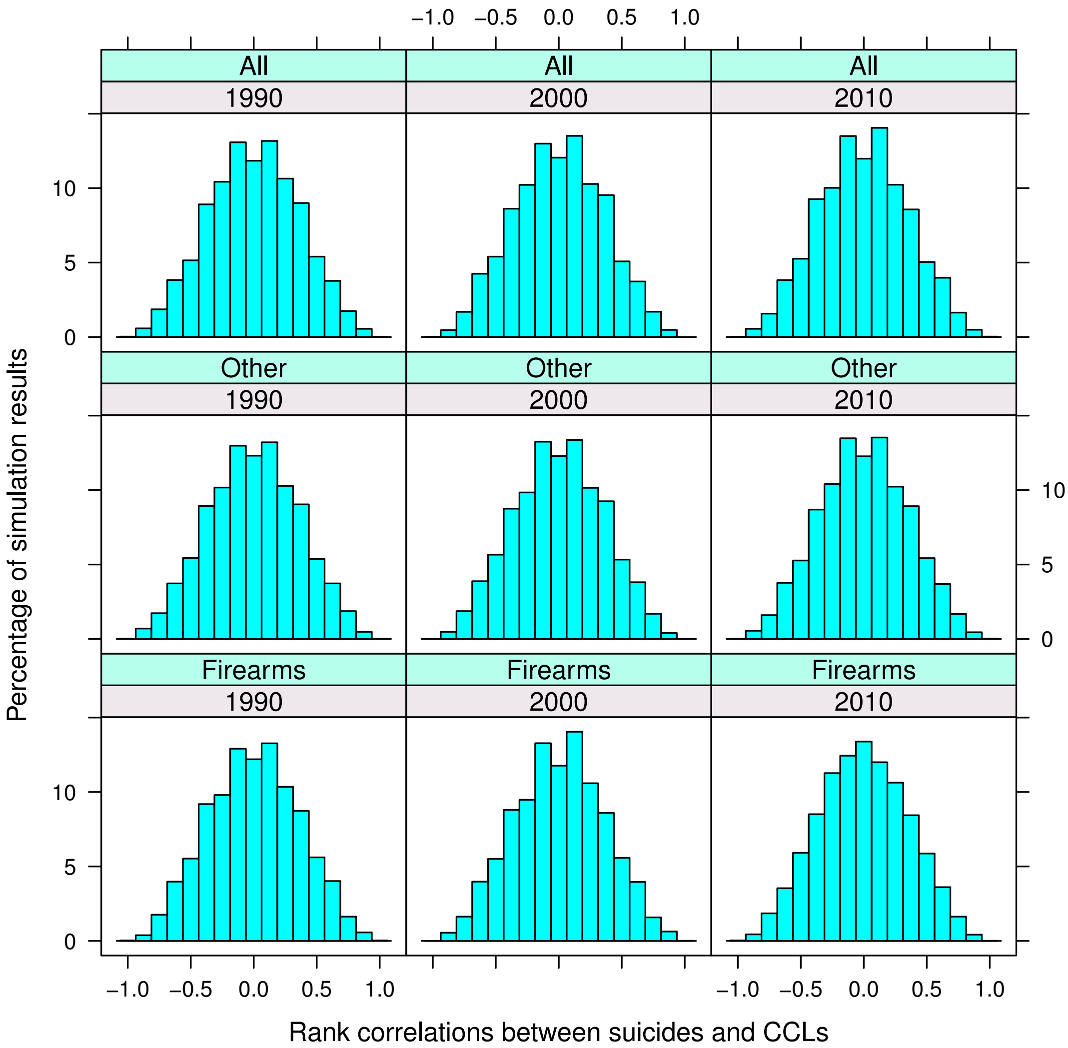

Figure 2: Histograms of simulated rank correlations between uncorrelated carry permits and suicides.

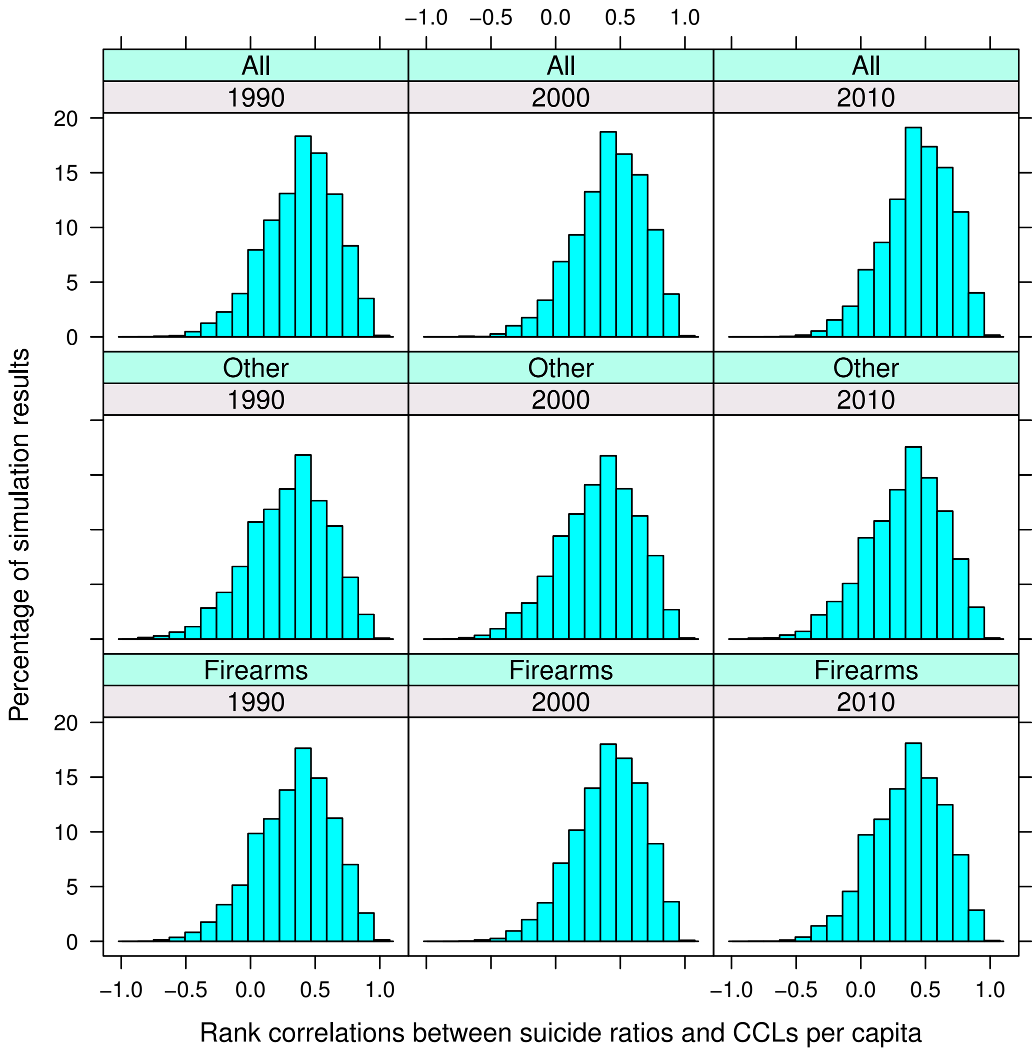

Figure 3: Histograms of simulated rank correlations between uncorrelated carry permits and suicide *rates*.

	firearms suicides		other suicides		all suicides	
	simulated	rescaled	simulated	rescaled	simulated	rescaled
year t	r^0	$\mathrm{rg}\, r^*$	r^0	$\mathrm{rg}\, r^*$	r^0	$\mathrm{rg}\, r^*$
1982	0.346	−0.072	0.277	0.055	0.362	−0.033
1985	0.398	0.447	0.321	−0.066	0.399	−0.142
1987	0.376	0.386	0.291	0.154	0.369	0.260
1990	0.342	0.797	0.295	−0.292	0.389	−0.064
1992	0.380	0.220	0.297	−0.075	0.382	−0.144
1994	0.343	0.594	0.278	−0.152	0.358	−0.055
1995	0.339	0.546	0.300	−0.320	0.397	−0.093
1996	0.313	−0.010	0.277	−0.269	0.358	−0.116
1997	0.343	0.417	0.280	−0.245	0.368	−0.087
1998	0.298	0.169	0.285	0.021	0.362	0.060
1999	0.322	0.803	0.302	−0.091	0.388	0.292
2000	0.406	0.691	0.335	−0.314	0.417	−0.118
2001	0.409	0.436	0.339	0.244	0.433	0.501
2002	0.432	0.266	0.316	−0.113	0.396	−0.188
2003	0.419	0.426	0.328	−0.498	0.414	−0.481
2004	0.416	0.971	0.330	−0.311	0.405	−0.110
2005	0.300	0.095	0.314	−0.531	0.391	−0.473
2006	0.329	0.603	0.299	−0.051	0.385	−0.074
2007	0.393	−0.067	0.331	−0.549	0.405	−0.514
2008	0.375	0.627	0.367	0.078	0.457	0.232
2009	0.418	0.484	0.327	−0.397	0.414	−0.316
2010	0.368	0.763	0.345	−0.492	0.441	−0.236
2011	0.317	0.609	0.277	−0.152	0.358	−0.141
average over time	0.364	0.444	0.309	−0.190	0.393	−0.089

Table 6: Rank correlations rescaled ($\mathrm{rg}\, r^*$) with simulated reference points of no correlation (r^0) between carry permit rates and suicide rates

	firearms suicides		other suicides		all suicides	
	simulated	rescaled	simulated	rescaled	simulated	rescaled
Year	r^0	r^*	r^0	r^*	r^0	r^*
1982	0.383	−0.086	0.342	0.050	0.392	−0.021
1985	0.408	0.551	0.367	−0.173	0.417	−0.061
1987	0.383	0.299	0.356	0.235	0.406	0.364
1990	0.374	0.792	0.358	−0.254	0.420	−0.053
1992	0.402	0.427	0.355	−0.258	0.416	−0.146
1994	0.381	0.605	0.347	−0.340	0.409	−0.188
1995	0.377	0.439	0.364	−0.295	0.431	−0.137
1996	0.359	0.034	0.349	−0.346	0.417	−0.263
1997	0.403	0.611	0.359	−0.397	0.426	−0.198
1998	0.340	0.260	0.356	−0.104	0.422	−0.048
1999	0.373	0.620	0.363	−0.200	0.428	−0.004
2000	0.439	0.546	0.399	−0.337	0.458	−0.208
2001	0.412	0.393	0.398	−0.153	0.460	−0.070
2002	0.436	0.163	0.383	−0.242	0.439	−0.169
2003	0.442	0.512	0.385	−0.390	0.439	−0.245
2004	0.429	0.678	0.393	−0.391	0.445	−0.276
2005	0.353	−0.103	0.375	−0.646	0.435	−0.588
2006	0.368	0.476	0.368	−0.266	0.432	−0.160
2007	0.409	0.092	0.390	−0.726	0.443	−0.678
2008	0.368	0.710	0.401	−0.121	0.456	0.240
2009	0.410	0.605	0.383	−0.493	0.439	−0.355
2010	0.378	0.769	0.393	−0.463	0.456	−0.254
2011	0.346	0.560	0.337	−0.642	0.394	−0.495
average over time	0.390	0.433	0.370	−0.302	0.429	−0.175

Table 7: Correlations rescaled (r^*) with simulated reference points of no correlation (r^0) between carry permit rates and suicide rates

C Further Regression Analysis Results

The results for the different models need to be viewed with caution. Joint significance is weak to nonexistent for all of them. The value of reporting the results lies in the robustness of the positive coefficient on the various transformations of the firearms proxy.

The time-pooled models are estimated with a constant as shown in Equations (10) and (12). In theory, this constant should be zero. However, for some of the pooled models, this constant tests weakly significant. This could indicate an ignored time effect, thereby contradicting the result from the respective Chow tests for timepoolability. Results for these models do not differ much when they are estimated without time pooling.

exogenous variable	Dependent variable all suicides	firearm suicides	other suicides
Log growth rates (pooled), Equation (10)			
$\ln(pop_t/pop_{t-1})$	0.1442	$-2.2778^{\dagger}$	0.5507
	(0.590)	(1.572)	(0.633)
$\ln(CCL_t/CCL_{t-1})$	-0.0044	0.7981^{**}	-0.1516
	(0.209)	(0.312)	(0.178)
R^2_{pooled}	0.0000	0.01411	0.0024
Estimation on ratios, Equation (11)			
pop^{-1}	0.4916	0.1870	0.3046
	(2.001)	(0.744)	(1.483)
$\Delta CCLR$	$0.0038^{\dagger}$	0.0030^{**}	0.0008
	(0.003)	(0.001)	(0.002)
R^2_{within}	0.0078	0.0254	0.0006
Elasticity estimation (pooled), Equation (12)			
$\Delta \ln pop$	$-0.8602^{\dagger}$	-2.4800^{*}	-0.6009
	(0.620)	(1.427)	(0.633)
$\Delta \ln CCLR$	-0.0044	0.7981^{**}	-0.1516
	(0.148)	(0.312)	(0.178)
R^2_{pooled}	0.0015	0.0152	0.0121

Table 8: Estimation results for panel models, standard errors in parentheses, $^{\dagger}/^{*}/^{**}/^{***}$ indicating two sided significance on 20/10/5/1% levels, robust standard errors according to Driscoll and Kraay (1998) computed with `vcovSCC` from Croissant and Millo (2008), $N = 198$ observations, R^2 adjusted