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Competition Policy and Exit Rates: Evidence from Switzerland*

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Abstract

This paper provides evidence on the relation between the intensity of product market competition and the probability of exit. We adopt a natural experiment approach towards analyzing the impact of a tightening of Swiss antitrust legislation on exit probabilities. Based on a sample of more than 68,000 firms from all major sectors of the Swiss economy, we find that the exit probability of non-exporting firms increased significantly, whereas the exit probability of exporting firms remained largely unaffected. Our results support the notion that there is a positive relationship between the intensity of product market competition and the probability of exit.

JEL Classification: D43, L23, L40.

Keywords: competition intensity, exit, natural experiment.

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1 Introduction

There appears to be a broad consensus that consumers typically benefit from the enforcement of antitrust or 'cartel laws'. Surprisingly, it it less clear whether producers actually suffer from the reduced opportunity to increase profits by collusion. For instance, Selten (1984) shows that, under free entry and exit, cartel laws prohibiting collusion are not necessarily bad for business: When collusion is effectively prevented, there will be fewer competitors around, so that active producers make higher profits on average than when they collude.[1]

More recent work by Sutton (1991, 1998) also emphasizes that the equilibrium number of firms tends to decrease when the intensity of product market competition increases. In particular, Sutton's analysis demonstrates that there is a robust relation between the intensity of product market competition and concentration in industries where sunk costs are *exogenous*: There is a lower bound to concentration that unambiguously increases with the intensity of product market competition. That is, the higher the intensity of product market competition, the lower the equilibrium number of firms that may be supported by this market. The picture is less clear, however, for industries with *endogenous* sunk costs, where firms bear significant costs for advertising or research and development (R&D) before competing in the product market. In this case, the effect of more intense product market competition on concentration can go either way.[2]

It is probably fair to say that the empirical evidence on the relation of product market competition and concentration is fairly scant and has produced mixed results (see Bittlingmayer 1985, Elliot and Gribbin 1977, and O'Brien et al. 1979). The lack of clear evidence is unsurprising, given the difficulties associated with measuring the intensity of product market competition and handling the notorious endogeneity problems of industry studies. In a recent analysis of the introduction of anti-cartel policy on concentration in the UK, Symeonidis (2000b) has circumvented these problems by adopting a "natural experiment" framework, viewing this policy change as exogenous.[3] His results support the notion that more intense product market competition increases concentration in both exogenous and endogenous sunk cost industries.

In the present paper, we focus on a closely related question that has largely been ignored in the previous literature: How does an (exogenous) increase in the intensity of competition affect a firm's *exit probability*? We think this is a natural question to ask, as

[1] See Phlips (1995, Ch. 3) for a textbook discussion of Selten's argument.

[2] Symeonidis (2000a, property 1) illustrates this ambiguity in the setting of a linear Cournot model with quality indices, where concentration is likely to decrease (increase) in the neighborhood of perfect collusion (the non-collusive Nash equilibrium, respectively).

[3] See Meyer (1995) on the use of natural and quasi-experiments in economics.

one might expect exits to increase (at least temporarily) when product market competition becomes more intense.[4] Also, taking the perspective of an individual firm allows us to sidestep the nontrivial problem of constructing useful concentration measures across a large number of diverse industries.[5] Following Symeonidis (2000b), we adopt a natural experiment approach towards analyzing the impact of a major change in Swiss antitrust law enacted on July 1, 1996. More specifically, we compare the impact on the 'treatment group' of non-exporting firms facing little competition in domestic markets with the impact on a 'control group' of exporting firms operating under international competition. In doing this, we exploit the dualistic nature of the Swiss economy with competitive export industries and highly cartelized domestic industries (see Borner et al. 1995).

Earlier contributions by Klepper and Graddy (1990), Agarwal and Gort (1996), and Van Kranenburg et al. (2002) suggest that a firm's exit probability should be expected to depend on firm-specific characteristics—such as a firm's size, age, geographical location etc.—as well as on industry-specific and macroeconomic conditions. To our knowledge, Van Kranenburg et al. (2002) is the only paper that has analyzed the relationship between the intensity of product market competition and exit rates. These authors use the (lagged) number of competitors in the daily newspaper industry under consideration as a proxy for the intensity of product market competition, and they find that exit rates tend to increase with the number of competitors.

From the perspective of the literature discussed above, the number of competitors in any given industry is likely to be a good proxy for the intensity of product market competition only if the number of firms may be treated as an exogenous variable, i.e., if it is safe to assume that firms cannot decide about entry or exit. In contrast, if the number of firms is an endogenous variable, a high intensity of product market competition is associated with a small equilibrium number of firms, as only a small number of firms may be supported by the profits available in this market. That is, relative to the case where the number of firms is exogenous, the chain of causation between the number of firms and the intensity of downstream competition is reversed. In particular, we should expect that the intensity of product market competition is high (rather than low) when

[4]For instance, Sutton (1991, 43) notes that, in an off-equilibrium configuration, the only way to recover sunk costs lies in a change of market structure, which may come about in two ways:

> "Either consolidation of ownership—whether by means of acquisition or merger—may bring about a rise in margins; or else in the longer run the failure to recover sunk costs will lead to an unwillingness to renew [a] plant as it becomes obsolete, so that concentration rises *as exit occurs*." [emphasis added]

[5]For instance, the well-known m-firm concentration ratio, which adds up the m highest shares in the industry (see e.g. Tirole 1988, 221), will not react to liquidations or mergers of smaller firms.

the number of firms is small. We think that, in a study of firm exits, it is natural to view the number of firms as being endogenous.

We use a large combined data set that has become available only recently. Part of the data stem from the Swiss business census, which is a complete inventory count of all firms active at the time of observation, encompassing approximately 297,000 firms. This census contains information on the characteristics of firms, such as their age, location, legal form, number of employees, etc. The data on the firms' exits were provided by Dun & Bradstreet, which has compiled a comprehensive database covering three different types of exits: (i) bankruptcy, (ii) voluntary liquidations, and (iii) mergers.

We employ a Cox (1972, 1975) model with time-varying covariates to characterize hazard rates. Our main results are the following. *First*, the change in Swiss antitrust law in 1996 had a strong impact on firm conduct, raising hazard rates for the full sample significantly. *Second*, whereas non-exporting firms experienced a significant increase of hazard rates, exporting firms already exposed to international competition were not significantly affected by the tightening of antitrust law. Taken together, these results suggest that the tigthening of antitrust law in 1996 led to an increase of the intensity of competition within Switzerland, raising exit rates significantly. *Third*, our findings with respect to the remaining determinants of hazard rates—firm- and industry-specific properties as well as macroeconomic conditions—are in line with the previous literature. In particular, we find that hazard rates tend to decrease in age and size.

The remainder of the paper is organized as follows. Section 2 provides a brief discussion of antitrust policy in Switzerland. Section 3 describes the data set and variables. Section 4 sets out the empirical model and discusses our main results. Section 5 concludes.

2 Antitrust Policy in Switzerland

Swiss antitrust policy has traditionally been perceived as being very permissive (see e.g. Porter 1990, 714). In part, the lax attitude of Swiss antitrust authorities towards anticompetitive conduct is explained by the fact that, until today, the Swiss constitution makes it difficult to declare cartels per se as unlawful. Before the revision of Swiss antitrust law in 1996, antitrust authorities were required to go through a rather opaque process called "balance method" (Saldomethode) to evaluate the costs and benefits of a particular (mis)conduct, with considerations such as the impact on the labor market or specific regions routinely playing an important role. Since it was generally very difficult to prove that a cartel actually had a negative 'net benefit', cartels were rarely prohibited. Neven and Ungern-Sternberg (1997, 36) describe the performance of Swiss competition

policy up to the mid 1990s as follows:

> "In the past, the [Cartel] Commission has relied far too much on judgements and far too little on sound analysis. In various dimensions (definition of relevant markets, evaluation of dominance, evaluation of countervailing benefits, imposition of remedies), the analysis is rather poor by the standards of other jurisdictions. It lacks organising principles, fails to bring appropriate evidence and often relies on highly judgmental evaluations."

Emphasizing the need for a thorough revision of Swiss antitrust law, Borner et al. (1995) further pointed out that the Swiss economy featured *dualistic characteristics*: On the one hand, there was the competitive export sector serving the world markets; on the other hand, there was a highly subsidized domestic sector facing little competition due to a mixture of public regulations and both horizontal and vertical collusion.

The revision of the antitrust law, enacted on July 1, 1996, finally led to the per se prohibition of so-called 'hard' cartels that eliminate 'effective competition' by fixing prices, restricting quantity or dividing up markets (Art. 7). Furthermore, the notorious balance method was abolished. Taken together, these improvements were expected to intensify competition in domestic markets considerably, even though there arguably remained a number of relevant shortcomings—such as the lack of power of competition authorities to penalize parties restricting competition without delay and to confiscate extra profits from unlawful behavior (OECD 2000).[6]

This rather drastic change in Swiss antitrust legislation allows us to study the effect of intensifying product market competition on firms' exit behavior using a natural experiment framework: The change in antitrust legislation generated variation in the intensity of product market competition that is plausibly exogenous (Meyer 1995). We can thus sidestep the well-known endogeneity problem of structure and performance studies (Schmalensee 1989).

To evaluate the impact of the change in antitrust law on the firms' exit behavior, we will refer to the dichotomy of the Swiss economy discussed above and distinguish the following types of firms:

- The vast majority of Swiss firms was solely active on domestic markets, i.e., they did not export to foreign countries ("*non-exporting firms*"). These firms should have been significantly affected by the change in antitrust legislation.

[6] Eliminating these and other shortcomings was the objective of yet another revision of the antitrust law enacted in April 2004.

- A smaller but non-negligable number of Swiss firms exported at least part of their output to the world market ("*exporting firms*") and had thus already faced intense competition on international markets. The impact of the change in Swiss antitrust law on these firms should thus be expected to be much smaller. In particular, it is reasonable to hypothesize that firms which were located in Switzerland but did export more than 2/3 of their output to foreign markets were not much affected by the change in Swiss antitrust legislation.

3 Data and Variables

In this section, we briefly discuss our data and the variables used to estimate the impact of the change in antitrust law on the firms' exit probability.

3.1 Data Source and Sample Composition

For the purpose of this study, we merged the following databases:

(1) *Swiss Business Census* (SBC 95). The SBC 95 is a complete inventory count compiled by the Swiss Federal Statistical Office (BFS), which contains all firms with more than 20 weekly aggregate working hours existing in September 1995, excluding the agricultural sector. The SBC 95 provides numerous variables that characterize the attributes of these firms.

(2) *Dun & Bradstreet Exit Database* (DBED). The DBED contains all exits of firms located in Switzerland from January 1994 to December 2000. It distinguishes the following types of exit: (i) bankruptcies, (ii) voluntary liquidations, and (iii) mergers.

The merged database covers an observation period between October 1995 and December 2000. After deleting all firms that were non-profit oriented according to their legal status—such as cooperatives ("Genossenschaften"), associations and clubs ("Vereine"), foundations ("Stiftungen"), churches, embassies and international organizations—, the merged database contained 276,123 firms. Since for sole proprietorships, the DBED does not fully cover voluntary liquidations and mergers, we dropped all sole proprietorships. Furthermore, we dropped all firms established prior to 1970, since no information on their founding dates was available.

After dropping these firms, our sample still includes more than 68,000 firms and is thus comparably large and comprehensive. In particular, we have firms of all size ranges

and ages up to 25 years in our sample, which has rarely been the case in previous studies. Furthermore, with the exception of the agricultural sector, our sample contains all industries represented in Switzerland (including services), whereas earlier work typically focussed on only a few industries and did not cover services due to data limitations.

However, we are also aware of two disadvantages of our database associated with the way entries and exits were recorded. First, whereas the DBED records exit times as exact dates (day/month/year), the SBC 95 gives entry dates in intervals only (various time spans).[7] One approach towards dealing with this problem in survival analysis is the use of interval-censored models. However, these models are not designed to handle time-varying covariates (changes within the intervals), which will be crucial for our analysis below. We therefore adopted the alternative approach of assuming a uniform distribution of entries within these intervals (since no further information was available) and simulating the date of entry, which yielded survival times measured in quarters (see section 3.2 below). Second, it is well-known that firms tend to announce voluntary liquidations with some delay (i.e., after closing down operations), giving rise to delays in registration. This is a general problem of business failures studies.

It is instructive to compare the composition of our sample with that of a related study by Harhoff et al. (1998) for West Germany, which is similar to ours in a number of respects: First, it contains manufacturing as well as service firms, unlike the vast majority of other studies. Second, these authors distinguish bankruptcies and voluntary liquidations. Third, they have older firms in the sample that were at risk before the survey period ("delayed entry"). Finally, their observation period is of a similar length as ours.[8]

We now want to highlight some special properties of our sample. The descriptive statistics given in Table 1 indicate that small firms make up a large proportion of Swiss firms: Using the number of employees (*Emp*) to measure firm size, average firm size is about 11 employees. The median firm size amounts to only 4 employees. Harhoff et al., in contrast, found an average size of roughly 276 employees, and a median size of slightly less than 8 for West Germany.

Recall from our above discussion that firms founded before 1970 had to be excluded from our sample due to data limitations. It is thus unsurprising that, on average, firms were only about 8 years old when they entered the survey period, whereas they were roughly 29 years old in Harhoff et al. The firms in our sample are thus relatively young.

[7] Entries were recorded in the following nine intervals (..., 1969], [1970,1980], [1981,1985], [1986,1990], [1991], [1992], [1993], [1994], [1995].

[8] Harhoff et al. (1998) in turn compare their results to the study of Wagner (1994) for young firms and find that "the difference is small enough to be accounted for by differences in industry composition" (p. 467).

We further find that, after excluding sole proprietorships, stock corporations clearly dominate in our sample, with close to 80% of all firms belonging to this group of firms.[9] This share looks surprisingly large compared to Harhoff et al., where only about 4% of the firms are stock companies. However, next to the fact that, unlike Harhoff et al., we exclude the largest group of sole proprietorships, it is largely explained by the fact that Swiss legislation poses very few obstacles to small firms to become stock companies.

Finally, taking a look at industries, we observe that more than two thirds of the firms in our sample belong to the service sector (in the SBC 95 this share amounts to three quarters of all firms). The vast majority of previous studies had—if at all—a somewhat limited access to data on firms in the service sector. In the sample of Harhoff et al., for instance, only 30% of the firms belong to the service sector.

3.2 Variables and Descriptive Statistics

We first explain the construction of our dependent variable and then discuss the explanatory variables (see Table 1 for the descriptive statistics).

<Table 1 around here>

3.2.1 Dependent Variable

Our dependent variable is the firm's lifetime, measured by how many quarters a firm stayed in business. As noted above, the DBED contains three different types of exit: (i) bankruptcies, (ii) voluntary liquidations, and (iii) mergers. In some studies, all three types of exit are pooled (e.g. Dunne and Hughes 1994). Other studies use a more narrow definition of exits—closely related to the concept of "failures"—excluding mergers, which do not necessarily imply a low profitability of the involved firms. In the following, we shall use a broad definition of exits and pool all three types of exit.[10]

In our sample, exits were recorded as bankruptcies if the firms filed for bankruptcy between October 2, 1995 and December 31, 2000. For the exact time of exit, we used the first available date which, in the bulk of bankruptcy cases, is the time when the court instituted bankruptcy proceedings (as opposed to the time when the firm finally closed down). This is due to the fact that the time spread between the opening of bankruptcy

[9] In the SBC 95, where sole proprietorships are included, we find this group to have the largest share (62.49%), at least in terms of their numbers; however, even there the share of stock corporations is with 30% still considerable.

[10] To check the robustness of our results, we have also used the more narrow definition of exits *excluding mergers*. The results of these estimations are similar to those presented in Table 3; they are available on request from the authors.

proceedings and the actual close down varies considerably, depending, for instance, on the size and the legal form of the firm. The other types of exit, voluntary liquidations and mergers, were recorded at the time when the respective firms were deleted from the commercial register. The time when they actually closed down would have been preferable, but was not available.

Using the founding and exit times, it is straightforward to calcuate the duration of a firm's presence in the market. Note that the resulting duration data is *right censored*, i.e. there are (many) firms that have not left the pool during the survey period. For these firms, we know that true duration is at least as large as observed duration. Furthermore, the data is *left truncated*, as all the firms covered by the SBC 95 must have been founded prior to October, 1995 and thus have been at risk before coming under observation (delayed entry). Both right censoring and left truncation will have to be taken into account when modelling the probability of exit.

3.2.2 Explanatory Variables

In our sample, all values of firm attributes refer to the time when the firm filed the relevant information for the business census, similar to Harhoff et al. (1998) and Konings and Xavier (2003). In most other studies, firm attributes refer to the time of the firm's founding. We are aware that some of these attributes—such as the firm's size—may change over lifetime. Yet, as in virtually all other studies on business failures, time-varying firm attributes are not available. To accommodate this problem, we control for the firm's age at the time of entering the sample. More specifically, we include the variable *Age 95,* which indicates the firm's age in September 1995 when it entered the SBC 95.

In virtually all previous studies, the size of a firm is operationalized by its assets (Dunne and Hughes 1994, Ranger-Moore 1997) or by the number of employees (Brüderl et al. 1992, Audretsch 1995, Harhoff et al. 1998). It is common to log transform these variables as it is natural to assume that marginal effects of size on exit probabilities decrease. In the present study, we describe size by the natural log of the number of employees (*LnEmp*); additionally, we include the square of this variable in order to test for non-monotonicity ($(LnEmp)^2$).[11]

We classify the legal form of firms into four groups, which differ with respect to the initial capital requirements (Brüderl and Schüssler 1990), ease of ownership transfer and liability rules (Harhoff et al. 1998) as well as tax treatment: (i) Partnerships (*Partner*), (ii) Limited liability companies (*Lim Liab*), and (iii) stock corporations (*Stock Corp*).

[11] Alternative specifications of firm size, e.g. the number of apprentices or the sales area, can be found in Kaiser (2004).

As noted above, sole proprietorships had to be dropped because the DBED does not fully cover the failures of these firms.

The SBC 95 further contains information on the nature of the firms' businesses. In our sample, we use dummy variables to identify the industry sector to which a firm belongs: Manufacturing (*Manufact*), construction (*Construct*), trade (*Trade*), or services (*Service*). Furthermore, we use a classification issued by BFS (1997) to control for geographical idiosyncrasies. That is, we use the following five dummy variables to indicate where a firm is domiciled: Eastern Switzerland, including the greater Zurich area and Graubünden (*Eastern CH*), Northwestern Switzerland (*NW CH*), Central Switzerland (*Central CH*), the French speaking area (*French CH*), and the Italian speaking area (*Ticino*). *Eastern CH* will serve as the reference variable.

Since a crucial aspect of our study will be to compare the impact of the change in antitrust law on non-exporting and exporting firms, we control for the export share of a firm, calculated as the ratio of exports and turnover, as indicated in the SBC 95. The database distinguishes the following firm types with respect to export activity:

(i) Non-exporting firms (*Exposh 1*); these firms serve as the reference group.

(ii) Exporting firms with export shares below one third (*Exposh 2*).

(iii) Exporting firms with export shares between one third and two thirds (*Exposh 3*).

(iv) Exporting firms with export shares above two thirds (*Exposh 4*).

Finally, we use a number of time-varying explanatory variables. The most important of these is a dummy variable controlling for the change in antitrust law (AL) in the third quarter of 1996; AL is zero up to the second quarter of 1996 and one after that. Next, we control for the external value of the Swiss currency (the Swiss Franc), using an index constructed by the Swiss National Bank (SNB) (2003) based on the real exchange rates with the 24 most important trade partners (*Ext Val*).[12] We further use a variable controlling for the aggregate movement of the economy in previous years, as in other studies with time-varying covariates. For instance, Van Kranenburg et al. (2002) use the lagged total number of firms while Ranger-Moore (1997) and Roberts and Thompson (2003) use the lagged aggregate number of failures or exits, respectively. We include the lagged number of bankruptcies (*Bankrupt*), generated by aggregating the yearly bankruptcies listed in the DBED.[13] We would expect that a higher number of bankruptcies in the

[12] As we use this variable with a one year lag, values between 1994 and 1999 enter the equation.

[13] This number is based on all firms, i.e. also on those which do not meet the requirements of the SBC 95.

previous year will increase hazard rates because of 'chain effects' (at work both within and across industries) that trigger further exits.[14]

3.3 Preliminaries on the Impact of Size and Age

Table 2 provides the sample frequencies of exits (in %)—broken down into failures and mergers—by firm size and firm age. It largely supports the finding of the previous literature that exit rates tend to decrease with *age* (Stinchcombe 1965, Caroll 1983, Amburgey et al. 1993, Olzak and West 1991, Mata and Portugal 1994, Audretsch et al. 2000) and *size* (Brüderl et al. 1992, Barron 1999, Audretsch et al. 2000, Agarwal and Audretsch 2001, Segarra and Callejón 2002).

<Table 2 around here>

More specifically, looking at exits by age (rightmost column), we find that exit rates decrease monotonically, with the exception of a negligible rebound for three-year-old firms. The overall decrease is more than 50%, from 21.1% for firms younger than two years to 10.4% for firms up to 25 years old. For size (bottom row) the decrease is strictly monotonic and amounts to about 45% from the smallest to the largest size class. Since the vast majority of exits are failures rather than mergers, these findings similarly apply to failures.

Our figures for age dependence resemble those of Harhoff et al. (1998). Our total average failure rate is 14.3% compared with theirs of 10.1%. The difference can be explained by our slightly longer observational period as well as by the fact that we excluded the oldest firms which should be expected to experience below-average failure rates. Moreover, our pattern of failure rates by size resembles theirs in the sense that an age-dependent decline can be observed for the smallest firms, whereas for larger firms, failure rates vary non-monotonically with age and do not show a clear pattern. Hence, while for small firms, getting older clearly lowers exit rates, for large firms advantages of age are less obvious. However, the bulk of firms in our sample are small, so that their negative duration dependence dominates our findings for the full sample.

As noted above, in addition to failures, exits as defined in the present study include mergers. Our figures show that the propensity to merge rises with the size of the firm, whereas the firm's age does not appear to make substantial differences. If we look at particular cells in Table 2, we find that firms that are both large and rather young are likely to merge.

[14] We use bankruptcies instead of failures (including voluntary liquidations) because we believe that detrimental chain effects are more strongly exerted by bankruptcies than by voluntary liquidations.

4 Empirical Model and Results

Duration models provide a suitable framework for characterizing the probability of exit. Let $T_i, i = 1, ..., n$, denote the continuous duration of firm i's survival in the market. The probability distribution of firm i's duration is characterized by $F_i(t) = \Pr(T_i < t)$, which determines the probability that firm i exits before some t. The corresponding density function is $f_i(t)$. Let $S_i(t) = \Pr(T_i \geq t) = 1 - F_i(t)$ denote the survivor function, which determines the probability that T_i is equal or larger than t. In the following we shall often refer to the *hazard function*

$$h_i(t) = \lim_{dt \to 0} \frac{\Pr\left(T_i \in [t, t+dt] | \, T_i \geq t\right)}{dt} = \frac{f_i(t)}{S_i(t)},$$

which, somewhat loosely, is the rate at which firm i exits at time t, given that it has not exited before, as a function of t. The value of this function is called the "hazard rate" or simply the "hazard" (Kiefer 1988; van den Berg 2001).

4.1 The Cox Model

The Cox proportional hazards model (Cox 1972; 1975) is the most popular approach towards characterizing the hazard function $h_i(t)$ by a vector of observed explanatory variables or covariates. Following Therneau and Grambsch (2000, p. 39), we use $x_{ij}, j = 1, ..., p$, to denote the jth covariate of firm i, denote the set of covariates by the $n \times p$ matrix $\mathbf{x}$, and let $\mathbf{x}_i$ denote the row vector of firm i's covariates. The Cox model then specifies the hazard function for firm i as

$$h_i(t | \, \mathbf{x}_i) = h_0(t) \exp(\mathbf{x}_i \beta),$$

where $h_0(t)$ is an unspecified non-negative "baseline hazard" which gives the shape of firm i's hazard function, $\exp(\mathbf{x}_i \beta)$ is the systematic part of the hazard, and β is the coefficient vector. This model is known as the "proportional hazards" model since, for any two firms i and $k, k \neq i$, and *fixed* covariates $\mathbf{x}_i$ and $\mathbf{x}_k$, we immediately have that

$$\frac{h_i(t | \, \mathbf{x}_i)}{h_k(t | \, \mathbf{x}_k)} = \frac{\exp(\mathbf{x}_i \beta)}{\exp(\mathbf{x}_k \beta)}$$

is constant over time. The parameters β may be estimated consistently by maximizing a partial likelihood function that does not depend on the baseline hazard (Kalbfleisch and Prentice 1980).

Clearly, the proportional hazards property no longer holds when the covariates *vary over time*, as the variables *AL*, *Bankrupt* and *Ext Val* in our study. Nevertheless, we can still derive valid econometric inference using the standard Cox model provided that the following conditions are satisfied (van den Berg 2001, p. 3398):

(i) $\mathbf{x}(t)$ is a *predictable* stochastic process. The concept of predictability stems from the counting process literature and essentially requires that the explaining variables are weakly exogenous (Ridder and Tunalı 1999, 196). More specifically, predictability implies that the value of $\mathbf{x}_i(t)$ is known infinitesimally before t, at time t^- or even earlier. Put differently, information on the value of $\mathbf{x}_i$ at time t does not help to predict a transition at t.

(ii) The realizations of $\mathbf{x}(t)$ and $\exp(\mathbf{x}_i(t)\beta), i = 1, ..., n$, are *bounded*.

In the present context, it is natural to assume that conditions (i) and (ii) are satisfied. First, consider condition (i). Predictability is satisfied, since our event time scale is discrete and we generally use lagged time-varying covariates (time-invariant covariates are trivially predictable). Next, consider condition (ii). Our time-varying covariates are the dummy variable for the change in antitrust law (AL), the number of bankruptcies (*Bankrupt*), and the external value of the Swiss Franc (*Ext Val*). Clearly, both AL and *Bankrupt* are bounded below and above by definition.[15] Finally, basic economic intuition suggests that *Ext Val* is bounded below and above, too.[16]

We shall therefore apply the standard Cox model below to estimate the impact of the change in antitrust law. With time-varying covariates, the Cox model is given by

$$h_i(t|\,\mathbf{x}_i(t)) = \lim_{dt\to 0} \frac{\Pr\left(T_i \in [t, t+dt] \middle|\, T_i \geq t, \{\mathbf{x}_i(u)\}_0^t\right)}{dt} = h_0(t)\exp(\mathbf{x}_i(t)\beta),$$

where $\{\mathbf{x}_i(u)\}_0^t$ denotes the time path of $\mathbf{x}_i$ up to t, i.e. $\mathbf{x}_i$ is simply replaced by $\mathbf{x}_i(t)$ (see van den Berg 2001, pp. 3397).

4.2 Results

Table 3 provides our estimation results. To interpret these results, observe that we do not report the estimated coefficients $\widehat{\beta}_j, j = 1, ..., p$, but the estimated hazard ratios

$$\widehat{HR}_j = \frac{\widehat{h}(t|\,x_j = x_j^0(t)+1, \mathbf{x}_{-j}(t))}{\widehat{h}(t|\,x_j = x_j^0(t), \mathbf{x}_{-j}(t))} = \exp(\widehat{\beta}_j), \quad j = 1, ..., p,$$

where $\mathbf{x}_{-j}(t) = (x_1(t), ..., x_{j-1}(t), x_{j+1}(t), ..., x_p(t))$. The hazard ratio is the factor by which the hazard function is multiplied if the covariate x_j increases by one unit. That

[15] The dummy variable AL is either zero or one. The number of bankruptcies (*Bankrupt*), in turn, is zero at the minimum; at the maximum, it equals the number of firms in the market.

[16] The minium of *Ext Val* is zero by definition. As to the maximum, observe that for *Ext Val* to go to infinity, the currency values of the most important trade partners (measured in Swiss Francs) would have to approach zero.

is, if $\widehat{HR}_j = 1$, the hazard rate does not change in response to a change in covariate j, whereas the hazard increases (decreases) if $\widehat{HR}_j > 1$ ($\widehat{HR}_j < 1$, respectively).

<Table 3 around here>

The estimated models given in Table 3 differ with respect to sample composition. The left-most column model reports the hazard ratios for the full sample where all firms are pooled. The other models are based on subsamples where only firms with the indicated export shares are included.

Our main interest lies in examining the impact of the change in antitrust law in 1996—represented by the dummy variable *AL*—on hazard rates. The pooled regression suggests that the change in antitrust law produced a significant overall increase of hazard rates of roughly 30%. This significant overall increase of hazard rates appears to be consistent with the idea that the tightening of antitrust law served to lower the degree of collusion in domestic markets. However, to substantiate the claim that the increase of hazard rates was generated by more intense competition—rather than some other exogenous "shock" occurring at the same time—it is necessary to compare the effects of the change in antitrust law on firms that are likely to be affected and firms that are not. As pointed out above, this is done by comparing the impact on non-exporting firms (*Exposh 1*) with the impact on exporting firms (*Exposh 2*, *Exposh 3*, *Exposh 4*) already facing competition on international markets. The estimated hazard ratios for these models show that the non-exporting firms were the only ones significantly affected by the change in antitrust law. They suffered a significant increase of hazard rates of more than 30%. Exporting firms, in contrast, were not significantly affected by the change in antitrust law.[17] Together, these findings suggest that the change in antitrust law indeed raised the intensity of competition in domestic markets, whereas competition in international markets remained largely unaffected. Since the vast majority of firms in our sample are non-exporting firms, the overall impact estimated by the pooled regression is nevertheless large and significant.

The pooled regression further indicates that the firm's *export activity* is an important determinant of the hazard rate in its own right. Intuitively, one would expect that a firm's survival probability should increase the more it diversifies, so that firms with medium export activity tend to have low hazard rates. Our results are consistent with this view, as the hazard ratios for firms exporting up to two thirds of their outputs

[17] The effect on exporting firms is highest for firms with small export shares (*Exposh 2*). It is roughly the same for firms with medium and large export shares (*Exposh 3* and *Exposh 4*). However, the effect is insignificant for all exporting firms. The latter finding still holds when combinations of *Exposh 2*, *Exposh 3* and *Exposh 4* are pooled to enlarge sample size.

(*Exposh 2* and *Exposh 3*) are less than one (if not necessarily significant). Firms that export more than two thirds of their output (*Exposh 4*), however, have the highest hazard rates.

Concerning *size*, the pooled regression shows that both the natural log of the number of employees (*LnEmp*) and its square ($(LnEmp)^2$) are significant. The hazard-reducing effect of *LnEmp* is countered by a hazard-increasing effect of $(LnEmp)^2$. This indicates that size advantages decrease up to an 'optimal size'. Further increases of size lead to an increase of hazard rates, giving rise to a ∪-shaped relationship between the number of employees and hazard rates.[18] This result, which is supported by the estimation results for the other models, is remarkable, as it is commonly accepted that size is positively related to the likelihood of survival.[19] However, many earlier studies have not really addressed the question whether there is a monotone relation between size and survival, using only one size term. Our results add to studies by Wholey et al. (1992), Dunne and Hughes (1994), Ranger-Moore (1997), Harhoff et al. (1998) and Chen (2002), which suggest that the relation between size and survival may be non-monotone at least for some industries. In line with the bulk of the literature, we also find that hazard rates significantly decrease with age (*Age*) (see Carroll (1983), Amburgey et al. (1993), Olzak and West (1991), Mata and Portugal (1994), and Audretsch et al. (2000)).

Our estimates further suggest that *legal form* is an important determinant of firm survival. Stock corporations have lower hazard rates than non-corporate firms. Partnerships (reference variable) are generally most likely to fail, followed by limited liability companies (with a hazard ratio of 0.734) and stock corporations (with a hazard ratio of 0.686). This suggests that the advantages of corporations, such as higher initial capital requirements, better access to financial capital and easier transfer of ownership, apparently dominate their disadvantages due to more risk averse behavior resulting from limited liability. Our results generally support the rankings of those of previous studies, such as King and Wicker (1988) and Harhoff et al. (1998).

We distinguish four main *industry sectors* on the one-digit level: manufacturing, construction, trade and services. It stands out that hazard rates for construction are much higher than those for other industries. Harhoff et al. (1998), in contrast, obtained

[18] We can calculate a "total hazard ratio" function $THR(Emp)$ as the product of the hazard ratios of *LnEmp* and $LnEmp^2$. This function shows a sharp initial drop and reaches its minimum of 0.522 at about 80 employees. That is, the hazard rates of firms with 80 employees are roughly half of those of firms with only one employee. For larger firms, the hazard ratio starts to increase again, but only very moderately. For firms with 250 employees its value is 0.545, for firms with 500 employees it rises to 0.585.

[19] Agarwal and Audretsch (2001, 22) note that "virtually every study undertaken has found that size is positively related to the likelihood of survival".

the lowest hazard rates for construction.[20] A possible explanation for the high hazard rates in our case may be found in falling real estate prices and falling construction expenses during the survey period.[21]

Regarding the *regions* where firms are located, it is noteworthy that hazard rates in the non-German speaking regions are generally significantly higher than in the German speaking regions. More specifically, relative to the German speaking reference region *Eastern CH*, which includes the greater Zurich area, all regions (except German speaking *Central CH*) suffer from significantly higher hazard rates.[22]

Finally, we consider the impact of macroeconomic conditions on hazard rates. Here, we controlled for the external value of the Swiss currency (*Ext Val*), as its fluctuation influences the exits of firms disparately, subject to their export activity, and its omission could seriously distort findings on the change of antitrust law. We find the external value to be significant in all models (with the exception of *Exposh 3*). As expected, an increase in the external value of the Swiss currency raises hazard rates, as it deteriorates the competitiveness of Swiss firms on foreign markets, and increases the competitiveness of foreign firms on domestic markets. Furthermore, we included the number of bankruptcies in the previous year (*Bankrupt*) to control for 'chain effects' associated with the general business climate.[23] For all models, we find the expected result that the number of lagged bankruptcies raises the propensity to exit.

5 Conclusion

The analysis presented in this paper has produced three main results. First, the change in Swiss antitrust law in 1996 led to a significant overall increase of hazard rates for Swiss firms. Second, whereas non-exporting firms suffered from a significant increase of hazard rates, exporting firms were not significantly affected. Third, all other standard determinants of hazard rates generated effects largely in line with the previous literature. Taken together, these results suggest that the tightening of Swiss antitrust law in 1996 led to an increase of the intensity of competition in domestic markets, but did not much

[20] However, for bankruptcies they also obtained the highest hazard rates for construction, in line with results of Kaiser (2004), who uses the database employed in the present paper.

[21] According to figures published by SNB (2003), the price index for appartments fell by 23.31% from 1994 to 1999; other real estate prices also showed significant decreases. For instance, for one-family houses prices dropped by 13.54% and for sales areas by 16.64%.

[22] However, only considering firms with the highest export shares (*Exposh 4*), firms in the French speaking area (including Geneva) have the lowest hazard rates.

[23] An additional business cycle indicator released by the Swiss Institute for Business Cycle Research (KOF-ETH) turned out be insignificant and was thus excluded.

affect the fairly competitive export sector of the Swiss economy.

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Table 1: Descriptive Statistics (source: SBC95, DBED, SNB, own calculations)

Code	Variable Contents	Mean/Value	Std. Dev./Share
	CARDINAL VARIABLES	Mean	Std. Dev.
	Survival		
Duration	Lifetime of the firm (quarters)	censored/truncated	
	Size & Age		
Emp	Number of employees	11.41	53.80
Age	Age of the firm at the SBC95 (quarters)	33.85	28.31
	Macroeconomic conditions		
Ext Val	Swiss Franc's external value (index)	102.45	4.82
Bankrupt	Number of bankruptcies (per quarter)	45.70	4.78
	CATEGORICAL VARIABLES	Value	Share (%)
	Legal form		
Partner	Partnership	Reference var.	13.48
StockCorp	Stock corporation	0(no), 1(yes)	79.33
LimLiab	Limited liability firm	0(no), 1(yes)	7.20
			$\sum = 100$
	Industry		
Manufact	Manufacturing	Reference var.	16.86
Construct	Construction	0(no), 1(yes)	12.17
Trade	Trade	0(no), 1(yes)	29.98
Service	Service	0(no), 1(yes)	40.99
			$\sum = 100$
	Regions		
Eastern CH	Eastern CH, Zürich and Graubünden	Reference var.	31.91
NW CH	Northwestern CH and Bern	0(no), 1(yes)	24.88
Central CH	Central CH	0(no), 1(yes)	10.44
French CH	French CH	0(no), 1(yes)	21.52
Ticino	Ticino	0(no), 1(yes)	11.25
			$\sum = 100$
	Export Shares		
Exposh 1	no export	Reference var.	76.38
Exposh 2	$< \frac{1}{3}$	0(no), 1(yes)	12.02
Exposh 3	$\left[\frac{1}{3}, \frac{2}{3}\right]$	0(no), 1(yes)	4.17
Exposh 4	$> \frac{2}{3}$	0(no), 1(yes)	7.43
			$\sum = 100$
	Antitrust		
AL	After change of regime	0(no), 1(yes)	—

Table 2: Exit rates by Age and Size (source: SBC95, DBED, own calculations)

Exit Rates (Failure Rates %, Merger Rates %) Number of Observations					
	Firm Size in Sept. 1995				
Age	*1-19*	*20-49*	*50-99*	*>100*	*Total*
<2	21.5 (20.9,0.6) 13365	15.0 (13.1,2.0) 512	11.3 (7.8,3.5) 141	10.3 (6.9,3.4) 87	21.1 (20.4,0.7) 14105
2	17.6 (16.9,0.7) 5121	17.4 (13.8,3.6) 224	8.9 (6.3,2.5) 79	14.0 (9.3,4.7) 43	17.4 (16.5,0.9) 5467
3	18.2 (17.4,0.8) 4379	10.6 (8.6,2.0) 245	7.7 (7.7,0.0) 78	20.8 (14.6,6.3) 48	17.6 (16.8,0.9) 4750
4	15.7 (14.9,0.8) 4658	14.4 (11.9,2.5) 277	9.5 (5.4,4.1) 74	9.3 (7.0,2.3) 43	15.5 (14.5,1.0) 5052
5-9	14.8 (14.1,0.7) 16858	9.8 (8.6,1.1) 1261	10.0 (7.9,2.1) 331	7.4 (5.0,2.5) 202	14.3 (13.6,0.8) 18652
10-14	12.8 (12.1,0.6) 10277	10.0 (9.3,0.7) 1012	12.1 (10.2,1.9) 264	8.3 (6.9,1.4) 144	12.5 (11.8,0.7) 11697
15-25	10.9 (10.1,0.7) 12297	8.3 (7.1,1.2) 1683	10.4 (8.4,2.0) 537	6.6 (4.8,1.9) 377	10.4 (9.6,0.8) 14894
Tot.	15.6 (14.9,0.7) 66955	10.5 (9.1,1.4) 5214	10.4 (8.2,2.2) 1504	8.6 (6.0,2.4) 944	15.1 (14.3,0.8) 74617

Table 3: Estimated hazard ratios ($\exp(\widehat{\beta}_j), j = 1, ..., p$)

Variable	**All**	**Exposh 1**	**Exposh 2**	**Exposh 3**	**Exposh 4**
Antitrust					
AL	1.3122***	1.3685***	1.0915	1.0461	1.0486
	(0.0750)	(0.0869)	(0.2050)	(0.3054)	(0.2447)
Export Shares					
Exposh 2	0.8649***				
	(0.0296)				
Exposh 3	0.9652				
	(0.0502)				
Exposh 4	1.0806**				
	(0.0399)				
Size & Age					
LnEmp	0.7426***	0.7296***	0.7300***	0.7924**	0.8050***
	(0.0165)	(0.0184)	(0.0527)	(0.0832)	(0.0631)
$(LnEmp)^2$	1.0346***	1.0434***	1.0333**	1.0190	0.9943
	(0.0053)	(0.0061)	(0.0163)	(0.0224)	(0.0199)
Age	0.9513***	0.9560***	0.9122***	0.9772	0.9240***
	(0.0036)	(0.0040)	(0.0124)	(0.0198)	(0.0137)
Legal Form					
Stock Corp	0.6860***	0.6858***	0.6004***	0.8119	0.7923
	(0.0179)	(0.0191)	(0.0604)	(0.1608)	(0.1245)
Lim Liab	0.7345***	0.7355***	0.6980**	1.0102	0.6466**
	(0.0300)	(0.0329)	(0.1001)	(0.2538)	(0.1319)
Industry					
Construct	1.2382***	1.1728***	1.8487***	1.9831***	2.4552***
	(0.0476)	(0.0491)	(0.2989)	(0.5686)	(0.7706)
Trade	1.0959***	1.0733*	1.1177	1.1087	1.0686
	(0.0343)	(0.0395)	(0.0967)	(0.1571)	(0.1181)
Services	0.9562	0.9191*	1.0280	0.8598	1.1315

Table 3 – Estimated hazard ratios (continued)

Variable	**All**	**Exposh 1**	**Exposh 2**	**Exposh 3**	**Exposh 4**
	(0.0293)	(0.0328)	(0.0910)	(0.1141)	(0.1197)
Region					
NW CH	1.0681**	1.0687**	1.1914**	0.8651	1.0731
	(0.0286)	(0.0324)	(0.0947)	(0.1176)	(0.1118)
Central CH	1.0190	1.0506	1.0517	0.7574	0.8485
	(0.0369)	(0.0435)	(0.1199)	(0.1412)	(0.0981)
French CH	1.1947***	1.2356***	1.2449**	1.1418	0.8126**
	(0.0309)	(0.0359)	(0.1093)	(0.1517)	(0.0742)
Ticino	1.2192***	1.2116***	1.2792*	1.1506	1.1639
	(0.0488)	(0.0554)	(0.1862)	(0.2278)	(0.1394)
Macroeconomic Conditions					
Ext Val	1.0605***	1.0589***	1.1127***	1.0349	1.0547*
	(0.0069)	(0.0077)	(0.0257)	(0.0329)	(0.0270)
Bankrupt	1.0651***	1.0643***	1.0792***	1.0683**	1.0705***
	(0.0057)	(0.0064)	(0.0195)	(0.0287)	(0.0220)
Test Statistics					
χ^2					
All variables (18)	1982.08***	1447.06***	268.76***	68.42***	233.5***
Legal Form (2)	208.37***	184.32***	25.87***	2.44	4.57
Industry (3)	78.77***	65.02***	15.89***	11.39***	8.83**
Region (4)	62.39***	62.22***	9.33*	6.66	13.84***
Export (3)	25.29***				
Log Likelihood	−97317.1	−74157.3	−7289.4	−2491.1	−5700.2
Number of Obs.	68681	52463	8250	2869	5100

Figures in parentheses are standard errors.

*, **, *** Coefficients are significant at the 10%, 5% and 1% level, respectively.

Dummy Coding of Categorical/Ordinal Variables

Legal Form: *Partnership* (ref. var.), *Stock Corp*, *Lim Liab*;

Industry: *Manufact* (ref. var.), *Construct*, *Trade*, *Services*;

Region: *Eastern CH* (ref. var.), *NW CH*, *Central CH*, *French CH*, *Ticino*;

Export: *Exposh 1* (ref. var.), *Exposh 2*, *Exposh 3*, *Exposh 4*.

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