KATHOLIEKE
UNIVERSITEIT
LEUVEN

DEPARTEMENT TOEGEPASTE ECONOMISCHE WETENSCHAPPEN

RESEARCH REPORT 0014

DEFINING FIRM EXIT: THE IMPACT OF SIZE AND AGE REVISITED

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D/2000/2376/14

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March 2000

This research was supported by an FWO grant (G.0149.97) and an OT grant of the Katholieke Universiteit Leuven (OT/96/3). The authors are indebted for the useful comments from David Audretsch, Peter Campbell III, Marnik Dekimpe, Patrick Van Cayseele, Paul Walsh, Leo Sleuwaegen, Nancy Huyghebaert and participants at seminars at Dartmouth College, Trinity College Dublin and Katholieke Universiteit Leuven.

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Abstract

This paper demonstrates that (1) the negative relationship between initial firm size and failure probability, and (2) the aging pattern of the failure rate are sensitive to the adopted definitions of entry and exit. We use two definitions to measure the timing of entry and exit: an economic definition, based on employment levels, and a legal definition, based on the firm's legal status. While initial size is negatively related to the exit rate under the economic definition, the relation becomes positive under the legal definition. The aging effect is much steeper under the legal than under the economic definition.

JEL classifications: G33, L10, C41

I. Introduction

The increased availability of firm level data sets in recent years has led to a series of papers that investigate the post-entry performance of new firms (e.g. Agarwal and Audretsch, 1998; Audretsch and Mahmood, 1995; Mata and Portugal, 1994). One of the "stylised facts" that comes out of this literature is that start-up size is negatively related to the probability of firm exit.[1] The explanation commonly given in the literature (e.g. Caves, 1998) for this finding is intuitively appealing. Firms enter at different sizes, which suggests they have different expectations and options about their costs or product markets. Firms with more positive expectations make larger initial commitments than firms with more negative expectations. The latter, however, incur a larger initial unit cost by limiting their entry size. Therefore, we may expect that small entrants are more likely to fail than large ones. A number of theories that investigate firm dynamics can generate such a negative relationship between start-up size and the probability of exit. For example, the learning models of Jovanovic (1982) and Pakes and Erikson (1989) model the growth process as a stochastic process where firms learn their true cost structure over time. These learning models establish a positive correlation between efficiency and size, which is the driving force behind the empirical predictions on growth, survival and size.

The main contribution of this paper lies in demonstrating that the aforementioned empirical "stylised fact" of the negative relationship between initial size and the probability of failure depends very much on the way in which firm exit is defined.[2] We use both an *economic* and *legal* definition of firm entry and exit, and find that the impact of start-up size on firm exit is highly sensitive to the definition that is used. Furthermore, our results show that this choice of definition also affects the resulting aging patterns associated with the exit probability, as well as the impact of other often-included cross-sectional control variables.

The first definition of exit that we consider, based on the firm's reported employment figures, has been used extensively in the industrial organisation literature that studies firm growth and survival. Under this definition, firms disappear from the

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[1] For overviews see Audretsch (1995), Caves (1998), and Sutton (1997).

[2] Recently, Audretsch et al. (1997) find that the effect of initial size on firm survival differs also between the service and manufacturing sectors. In particular, they find that firms with a start-up size of

data set, i.e., exit, when employment falls to zero and does not become positive again during the observation period.[3] We call this the *economic* definition of exit, which has also been adopted by Audretsch, Houweling and Thurik (1997), Audretsch, Klomp and Thurik (1997), Audretsch and Mahmood (1995), Mata and Portugal (1994), and Wagner (1994), among others. The reasoning behind this operationalisation is that the firm ceases to engage in economically meaningful activities when it must operate without employees. Empirical studies in the accounting and finance literature, in contrast, tend to use the *legal* dates of failure to determine the timing of exit. For example, either the bankruptcy (chapter 11) filing date, the liquidation (chapter 7) date, or the last date for which financial data was published were used as the exit date in Altman (1968), Casey and Bartczak (1985), Gentry, Newbold and Whitford (1985,1987), Platt and Platt (1990) and Zavgren (1985), among others. This is the second definition of exit that we consider.

The timing of entry is usually more arbitrary and less clearly defined in the empirical literature. Many studies give no details about how entry was determined, and those that do are characterised by a lack of uniformity. In the entrepreneurship literature, for example, the timing of entry has been established by the date on which firms have first opened a bank account (Cressy, 1996) or through questionnaires (Coopers, Gimeno-Gason and Woo, 1994). Most papers in industrial organisation either use official registries (Audretsch, 1995b) or define an entry as a firm that appears with positive employment for the first time in the data set (e.g. Audretsch, Klomp and Thurik, 1997; Mata and Portugal, 1994).

In this paper, we illustrate the importance of entry and exit definitions on the obtained empirical results regarding the exit behaviour of firms. Section II describes the data. In section III we discuss the empirical implications of the two definitions. Section IV concludes our study.

II. Sample description

Our data cover 14,584 manufacturing firms located in the Northern part of Belgium, Flanders, that started operations between 1985 and 1994, and reported positive employment before 12/1994. The data are retrieved from the registry of the

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less than 5 employees are less likely to survive, but this negative relationship between initial size and exit becomes insignificant for service firms that enter with more than 5 employees.

[3] Note that a firm with an observed employment pattern 2-2-0-1-2-0 has an economic exit in year 6 only.

Central Bank, where all firms are legally required to submit their company accounts. Hence, we cover virtually the entire population of manufacturing firms that report positive employment within our observation period.[4] Each of these firms was tracked through 12/1994 to determine whether they failed within this period. This data set allows us to analyse all cohorts of new entrants over this period, rather than focusing only on the cohort say of 1985, which is a common, yet more restrictive, procedure in firm survival studies.

In our data set, economic entry occurs in the year when positive employment is first reported. The legal entry year is defined as the year in which the firm received its VAT number. For each of the firms in our dataset, both economic and legal entry were observed. Economic exit occurs in the year when a firm no longer reports positive employment, while legal exit occurs when the courts declare that the firm is bankrupt or must be liquidated. Hence, we use the same set of firms to assess the potentially confounding impact of entry and exit definitions on some stylised empirical facts in the firm-survival literature.

A first indication of the importance of the definition adopted become apparent from some simple summary statistics. Under the economic definition, 20.43% of the firms exit, while only 8.79% of the firms exit under the legal definition. Hence, the economic definition significantly exceeds the legal exit. Only 8.69% of the firms exit under both definitions. While 98.8% of the firms with legal exit also experience economic exit, 42.5% of firms with economic exit have no legal exit. Economic exit tends to occur either before or in the same year as legal exit: 81.3% of all exits experience legal exit in the same year or the year after economic exit. With respect to the entry definitions, we find that economic and legal entry occur within 2 years of each other for 92.6% of the observations. Economic entry never occurs before legal entry, and 50.3% of the observations have the same entry year under both definitions. These figures therefore suggest that a substantial number of firms operate with zero employees in the first or last years of their existence. An explanation for this finding might be found in the use of interim labour and outsourcing. We explore this issue further in the next section where we analyse whether the difference between the legal

[4] The dataset at our disposal consists of all Belgian firms that are subject to the Belgian accounting law, which obliges them to publish annual accounts. All incorporated enterprises that carry on a commercial activity, whether or not they are listed on a stock exchange, publish full or abbreviated accounts. This means that even small companies, with some minor exceptions, are subject to the accounting law concerning the preparation and publication of accounts. Belgium has indeed gone

and economic definition of exit affects the empirical relationship between firm size and the exit probability.

III. Empirical results

We investigate the exit rate of new firms using a hazard approach. We start by estimating a Cox (1972) model with a parametric Weibull specification for the aging pattern. Next, we examine the robustness of our findings: first, we consider whether the assumption of a parametric time dependence affects our results, and then we examine whether the presence of unobserved firm heterogeneity has a substantive impact on our findings. The methodology is explained in detail in the appendix.

We follow the literature, and assume that the firm level hazard is a function of firm specific variables, sector variables and aggregate economic conditions, all measured during the start-up year. Firm start-up size is measured by the initial employment level, as is common in this literature (e.g., Audretsch and Mahmood, 1995, and Mata and Portugal, 1994). The sector variables, included as control variables, are (1) the minimum efficient scale (MES) proxied by the median employment for the industry at the three digit level of the NACE classification, (2) the capital intensity of the industry proxied by total tangible fixed assets[5] of the industry normalised by industry sales, (3) the herfindahl index of concentration, and (4) an indicator of the homogeneity of the sector. The latter is a dummy variable equal to 1 if the sector is neither advertising nor R&D intensive, and zero otherwise. The source for this variable, which proxies for the absence of endogenous sunk costs in the sector, is Davies and Lyons (1996), who classify European sectors in R&D intensive, advertising intensive and homogeneous sectors.

Firms starting in different years may face different aggregate economic conditions during their start-up year. We control for this difference through the corresponding growth in real GDP. We also take into account a potential aging effect. As in Dekimpe and Morrison (1991), who study the survival of retail firms, our first model assumes a Weibull distribution for the aging effect. A more flexible distribution is adopted later in the paper.

much further than was required under the terms of the European Union's Fourth Directive. Financial intermediaries and hospitals are subject to special rules, and are therefore excluded from the sample.

[5] Tangible fixed assets are included in the balance sheet at their historic cost after depreciation. However, they can be revalued at regular intervals. If this is the case, the depreciation charge is based on the revalued amount.

Table 1 gives the results of the Weibull model. Model A uses the economic definitions of entry and exit, while model B uses the legal definition. Our findings indicate that the choice of definitions affects several substantive conclusions about the exit behaviour of firms: (1) the aging pattern of the exit rate, (2) the relationship between firm size and exit, and (3) the impact of the control variables. [6]

First, and conform to the summary statistics mentioned before, we find that the exit rate of the base group[7] in the year of entry is considerably lower under the legal definition (0.23%) compared to the economic definition (2.18%). Second, even though both aging patterns show a positive slope, we find the increase with age to be more slowly under the economic definition (*b*=0.6775) than under the legal definition (*b*=1.3091). Neither result is surprising. First, the lower exit rate under the legal definition can be attributed to the fact that firms may continue to operate, albeit under financial difficulties, without employees to postpone bankruptcy. For instance, firms may replace employees with temporary labour (which is not included in the number of employees reported) or outsource some of the activities to external suppliers, and thereby postpone or avoid legal exit while still being classified as an economic exit. As for the sharper increase under the legal definition after the first few years, this may be attributed to the time-consuming nature of the legal proceedings of bankruptcy. Weiss (1990), for example, reports an average time from filing the bankruptcy petition to resolution of 2.5 years in the U.S., with a standard deviation of 1.4 years. For Belgian firms, the time from filing to resolution can take up to 5 years.[8] As a consequence, one may not observe many legal exits in the first few year(s), even when firms file for bankruptcy during the first months after entry, but as time progresses, more and more cases have moved through the legal system and reach actual legal exit.

<<INSERT TABLE 1 HERE>>

Third, also the impact of start-up size varies with the definition being used. Model A reports a negative and significant impact, which is consistent with the

[6] To determine whether the difference in the results between models A and B is due to the differences in the entry or differences in the exit definitions, we also estimated a model using the legal entry definition with the economic exit definition. The results are very similar to model A, suggesting that the exit definition is the driving force.

[7] The base exit rate is obtained when all explanatory variables are zero.

aforementioned "stylised" empirical fact. However, model B, which uses the legal definitions, reports a positive and significant effect. Even though this result may look surprising at first, there is an intuitively appealing explanation for this finding.

Firms that face financial difficulties will try to realise cost savings. These may be more easily realised for flexible resources than for fixed (committed) resources (Kaplan and Cooper, 1998). Typical examples of flexible resources, which can be acquired on a needs basis from external suppliers, are interim (temporary) labour and subcontracting. A reduction in the use of these resources generates an immediate cost saving. With respect to committed resources, on the other hand, the firm acquires the capacity before it is actually used. These costs are fixed, and a typical example is fixed assets. Employees who are 'permanent' also represent a fixed cost, especially when firms face a strict employment protection legislation, as is the case in Belgium and many other European countries.[9] Firms facing liquidity problems have an incentive to replace committed resources by flexible resources; i.e., to replace employees with interim labour and replace fixed assets with subcontracting. These actions will reduce unused capacity, thereby reducing total costs for the firm, and in turn postponing legal exit.

How is this related to our discussion on the impact of the size factor on exit behaviour? Two issues are important. First, since interim labour is not included in the number of employees reported in the financial statements, a switch to interim labour is useful in postponing legal exit, but not economic exit. Hence, the motivations for the negative relationship between size and exit behaviour as discussed in Section 1 can explain economic exit while these same motivations need to be augmented with the flexible resources argument to explain legal exits. Second, especially smaller entrants are expected to benefit from interim labour and subcontracting. Small firms not only have fewer alternatives to use excess capacity in employees or fixed assets, but they also face less stringent social protection regulations than large firms. The latter may allow them to be faster in responding to shocks (less x-inefficiency). Put differently, smaller firms are more apt at using interim labour to postpone legal exit than large firms, which may explain the observed sign switch.

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[8] Belgian National Institute for Statistics.
[9] The implied lower flexibility in terms of job turnover is clearly demonstrated in labour economics (e.g. Garibaldi, Konings and Pissarides, 1997).

While intuitively appealing, this explanation can only have face validity when subcontracting and interim labour are indeed significant factors. With respect to subcontracting, Chalos (1995) states that functions and services that have traditionally been performed in-house are increasingly being obtained from outside vendors. Furthermore, it is interesting to see that interim labour has witnessed an enormous growth in Belgium and other European countries over recent years. In particular, between 1985 and 1998, there was a five-fold increase in the number of hours worked by interim labour, and a more than five-fold increase in the number of people that worked under interim contracts in Belgium (Peeters, 1999).

Finally, the signs and significance of the coefficients of the control variables also vary with the definition used. Under the economic definition, the sector variables homogeneity, capital intensity and concentration index have a highly significant impact on exit rate. Under the legal exit definitions, however, new firms have significantly better survival chances in industries with a low MES while the other sector characteristics do not have a significant impact. The impact of economic conditions is as expected under both definitions: the more favourable the economic conditions during the start-up phase, the lower the exit rate. Not surprisingly, the impact of economic conditions is larger under the economic definition.

Table 2 adds the impact of average industry wages per employee in the entry year to the set of explanatory variables. This variable indicates whether the firm operates in a high labour cost industry. In terms of the interpretation offered above, the variable indicates to what extent the industry is characterised by high or low fixed costs. As expected, the addition of this variable does not affect our results for the legal definition, as interim labour offers a way around high fixed labour costs. For the economic definition, in contrast, we expect that high wage costs deter survival, and thus contribute to the firm's economic exit. We indeed find a positive and highly significant effect in model A of Table 2. This effect dominates the size effect, which becomes insignificant (yet stays negative).

<<INSERT TABLE 2 HERE>>

IV. Robustness checks

Two of our main conclusions deal with the impact of the chosen definition on (1) the aging pattern, and (2) the sign reversal for the size variable. In the following section, we consider two alternative explanations for these phenomena. Specifically, we examine whether they are "model-induced" by the specific choice of the parametric baseline hazard (Weibull) and/or the absence of corrections for unobserved heterogeneity. Both issues have been shown to potentially affect the inferences on the time dependence of exit behaviour and the consistency of the cross-sectional explanatory variables.

IV.a. Aging pattern

So far, the time dependence of the exit rate has been modelled with the Weibull distribution. However, the true time dependence may be neither monotonically increasing nor monotonically decreasing, which are the two forms allowed under the Weibull distribution. Agarwal and Gort (1996), for example, find that hazard rates first increase, and then eventually fall. They explain this pattern through the industry life cycle. In Table 3, we model the time dependence semi-parametrically. In the absence of a theoretical justification for choosing a particular distribution, this approach is preferred over imposing a particular (potentially incorrect) distribution, which would result in inconsistent parameter estimates. The semi-parametric approach we adopt (see McDonald and Van de Gucht, 1999 and the appendix for details), on the other hand, results in consistent estimates even when the true form of the underlying baseline hazard is unknown, as shown in Meyer (1986,1990). Under this approach, the aging effect is measured as a piecewise approximation of an underlying, possibly very complex, continuous time-dependence pattern. The initial exit rate gives the exit rate of the base group in the first period. Positive (negative) jumps in the subsequent years indicate a higher (lower) exit rate compared to the first period. Given the small number of exits in the first year (only 2 legal exits and no economic exits), the first period in our estimations consists of the first 2 years after entry, with an imposed equal annual exit rate for years 1 and 2.

<<INSERT TABLE 3 HERE>>

With the semi-parametric specification in Table 3, the initial exit rate under the economic definition (model A) increases to 3.19%, which again is higher than the initial exit rate under the legal definition (0.17% in model B). As with the Weibull models, the aging pattern is upward sloping for both definitions: the annual conditional exit rate in years 3 through 10 is always significantly higher than in the initial years, as is apparent from the positive and significant jumps in the exit rate. The two time patterns behave differently, however, as is illustrated in Figure 1. The annual exit rate is consistently higher under the economic definition, where the exit rate significantly increases in year 3, then seems relatively stable through year 6, after which it increases again, with a decrease in year 9 (ignoring year 10, which is based on only a small number of observations). Under the legal definition, on the other hand, the exit rate increases through year 5 after which it seems to stabilise with a small upward jump in year 9.

<<INSERT FIGURE 1 HERE>>

The signs and significance of the other coefficients are mainly unaffected by how the time dependence is modelled. Start-up size remains negative (-0.4089) under the economic definition, and positive (2.5459) and significant under the legal definition. The signs and significance of the industry control variables and economic conditions are also similar to those reported for the Weibull model.

IV.b. Unobserved heterogeneity

The specification used thus far may not fully account for all the heterogeneity in the exit rates. Indeed, explicit control for unobserved heterogeneity is necessary to avoid incorrect inferences as to the true time-dependency of the exit rate, and to avoid inconsistent estimates for the other parameters (which would be an alternative explanation for the sign switch of the size variable). The spurious influences that can be made when this is not done are illustrated when assuming a situation where every firm has a *constant* exit probability over time, but some firms have a high probability while others have a low probability. In such a case the firms with a high exit probability will likely exit after a few periods, leaving a higher proportion of firms with low exit probabilities. Thus, fewer exits will be observed as time passes giving the incorrect impression that exit probabilities are *decreasing* over time. Not accounting for unobserved heterogeneity may not only yield a spurious negative duration dependence, but may also give inconsistent parameter estimates (Heckman and Singer (1984a, 1984b), and Manton, Singer and Woodbury (1992)). We follow the approach of Dekimpe et al. (1998) and Schmittlein and Morrison (1983) and let the initial exit rate vary across firms according to a certain distribution, namely the gamma distribution. The details of this approach can be found in the appendix.

Table 4 reports the results. Under the economic definition, accounting for unobserved heterogeneity affects neither the estimated time dependence nor the impact of the explanatory variables (compare model A of Table 4 with the same model in Table 3). The coefficient of variation, which measures the degree of heterogeneity of the initial exit rate across firms, indicates, however, that the unobserved heterogeneity is much larger under the legal definition (2.9151 compared to 0.0026 in model A). With respect to the time dependence under the legal definition, this results in a much steeper aging pattern, while the initial exit rate (0.16%) remains unaffected.

<<INSERT TABLE 4 HERE>>

Figure 2 illustrates the time dependence when accounting for unobserved heterogeneity under both definitions. As was apparent from Table 4, accounting for unobserved heterogeneity has little impact on the time dependence under the economic definition. However, ignoring unobserved heterogeneity when it is high, as is the case under the legal definition, results in an understatement of the time dependence, resulting in a lower annual conditional exit rate. This figure illustrates once more that the definitions have a substantial impact on the conclusions that are drawn about the time dependence. Under the economic definition, the time dependence appears much more stable: the annual exit rate ranges from 3.19% in year 1 to 6.33% in year 8 (ignoring year 10). Under the legal definition, the time dependence is much steeper, ranging from 0.16% in year 1 to 13.88% in year 9 (again ignoring year 10). Here, the conditional exit probability rises each year.

<<INSERT FIGURE 2 HERE>>

Accounting for unobserved heterogeneity also affects some of the control variables under the legal definition. For example, MES becomes insignificant, and the impact of macroeconomic conditions (-9.3240) is now comparable in size to the economic definition (-9.8622). Yet, our substantive conclusion of a positive firm-size effect still holds.

V. Conclusions

This paper illustrates the sensitivity of empirical findings about firm exit behaviour to the measurement of entry and exit. For a dataset of 14,584 entrants we identify both the *economic* entry and exit, as well as the *legal* entry and exit. The economic definition is based on the number of employees reported by the firm; the firm's legal life starts when its VAT number is assigned and ends with the pronouncement of bankruptcy or liquidation.

Our results show that these definitions matter: we find significant differences in both the time pattern of the exit probability and the impact of cross-sectional variables. Firms are more likely to exit under the economic definition, but the age of the firm has greater impact under the legal definition. Our results show furthermore that the negative relation between start-up size and exit probability holds only under the economic definition. One potential explanation that we offered for the observed positive relation under the legal definition was related to firms switching from fixed to more flexible resources such as interim labour and outsourcing to avoid or delay bankruptcy or liquidation.

Which definition, the economic or the legal, is most relevant depends upon the questions one attempts to answer. The economic definition may be more appropriate when researching topics related to industrial organisation such as the impact of changing competitive conditions. The legal definition is more appropriate when examining questions related to start-up financing, cost of capital and default risk of financial instruments. An analysis of the time span between economic and legal exit may provide insights into the ability of firms to exchange fixed for flexible resources. While both definitions obviously have their value, our findings clearly warn applied researchers about the potential pitfalls of generalising empirical findings from one research stream to another research tradition. Indeed, some 'stylised' facts may be highly dependent upon the operationalisation of entry and exit definitions.

Our discussion has also highlighted the role of temporary labour and outsourcing as a potential explanation for the difference between economic and legal exit, at least in a European context. Whether our findings also hold for other markets, like the US, needs to be investigated with similar data sets.

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Table 1: The Weibull model

	A Economic definition	B Legal definition
Aging pattern:		
initial exit rate	0.0218	0.0023
b	0.6775***	1.3091***
Start-up size:		
Size	-0.4762*	2.5315***
Industry variables:		
MES	-0.0099	0.3252**
Homogenous	-0.1423***	-0.0265
Capital intensity	1.5138***	0.4271
Concentration	0.6638**	-0.5082
Economic conditions:		
GDP	-12.8171***	-3.5972**
LL	-11843.983	-6189.702

***, **, * indicate significance at the 1%, 5% and 10%, respectively, using 2-sided t-tests.

Table 2: The Weibull model: impact of industry wages

	A Economic definition	B Legal definition
Aging pattern:		
initial exit rate	0.0196	0.0022
b	0.6850***	1.3110***
Start-up size:		
Size	-0.4634	2.5370***
Industry variables:		
MES	-0.1605	0.3245**
Homogenous	-0.1300***	-0.0214
Capital intensity	1.4888***	0.4116
Concentration	0.5607*	-0.5451
Wages	0.0848***	0.0307
Economic conditions:		
GDP	-12.6241***	-3.5218**
LL	-11838.443	-6189.430

***,**,* indicate significance at the 1%, 5% and 10%, respectively, using 2-sided t-tests.

Table 3: Hazard model with semi-parametric aging effect

	A Economic definition	B Legal definition
Aging pattern:		
initial exit rate	0.0319	0.0017
jump in year 3	0.5133***	2.0727***
jump in year 4	0.4576***	2.4329***
jump in year 5	0.5485***	2.6529***
jump in year 6	0.5390***	2.6811***
jump in year 7	0.6693***	2.6957***
jump in year 8	0.6841***	2.7801***
jump in year 9	0.6062***	2.8773***
jump in year 10+	1.005***	2.7094***
Start-up size:		
Size	-0.4089	2.5459***
Industry variables:		
MES	0.0057	0.3197**
Homogenous	-0.0815*	0.0126
Capital intensity	0.9566***	0.0081
Concentration	0.5646**	-0.4874
Wages	0.0655***	0.0229
Economic conditions:		
GDP	-9.8638***	-3.4481***
LL	-12003.287	-6137.314

***,**,* indicate significance at the 1%, 5% and 10%, respectively, using 2-sided t-tests.

Table 4: Semi-parametric model: accounting for unobserved heterogeneity

	A Economic definition	B Legal definition
Aging pattern:		
initial exit rate	0.0319	0.0016
jump in year 3	0.5134***	2.1871***
jump in year 4	0.4577***	2.7379***
jump in year 5	0.5486***	3.1894***
jump in year 6	0.5390***	3.4684***
jump in year 7	0.6694***	3.7257***
jump in year 8	0.6842***	4.0652***
jump in year 9	0.6062***	4.4474***
jump in year 10	1.0049***	4.5689***
Start-up size:		
Size	-0.4091	4.6959***
Industry variables:		
MES	0.0055	0.7018
Homogenous	-0.0815*	0.0426
Capital intensity	0.9571***	0.0015
Concentration	0.5646*	-0.7237
Wages	0.0654***	0.0319
Economic conditions:		
GDP	-9.8622***	-9.3240***
CV	0.0026	2.9151
LL	-12003.288	-6131.084

***,**,* indicate significance at the 1%, 5% and 10%, respectively, using 2-sided t-tests.

Figure 1: Semi-parametric model: time dependence of the exit rate

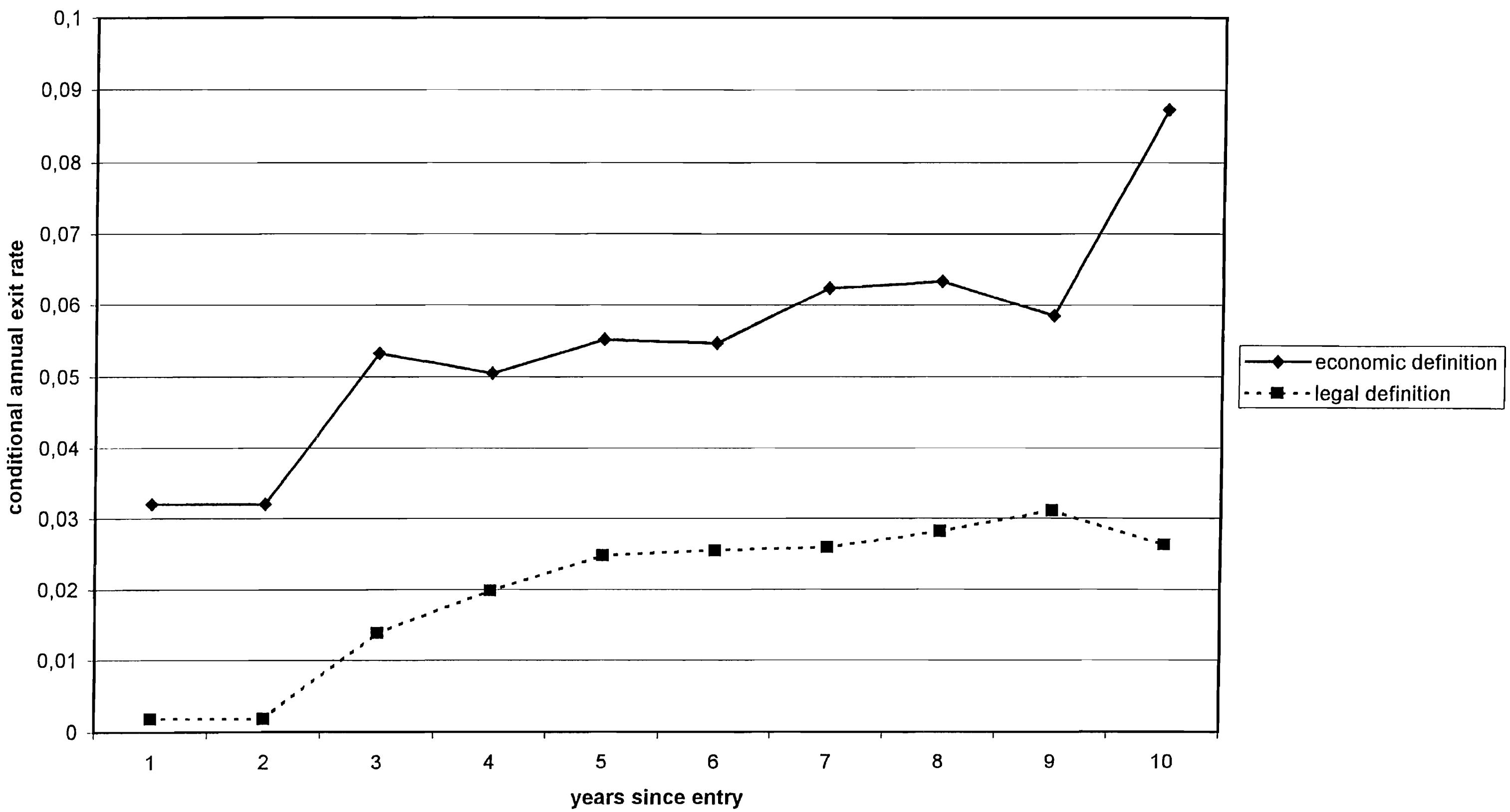
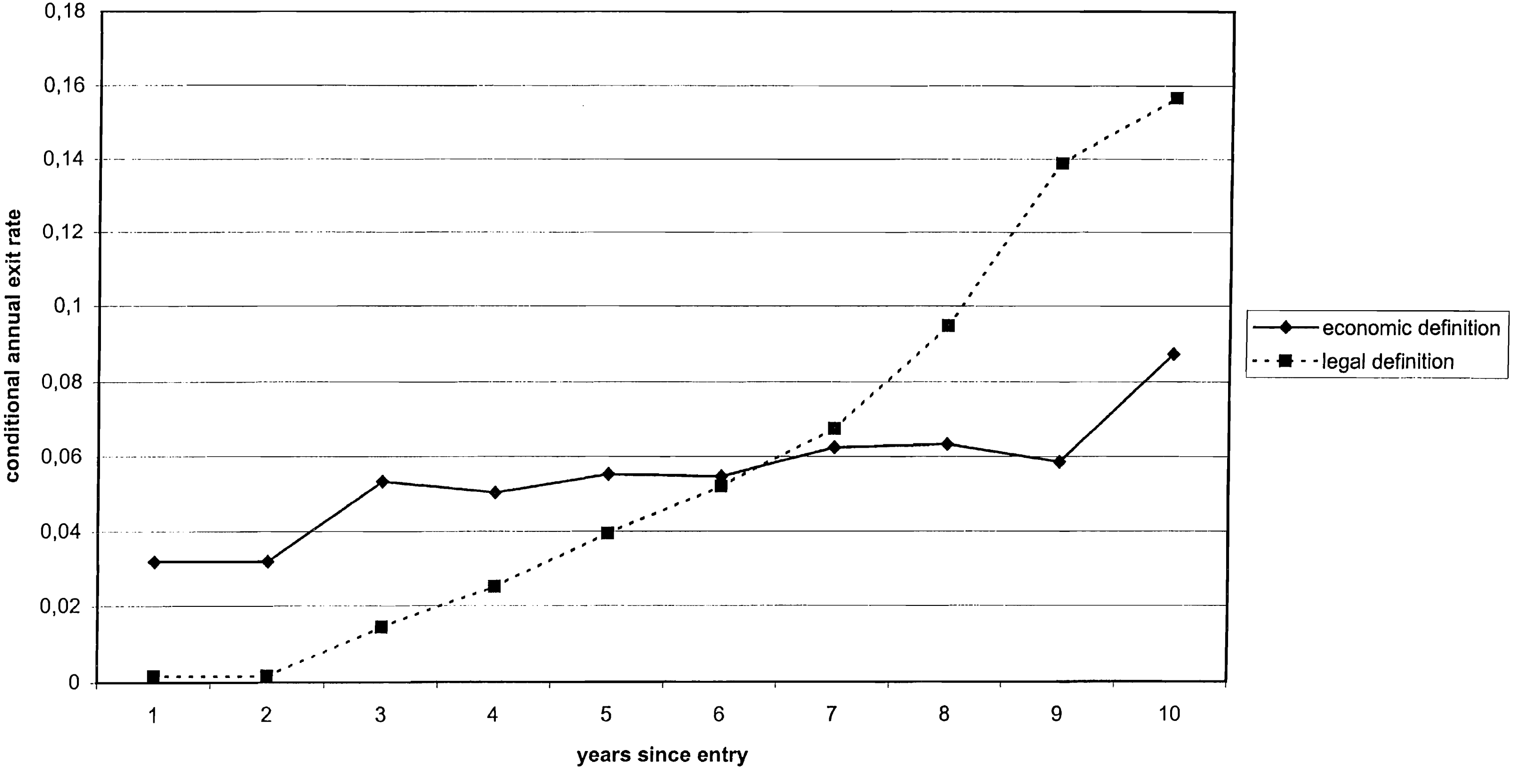


Figure 2: Semi-parametric model: time dependence of the exit rate when accounting for unobserved heterogeneity



Appendix: Methodology

Let the random variable T denote the time between firm entry and exit (failure), with associated probability density function $f(t)$ and cumulative density function $F(t)$. Then,

$$h(t) = \frac{f(t)}{1 - F(t)} \qquad (1)$$

is the hazard, i.e. , the rate at which firms exit during period t given that they have not done so in the previous t-1 periods since entry. The survival function $S(t)=1-F(t)$ denotes the probability that exit does not occur for at least t periods. To model the hazard rate as a function of covariates, we use the Cox (1972) formulation and let

$$h(t) = h_0 \, e^{\beta X}, \qquad (2)$$

where X is a vector of explanatory variables, ß is a vector of parameters, and h_0 is the exit rate of the base group (referred to as the 'initial exit rate' in our results). The base group consists of those firms for which all explanatory variables equal zero. A positive ß-coefficient implies that a positive value of the associated variable augments the exit rate.

When T follows a Weibull distribution, the exit rate varies as a power of t and the hazard rate becomes

$$h(t) = h_0 \, e^{b \ln(t)} \, e^{\beta X}, \qquad (3)$$

where b is the shape parameter. When b>0 (b<0) the hazard increases (decreases) with time. When b=0, the hazard rate is constant (and the exponential model is obtained).

The semi-parametric approach involves adding a vector of time-varying dummy variables, $D(t)$, to the model:[1]

$$h(t) = h_0 \, e^{\beta X} \, e^{c D(t)}, \qquad (4)$$

where c is a vector of coefficients. A separate variable is used for each period; for example, D(3) takes on the values (0 0 1 0 0 ...).[2] h_0 then gives the exit rate of the base group in the first year. Positive (negative) c-coefficients indicate a higher (lower) exit rate compared to the first period (these coefficients are referred to as the 'jumps'

[1] Examples of applications of this semi-parametric approach include McDonald and Van de Gucht (1999), Dekimpe, Van de Gucht, Hanssens and Powers (1998) and Han and Hausman (1990).
[2] To avoid identification problems when simultaneously estimating c_1 and h_0, no separate indicator variable is added for the first period.

in our results). To estimate the parameters h_0, c and ß, maximum likelihood estimation is used. The contribution of the i-th firm to the likelihood function is given by the following expression :

$$L_i(t_i|h_0) = [S_i(t_i - 1|h_0) - S_i(t_i|h_0)]^{1-d_i} [S_i(t_i - 1|h_0)]^{d_i} \tag{5}$$

where t_i is the number of observed periods until exit or censoring, and d_i is an indicator variable equal to one for censored observations (survivors) and zero for completed observations (firms that exit). The expression for the hazard function given in equation (4) can be substituted into equation (5) through the following general relationship (Lancaster, 1990; Ross, 1980):

$$S_i(t_i) = e^{-\theta_i(t_i)} \text{ where } \theta_i(t_i) = \int_0^{t_i} h_i(u)du \tag{6}$$

To account for unobserved heterogeneity we follow Schmittlein and Morrison (1983) by weighting the conditional likelihood by the relative occurrence of its h_0 value by means of a gamma mixing distribution $g(h_0)$, in which case the unconditional likelihood becomes

$$L_i(t_i) = \int_0^{\infty} L_i(t_i|h_0)\, g(h_0)\, dh_0 \tag{7}$$

The log-likelihood function for N firms is then

$$LL = \sum_{i=1}^{N} \ln\left\{ (1 + d_i) \left[\frac{a}{B_i(t_i - 1) + a} \right]^r - \left[\frac{a}{B_i(t_i - 1) + (1 - d_i) e^{\beta X_i + cD_i(t_i)} + a} \right]^r \right\} \tag{8}$$

$$\text{where } B_i(t) = \sum_{j=1}^{t} e^{\beta X_i(j) + cD_i(j)} \tag{9}$$

and a and r are the location and shape parameters, respectively, of the gamma distribution. The mean of the gamma distribution is given by $E(h_0)=r/a$. The coefficient estimates of the β and c vectors should then be interpreted relative to r/a. The coefficient of variation, CV, can be interpreted as an indication of the degree of heterogeneity of h_0 across firms:

$$CV = \frac{\sqrt{\text{var}(h_0)}}{E(h_0)} = r^{-1/2} \tag{10}$$

A low r-value indicates a large amount of heterogeneity, whereas high values are obtained for relatively homogeneous populations.