

Export-market dynamics and the probability of firm closure:

Evidence for the UK¹

By

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Abstract

This study presents the first empirical analysis of the determinants of firm closure in the UK with an emphasis on the role of export-market dynamics, using panel data for a nationally representative group of firms operating in all-market based sectors during 1997-2003. Our findings show that the probability of closure is (*cet. par.*) significantly lower for exporters, particularly those experiencing export-market entry and exit. Having controlled for other attributes associated with productivity (such as size and export status), the following factors are found to increase the firm's survival prospects: higher capital intensity and TFP, foreign ownership, young age, displacement effects (through relatively high rates of entry of firms in each industry), and belonging to certain industries. Interestingly, increased import penetration (a proxy for lower trade costs) leads to a lower hazard rate for exporting entrants and continuous exporters, whilst inducing a higher hazard rate for domestic producers or those that quit exporting.

1. Introduction

Substantial evidence of the benefits from international trade has been well documented in both theoretical and empirical literature, which often provides governments in various countries with a rationale for intervention to help firms develop their exporting activities when market failures are present. These benefits are largely linked to the higher productivity of exporters, which then contribute to overall UK productivity growth through various channels, such as the entry of higher productivity exporters (including the so-called 'born global' companies; see Oviatt and McDougal, 1995); existing exporters becoming more productive over time and/or intra-industry resources are reallocated to higher productivity exporters; and the shutdown of lower productivity firms - especially non-exporters with the lowest productivity levels, as predicted by some recent theoretical models (Bernard *et. al.*, 2003; Melitz, 2003).

Thus it is not surprising that many government programmes aimed at export promotion help firms enter export markets and, given this, it is relevant to ask if (new) exporters will be able to enjoy better survival prospects (*vis-à-vis* those having not entered such international markets) where risk, uncertainty and competition are all likely to be higher. Understanding which factors determine the firm's risk of closure in international markets holds the key to evaluating the efficacy of export-promotion policies. As pointed out by Alvarez and Lopez (2006), if survivability of businesses depends on trade costs, public policies might concentrate on reducing these costs. By contrast, if firms' hazard rates of closure in export markets are the result of large differences in productivity between exporters and non exporters, then policies that concentrate on facilitating entry may not generate lasting increases in export participation if they are not accompanied by improvements in firms' productivity.

The extant theoretical and empirical literature on the determinants of plant/firm closure, tend to suggest that productivity and sunk costs have a major role in explaining shut-down decisions. Early theoretical work was particularly concerned with how productivity was related to initial size, the learning-by-doing effect associated with the age of a new entrant, and the likelihood of survival. Both initial size and age are positively related to survival, both in theoretical models and in the overwhelming majority of empirical results obtained. In addition, firms that ‘actively learn’ by investing in intangible assets (and consequently increase their specific internal capabilities and ability to absorb external knowledge) are expected to be more likely to survive.

From a global perspective, firms can also acquire (external) knowledge through participating in export markets, so those operating in overseas markets are expected to have better (*cet. par.*) survival prospects. Exporting can also signal positive information about the firm, beyond measured productivity, so exporters should have a lower probability of failure. In contrast, the literature reviewed in the next section shows that higher import penetration increases the probability of closure of the least efficient producers, particularly those supplying domestic markets, but lowers the hazard rate for those firms that export (even after controlling for their higher productivity levels).

Previous literature has provided both theoretical and empirical evidence on the determinants of firm closure and we begin with a brief overview of this literature, in particular emphasising the role of exporting and/or export-market dynamics in influencing the survivability of firms from an international perspective. Nevertheless, as will be seen, there is a distinctive lack of research on the impact of changing exporting status on the firm’ rate of survival, and to the best of our knowledge, this is

the first micro-based investigation in this regard using comprehensive and nationally representative data for the UK firms⁴, covering all market-based sectors.

In particular, this paper considers whether exporting and the lowering of trade costs more generally impact on firm survival rates. In an attempt to add to the rather limited body of evidence on firm closure in the context of international trade, our results show that exporting firms have lower probabilities of closure, conditional on controlling for other factors linked to productivity. In addition, higher import penetration increases the probability of closure for domestic producers, but lowers the hazard rate for those firms that export (even after controlling for their higher productivity levels).

The rest of the paper is organised as follows. Section 2 reviews the existing literature on the determinants of firm closure, with emphasis on the role of participation in export markets. Section 3 describes the data sources used and presents some descriptive statistics, followed by an introduction to our empirical model. Our estimation results are discussed in Section 4, along with some further analytical results. The last section summarises and discusses its relevance to policy making.

2. Literature review

There are a number of theoretical and empirical models of the decision of the firm to shutdown some or all of its capacity. In all the decision depends fundamentally on the firm's prospects for profits, and this in turn is dependent on its productivity and

⁴ Ideally we would wish to consider plant closure, as this would allow us to look at the behaviour of multi-plant enterprises and thus how decisions are made about changes in output capacity (including product lines) without necessarily ceasing production altogether. However, we do not have plant level data in FAME, and must therefore concentrate on an analysis at the firm level – recognising that the vast majority of firms are single-plant enterprises. Thus in what follows we usually refer to the firm as our 'unit of analysis'.

whether this is above a certain shut-down threshold (defined as the lowest level of productivity that would enable the firm to have positive discounted expected profits greater than its liquidation value over future periods). Such a framework leads to productivity and also sunk costs having a major role in explaining closure decisions, and thus the internal and external factors, attached to the firm and the industry in which it operates, that impact on productivity and sunk costs.

Initial, current size and age of the firm

Early theoretical work was particularly concerned with how productivity was related to initial size, the learning-by-doing effect associated with the age of a new entrant, and the likelihood of survival. The models developed by Jovanovic (1982) and Pakes and Ericson (1998) are based on new enterprises entering a perfectly competitive industry with a time-invariant productivity level but without knowing this level and therefore their ‘ability’ to compete effectively; thus they face an immediate cost disadvantage because they typically enter below minimum efficient scale (MES), and then learn whether they can survive. Thus, initial size is positively related to survival since the greater the initial size at start-up the less the need to grow to attain the industry MES. These models also imply that the longer the firm stays in the market the more likely it is to continue to survive through learning-by-doing, and consequently age is negatively related to risk of closure. This relationship between size, age and survival is modified in Ericson and Pakes (1992) to allow firms to actually know their initial ability level and then post-entry invest to improve their performance; this strengthens the relationship between age and survival, and suggests that the effect of the firm’s initial size on its current size should decrease over time and consequently current size should be a better predictor of survival than initial size,

the older is the firm.⁵ Most empirical studies confirm the predictions of the ‘learning-by-doing’ models. For example, Lieberman (1990) found that small firms were more likely to close in declining industries, while Doms *et. al.* (1995) and Geroski (1995) specifically found that larger initial size had a large, statistically significant negative effect on firm closure, while age was strongly negatively related to closure. Dunne *et. al.* (2005) also showed that size was negatively related to closure (but positively related to product-line exit). Audretsch (1994, 1995) and Bernard and Jensen (2002) also found initial size and the age of firms to have the predicted relationship. The age-closure relationship is confirmed in such studies as Dunne *et. al.* (1989); Mata (1994); Mata and Portugal (1994); Mata *et. al.* (1995); Boeri and Bellman (1995); Cefis and Marsili (2006); and for the UK see Harris and Hassaszadeh (2002) and Disney *et. al.* (2003a). Few studies have tended to include both initial size and current size as determinants of closure probabilities. Mata *et. al.* (1995) found that current firm size improves the chances of survival but initial size was positively related to firm exit. They explained this counterintuitive result by stating that if a firm survives but starts smaller and therefore experiences faster post-entry growth, then it has a higher probability of survival, so the initial size effect is positive. Harris and Hassaszadeh (*op. cit.*) and Disney *et. al.* (*op. cit.*) also found that initial size had a positive effect, while current size had a negative impact, on firm closure. Like Mata *et. al.* (*op. cit.*), both studies show that the impact of initial size and current size is not independent of the relationship between these two variables (i.e., subsequent growth following entry) and therefore the probability of closure depends on whether the firm is growing or

⁵ Ghemawat and Nalebuff (1985) take a different approach: they are interested in declining industries and exiting, and so assume all producers have equal costs but as demand diminishes the smaller producer, having lower output, can operate as a profitable monopolist over a longer period of time as demand falls. Hence, larger plants/firms exit first. However, a positive relationship between firm size and exit is more likely when multi-plant operations are taken into consideration (see the discussion below on the role of ownership and branch plants).

declining from their initial size (growth reduces the hazard of closing, while declining increases the hazard). They go even further and consider the interaction of initial and current size with the age of the establishment, showing that the effect of initial size wears off as the firm gets older, the negative impact of firm growth on the hazard of closure increases with age but this relationship between growth and age is non-linear and tapers off as the firm gets larger over time.

Intangible assets

Ericson and Pakes (1995) and Olley and Pakes (1996) extended Jovanovic's model to include the investments of individual firms (particularly on intangible assets) to allow for 'active learning', thus relaxing the assumption that firm productivity levels are constant over time. According to resource-based theories of the firm, firms that invest in intangible assets, such as advertising, goodwill, R&D and other innovative activities, and consequently increase their specific internal capabilities and ability to absorptive external knowledge, are more likely to survive. Aw *et. al.* (2005) also allow firms to generate (external) knowledge through participating in export markets, so that the evolution of firm productivity over time is determined by past productivity as well as investments in such knowledge acquiring activities as undertaking R&D and exporting.⁶ There have been few studies that provide empirical evidence on the relationship between firm closure and intangible assets possessed by firms; some exceptions are Hall (1987), Kimura and Fujii (2003), Esteve Pérez *et. al.* (2004), and Cefis and Marsili (2006), all of who found that spending on R&D and/or innovation reduced the probability of closure. However, investments in intangible assets are also associated with uncertainty and thus a higher risk of failure in a more technologically

⁶ Note internal knowledge generation (through in-house R&D) and acquiring external expertise through export market participation are not necessarily independent; indeed, the productivity effect of export market participation is likely to depend on the firm's in-house capability to assimilate new information – i.e. its absorptive capacity (see Harris and Li, 2008, for a discussion).

competitive environment, such that some of the extant literature has found that, taking R&D for instance, closure rates are higher in R&D intensive industries (Audretsch and Mahmood, 1995; Audretsch *et. al.*, 2000; Segarra and Callejón, 2002).

Productivity

In addition to the learning-by-doing model of Jovanovic (1982), the model of firm and market dynamics developed by Hopenhayn (1992) has provided important theoretical underpinnings on what determines firm closure. Similar to Jovanovic, Hopenhayn showed that survival in a perfectly competitive industry increases with the size and age of the firm, and that firms with the lowest levels of productivity close first. Each firm's productivity is determined by a random draw from a Markov process whereby higher future productivity is a function of current productivity (the distribution of productivity is represented by the function $F(\theta_{t+1}|\theta_t)$ which is strictly decreasing in θ_t), such that the higher is current productivity the more likely that a firm will experience high productivity levels in the future.

This implies that productivity levels are persistent, while those firms with the lowest levels exit the industry first (as their productivity will fall below the exit threshold). In addition, any cohort of continuing firms at time t will have a productivity distribution to the right of entering firms (the productivity distribution of the continuing cohort stochastically dominates), since older firms constitute a larger proportion of high productivity units. Thus the age of the firm and survival will be positively related. Similarly, firms with an initially low level of productivity experience low productivity in the future, and thus have a greater probability of closure; thus heterogeneity and persistence in the distribution of productivity implies that closing and initial productivity differences (as proxied by initial size) will be inversely related.

Sunk costs

In addition to these similar predictions between the Jovanovic and Hopenhayn models with respect to the relationship between size, age and productivity levels and the probability of firm closure, Hopenhayn also showed that increased sunk costs played a key role in lowering (entry and) exit. Potential entrants into an industry must pay a sunk entry cost (c_e), with higher levels of c_e providing a barrier to entry which insulate incumbents from the effects of market selection based on their productivity levels (given that incumbents do not have to pay c_e). Therefore, industries with higher c_e will have lower probabilities of exit (and entry, given that firms only learn their productivity level post-entry). This further implies that the productivity differential between surviving and closing firms will be greater in markets with higher sunk costs, as higher c_e reduces the closure threshold and thus firms shutting down are relatively more likely to be concentrated in the lowest end of the productivity distribution.

Other have directly looked at closure and sunk costs. Dixit (1989) showed that if capital investments had a substantial sunk (as opposed to just a fixed) cost component and demand was uncertain then firms would wait longer before entering a market and once there would wait longer before closing, even if they were incurring losses (van Ewijk, 1997, provided corroborating evidence in his theoretical model of entry and exit). Doms *et. al.* (1995) also made the argument that plants with higher capital-labour ratios might have a lower ratio of variable to fixed costs, and would therefore remain in operation as long as variable costs were covered after experiencing negative demand shocks. There is supporting empirical evidence of a positive relationship between sunk costs and survival in a small number of studies, although Siegfried and Evans (1994) were not able to point to many when surveying early attempts at establishing the role of sunk costs. In contrast, Audretsch (1994), Doms *et. al.* (1995),

Kleijweg and Lever (1996) and Bernard and Jensen (2002) all reported a significant negative relationship between capital intensity and firm closure. Harris and Hassaszadeh (2002) showed that the impact of sunk costs (as proxied by the capital-labour ratio) was important but needed to take account of age effects; i.e. there was only a negative relationship between sunk costs and closure as a plant aged. This accorded with the results presented in Colombo and Delmastro, (2000, 2001) and suggested that shut-down decisions were much more irreversible in older plants who had become attached to an industry, while younger, more capital intensive plants might be exhibiting higher fixed (as opposed to sunk) costs and therefore had more flexibility to leave.⁷ Lastly, Hölzl (2005) had also tested for the impact of sunk costs associated with intangible assets (proxied by advertising intensity) as well as tangible sunk costs (associated with capital intensity), finding (somewhat unexpectedly) that intangible assets were associated with a higher rate of failure (Kimura and Fujii, 2003, also found a positive relationship between advertising intensity and closure for Japan).

Exporting activities

The decision of closure of exporting firms, depends mainly on industrial characteristics such as the level of sunk costs, and as with the case of closure in normal markets, the minimum productivity or efficiency needed to secure non-negative profits (Das *et. al.*, 2001; Bernard and Jensen, 2004a). Hopenhayn's (1992) model has been adapted and extended to include trade (e.g. Melitz, 2003) and to allow for the impact of market size on firm turnover (e.g. Asplund and Nocke, 2006). Melitz allowed for monopolistic competition and introduced trade costs (and in particular that firms that exported faced additional sunk costs when entering international

⁷ The resale value of their capital equipment may be much higher, since irreversibility is known to increase with the product- and firm-specific nature of investments, which are likely to be strongly associated with the vintage of the plant. See below for more on the role of fixed (as opposed to sunk) costs.

markets), showing that this induced only the most productive firms to export and simultaneously caused firms with the lowest levels of productivity to exit. Thus, Bernard and Jensen (2002) noted that "... to the extent that exporting signals positive information about the plant, beyond measured productivity, we would expect current period exporters to have a lower probability of failure" (p.8). In related, and more recent work, Bernard *et. al.* (2005) developed a model of comparative advantage with exporting, with monopolistic competition and firms with heterogeneous levels of productivity. Again only the most productive firms were able to overcome sunk costs when producing for international markets, while trade liberalisation resulted in increased competition both home and abroad. The result was higher exit rates for firms that had the lowest productivity in domestic markets; some exporters with relatively low levels of productivity ceased to export (although the closure rate for former exporters would be lower than for non-exporters, given their on average higher levels of productivity vis-à-vis domestic producers); and some firms with high levels of productivity started to export (given access to larger markets).

Evidence on whether exporting firms have lower probabilities of closure, conditional on controlling for other factors linked to productivity, is beginning to emerge. Bernard and Jensen (2002) found that exporting reduced plant closure in the U.S. by as much as 15% (after accounting for the impacts of size, productivity, factor intensity, and ownership structure). Hölzl (2005) confirmed that exporting reduced closure for Austrian manufacturing, while exporting firms in Spain experienced a 28% lower probability of failure when compared to non-exporters (Esteve Pérez *et. al.*, 2004). Kimura and Fujii (2003) and Sabuhoro *et. al.* (2006) also confirm the positive relationship between exporting and survival for Japanese and Canadian manufacturing and service sector firms, respectively.

Trade costs and import penetration

As well as exporting impacting on the probability of closure, globalisation and liberalisation of markets has resulted in greater levels of import penetration. Reductions in trade costs results in both increases in exports and imports (Brander, 1995; Melitz, 2003; Bernard *et. al.*, 2005), with increased import competition increasing the probability of closure of the least efficient producers, particularly those supplying the domestic market. In addition to domestic producers facing greater competition, when there is higher intra-industry trade associated with differentiated products, this results in different survival thresholds in industries/countries with different efficiency levels, increasing the risk of closure for weaker firms in those sectors with the most heterogeneous products (Faley *et. al.*, 2004). It is also likely (based on the technology transfer literature and also anticipating our empirical results later on), that firms that export in these differentiated markets (with potentially higher levels of import penetration) may actually benefit from increased imports of (intermediate) goods and services (which presumably brings with it potentially higher levels of technology transfer through access to a wider stock of knowledge – cf. Ethier, 1982; Grossman and Helpman, 1991; and Eaton and Kortum, 2001, 2002). High import penetration may therefore increase the probability of closure for domestic producers, but lower the hazard rate for those firms that export (even after controlling for their higher productivity levels).

Evidence of the impact of higher import penetration on firm closure is also relatively new; e.g., Gullstrand (2005) found that imports increased the probability of closure in Swedish textiles, but the impact was moderated in those sub-groups with higher levels of intra-industry trade. Baggs (2005) found that Canadian tariff reductions increased the hazard rate of closure for Canadian firms (U.S. reductions had the opposite

effect); however, Hölzl (2005) found import growth had an insignificant effect on exit (and entry) in Austria. Hölzl attempted to rationalise his results by noting that "... a possible explanation ... may be the fact that in Austria manufacturers are often at the same time recipients and distributors of imported goods" (p. 2439). Earlier evidence of higher imports increasing closure rates is provided by Bernard and Jenson (2002) for U.S. manufacturers (although their significant effects related to imports from low wage countries increasing hazard rates); von der Fehr (1991) for Norway and Angostaki and Louri (1995) for Greece also found the expected impact of imports on closure rates.

Asplund and Nocke (2006) extended Hopenhayn's (1992) model by allowing for imperfect competition, finding that firm turnover (entry and closure) increase with market size, with the net effect of a higher rate of failure for less productive firms and a lower rate of closure for the most productive enterprises. They also were able to distinguish between the impact of sunk and fixed costs, finding that sunk costs lower the probability of closure (as above) while higher fixed costs lowers the probability of failure for the most efficient firms (as they gain more from higher prices) but increases closure for less efficient firms (who have to meet a higher close-down threshold and so more shutdown).

Foreign ownership

Moreover, the importance of ownership in determining the probability of survival has also been well documented in the literature⁸. With respect to country of ownership, distance effects may make it more likely that overseas companies find it less

⁸ Since our data is at the firm level, and in addition does not allow us to track ownership changes over time, we limit our discussion here to just those aspects that can be incorporated into our study. Thus the evidence on plant status (single-plant enterprises versus multi-plant) and changes in ownership (acquisitions by foreign- or domestically-owned firms) are not included in here (rather see Bernard and Jensen, 2002, for the U.S. and Harris and Hassaszadeh, 2002, for the UK).

problematic if needing to close capacity in host countries. In contrast, if foreign-owned subsidiaries have access to superior foreign technologies and other proprietary assets available from the multinational enterprise, this would suggest that such firms will have higher productivity and thus a higher probability of survival. In addition, the model of asset-seeking FDI (cf. Dunning, 1998; Wesson, 1999) suggests that foreign-owned firms hope to create advantages for themselves through acquiring and internalising valuable assets in the host nation and thus they will wish to establish capacity in the host nation by acquiring firms with comparatively superior productivity levels and with technological characteristics that match more closely with their own use of technology (e.g., capital- and intermediate input intensive). Otherwise, such FDI involves excessive costs when adapting and modifying the acquired firm's use of different technology, when gaining expertise and experience in the host market, and when building up trust in order to gain from the internal market operated by the larger post-acquisition enterprise.

In terms of empirical evidence, Dunning (1998) and Wesson (1999) looked at foreign-owned plants in the Northern region of England and found country of ownership was not important as a determinant of closure. Colombo and Delmastro (2000) found a strong, positive effect between foreign ownership and plant closure for Italian metalworking plants in 1986, while Görg and Strobl (2003) for Ireland, Bernard and Sjöholm (2003) for Indonesia, Esteve Pérez *et. al.* (2004) for Spain, and Gullstrand (2005) for Sweden all found a greater probability of closure for foreign-owned plants compared to domestically-owned ones. By contrast, both Li and Guisinger (1991) for the U.S. and Mata and Portugal (2002, 2004) for Portugal found the opposite result, while Harris and Hassaszadeh (2002) found that overall UK plants that belonged to the foreign-owned sector were generally less likely to shutdown, but that this

diminished with the age of the plant. Note, most of this evidence relates to plants and not firms, and thus may conflate the separate effects of foreign-owned firms having a higher likelihood of closing plants (especially in assembly units in peripheral, lower cost regions) but overall maintaining high levels of productivity within the firm.

Productivity

A number of studies have incorporated direct measures of profitability (or productivity) when attempting to explain closure, rather than just relying on size and age variables to proxy for relative costs. Siegfried and Evans (1994) found that many early studies (often using aggregated industry-level data) were not successful at linking profitability to closure, often because of collinearity with measures such as sunk costs. Kleijweg and Lever (1996) also found profitability insignificant. In contrast, total factor productivity was strongly, negatively related to closure in the studies by Doms *et. al.* (1995); Kovenock and Phillips (1997), Bernard and Jensen (2002), and Harris and Hassaszadeh (2002)^{9, 10} leading them to conclude that relatively less efficient firms closed first (or that firms closed relatively less efficient plants).

Displacement effect

Closure may also occur when firms are displaced by more efficient new entrants, resulting in the often reported stylised fact that entry and exit appear to be strongly correlated over time. Gabszewicz and Thisse (1980) provided a theoretical justification while, for example, Caves and Porter (1977), Geroski (1991), Baldwin (1995), Dunne *et. al.* (1988), Mata and Portugal (1994), Boeri and Bellman (1995)

⁹ Harris and Hassaszadeh (*op. cit.*) found that UK plants with lower technical efficiency (their proxy for productivity) had a higher probability of closure, but this negative relationship declined with the age of the plant.

¹⁰ Dunne *et. al.* (2005) for the U.S. and Gullstrand (2005) for textiles in Sweden used a labour productivity measure, with an expected negative sign obtained.

Ilmakunnas and Topi (1999), and Harris and Hassaszadeh (2002) all provided empirical support that (lagged) entry positively determined closures (although the effect was not always statistically significant).

Macroeconomic influences and output growth

Lastly, the state of the business cycle should impact on the hazard of plant closure (e.g., van Ewijk, 1997). Growth in real GDP is expected to reduce the number of closures as profitability rises, but downturns in aggregate demand may be asymmetric since entry and exit are strongly correlated. Nevertheless, falls in profitability are expected to increase the rate of bankruptcies (albeit in a biased direction towards smaller enterprises). The empirical evidence is mixed; for example, Kovenock and Phillips (1997) found that capacity utilisation had the expected negative and statistically significant relationship with plant closures, but changes in output demand were not significant. Mata *et. al.* (1995) found that high industry growth lowered hazard rates, but that industries with high growth and high entry rates had higher hazard rates. Lieberman (1990) did not find capacity utilisation to be important, while Boeri and Bellman (1995, p. 493) stated that "... all in all, *even after controlling for industry heterogeneity and displacement effects, cyclical factors do not seem to affect the pace of exit*" (italics in original). Ilmakunnas and Topi (1999) specifically compared the microeconomic and macroeconomic influences on entry and exit, and conclude that macroeconomic factors explain entry fairly well, but exits much less so. Harris and Hassaszadeh (2002) found that expansion within an industry provided UK plants with a cushion against the risk of closure, but only for plants that had been in operation for 10 years or more; younger plants faced a higher hazard rate of failure when industry growth was relatively high suggesting that such high growth may attract too many young plants that turn out not to be viable. Likewise, Disney *et. al.*

(2003b) found manufacturing output growth to be insignificant in determining the hazard of closure, except when it also entered as a joint-variable with plant age. Then increases in demand were still positively associated with closure, but the joint demand-age variable was statistically significantly negative (although smaller in value than the direct impact of demand on closing). Dunne *et. al.* (2005) for U.S. manufacturing found a negative relationship between growth in demand and plant closure, while Baggs (2005) reports a similar result for Canadian industries.

In summary, this selective review of what is a large literature shows that past theoretical and empirical studies have set out to test several hypotheses concerning the determinants of firm closure such as: the age of the enterprise; its size (initial and current); sunk costs (including intangible costs); productivity; ownership; the displacement effects of new entry; and the state of the business cycle.¹¹ However, there has been less attention given in the literature to the effects of exporting and the associated export-market dynamics, and this is the particular focus of the present study. In particular, we are interested in obtaining evidence for the UK as to whether firms that overcome entry barriers into overseas markets have lower rates of closure.

3. The data and empirical model

The *FAME* dataset for 1996-2004 is used for this study, which includes all firms operating in the UK that are required to make a return to Companies House. It contains basic information on firm-specific characteristics, such as turnover, intermediate expenditure, employment, assets, and most importantly, overseas sales.

Apart from financial information, *FAME* also has information on the year of

¹¹ Other factors also feature, such as changes in ownership, the relationship between firms and plants, product differentiation, market size, and the wider effects of trade liberalisation. However, we concentrate here on those factors we can control for using the *FAME* database.

incorporation of the company, postcodes, the 4-digit 2003 SIC industry code, and country of ownership. Note, we only use data containing unconsolidated accounts, to avoid double counting and within firm transfer effects. Our final dataset used for statistical analysis comprises of an unbalanced panel, containing 91,701 firms with 247,028 observations during 1997-2003.¹²

The *FAME* dataset is severely biased towards large enterprises, and thus is unrepresentative of UK firms. To obtain a distribution representative of the population of firms operating in the UK, we treat the firms in the *FAME* dataset as a sample of the nationally-representative Annual Respondents database (or *ARD*) population¹³, and consequently weight the *FAME* data to produce a representative database (by industry and firm size). In practice, we have obtained aggregated turnover data from the *ARD* sub-divided into 5 size-bands (based on turnover quintiles) and 3-digit industry SICs¹⁴. We then aggregate the *FAME* data into the same sub-groups, so that we can calculate weights using the total turnover data from the *ARD* divided by the comparable data from *FAME*. In essence, the *FAME* dataset is weighted to acquire the same distribution of turnover as those firms in the *ARD*.¹⁵ Table A.1 (in the Appendix) provides the relevant evidence and validates our weighting approach.

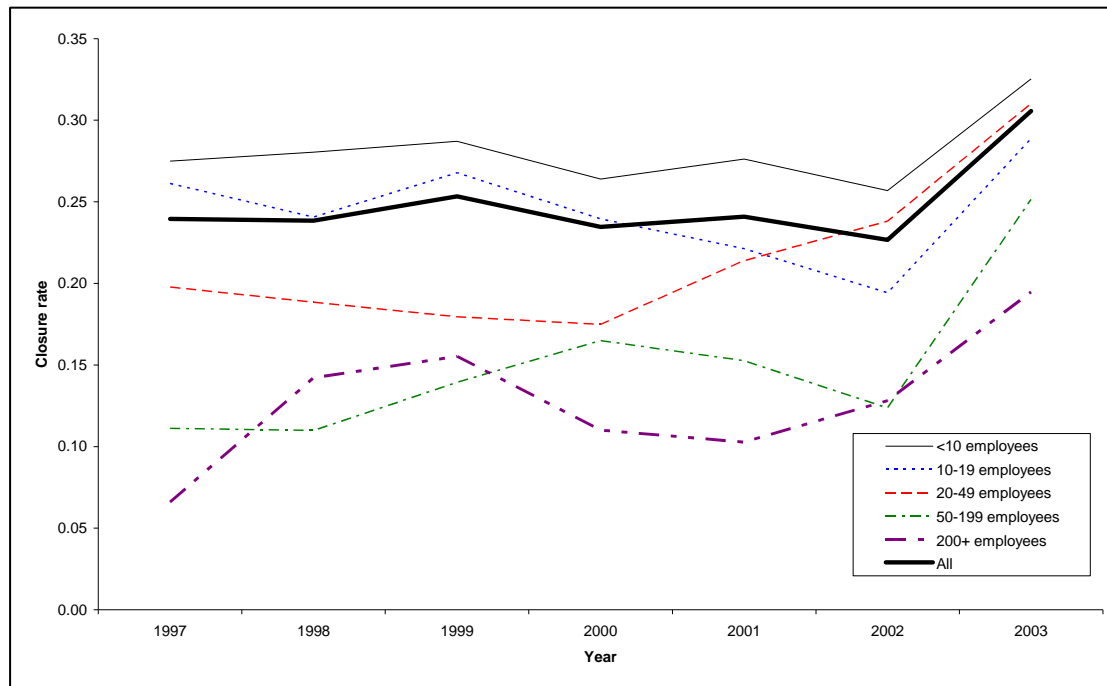
Figure 1: Closure rate^a by employment size-band in UK market-based sectors, 1997-2003

¹² The first year of data (1996) is 'lost' as we include industry growth rates as one of our determining variables (thus 1996 growth cannot be computed without 1995 data). The last year of data (2004) is also omitted as in 2004 all firms 'exit' in this year (i.e. the data is right-censored).

¹³ For a details description of the *ARD* (available at the ONS), see Oulton (1997), Griffith (1999) and Harris (2005).

¹⁴ Where there are fewer than 10 enterprises in any sub-group, these data are not used, so as to avoid disclosure of confidential information in these ONS data. This results in a loss of some 4% of the total turnover available in the *ARD*.

¹⁵ Note we do not weight the *FAME* data for 34 industries because the *FAME* data have better coverage in terms of total turnover than the *ARD*. These 34 industries (out of 215 in total) account for just 2.9% of total *FAME* turnover. Note also, the *ARD* does not contain data for Northern Ireland but since this region is rather small it will not have much of an effect in the weights used.



^a Proportion of firms closing per annum.

Note, a Cox regression-based test for the equality of the survival curves for these employment size-band groups rejects the null at > 1% significance level.

Source: weighted *FAME* data

Figure 1 presents the closure rate of firms in the dataset, with separate rates for different enterprise size. Overall, there was a relatively stable rate of closure of some 24% of firms between 1997-2002, rising to 31% in 2003. The smallest firms (in terms of their employment) had the highest rates of closure (28% for those employing under 10 vis-à-vis an average of 13% for firms with 200+ employees), and generally more stable patterns over time. Definitions of the variables used to estimate a model of closure are presented in Table 1. Information on (weighted) means and standard deviations for each variable are also provided, divided into 5 sub-groups depending on the exporting status of the firm. These show that firms that exported at some time during the period tended to be slightly younger (except for those that entered exporting); they were much more likely to be foreign-owned; slightly more likely to have non-zero intangible assets; operate in industries experiencing on average lower growth in output and less displacement (except for firms that both entered and exited from exporting); they had higher levels of both initial and current employment; higher

relative total factor productivity; and higher capital intensity (except for firms new to exporting). In all, there is evidence that exporting firms have many of the characteristics associated with lower unconditional hazard rates of closure.

Figure 2 confirms that the baseline survival functions ($h_0(t)$ in Equation (2) below¹⁶) for firms that enter and/or exit overseas markets are indeed considerably higher; however, there is little difference in the baseline survival functions of firms that are continuous exporters versus non-exporters. Thus, it is of interest to consider whether the covariates in the model to be estimated point towards any movement away from these baseline functions (and more particularly towards lower hazard ratios for the continuous exporter sub-group).

¹⁶ Note, the baseline survivor functions in Figure 2 are equivalent to “covariate-adjusted” Kaplan-Meier estimates; if $\mathbf{x}(t)$ in equation (2) below was zero, then we would have the usual Kaplan-Meier estimates functions (which do not have as smooth an appearance as in Figure 2). In fact, the functions shown used the estimated $\boldsymbol{\beta}$ to put all firms on the same level by adjusting for the covariates, before proceeding with the Kaplan-Meier calculations.

Table 1: Variable definitions and descriptive statistics ^a, by exporting status

Variable	Definition	Never exported		Always exported (EXP_always)		Entered exporting (EXP_entry)		Exited exporting (EXP_exit)		Entered/exited exporting (EXP_both)		All firms	
		Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD
AGE _{it}	Age of firm (<i>t</i> minus year opened + 1) in years	16.6	17.2	20.7 [†]	19.7	14.5	14.9	19.4 [†]	17.5	18.8 [†]	17.8	17.0	17.4
FO _{it}	Dummy coded 1 if firm <i>i</i> is foreign-owned at time <i>t</i>	0.039	0.194	0.206 [†]	0.405	0.198 [†]	0.399	0.171 [†]	0.377	0.213 [†]	0.409	0.062	0.242
INTGBL _{it}	Dummy coded 1 if firm <i>i</i> has non-zero intangible assets at time <i>t</i>	0.115	0.319	0.160 [†]	0.367	0.157 [†]	0.364	0.133 [†]	0.339	0.146 [†]	0.353	0.120	0.325
ln GROWTH _t ^b	Growth in industry real gross output, <i>t</i> -1 to <i>t</i>	-0.026	0.194	-0.028 [‡]	0.207	-0.030 [‡]	0.180	-0.029 [‡]	0.178	-0.036 [‡]	0.207	-0.026	0.195
ln DISPLACE _t ^b	Employment of new entrants ÷ employment of existing firms in time <i>t</i>	-4.51	3.33	-4.62 [‡]	3.20	-4.91 [‡]	3.04	-5.05 [‡]	3.16	-4.31 [‡]	3.18	-4.53	3.31
ln OPEN _i	Employment in firm <i>i</i> in year first observed	2.10	1.32	3.05 [†]	1.65	2.53 [†]	1.46	2.82 [†]	1.58	2.83 [†]	1.63	2.22	1.40
ln TFP _{it}	Relative TFP in firm <i>i</i> at time <i>t</i> (see appendix for details)	-0.172	0.452	-0.064 [†]	0.313	-0.120 [†]	0.450	-0.128 [†]	0.442	-0.142 [†]	0.472	-0.160	0.442
ln KL _{it}	Tangible assets-to-labour ratio for firm <i>i</i> in time <i>t</i>	1.93	1.90	1.94 [†]	1.50	1.87	1.51	2.05 [†]	1.93	2.12 [†]	1.74	1.94	1.86
ln EMP _{it}	Current employment in firm <i>i</i> in time <i>t</i>	2.14	1.34	3.08 [†]	1.66	2.70 [†]	1.51	2.48 [†]	1.48	2.70 [†]	1.58	2.25	1.41
ln HERF _{it}	Industry (2-digit) Herfindahl index in time <i>t</i>	-4.675	0.841	-4.295 [†]	0.895	-4.413 [†]	0.849	-4.481 [†]	0.827	-4.342 [†]	0.919	-4.626	0.856
ln IMP _{it}	Import penetration in time <i>t</i> ^c	0.234	2.140	1.563 [†]	2.182	1.568 [†]	2.064	1.372 [†]	1.926	1.410 [†]	1.993	0.420	2.185

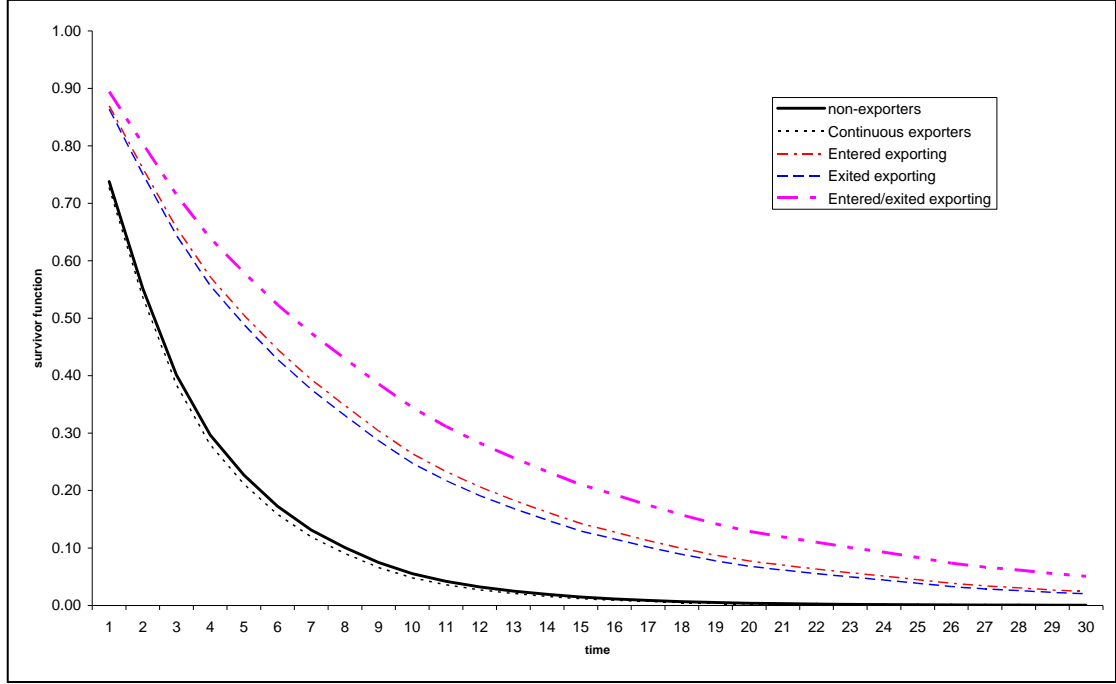
^a In addition the variables entering the model included 33 industry dummies at the 2-digit (SIC 2003) level and 11 Government Office region dummies (as well as composite dummies involving the AGE variable and export status)

^b Calculated separately for each 3-digit industrial sector.

^c Source: Table 3.6 in 2006 UK Input-Output Tables (ONS, 2006)

^{†(‡)} p-value > 0.01 of Kolmogorov-Smirnov test rejecting null that the distribution of 'never exported' sub-group of firms lies to the right (left) of tested sub-group.

Figure 2: Baseline survival functions^a for exporting and non-exporting firms in UK market-based sectors, 1997-2003



^a Obtained from estimating equation 2.

Note, a Cox regression-based test for the equality of the survival curves for these exporting sub-groups rejects the null at > 1% significance level.

In order to model the determinants of firm closure, and defining the hazard rate of firm i as the probability that it shuts down in time t having survived until t , the hazard function $h(\cdot)$ is given by:

$$h(t; \mathbf{X}(t)) = P[\text{firm at } t \mid \text{survival to } t; \mathbf{X}(t)] = P[T = t \mid T \geq t, \mathbf{X}(t)] \quad (1)$$

where $\mathbf{X}(t)$ is the covariate path of \mathbf{x} up to t . We choose to estimate a Cox proportional hazard model (Cox 1972, 1975):

$$h(t) = h_0(t) \exp(\mathbf{x}(t) \boldsymbol{\beta}) \quad (2)$$

which comprises a parameterised function of firm characteristics, $\exp(\mathbf{x}(t)\boldsymbol{\beta})$; and a non-parametric base-line hazard, $h_0(t)$, as this is preferable to specifying a functional

form such as the Weibull distribution, which may lead to misspecification of the baseline hazard function. Such a semi-parametric model is more flexible than other alternative specifications, and in our empirical work we found the results to be more plausible when compared to parametric modelling of the baseline hazard¹⁷. Given the need to meet the proportional hazard assumption, we allowed the baseline hazard to differ for 3 sub-groups based on allocating firms in time t to whether the industry they belonged to had >30%, 10-30% or <10% of firms engaged in exporting.

The Cox proportional hazards model in STATA (version 9) was used to obtain estimates, and the actual modelling strategy used was to start with a simple model (referred to as Model 1 and involving only the key variables, including exporting defined as a dummy variable taking on a value of 1 if the firm exported in time t), and then allow for different types of export status (Model 2), before entering variables multiplied by ‘age’,¹⁸ and excluding insignificant cross-product variables involving the age variable (Model 3). We then allowed all the variables in the model to enter for a third time having been multiplied by the 4 exporting sub-groups outlined in Table 1, and dropping all cross-product terms involving age and these export sub-groups that were insignificant (Model 4). The results for Models 1 – 4 are given in Table 2. The ‘restricted model’ approach was checked against ‘full model’ by testing whether the

¹⁷ In order to assess the appropriateness of various assumptions on the underlying distributions, as a preliminary visual analysis, we have plotted the estimated integrated hazard function for generalised residuals; and by comparing this with the actual integrated hazard for the assumed distribution, we are able to show that neither the Weibull nor exponential distribution is in line with our data specification. As robustness checks, we have also attempted modelling the baseline hazard in a parametric setting; nevertheless, alternative parametric models fail to generate plausible results. Therefore, we conclude the proportional hazard model is a valid approach to estimating our hazard function.

¹⁸ As Disney *et al.* (2003a) point out: “... with the Cox specification, we cannot enter age directly since it is collinear with the baseline hazard. We could enter age directly if we adopted a parametric specification for the baseline, but we would then be relying on identification of the age effect from the assumed functional form” (p.105).

Table 2: Parameter estimates of the (stratified) weighted hazard model for UK firms in market-based sectors, 1997-2003

Variable	Model 1		Model 2		Model 3		Model 4		$\exp(\hat{\beta})-1$
	$\hat{\beta}$	z-value ^a	$\hat{\beta}$	z-value	$\hat{\beta}$	z-value	$\hat{\beta}$	z-value	
FO _{it}	-0.223	-4.79	-0.195	-4.27	-0.193	-4.23	-0.131	-2.21	-0.123
INTGBL _{it}	0.032	0.72	0.034	0.76	0.035	0.79	0.040	0.90	0.041
<i>ln</i> GROWTH _t	0.221	2.59	0.221	2.62	0.103	0.92	0.097	0.85	0.102
<i>ln</i> KL _{it}	-0.039	-5.13	-0.038	-5.00	-0.050	-4.85	-0.048	-4.59	-0.047
<i>ln</i> DISPLACE _t	-0.009	-1.94	-0.010	-1.97	-0.010	-1.98	-0.009	-1.92	-0.009
<i>ln</i> OPEN _i	0.095	4.25	0.106	4.50	0.103	4.38	0.098	3.82	0.103
<i>ln</i> EMP _{it}	-0.237	-11.12	-0.246	-11.00	-0.242	-10.87	-0.223	-9.03	-0.198
<i>ln</i> TFP _{it}	-0.062	-1.93	-0.068	-2.12	-0.068	-2.11	-0.072	-2.46	-0.069
<i>ln</i> HERF _t	0.030	1.08	0.032	1.14	-0.000	-0.01	-0.004	-0.11	-0.004
<i>ln</i> IMP _t	0.007	0.60	0.010	0.93	0.010	0.91	0.029	2.62	0.019
EXPORT _{it}	-0.121	-3.76	–	–	–	–	–	–	–
EXP_always _{it}	–	–	-0.006	-0.19	-0.000	-0.02	0.343	4.53	0.409
EXP_entry _{it}	–	–	-0.669	-10.87	-0.810	-9.26	-0.756	-5.09	-0.530
EXP_exit _{it}	–	–	-0.525	-6.16	-0.681	-5.18	-0.256	-1.80	-0.226
EXP_both _{it}	–	–	-1.018	-9.54	-1.021	-9.56	-1.052	-9.83	-0.651
<i>ln</i> GROWTH _t × AGE _{it}	–	–	–	–	0.009	2.03	0.009	2.04	0.009
<i>ln</i> KL _{it} × AGE _{it}	–	–	–	–	0.001	2.11	0.001	2.36	0.001
<i>ln</i> HERF _t × AGE _{it}	–	–	–	–	0.002	2.41	0.002	2.48	0.002
EXP_entry _{it} × AGE _{it}	–	–	–	–	0.009	2.40	0.011	2.76	0.011
EXP_exit _{it} × AGE _{it}	–	–	–	–	0.008	2.49	0.010	2.82	0.010

$FO_{it} \times EXP_always_{it}$	-	-	-	-	-	-	-0.172	-2.04	-0.158
$\ln KL_{it} \times EXP_always_{it}$	-	-	-	-	-	-	-0.031	-1.69	-0.030
$\ln OPEN_i \times EXP_always_{it}$	-	-	-	-	-	-	0.066	1.48	0.068
$\ln EMP_{it} \times EXP_always_{it}$	-	-	-	-	-	-	-0.169	-3.83	-0.156
$\ln TFP_{it} \times EXP_always_{it}$	-	-	-	-	-	-	-0.213	-3.01	-0.192
$\ln IMP_t \times EXP_always_{it}$	-	-	-	-	-	-	-0.027	-2.09	-0.027
$FO_{it} \times EXP_exit_{it}$	-	-	-	-	-	-	-0.459	-1.90	-0.368
$\ln EMP_{it} \times EXP_exit_{it}$	-	-	-	-	-	-	-0.156	-3.40	-0.145
$\ln TFP_{it} \times EXP_entry_{it}$	-	-	-	-	-	-	-0.155	-2.29	-0.143
$\ln IMP_t \times EXP_entry_{it}$	-	-	-	-	-	-	-0.072	-2.29	-0.070
$\ln KL_{it} \times EXP_entry_{it}$	-	-	-	-	-	-	-0.107	-2.13	-0.101
$\ln OPEN_i \times EXP_entry_{it}$	-	-	-	-	-	-	-0.142	-1.73	-0.133
$\ln EMP_{it} \times EXP_entry_{it}$	-	-	-	-	-	-	0.177	2.14	0.194
$\ln IMP_t \times EXP_entry_{it}$	-	-	-	-	-	-	-0.030	-1.98	-0.029
$\ln GROWTH_t \times EXP_entry_{it}$	-	-	-	-	-	-	0.677	1.80	0.967
2-digit SIC industry dummies	YES		YES		YES		YES		
Log likelihood	-467,195.75		-466,799.48		-466,744.53		-466,592.94		
Cox-Snell Pseudo-R ²	0.243		0.244		0.245		0.246		
Link test ^b : $H_0: \beta_2 = 0$	0.97		0.60		0.27		0.65		
Sig. level. omitted cross-product terms $H_0: \Sigma\beta_i = 0$					0.61		0.15		

No. of observations = 247,028; no. of firms = 91,701; no. of closures observed = 61,482

^a Based on robust standard errors;

^b Tests of model misspecification are based on log relative hazard = $\beta_1 (\mathbf{x}\hat{\beta}_x) + \beta_2 (\mathbf{x}\hat{\beta}_x)^2$. Under the correct specification, $\beta_1 = 1$ and $\beta_2 = 0$; thus we test $\beta_2 = 0$. The significance level for rejecting the null is reported here.

omitted parameter estimates were jointly zero (significance levels are reported for Models 3 and 4 in Table 2, showing that dropping jointly insignificant variables is accepted). We also report a pseudo Cox-Snell R^2 test statistic, based on a likelihood ratio test that the remaining parameter estimates are jointly non-zero (Sabuhoro *et. al.*, 2006, p.63 provide more details), and the results from a ‘link test’ that tests whether squared values of the model’s explanatory variables are jointly significant (which would suggest that the model is mis-specified). Details of this test are provided in Cleves *et. al.* (2002, Chapter 11). In summary, the models estimated appear to be adequate for analysing the determinants of firm closure using the (weighted) *FAME* dataset.

4. The results

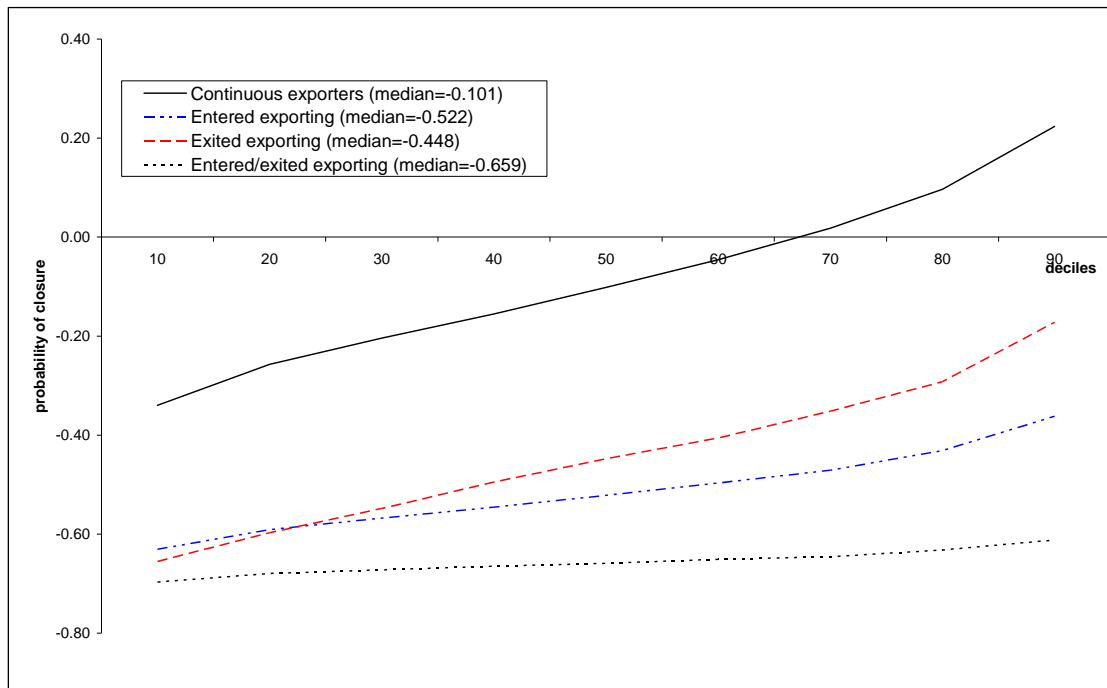
The results for Model 1 in Table 2 show that firms that exported at some time during 1997-2003 had a statistically significant lower probability of closure. However, the hazard rate ($e^{\hat{\beta}} - 1$) for this variable is not large; it shows that firms were 11.4% less likely to shutdown if they exported, conditional on having controlled for other covariates. Model 2 produces more details for the various exporting sub-groups, showing that those that exported continuously were not significantly more likely to survive, but firms that entered and/or exited export markets had much lower hazard rates. Model 3 confirms these differences across the various exporting sub-groups, and also shows that certain variables (including exporting status) impact on the probability of closure with differing effects depending on the age of the firm. Finally, Model 4 provides the most detailed set of results since it allows for ‘age’ and ‘exporting’ to have a differential impact combined with the other variables in the model. This is our preferred model.

Table 3: Changes in (marginal) weighted hazard rate of closure for continuous exporters by firm characteristics, 1997-2003

Hazard rate ^a (deciles)	<i>ln</i> KL	<i>ln</i> OPEN	<i>ln</i> TFP	<i>ln</i> EMP	<i>ln</i> IMP	FO
less than -0.327	2.89	5.67	0.14	5.96	2.46	0.45
-0.327 to -0.242	2.55	4.22	0.08	4.38	2.59	0.37
-0.242 to -0.191	2.29	3.82	0.04	3.93	2.44	0.25
-0.191 to -0.147	2.27	3.53	0.00	3.60	2.06	0.22
-0.147 to -0.098	2.09	3.06	-0.02	3.09	1.96	0.26
-0.098 to -0.048	1.97	2.79	-0.03	2.81	1.42	0.21
-0.048 to 0.019	1.98	2.56	-0.09	2.52	1.25	0.14
0.019 to 0.093	1.36	2.18	-0.11	2.15	1.14	0.11
0.093 to 0.210	1.24	1.80	-0.18	1.71	0.86	0.05
more than 0.021	0.85	1.32	-0.48	1.09	0.00	0.03

^a This is calculated as $e^{(0.343 - 0.17 \times FO - 0.031 \times \ln KL + 0.066 \times \ln OPEN - 0.213 \times \ln TFP - 0.169 \times \ln EMP - 0.027 \times \ln IMP)} - 1$

Figure 3: Changes in (marginal) weighted hazard rate^a of closure for exporters, 1997-2003



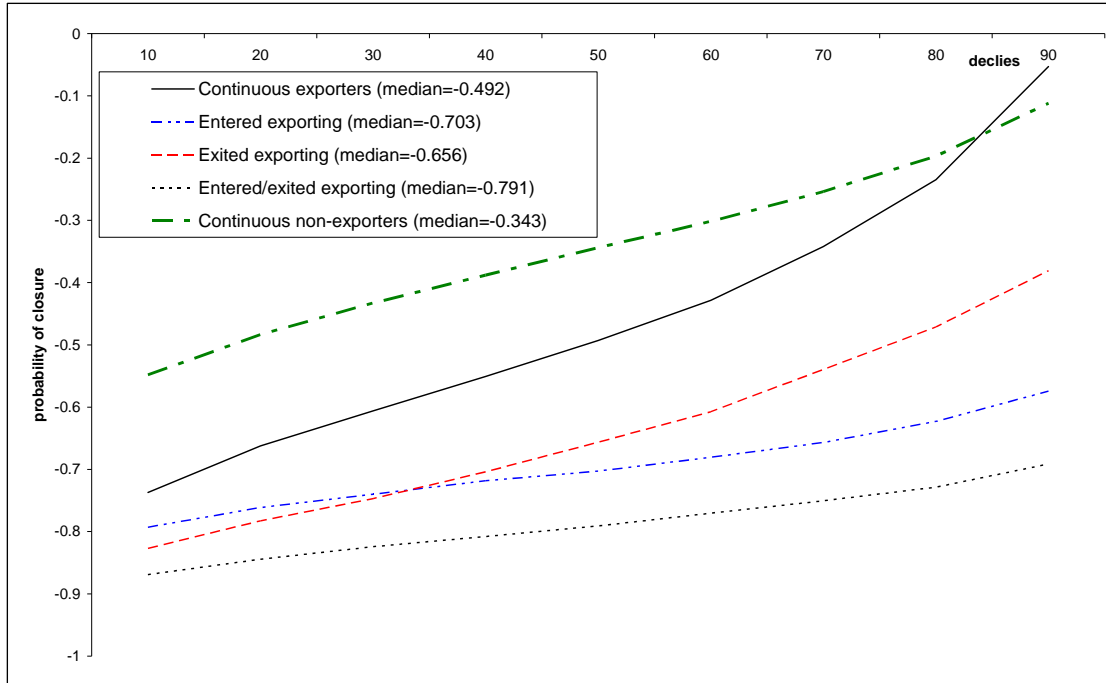
^a continuous exporters = $e^{(0.343 - 0.17 \times FO - 0.031 \times \ln KL + 0.066 \times \ln OPEN - 0.213 \times \ln TFP - 0.169 \times \ln EMP - 0.027 \times \ln IMP)} - 1$

entered exporting = $e^{(-0.756 + 0.011 \times AGE - 0.107 \times \ln KL - 0.142 \times \ln OPEN + 0.177 \times \ln EMP - 0.030 \times \ln IMP)} - 1$

exited exporting = $e^{(-0.256 + 0.010 \times AGE - 0.459 \times FO - 0.156 \times \ln EMP - 0.155 \times \ln TFP - 0.072 \times \ln IMP)} - 1$

entered/exited exporting = $e^{(-1.052 + 0.677 \times \ln GROWTH)} - 1$

Figure 4: Changes in weighted hazard rate^a of closure for various sub-groups, 1997-2003



$$^a \text{ continuous non-exporters} = e^a - 1 \text{ where } e^a = e^{[-0.131 \times FO + 0.040 \times INTGBL + 0.097 \times \ln GROWTH - 0.048 \times \ln KL - 0.009 \times \ln DISPLACE + 0.098 \times \ln OPEN - 0.072 \times \ln TFP - 0.223 \times \ln EMP - 0.004 \times \ln HERF - 0.019 \times \ln IMP + 0.009 \times (\ln GROWTH \times AGE) + 0.008 \times (\ln KL \times AGE) + 0.002 \times (\ln HERF \times AGE)]}$$

$$\text{continuous exporters} = e^{(a + 0.343 - 0.17 \times FO - 0.031 \times \ln KL + 0.066 \times \ln OPEN - 0.213 \times \ln TFP - 0.169 \times \ln EMP - 0.027 \times \ln IMP)} - 1$$

$$\text{entered exporting} = e^{(a - 0.756 + 0.011 \times AGE - 0.107 \times \ln KL - 0.142 \times \ln OPEN + 0.177 \times \ln EMP - 0.030 \times \ln IMP)} - 1$$

$$\text{exited exporting} = e^{(a - 0.256 + 0.010 \times AGE - 0.459 \times FO - 0.156 \times \ln EMP - 0.155 \times \ln TFP - 0.072 \times \ln IMP)} - 1$$

$$\text{entered/exited exporting} = e^{(a - 1.052 + 0.677 \times \ln GROWTH)} - 1$$

The first major result reported in Model 4 is that, *cet. par.*, the dummy variable for continuous exporters indicates a higher likelihood of firm closure (almost 41% higher, based on the last column which reports hazard rates). However, the variable *EXP_always* also enters the model jointly with a number of other covariates; thus Table 3 presents the deciles of the hazard rate for this variable, and provides for each decile the mean values of those variables linked to *EXP_always*. The (marginal) probability of closure is negative for firms up to the 70% decile, showing that for these sub-groups (covering nearly 70% of all continuous exporters) the negative impacts on the hazard rate of higher capital-intensity, larger employment size, higher relative TFP, and foreign ownership are

sufficiently strong to counterbalance the positive value attached to the *EXP_always* dummy variable.¹⁹ Put another way, for those sub-groups with negative hazard rates, the values of *ln* KL, *ln* OPEN, *ln* TFP, *ln* EMP, and FO are relatively large, ensuring that overall hazard rates are negative and thus the majority of firms engaged in continuous exporting have a lower probability of closure.

Other exporting sub-groups had unambiguously lower hazard rates: Figure 3 plots these showing that the average firm that entered overseas markets during 1997-2003 was 52.2% less likely to close; the average firm that ceased exporting was 43.9% less likely to close; while those firms that both entered and exited from export markets were 64.8% less likely to shutdown.

For completeness, we also report the hazard rates for all sub-groups of firms (aggregated into deciles). Figure 2 shows that the baseline survival functions for the continuous exporting and non-exporting groups are very similar. These (natural) survival rates represent the inverse of the baseline hazard, $h_0(t)$, while Figure 4 shows how the covariates shift firms away from these baseline hazards. Comparing continuous exporters and non-exporters, it can be seen that the former have relatively (much) lower hazard rates up to the 70% decile, after which there is little difference in the probability of closure. In summary, the overwhelming majority of firms that exported had lower hazard rates than non-exporters, and these rates were much lower for firms that entered and/or exited overseas markets. For new exporters, this may reflect not only the initial high productivity that comes with exporting, but also the increase in resources and capabilities

¹⁹ The role of larger ‘opening’ and ‘current’ employment size, leading to overall negative impacts on the hazard rate (based on Mata *et. al.* 1995) is discussed later. Note, also that the impact of import penetration for the continuous exporting sub-group is overall neutral (the composite dummy variable largely negates the effect shown through the parameter estimate of *ln* IMP_i).

that are needed prior to overcoming barriers to exporting; for firms that cease exporting, their lower hazard rate may reflect a positive outcome of engaging in a decision making process that results in the firm concentrating on domestic production.

The results from Model 4 in Table 2 also show that foreign-owned subsidiaries in the benchmark sub-group were (*cet. par.*) 12.3% less likely to close; those that always exported had a 28.1% lower hazard rate, while firms that ceased exporting and were also foreign-owned were 49.1% less likely to close. The latter sub-group include the UK subsidiaries of MNE's that ceased exporting once they had set up new subsidiaries in those countries they previously exported to from the UK. Overall, our results are therefore consistent with foreign-owned subsidiaries gaining from access to superior technologies and practices transferred from the parent MNE. In addition, we find that having positive intangible assets had no significant impact on the hazard rate²⁰. Recall that there have been mixed results in previous studies with respect to the impact of the intangibles, given that it can also be associated with failure attached to a higher risk, as well as enhancing the firm's internal capabilities (e.g R&D and innovation). With regard to displacement effects (through relatively high rates of entry of firms in each industry), a one standard deviation increase in \ln DISPLACE (see Table 1) actually decreases the probability of closure by nearly 3% rather than leading to displacement of existing firms. Having controlled for other attributes associated with productivity (such a size and export status), our results confirm previous studies which show that firms with higher total factor productivity are less likely to close; in this study, a one standard deviation increase in \ln TFP reduces the hazard rate by 3.1%.

²⁰ We also included \ln (1 + real value of intangible assets) instead of INTGBL, as a check on whether the volume of assets was more important, but the results were even less significant.

As to the impact of import penetration on the probability of closure, the results from Model 4 show that for non-exporters and those firms that exited from exporting increases in imports lead to a higher hazard rate (a one standard deviation in \ln IMP of around 1.9 for these firms increased the probability of closure by 3.6%). But for firms that always exported or entered overseas markets, the overall impact of increased import penetration was to lower the hazard rate (with a one standard deviation increase in \ln IMP, the effect was between 1.7% and 2.1% lower). Our results therefore suggest that for new and continuing exporters, higher levels of intra-industry trade are beneficial.

Certain other covariates impact on the probability of closure in conjunction with the age of the firm; Table 4 reports the parameters of the hazard rate ($e^{\hat{\beta}} - 1$) for just those parameters with significant values when age effects are important (four n -tiles were used for the AGE variable based on equal numbers of firms in each sub-group). Rather than provide a cushion to firms, output growth at the industry level increases the hazard rate of closure, but at a much higher rate for the oldest firms (over 60% higher) who are more vulnerable to an expansion of the market (which is consistent with the theoretical predictions on the role of market size in Asplund and Nocke, 2006).²¹ In line with theory and past empirical work, increases in capital intensity (a proxy here for sunk costs and thus barriers to entry and exit) reduced the hazard rate, but this effect diminished with the age of the firm (a one standard deviation increase in \ln KL reduced the hazard rate by 8.2% for the youngest firms, but by only 2.6% for the oldest). Older firms also had lower negative hazard rates associated with entering or exiting export markets. In all, Table 4 shows that when age impacts on hazard rates through other covariates, the general pattern

²¹ Note we also tried entering \ln DISPLACE \times \ln GROWTH, but the interaction term was not significant.

is that older firms have relatively higher rates of closure. Thus, conditional on having controlled for other covariates associated with productivity (such as size and TFP), we find in this study that the oldest UK firms (over 22 years old) are often more vulnerable to closure. This may seem at odds with other studies, but it should be noted (Disney *et. al.*, 2003a) that in most theoretical models the efficiency level attached to the exit threshold that firms must exceed to survive increases with age, and thus the dependence of hazard rates on age is therefore an empirical matter.²²

Table 4: Mean parameter estimates $e^{\beta}-1$ by age sub-group in years (based on Table 2)

	AGE <= 5	5 < AGE <= 11	11 < AGE <= 22	AGE > 22
$\ln \text{GROWTH}_t$	0.136	0.186	0.273	0.618
$\ln \text{KL}_{it}$	-0.045	-0.041	-0.034	-0.013
$\text{EXP}_{\text{entry}_{it}}$	-0.513	-0.486	-0.440	-0.244
$\text{EXP}_{\text{exit}_{it}}$	-0.200	-0.161	-0.094	-0.179
$\ln \text{HERF}_t$	0.004	0.015	0.034	0.096

Counter to expectations, initial size has a positive impact on the hazard rate of closure. However, it is possible (following Mata *et. al.* 1995) to reformulate the results and consider the relationship between initial size and growth in employment since the firm opened. This results in the relationships shown in Table 5, showing that initial size does have a negative impact and assuming that the firm grows from this initial level over time there is an additional negative impact on the probability of closure. These results are therefore in line with past studies. Current size is more important than initial size for all sub-groups (except firms that started exporting), indicating that *cet. par.* larger firms are less likely to close (with values for $(e^{\beta}-1)$ in Table 5 showing that for a one standard

²² Studies of family-owned firms (which make up the majority of SME's in the UK – see Harris and Reid, 2008) have shown that they often are susceptible to failure and closure when the founder retires (see Santarelli and Lotti, 2005).

deviation increase in \ln EMP, the probability of closure ranges from about 6.9% lower for entrants into export markets to 65.2% lower for firms that continuously export).²³

Table 5: Changes in weighted hazard rate of closure due to firm size (initial and current)

	$\hat{\beta}$ [\ln OPEN _i]	z-value	$\hat{\beta}$ [\ln (EMP _{it} ÷ OPEN _i)]	z-value
Benchmark ^a	-0.125	-10.88	-0.223	-9.03
EXP_always	-0.228	-13.32	-0.393	-10.57
EXP_entry	-0.090	-2.27	-0.046	-0.58
EXP_exit	-0.281	-6.22	-0.379	-7.74

^a Continuous non-exporters and firms that both entered/exited from export markets.

5. Summary and conclusions

In an attempt to add to the small but growing body of evidence on firm closure (particularly in an international context), the weighted *FAME* panel data for 1997 to 2003 has been used to estimate a Cox proportional hazard model of firm closure in an open economy. Our major findings show that exporting firms have better survival prospects, conditional on controlling for other factors linked to productivity: the majority of firms engaged in continuous exporting have a lower probability of closure (vis-à-vis those producing for domestic market solely); while the hazard rates are unambiguously much lower for firms that have entered and/or exited overseas markets. We also find that foreign-owned subsidiaries are less likely to close, while having positive intangible assets has no significant impact on the hazard rate; unexpectedly, displacement effects (through relatively high rates of entry of firms in each industry) decrease the probability of closure

²³ Note, the intercept dummies for each 3-digit industry are not reported here but these show that there are significant and large differences across sectors. For instance, firms belonging to the textile, leather & clothing and repair/sale of motor vehicles sectors were 64.6% and 35.4%, respectively, more likely to exit; *cet. par*, firms in the non-metallic minerals and financial intermediation sectors were 28.4% and 47.6%, respectively, less likely to close during the 1997-2003 period. Firms located in Scotland, Wales and Northern Ireland were between 24.1 – 52.1% more likely to close.

rather than leading to displacement of existing firms. Having controlled for other attributes associated with productivity (such a size and export status), firms with higher total factor productivity are less likely to close. While increased import penetration (which is associated with lower trade costs) leads to a higher hazard rate for those firms that have never participated in international markets or that exit from exporting, interestingly, for firms that have always been exporting or just entered overseas markets, increased import penetration leads to a lower hazard rate. Output growth at the industry level increases the hazard rate of closure, but at a much higher rate for the oldest firms who are more vulnerable to an expansion of the market. In addition, increases in capital intensity (a proxy here for sunk costs and thus barriers to entry and exit) reduce the hazard rate.

Overall, and conditional on having controlled for other covariates associated with productivity (e.g. size and capital intensity), we find in this study that the oldest UK firms (over 22 years old) are often more vulnerable to closure. This is consistent with theoretical models showing that the efficiency level attached to the shut-down threshold that firms must exceed to survive increases with age; and the initial employment size of a firm when it starts-up has a negative impact and assuming that the firm grows from this initial level over time there is an additional negative impact on the probability of closure. Lastly, industry differences (although not explored in any detail here) are also very important in determining the probability of which firms close and which continue in business.

As to the policy dimension, it was stated in the introduction that if closure depends on trade costs, public policies might concentrate on reducing these costs. By contrast, if

firm closure is the result of large differences in productivity between exporters and non exporters, then policies that concentrate on facilitating entry may not generate lasting increases in export participation if they are not accompanied by improvements in firms' productivity. Overall, our results point to firm heterogeneity and thus differences in productivity (as measured through a number of variables) being one of the main determinants of firm survival, suggesting that export promotion will be most successful if it can help firms to become more productive (rather than to just combat market failures associated with a lack of information on export markets, or similar barriers to internationalisation). However, we have also found that lowering trade costs (and thus increasing import penetration) would both encourage (better) exporters and increase the probability of non-exporters (with lower productivity) leaving the market place. That is, increased global competition should overall push-up UK aggregate productivity (through a so-called 'batting-average' effect).

Appendix

Weighting the FAME data

Table A.1: GB^a turnover (£m) in 2003 based on *FAME* and *ARD* by size-bands

Size-band ^b	<i>FAME</i>				<i>ARD</i>	<i>FAME</i>	
	Unweighted	%	Weighted	%		%	<i>ARD</i>
	(1)		(2)		(3)		(2) ÷ (3)
<44	26.9	0.0	460.1	0.0	509.7	0.0	90.0
44 to 227	380.7	0.0	7,969.6	0.7	8,509.4	0.8	93.5
>227 to 1184	3,578.5	0.2	60,892.8	5.3	63,751.8	5.9	95.5
>1184 to 7244	49,683.9	2.4	198,417.1	17.4	200,988.3	18.5	98.7
>7244	1,988,609.3	97.4	874,543.9	76.5	812,157.0	74.8	107.5
All	2,042,279.4	100.0	1,142,283.4	100.0	1,085,916.2	100.0	105.0

^a Unweighted *FAME* data covers the UK ^b Size-bands are in £'000 Source: *ARD* and *FAME* databases

Obtaining estimates of TFP

To obtain estimate of total factor productivity (TFP), we estimate (using a system-GMM panel approach that takes into account endogeneity and sample-selection issues through appropriate use of instrumental variables) an augmented production function as follows:

$$y_{it} = \alpha_0 + \alpha_E e_{it} + \alpha_M m_{it} + \alpha_K k_{it} + \alpha_T t + \gamma X_{it} + \varepsilon_t \quad (\text{A.1})$$

where y , e , m , and k refer to the logarithms of real gross output, employment, intermediate inputs and tangible assets in firm i in time t . We have also included a vector of variables, X , that determine TFP; hence TFP growth in this instance is defined as (dropping sub-scripts):

$$\ln TFP = \hat{\alpha}_T + \hat{\gamma} \dot{X} \equiv \dot{y} - \hat{\alpha}_E \dot{e} - \hat{\alpha}_M \dot{m} - \hat{\alpha}_K \dot{k} \quad (\text{A.2})$$

Details (and results) are provided in Harris and Li (2007).

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