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Foreign-owned Plants and Job Security

by

Martyn Andrews, Lutz Bellmann, Thorsten Schank and Richard Upward

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The Authors

Martyn Andrews is a Senior Lecturer in Economics at the University of Manchester. Lutz Bellmann is a Senior Researcher at the Institut für Arbeitsmarkt und Berufsforschung, Nürnberg. Thorsten Schank is Researcher at the Chair of Labour and Regional Economics at the Friedrich-Alexander-University, Erlangen-Nürnberg. Richard Upward is an Associate Professor in the School of Economics and a GEP Research Fellow.

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Abstract

We investigate the hypothesis that workers in foreign-owned establishments face greater job insecurity. Using linked employer employee data from Germany, we examine whether foreign-owned establishments are more likely to exit production, and whether workers in foreign-owned establishments face higher separation rates, net of establishment exit. We find that, after controlling for the different characteristics of foreign and domestic establishments, foreign establishments have higher exit rates and higher separation rates, but the effect is quantitatively small. In contrast, foreign-owned establishments which do not export appear to have considerably higher exit rates and separation rates.

JEL classification: F23, J63, C41

Keywords: Foreign ownership, job security, linked employer-employee data

Outline

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2. *Previous literature*
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4. *Methods*
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Non-Technical Summary

One often hears claims that the labour market has become more insecure. It has been suggested, for example, that instead of a "job for life", workers face the prospect of having a string of short-lived jobs over the course of their careers. One possible explanation for this reduction in job security is the increasing importance of foreign-owned firms in OECD labour markets.

For example, it has been argued that employment in foreign-owned firms is more volatile than employment in domestic firms because foreign-owned firms have more flexibility in choosing where to locate production. If a firm has production plants in Germany and China, an increase in the costs of production in Germany might cause the firm to lay off workers in the UK and expand production abroad. In the extreme case, a multinational might actually shut-down an entire plant in one country. If this were the case, we would find that (a) foreign-owned plants are more likely to close and (b) jobs in foreign-owned plants are more likely to end.

In this paper we investigate both these issues. We follow a large sample of plants in West Germany from 2000 until 2005. We also follow the workers in those plants over the same period. We measure the probability that plants shut down, and the probability that workers leave their jobs.

We find that foreign-owned plants actually have *lower* closure rates, and *lower* job separation rates. But this is partly because foreign-owned plants are larger and more profitable. When we compare plants of equal size and profitability, we find that foreign-owned plants have slightly higher closure and job separation rates. However, the effect is very small, and does not seem to justify growing fears that globalisation makes the labour market a more insecure environment.

1 Introduction

Scheve & Slaughter (2004) argue that "foreign direct investment by multinational enterprises is the key aspect of [international] integration generating risk." Essentially, employment in foreign-owned firms is thought to be more volatile because foreign-owned firms can more easily shift production between locations. A related argument is made by Görg & Strobl (2003), who shows that foreign-owned plants are more likely to exit the market than similar domestically-owned firms.

Employment volatility and plant exit inevitably cause worker turnover. Workers who lose their jobs in foreign-owned plants are forced to move to other plants, many of which will be domestically-owned. A separate literature has suggested that such worker turnover can benefit the host economy, because human capital gained in the foreign-owned firm is transferred to domestic firms. For example, Blomstrom & Kokko (2003) suggest that the movement of workers is an important channel by which productivity "spillovers" occur. However, as noted by Fosfuri, Motta & Ronde (2001), evidence on spillovers due to workers' mobility is "scarce and far from conclusive."

In this paper we investigate both these issues. Using data from a large survey of German establishments linked to the workers in those plants, we examine: first, whether foreign-owned establishments are more likely to exit; second, whether workers in foreign-owned establishments have higher separation rates.

We find that, after controlling for the different characteristics of foreign and domestic establishments, foreign establishments have higher exit rates and higher separation rates, but the effect is quantitatively small. In contrast, foreign-owned establishments which do not export appear to have considerably higher exit rates and separation rates.

In Section 2 we discuss the related literature in more detail, and explain how our data and methods contribute. In Section 3 we describe the data and how we construct our measures of establishment exit and worker turnover. The econometric method we use for both measures is discussed in Section 4. Our results are presented in Section 5, and Section 6 concludes.

2 Previous literature

2.1 Foreign ownership and plant exit

There is a substantial literature on the determinants of firm (or establishment) success and failure, where failure is defined in terms of exit from the market. Various theoretical models suggest that larger firms and older firms will have lower hazard rates (Jovanovic 1982, Hopenhayn 1992), and this is largely borne out in the empirical findings.

Early studies by Dunne, Roberts & Samuelson (1988, 1989) provide descriptive evidence on the proportion of plants which exit over a five year period as a function of their industry and size. They find that failure rates decline with current size and the age of the plant. Studies such as Wagner (1994), Mata & Portugal (1994), Audretsch & Mahmood (1995) and Disney, Haskel & Heden (2003) have used duration models to estimate the probability of plant exit per period conditional on survival up to that period. The failure rate of firms is found to be declining in duration (new firms are most likely to fail). Important explanatory variables for exit include size, whether the firm has multiple plants and various measures of market structure (such as industry concentration).

A small number of studies have included the nationality of ownership as a regressor in a model of exit. Görg & Strobl (2003) find that manufacturing plants in Ireland which are owned by foreign multinationals actually have *lower* exit rates. However, this does not account for the fact that foreign-owned plants have characteristics typically associated with lower exit rates. For example, foreign-owned plants tend to be larger. Once these factors are accounted for, Görg & Strobl find that foreign-owned plants have higher exit rates.

Bernard & Sjöholm (2003) find similar results for a sample of plants in Indonesia. In the raw data, foreign-owned plants are far less likely to shut down than domestic plants. Once again, however, this is because foreign-owned plants tend to be larger and more productive. Controlling for plant size and productivity, foreign-owned plants are significantly more likely to close than domestic plants.

These results are also confirmed by Bernard & Jensen (2007) for the U.S. Plants which are part of multinational enterprises have lower exit rates, but tend to be larger, older and more productive. Once these characteristics are taken into account, plants belonging to multinationals have higher exit rates.

Alvarez & Görg (2005) use plant-level data for Chile, and find that although foreign plants are more likely to exit, controlling for other firm and industry characteristics, it is only domestic market oriented multinationals which are more “footloose”. Foreign-owned plants which export do not have higher exit rates.

An important question raised by these studies is the extent to which the characteristics of foreign-owned plants which are associated with lower exit rates (e.g. size and productivity) are the result of an earlier change of ownership, or whether multinationals simply takeover plants which already had these characteristics. Clearly, the benefits of foreign ownership to the domestic economy depends very much on which of these is the more important.

All of these studies use firm- or plant-level data, and therefore are not able to examine whether the characteristics of a firm’s workforce are important determinants of exit. One contribution we are able to make by using linked worker-firm data is to include detailed measures of the firm’s human capital in the exit equation.[1]

2.2 Foreign ownership and worker turnover

To the best of our knowledge there are no studies which use individual-level data on worker separation rates to examine whether turnover is higher in foreign-owned establishments. There is, however, a large literature which examines the determinants of worker separation rates more generally. Two important theoretical frameworks are the job matching literature (e.g. Jovanovic (1979)) and the literature on firm-specific human capital, which dates back at least to Becker (1962). Both of these frameworks predict that the probability of separating declines with job tenure, although for different reasons. This prediction is consistently borne out by the empirical evidence, see for example Anderson & Meyer (1994, Table 7).

There is also a recent theoretical literature on human capital spillovers between foreign and domestic firms. Human capital spillovers can be defined as those spillovers which occur because of training of workers in foreign-owned plants, and the subsequent movement of workers between plants. In addition Blomstrom & Kokko (2003) notes that beneficiaries might also include employees of suppliers, subcontractors and customers of MNCs.

Fosfuri *et al.* (2001) study the theoretical conditions under which spillovers occur. Spillovers occur when foreign-owned plants train workers who later join domestic plants. But they note that “. . . evidence on spillovers due to workers’ mobility is scarce and far from conclusive.” Glass & Saggi (2002) develop a model in which foreign-owned firms offer higher wages to *prevent* turnover. They argue therefore that turnover (and hence economic insecurity)

[1]Abowd, McKinney & Vilhuber (2007) is another recent paper which uses linked employer-employee data to relate the probability of plant exit to the human capital of its workforce.

will be lower in foreign-owned plants. There is also some evidence from case studies in developing countries (Gershenberg 1987) that foreign-owned firms offer more training to technical workers and managers than local firms, and that the mobility of managers in these plants is actually much lower, reducing the potential for spillovers.

One of the few papers to (even indirectly) estimate the extent to which worker mobility contributes to the productivity of domestic firms is Görg & Strobl (2005). Using data from a sample of manufacturing firms in Ghana, they find that firms whose entrepreneur previously worked in multinationals in the same industry are more productive than other domestic firms.

In contrast, however, there is also a distinct literature which suggests that the actions of multinational firms may be associated with greater job turnover. As noted, Scheve & Slaughter (2004) provide some general evidence that foreign direct investment activity (both inward and outward) is positively associated with workers' perceived job insecurity.

Fabbri, Haskel & Slaughter (2003) argue that multinational firms have more elastic labor demands than domestic firms, which would be consistent with higher worker turnover rates. They present industry-level evidence for the U.K. and U.S. which shows that the labor demand elasticity for unskilled workers has increased over a period in which multinational activity has also expanded. Firm-level evidence is provided by Navaretti, Turrini & Checchi (2003), who estimate dynamic labour demand equations across 11 European countries. They show that although foreign-owned firms adjust labour faster than domestic firms, the total size of the adjustment is actually smaller. This may however, reflect the fact that foreign-owned firms have a more skilled labour force, and hence a lower labour demand elasticity.

Our contribution is to examine whether individual workers in foreign-owned firms are actually more "insecure" than similar workers in domestic firms. This is important because it relates both to whether human capital spillovers are an important feature of multinational activity, and to whether multinational firms contribute to greater economic insecurity.

3 Data description

There are two data sources. The first is the *Institut für Arbeitsmarkt- und Berufsforschung (IAB) Establishment Panel*, an annual survey of approximately 8,250 plants located in the former West Germany and an additional 7,900 plants in the former East Germany. The survey started in 1993 and is ongoing. It covers 1% of all plants and 7% of all employment in Germany, and is therefore a sample weighted toward larger plants. Information is obtained by personal interviews with plant managers, and comprises about 80 questions per year, giving us information on, for example, total employment, bargaining arrangements, total sales, exports, investment, wage bill, location, industry, profit level and nationality of ownership. Ownership is defined as either West German, East German, foreign, or public.[2] Complete information on plant ownership is available for all plants only in 2000 and 2004. Plants which enter the sample between 2000 and 2004 also have information on plant ownership recorded. A detailed description of the IAB Establishment Panel can be found in Kölling (2000).

The second source of data is the employment statistics register of the German Federal Office of Labour (*Beschäftigtenstatistik* or BS), which covers all workers or trainees registered by the social insurance system. The register covers about 80% of workers in West Germany and about 85% in East Germany. Information on workers includes basic demographics, start and end dates of employment spells, occupation and industry, earnings, qualifications (school and post-school), and a plant identification number. A detailed description of the employment data can be found in Bender, Haas & Klose (2000).

We restrict the analysis to plants in the private sector. Almost all workers in the private sector will be covered by the social insurance system, so we would expect that the linked data covers nearly 100% of workers.

By using the plant identification number we can associate (or 'link') each worker with a plant in the panel. We therefore observe almost all workers in about 14,000 plants each year. Because the employment register is spell-based (one record for each employment spell), the combined data is potentially complex. To simplify, we select all workers in the employment register who are employed by the surveyed plants on June 30th each year. This yields an unbalanced annual panel of workers together with detailed information on the plants in which they work. We refer to the linked data as the Linked IAB Panel, or LIAB; the data discussed below is our regression sample and not the LIAB in general.[3]

Figure 1 presents a stylised picture of the data. It shows four plants, and some workers in each plant.

[2]The relevant question is: "Is the establishment mainly or solely in: (a) West German ownership (b) East German ownership (c) Foreign ownership (d) Public ownership (e) No single owner which holds majority?" Our analysis considers only plants under (a)-(c).

[3]In this draft, we only analyse West German plants.

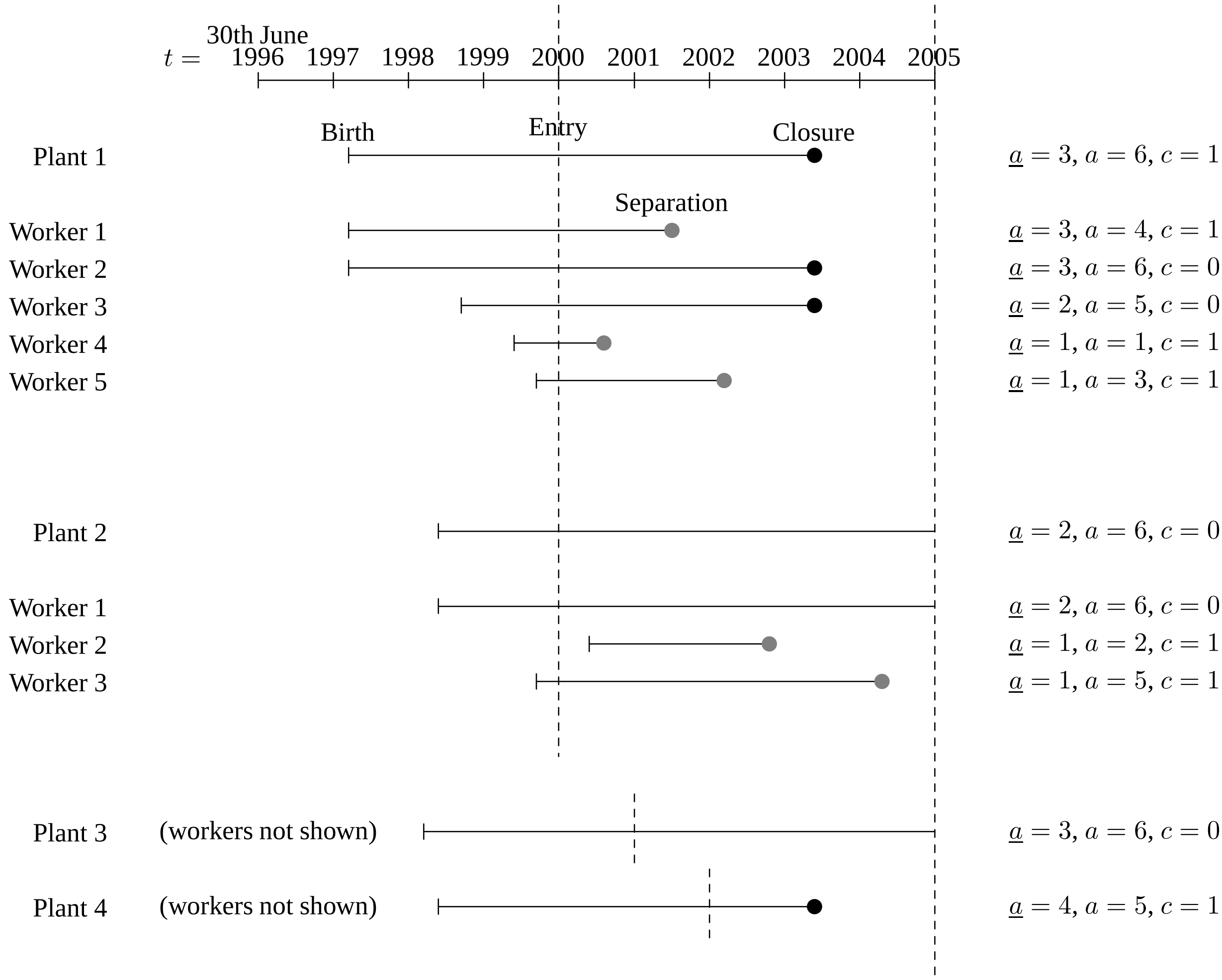


Figure 1: Structure of the data

3.1 Plants

In the regression sample, plants can either shut-down (or "close") or be right-censored by the end of sample (30/6/2005). By construction, there is no attrition (another form of right-censoring) in that no plants leave the sample before 30/6/2005. In Figure 1, Plants 1 and 4 close whereas Plants 2 and 3 are right-censored on 30/6/2005.

We ensure that there is no attrition from the sample by making use of the linked data. When a plant attrits from the IAB panel we search the BS to find out if it is still employing any workers. A plant which attrits is thus followed until it stops employing any workers in the BS, or until 30/6/2005, when it is censored. This is potentially important, because it seems likely that attrition from the IAB panel is non-random with respect to the durations of interest i.e. plant closure and worker separations.

It is the duration to plant closure that we seek to model in this paper (together with worker separations, discussed below). However, identifying such plants is not straightforward. In administrative databases it is usual to assume that a plant (or a firm) has closed when its identification number disappears from the data. But this is problematic because plants may change their identification number while remaining in production. This may occur because of some re-organisation, perhaps because plants are sold from one firm to another. However, because our data are a combination of a detailed plant-level survey and an administrative population, we can be more precise about the definition of plant closure.

There are two pieces of information which are informative regarding plant closure. The first is whether or not the plant identification number remains in the Beschäftigtenstatistik. The second is the interview outcome from the IAB panel. The interview outcome essentially takes three values: (1) "Same plant as last year" (2) "Plant still exists but has left panel" (3) "Plant has closed".

If the interview outcome indicates that the plant has closed (3), we assume that the plant has genuinely closed whether or not a plant with the same identification number continues to exist in the Beschäftigtenstatistik. Similarly, if the interview outcome indicates the "same plant as last year" (1), we assume that the plant has not closed even if the plant identification number disappears from the Beschäftigtenstatistik. The third possibility is that the plant still exists but has left the panel (2); here we are forced to use the existence of the plant identification number to determine whether the plant has closed or not. The dummy variable c_j takes the value 1 if the plant closes during the sample period and 0 otherwise. We also denote the age at which it exits the data as a_j, the integer value observed on the preceeding June 30th. In the figure, Plant 1 closes between 30/6/2003 and 30/6/2004; Plant 2 is right-censored on 30/6/2005.

Plants can either be observed on 30/6/2000 (e.g. Plants 1 and 2) or they enter the data afterwards (e.g. Plants 3 and 4). The IAB increased the sample size considerably over time, so for many plants the observation period starts after 30/6/2000.

The age of a plant is calculated by finding the year in which the plant first appears in the Beschäftigtenstatistik.[4] We denote the age (or elapsed duration) at which the plant enters the sample as $\underline{a}_j$. Most plants are born some time before the start of the sample period, and for these plants $\underline{a}_j > 1$ when they are first observed. Thus we have left-censoring or *late entry*.

[4]The earliest year in the Beschäftigtenstatistik is 1977, and so age will be underestimated.

Figure 2 clarifies exactly how completed duration is measured. Plant 1 is born at some point between 1/7/1997 and 30/6/1998. It is sampled in the IAB panel on 30/6/2000, at which point its age (elapsed duration) is between 2 and 3 years, which is rounded up to the nearest integer ($\underline{a}_1 = 3$). The plant closes ($c_1 = 1$) at some point between 1/7/2003 and 30/6/2004, so the last observation of the plant is on 30/6/2003 when its completed duration is $a_1 = 6$. The corresponding data for the other three plants are given in Figure 1. For Plant 2, it is ($\underline{a}_2 = 2, a_2 = 6, c_2 = 0$). Plant 3 is different, in that it enters the data on 30/6/2001. It was born between 1/7/1998 and 30/6/1999 and so its age on entering the sample is $\underline{a}_3 = 3$. It is right-censored ($c_3 = 0$) at the end of the sample, which means that its completed duration is $a_3 = 6$. The corresponding data for Plant 4 is ($\underline{a}_4 = 4, a_4 = 5, c_4 = 1$).

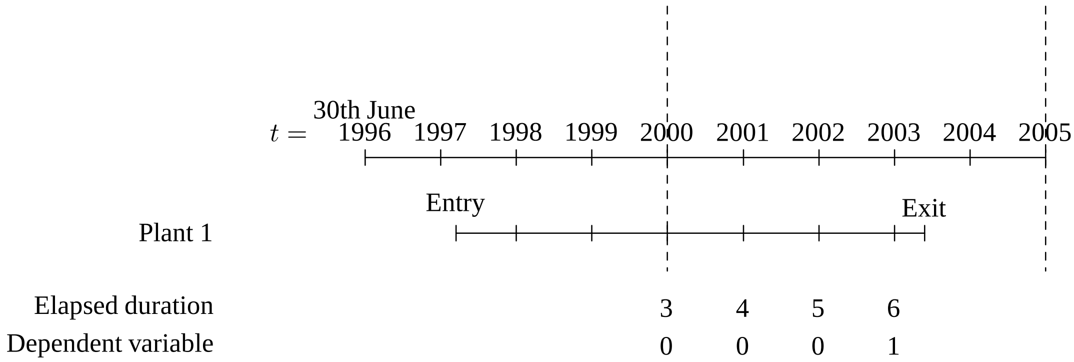


Figure 2: Definition of elapsed duration

Table 1 shows the basic movement of plants in and out of the sample. The number of plants that enter the panel each year is much bigger than the number that leave it. The total number of entrants is 3,378 whereas the number of closures is 1,399. As noted, the sample grows over time. When the 1,399 closures is divided by 31,855 plant-years at risk, this gives a raw hazard to closure of 0.0439. In other words, on average, of all plants alive at a certain date, within a year 4.4% of them will have closed. 1/0.0439 gives a mean completed duration of 22.8 years: this is measured from entry into the panel.

TABLE 1 HERE

There are 4,209 plants sampled on 30/6/2000 and another 4,738 enter afterwards. Mean completed duration can be computed for the last year in which these 8,947 plants are observed: it is 14.9 years for those that close and 17.6 for those who do not.[5]

[5] As noted earlier, this will be an underestimate because some plants are born before the start of the Beschäftigtenstatistik.

3.2 Workers

For each plant j we observe all workers $i = 1, \ldots, N_j$ who are in the Beschäftigtenstatistik. Workers can either exit the data or they can be censored by end of sample. Because there is no attrition for plants, there is no attrition in the worker data. In other words, when a worker exits the data, it is because his tenure with his employing plant has come to an end. This can happen for one of two reasons, namely the plant closes or because the worker *separates* from the plant. It is a feature of these linked data that we are able to distinguish between these two events. Of course, a separation can either be a quit or a fire, but we do not observe this.

For example, in Figure 1, Plant 1 employs 5 workers on 30/6/2000. Three of these workers ($i = 1, 4, 5$) separate from the plant during the sample period. The two remaining workers leave the plant because the plant closes. As with the plant, the variable $\underline{a}_i$ records the elapsed duration of the worker employed in the plant when she enters the sample, and a_i records completed duration. For workers, however, the relevant durations are the length of time since joining the plant (tenure) rather than the age of the plant. The dummy variable c_i takes the value 1 if the worker separates from the plant during the sample period and 0 otherwise. Note that the completed duration to separation for workers 2 and 3 in plant 1 is censored because the plant closes before this event can occur.

Workers may join plants when the plants first appear in the Beschäftigtenstatistik, or they may be hired afterwards. For example, in Figure 1, workers 1 and 2 in plant 1 join the plant when it is born, while workers 3, 4 and 5 are hired later.[6]

TABLE 2 HERE

Table 2 shows that 517,909 workers exited the data; of these, 474,749 workers separated from their plants (quits or fires) and 43,160 workers exited because their plant closed. In other words, 8.3% of exits are because of plant closure. It is the hazard to separation that we seek to model. Given 474,749 worker separations, when divided by 3,950,506 worker-years at risk, the raw worker hazard to separating is 0.1202. Similarly, the raw worker hazard to plant closure is 0.0109. 758,556 workers entered the panel: 328,028 of them were hired and the worker's plant entering the panel accounts for the rest. This gives a worker hiring rate

[6] Note that whether or not workers are hired before 30/6/2000 makes no difference to their elapsed duration when they are first observed in a plant because we observe their tenure from the Beschäftigtenstatistik. Worker 2 in plant 2 joins the plant after 30/6/2000 and first appears in the data in 2001 with $\underline{a}$=1. Worker 3 in plant joins the plant just before 30/6/2000, but still has $\underline{a}$=1 in 2000.

of 328,028/(3,950,506−570,949) or 0.0971.[7] This suggests that the sample period is one of contraction amongst existing firms in the West German labour market. Note that there we observe 1,329,505 workers, ie the 758,556 workers observed on 30/6/2000 and the 570,949 workers who enter afterwards.

4 Methods

Plant hazards

We wish to model the probability that a plant closes as a function of its age (elapsed duration) and whether it is foreign-owned. Our data comprise a panel of plants $j = 1, \ldots, J_t$ observed annually on 30 June for $t = 2000 \ldots 2004$. The appropriate econometric framework is that of discrete-time duration models, because a plant may close at any point between 1st July in year t and 30th June in year $t + 1$, but we do not observe the precise date on which this happens.

The fundamental concept relating plant age to closure is the hazard function. This has been used both in the general analysis of plant closure (Audretsch & Mahmood 1995) and in the analysis of foreign ownership on plant closure (Görg & Strobl 2003). The hazard for plant j, h_{ja}, is defined as the probability that a plant closes at some point between age (elapsed duration) $a - 1$ and a, conditional on having survived to age $a - 1$:

$$h_{ja} = \Pr(a - 1 < A_j \leq a \mid A_j > a - 1) \quad a = 1, 2, \ldots$$

where A_j is the latent age of the plant j. Recall that we denote the completed duration for plant j as a_j. Also recall that for most plants in the sample we have delayed entry (i.e. $\underline{a}_j > 1$). Our econometric methods need to take account of this delayed entry. We also need to deal with the more common problem that the sample ends before all plants close (right-censoring). Standard references are Wooldridge (2002), Cameron & Trivedi (2005) and Jenkins (2005). It is Jenkins that we use here.

To start, consider the standard case where all plants are observed from birth, but for some plants there is right censoring. The log-likelihood function for plant j is given by (see Jenkins

[7] $N_{2000} = 570,949$ is excluded from the numerator because it is not part of the risk set. This is because we do not observe hires in 2000.

2005, Eqn (6.10)):

$$L_j = h_{ja_j}^{c_j}(1 - h_{ja_j})^{(1-c_j)} \prod_{k=1}^{a_j-1}(1 - h_{jk}) \tag{1}$$

Recall that the indicator variable $c_j = 1$ if a plant closes and zero otherwise. The likelihood for a plant which closes at age a_j is $(1 - h_{j1})(1 - h_{j2}) \ldots (1 - h_{j,a_j-1})h_{ja_j}$, whereas the likelihood for a plant which does not close at age a_j is $(1 - h_{j1})(1 - h_{j2}) \ldots (1 - h_{ja_j})$. (Suppose we observed Plant 1 from birth in Figure 2, then $L_1 = (1 - h_{11})(1 - h_{12})(1 - h_{13})(1 - h_{14})(1 - h_{15})h_{16}$.)

A standard approach for estimating this model is to expand the data so that each plant contributes a_j rows. Define a binary indicator variable y_{jk} which equals one if plant j closes at age k and zero otherwise. If $c_j = 1$ then $y_{ik} = 1$ in the last row $k = a_j$ and zero in all other rows. If $c_j = 0$ then y_{ik} equals zero in all rows. (If Plant 1 were observed from birth, it would have 6 rows of data, with y_{1k} as $\{0,0,0,0,0,1\}$.) We can then write the log-likelihood for plant j as

$$\log L_j = \sum_{k=1}^{a_j}[y_{jk} \log h_{jk} + (1 - y_{jk}) \log(1 - h_{jk})].$$

This is the likelihood for a binary dependent variable, and models can be estimated using standard software. To model the effect of covariates on the hazard rate, it is usual to adopt the proportional hazards assumption. Define a vector of covariates $\mathbf{x}_{jk}$ which characterise each plant. $\mathbf{x}_{jk}$ includes a measure of foreign ownership. It also includes some worker-level covariates that have been averaged to the plant-level. Under the proportional hazards assumption, the precise form of the hazard is given by the complementary log-log link function:

$$h_{jk} = 1 - \exp(-\exp(\mathbf{x}'_{jk}\boldsymbol{\beta} + \gamma_k)) \quad k = 1, \ldots, a_k.$$

The γ_k terms are interpreted as the log of a non-parametric piecewise linear baseline hazard. The notation $\mathbf{x}'_{jk}$ explicitly acknowledges that the covariates may vary with elapsed duration or calendar time.

We now deal with the problem of late entry. As noted, most plants in our sample have $\underline{a}_j > 1$ and so have already been at risk of closing for some time, depending on their age. The implication of this is that one is more likely to observe long rather than short durations. This is a classic sample selection problem. To deal with this, the likelihood needs amending. The likelihood contribution for plant j is now

$$L_j = \frac{h_{ja_j}^{c_j}(1-h_{ja_j})^{(1-c_j)}\prod_{k=1}^{a_j-1}(1-h_{jk})}{S_j(\underline{a}_j - 1)}$$

This is the likelihood given in Equation (1), but also conditions on survival up to age $\underline{a}_j - 1$, where $\underline{a}_j$ is the age at which plant j enters the data. But the denominator can be written as

$$S_j(\underline{a}_j - 1) = \prod_{k=1}^{\underline{a}_j - 1}(1-h_{jk})$$

and this leads to the convenient cancelling result (Jenkins 2005) so that

$$L_j = h_{ja_j}^{c_j}(1-h_{ja_j})^{(1-c_j)}\prod_{k=\underline{a}_j}^{a_j-1}(1-h_{jk})$$

and so the log-likelihood becomes

$$\log L_j = \sum_{k=\underline{a}_j}^{a_j}[y_{jk}\log h_{jk} + (1-y_{jk})\log(1-h_{jk})]. \tag{2}$$

This is very similar to the standard expression, except that the summation runs from the age of the plant when it enters the data. (Because Plant 1 survived until 30/6/2006, its likelihood contribution is $L_1 = (1-h_{13})(1-h_{14})(1-h_{15})h_{16}$ and $y_{1k} = \{0,0,0,1\}$.)

Worker hazards

The great advantage of linked employer-employee data is that we can also observe separations at the employee level, as well as plant closure. This is important because separations occur even when plants do not close. As noted in Section 2, various theories suggest that foreign-owned plants might have higher (or lower) labour turnover as well as differential closure rates. We define a worker separation to occur when worker i leaves plant j, but plant j does not close.

The appropriate econometric framework for worker separations is one which relates the probability of separating from plant j to elapsed time in that plant, or the worker's tenure. Thus, the econometric model is almost the same as that used for plant closure, except that elapsed and completed durations are measured by tenure rather than plant age. As already discussed (see also Figure 1), one important additional feature is that the duration to separation can be censored by two possible events. As with plants, the first censoring event is the end of the sample period. The second censoring event for workers is that plants may close before a separation can occur.

Workers' tenure is not zero when a plant enters the data for most workers, because they joined their plant earlier. Thus we also have late entry in the worker data, as well as a standard case where the worker is hired by the plant after the plant enters the sample. The likelihood developed earlier for plants applies to both types of worker. To estimate the hazard to separation the log-likelihood for worker i is therefore

$$\log L_i = \sum_{k=\underline{a}_i}^{a_i} [y_{ik} \log h_{ik} + (1 - y_{ik}) \log(1 - h_{ik})]$$

where $\underline{a}_i$ is worker i's elapsed tenure when the plant is first observed, a_i is his completed tenure at the time of separation (or the end of the sample), y_{ik} is the dummy indicating separation in row k and h_{ik} is the corresponding hazard function:

$$h_{ik} = 1 - \exp(-\exp(\mathbf{x}'_{ik}\boldsymbol{\beta} + \gamma_k)) \quad k = 1, \ldots, a_k.$$

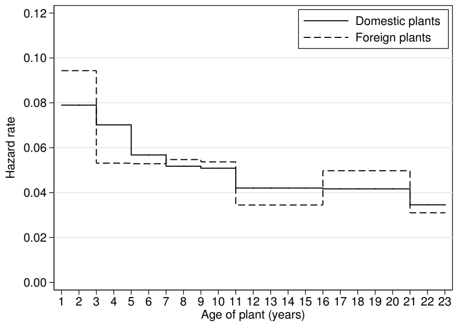
$\mathbf{x}'_{ik}$ denotes worker covariates, which can vary by elapsed duration and/or calendar time, and can also contain plant-level information.

In principle, one could also model the duration until plant closure using the worker-level data. In the context of Figure 1, we could use the eight workers to estimate the hazard to plant closure. Two workers experience plant closure, while the other six workers have durations censored either by the end of the sample period, or by separating before the plant closes. However, it does not make sense to model the duration to plant closure as a function of *worker* tenure. The appropriate measure of duration is plant age. Using plant age in a worker-level duration model, however, is just a re-weighting of the plant-level duration model.

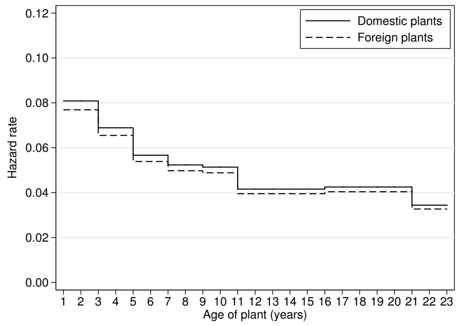
5 Results

5.1 Foreign ownership and plant closure

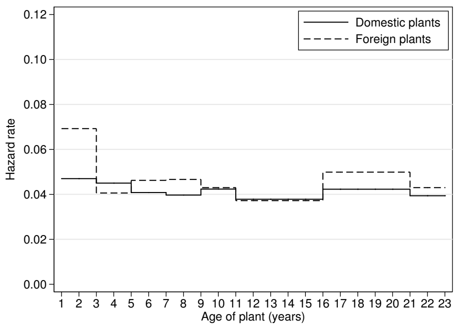
In Figure 3 we plot estimates of the hazard to plant closure. The raw data is plotted in panel (a), although we have grouped the hazard into wider age bands to ensure that a sufficient



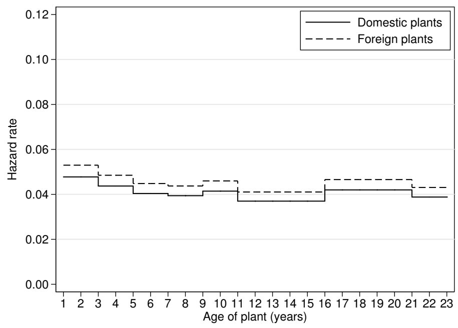
(a) Raw hazard



(b) Raw hazard, imposing PH



(c) Hazard conditional on $\mathbf{x}_{jt}$



(d) Hazard conditional on $\mathbf{x}_{jt}$, imposing PH

Figure 3: Plant closure hazards

number of plant closures occur in each band. The solid line gives the hazard rate for domestic plants.[8] As with the existing literature, surveyed in Section 2, the raw hazard to closing for a domestic plant declines with its age, falling from about 0.08 in the first two years to less than 0.04 after 20 years. (This is consistent with the raw hazard of 0.044 noted above, because the plant-age distribution is skewed in favour of older plants.) This result may occur because of selection effects, or because plants' productivities genuinely improve over time.

The dashed line in panel (a) shows the equivalent raw closure hazard for foreign-owned plants. A test of the equality of the two hazard rates cannot be rejected ($\chi^2(8) = 2.54$, p-value 0.96), so in the raw data there is no significant difference in the closure rates of the domestic and foreign plants.

[8]Throughout, all plant-level analyses are based on 31,555 plant-years and 8,947 plants. See Table 1. Similarly, we use 3,950,506 worker-years and 1,329,505 workers. See Table 2.

The estimation methods described in Section 4 rely on the proportional hazards assumption to model the effect of any covariate, such as foreign ownership. This restriction might be unwarranted if the effect of ownership on closure probability varies with plant age. For example, the initially high closure probability of domestic plants might be because they have not had time to learn about the most productive production methods. A foreign-owned plant which is part of a larger multinational might be given access to more productive techniques initially, in which case its closure probability would be lower initially. This advantage might diminish with plant age, in which case the difference in closure hazards of foreign and domestic plants would not be proportional.

However, the proportional hazards assumption imposed in panel (b) of Figure 3 cannot be rejected ($\chi^2(7) = 2.34$, p-value 0.94). The estimated "foreign ownership effect" in panel (b) is just -0.052 with a standard error of 0.097. As well as being statistically insignificant, this is a qualitatively small effect, as can be seen in panel (b).[9]

We now consider what happens to the hazard rates for foreign and domestic plants when we control for their observable characteristics. The regressors included in the vector $\mathbf{x}_{jk}$ are summarised in Table 3 and 4. Note that because we have linked employer-employee data, we are able to include the average characteristics of a plant's workforce as well as characteristics of the plant itself.

TABLES 3 AND 4 HERE

Table 3 shows that foreign-owned plants are more than twice as large, on average, as domestic plants. They are much more likely to operate a works council,[10] are more likely to export their output, more likely to be part of a larger firm and more likely to engage in sectoral and firm-level bargaining. They have higher levels of investment and they are more likely to report "very good" profits.[11] They are more likely to be located in the centre of large urban areas, and are more likely to be in the producer goods and investment goods industries.

Table 4 shows that workers in foreign-owned plants have slightly lower tenure, are more likely to be male and less likely to be an apprentice. As is well known, they are paid higher wages,[12]

[9]It should be emphasised that -0.052 is the difference in log-hazard rates between domestic and foreign plants and can be interpreted as an approximate percentage difference in the usual way.

[10]This is closely linked to their larger size: see Addison, Schnabel & Wagner (2001) for a description and analysis of German works councils.

[11]The profitability variable comes from the question "How was the profit situation in the last business year?"

[12]Andrews, Bellmann, Schank & Upward (2007) analyse the wage effects of foreign ownership using these data.

they have higher qualification levels and they are more likely to be managers, engineers but also in basic manual occupations.

So, the fact that the raw closure hazards in panels (a) and (b) are very similar could be misleading. Including a detailed set of plant- and worker-characteristics, using Equation (2), gives us the hazards plotted in panels (c) and (d). The result is to shift up the hazard of foreign-owned plants up relative to domestic plants. After imposing proportional hazards the foreign-ownership effect is estimated to be 0.106, but with a standard error of 0.106. We cannot reject the proportional hazards assumption, so panel (d) represents our preferred model. We label this as our base model; full estimates are reported in Table 5.

The other striking effect of including $\mathbf{x}_{jk}$ is that the hazard becomes much flatter: compare panel (d) with panel (b), for example. This strongly suggests that the apparent negative duration dependence observed in the raw hazard is primarily a selection effect. Suppose that each plant actually faces a constant risk of closure which does not change with a plant's age. One could think of this as a productivity shock which arrives at each plant in each period with constant mean and variance. However, some plants have a higher (fixed) productivity advantage which means that they can withstand greater negative shocks to their productivity. Plants with higher fixed productivities will therefore survive for longer, on average. The average productivity of the sample will therefore increase as elapsed age increases, leading to the apparent downward-sloping hazard shown in panel (b). If our observable characteristics are a good proxy for productivity, then their inclusion will make the hazard flatter.

TABLE 5 HERE

Consider now the estimates for all the other covariates, reported in Table 5. Recall that this is an estimate of a hazard to closing, so a positive coefficient means that that characteristic is associated with a greater risk of plant failure. Estimates on dummy variables should be interpreted as a proportional shift in the hazard (as with foreign ownership above); if logged, estimates on continuously measured covariates should be interpreted as elasticities. The hazard to a plant closing is declining in firm size and profitability, but is higher for plants with a works council.[13] The effect of plant-size is large. Small plants have double the closure rate of medium-sized plants [$100(e^{-0.344+1.103} - 1) = 113.6\%$], whereas very large plants have a closure rate 65.1% lower than medium-sized plants [$100(e^{-2.155+1.103} - 1)$].

[13] See Addison, Bellman & Kölling (2004) for evidence that works councils and plant closings are positively associated.

Worker characteristics generally seem less important, but two results stand out. First, there is a significant non-linear relationship between plant closure and average wages. The estimated turning point is within the range of the data, but below the mean. A one standard deviation in average log wages increases the closure hazard by about 0.2, so for most plants an increase in wages is associated with an increase in the closure probability. If we believe that plants with higher unobserved productivity will be less likely to close and more likely to pay higher wages for its workers then this coefficient is actually biased downwards. However, disentangling the causal effect of wages on plant closure is beyond the scope of this paper. Second, there is a significant relationship between plant closure and average tenure which is negative within the range of the data. A one standard deviation increase in the average tenure of a plant's workforce reduces its closure hazard by about 0.3.

In Table 6 we summarise the estimated coefficient on foreign ownership for a variety of specifications. As already noted, the raw effect is negative and insignificantly different from zero; this becomes positive (but still small and insignificant) when controlling for a full set of plant and worker characteristics.

TABLE 6 HERE

One issue which arises when estimating effects at the plant level is whether one should weight by plant size. This is important because the size distribution of plants is so skewed. Plants with more than 500 workers, for example, account for more than 60% of all worker-years in the data, but account for only 7% of plant-years. Therefore if the effect of foreign ownership varies with plant size, weighting could substantially alter our conclusions about the effect of foreign ownership on the labour market. In Figure 4 we plot the estimated coefficient, together with standard error bars, for four samples split by plant size. Figure 4 reveals that small foreign-owned plants are significantly more likely to close than small domestic firms, while the effect for the three other size groups is poorly determined and rather small. Weighting by plant size (reported in Table 6) increases the estimated foreign ownership effect substantially, but the effect is still insignificantly different from zero.

A key finding of Alvarez & Görg (2005) is that foreign-owned plants are only more "foot-loose" if their production is oriented towards domestic markets. In Figure 5 we plot the probability of plant closure against the export intensity of the plant. As our results in Table 5 showed, exporting plants are less likely to close, but there appears to be a higher closure rate for foreign non-exporting plants. We can test this proposition by interacting the foreign

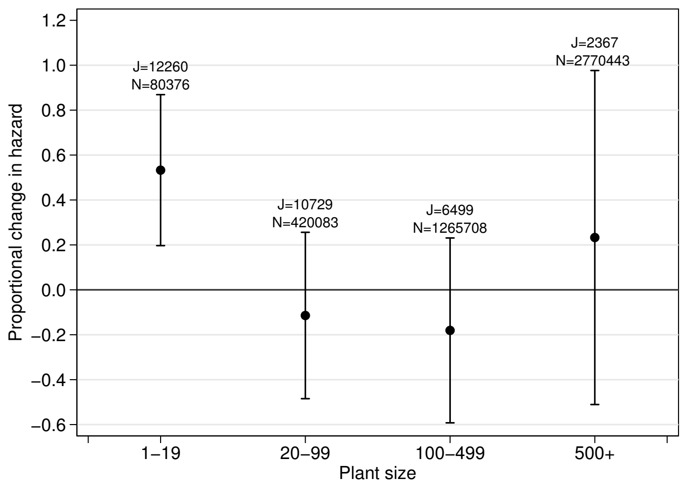


Figure 4: The effect of foreign ownership on plant closure by plant size

ownership dummy with the exporting dummy. The results, also shown in Table 6 confirm Alvarez & Görg's finding, but for Germany. Foreign-owned exporting plants have a significantly lower closure hazard than foreign-owned plants which do not export.

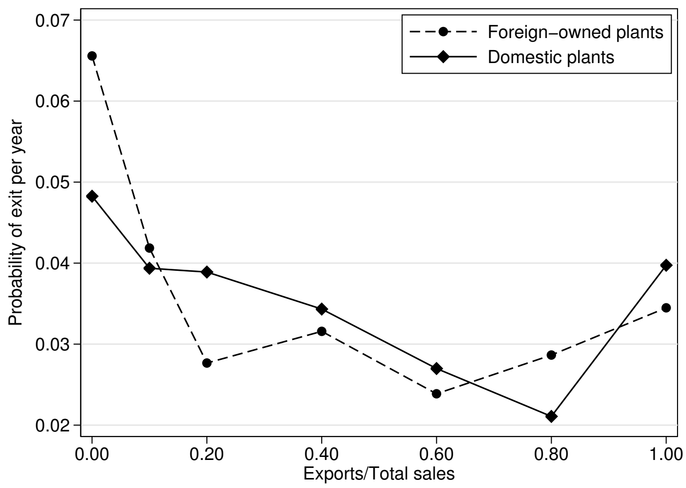


Figure 5: Plant closure and exports

5.2 Foreign ownership and worker separations

Even if foreign-owned plants do not have higher closure rates, it is still possible that they contribute to greater employment insecurity by having greater turnover of workers. In this section we therefore examine the separation rates of workers. Note that separation which occur because of plant closure are treated as censoring (see Section 4). However, with these administrative data we cannot distinguish separations which are initiated by the employer ("layoffs") from separations which are initiated by the worker ("quits").

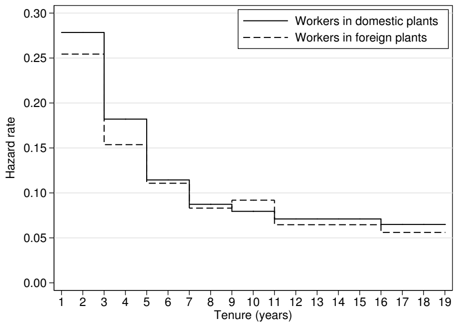
Figure 6 draws the estimated hazards for worker separations using the same four specifications used for the plant closure hazards.[14] Panel (a) plots the raw hazard of separating for workers in foreign- and domestically-owned plants. As is well known, the separation hazard exhibits negative duration dependence. A number of theories are consistent with this finding. As with the plant hazards, these theories suggest either that the downward sloping hazard is the result of selection, or the result of genuine changes in the probability of separation over elapsed time.

Matching models, for example, suggest that good matches between workers and firms are likely to endure, while bad matches are likely to end early. Thus, as tenure increases, the sample of remaining matches tends to improve, and the average separation rate of the remaining matches falls.

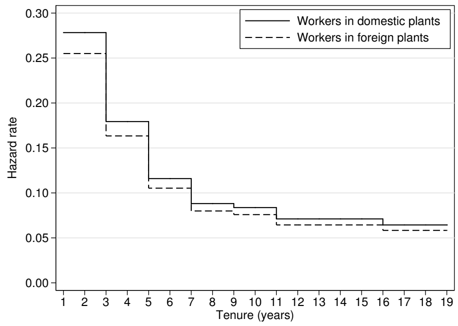
Human capital models, on the other hand, suggest that workers accumulate firm-specific human capital which increases their marginal product as tenure accumulates. If their wage increases by less than their marginal product, both the worker and the firm will have more incentive to maintain the employment relationship, and the separation rate will fall.

The hazard for workers in foreign-owned plants is below that for workers in domestically-owned plants for every tenure band except for one. Although the two hazards are not very far apart, the large sample size means we can easily reject the null that the hazard rates are equal ($\chi^2(7) = 1239.56$, p-value < 0.001). In addition, we can also reject the null that the difference between the two hazards is a constant proportion ($\chi^2(6) = 512.60$, p-value < 0.001). However, the resulting hazard after imposing the proportional hazards assumption, plotted in panel (b), look reasonably similar to those in panel (a). The average foreign ownership effect is estimated to be -0.1 with a standard error of 0.004. Although this is highly significant, it is still quite a small effect, as shown in panel (b) of Figure 6.

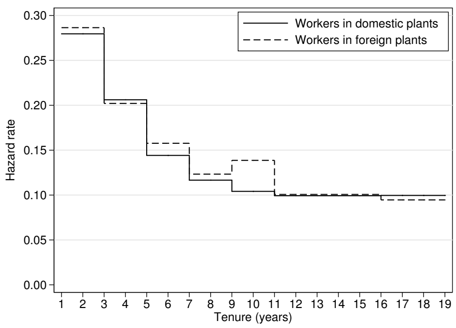
[14] In fact, the much greater number of observations at the worker level implies that we do not have to group the hazard as we did for the plant closure hazards. We have done this purely for comparability.



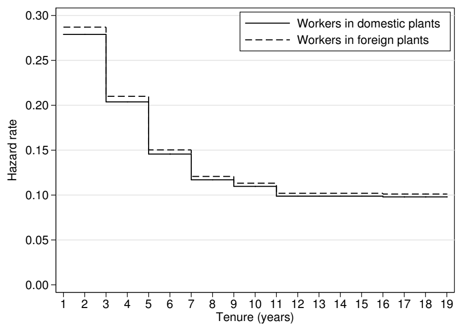
(a) Raw hazard



(b) Raw hazard, imposing PH



(c) Hazard conditional on $\mathbf{x}_{it}$



(d) Hazard conditional on $\mathbf{x}_{it}$, imposing PH

Figure 6: Worker separation hazards

In panel (c) we estimate the hazard after controlling for a full set of covariates, which are essentially the same as those used in the plant closure hazards. Interestingly, the effect of including covariates in the worker separation model is very similar to the effect of including covariates in the plant closure model. The hazard for workers in foreign-owned plants shifts up relative to the hazard for worker in domestically-owned plants. However, in contrast to the plant closure results, the introduction of covariates does not make the hazard "flatter", as would be expected if the shape of the hazard were drive entirely by selection effects.

In panel (d) we impose the proportional hazards assumption, which is again easily rejected ($\chi^2(6) = 493.78$, p-value < 0.001). The estimated foreign ownership effect jumps from -0.102 (panel b) to 0.036 (panel d) when we include covariates. In other words, once we control for the differences in observable characteristics between workers in foreign and domestic plants, workers in foreign plants actually have *higher* separation rates. However, once again the size of the effect is very small.

TABLE 7 HERE

The results of the proportional hazard base model are reported in Table 7. The separation rate is non-linear in firm size: very small and very large firms appear to have the lowest separation rates. Separations are decreasing in profitability: plants with “bad” profits have a separation hazard about 0.2 higher than plants with “very good” profits (the base group). The separation hazard is also non-linear in plant age (in contrast to the plant closure hazard). The separation hazard is lowest for young plants and plants aged over 20 years. Plants which have a works council, which export, and which are not part of a larger firm have a lower separation hazard.

A summary of the foreign ownership effect is reported in Table 8. In the raw data foreign-owned firms have lower separation rates, but once we control for covariates this becomes small and positive. Because of the strong interaction effect of exporting and foreign ownership in the plant closure hazards, we repeated this exercise for worker separations. We find a very similar result: foreign plants which do not export have significantly higher separation rates than either domestic plants or foreign plants which do export.

TABLE 8 HERE

6 Conclusions

A number of empirical studies, and some theory, has suggested that multinational firms may contribute to job insecurity, either because they are more “footloose” than domestic producers, or because they can adjust their employment more rapidly. In contrast, other theories have proposed that foreign producers may actually wish to have lower worker turnover in order to prevent the loss of skilled workers.

Using a large linked employer-employee dataset we are able to investigate these issues empirically. These data allow us to measure whether foreign ownership affects the probability of plant closure, and whether workers in foreign-owned plants have greater separation rates. In contrast to previous studies in this area, our econometric methodology accounts for a discrete measure of duration, and for the fact that we have left-censoring.

We find that foreign-owned plants actually have slightly lower hazards to closure, but the difference between foreign and domestic plants is not statistically significant. Controlling for a wide range of plant- and worker-level characteristics reverses the result, so that foreign-owned plants have slightly higher hazards. But again the effect is insignificant. We are also able to show that certain characteristics of the workforce can have a significant influence on the survival of a plant. In particular, plants with a more experienced workforce have significantly lower closure hazards.

We then examine the separation rates of workers in the same sample of plants. Our results are very similar to the plant closure results. In the raw data, foreign-owned plants have lower separation rates, but once we control for observable characteristics this conclusion is reversed.

Finally, we should note that the "foreign-ownership effect" does not appear to be constant across different types of plants. In particular, foreign-owned plants which are small and which do not export their output have significantly higher closure rates and worker separation rates. Overall, however, our results do not support the claim that the actions of foreign-owned companies are significantly changing the stability of employment.

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7 Tables

Table 1: Plant entry and exit*

Year	*No. of plants* J_t *on 30 June*	*Plants exiting* $O_{t,t+1}$	*Plants entering* $I_{t-1,t}$
2000	4,209	129	—
2001	5,627	238	1,547
2002	6,636	323	1,247
2003	7,477	351	1,164
2004	7,906	358	780
total	31,855	1,399	4,738

* The table displays the standard stock-flow identity: $J_t = J_{t-1} + I_{t-1,t} - O_{t-1,t}$. Plants who exit are genuinely those that close (there is no attrition in these data); plants who enter are left-censored.

Table 2: Worker entry and exit*

Year	*No. of workers* N_t *on 30 June*	*No. of workers exiting* $o_{t,t+1}$*, because*		*No. of workers joining* $i_{t-1,t}$*, because*	
		firm closes	*worker separates*	*firm joins sample*	*worker hired*
2000	570,949	2,987	70,845		
2001	764,126	10,678	105,131	176,474	90,535
2002	826,432	10,107	92,311	93,867	84,248
2003	853,446	8,966	92,927	49,645	79,787
2004	935,553	10,422	113,535	110,542	73,458
total	3,950,506	43,160	474,749	430,528	328,028

* The table displays the standard stock-flow identity: $N_t = N_{t-1} + i_{t-1,t} - o_{t-1,t}$.

Table 3: Means for plant-level covariates[a]

	Domestic	*Foreign*
Firm age (years)	15.896	14.854**
Firm size (number of workers)	127.481	343.656**
Plant size bigger than initial size[c]	0.776	0.673**
Works council	0.300	0.642**
Firm exports	0.292	0.617**
Plant not part of larger firm	0.752	0.339**
Sectoral bargaining agreement	0.521	0.562**
Firm-level bargaining agreement	0.050	0.079**
Investment (relative to median)	46.525	136.535**
Profits "very good"	0.054	0.089**
Profits "good"	0.276	0.287
Profits "Satisfactory"	0.335	0.321
Profits "Just sufficient"	0.200	0.166**
Profits "Bad"	0.135	0.138
Population >500,000 (central)	0.359	0.441**
Population >500,000 (outskirts)	0.064	0.056
Population 100,000-500,000 (central)	0.186	0.175
Population 100,000-500,000 (outskirts)	0.116	0.098
Population 50,000-100,000 (central)	0.018	0.015
Population 50,000-100,000 (outskirts)	0.048	0.046
Population 20,000-50,000	0.094	0.086
Population 5,000-20,000	0.078	0.068
Population 2,000-5,000	0.022	0.013
Population <2,000	0.016	0.003**
Mining, energy	0.016	0.015
Food	0.042	0.030
Consumer goods	0.058	0.053
Producer goods	0.082	0.179**
Investment goods	0.140	0.236**
Construction	0.139	0.033**
Trade	0.210	0.195
Transport & communications	0.058	0.056
Catering	0.037	0.056**
Business services	0.181	0.125**
Other services	0.038	0.024*
Local unemployment rate[d]	8.796	8.504**
Herfindahl 3-digit concentration index[e]	0.004	0.008**
Industry employment growth in previous year	−0.001	0.012**

[a] Table shows means pooled across 2000–2004. These are the means of the variable when the plant entered the sample, ie one observation per plant. There are 8,947 plants (8,192 domestic and 755 foreign-owned).

[b] Asterisks indicates the significance of a t-test of the difference in the means; * is significant at the 5% level; ** is significant at the 5% level.

[c] Initial size is based on the employment of the plant in the first year in which it appears in the Beschäftigtenstatistik

[d] Unemployment rate varies at the three-digit *Kreis* level, an intermediate level of administration between the *Länder* (German states) and the local/municipal levels (*Gemeinden*).

[e] The concentration index is based on employment.

Table 4: Means for individual-level covariates[a]

	Domestic	*Foreign*
Tenure (years)	5.971	5.789
Non-German	0.103	0.127
Female	0.279	0.244
Apprentice	0.086	0.059
Part-time worker	0.082	0.050
Home worker	0.001	0.000
Daily wage €	80.854	93.376
Age	36.050	36.352
Without apprenticeship or Abitur	0.215	0.201
Apprenticeship, no Abitur	0.541	0.522
No apprenticeship, with Abitur	0.027	0.031
With apprenticeship and Abitur	0.037	0.046
Technical college degree	0.049	0.065
University education	0.047	0.078
Education unknown	0.084	0.057
Basic manual occupation	0.272	0.341
Qualified manual occupation	0.197	0.152
Engineers and technicians	0.128	0.152
Basic service occupation	0.122	0.074
Qualified service occupation	0.009	0.003
Semi-professional	0.004	0.003
Professional	0.008	0.007
Basic business occupation	0.062	0.060
Qualified business occupation	0.176	0.174
Manager	0.021	0.034

[a] Table shows means pooled across 2000–2004. These are the means of the variable when the worker entered the sample, ie one observation per worker. There are 1,329,505 workers (1,067,691 in domestic plants and 261,814 in foreign-owned plants).

[b] All means are significantly different at the 5% level.

Table 5: Plant-level closure hazard, base model

	Coeff.	*Std. Err.*
Plant is foreign-owned	0.106	(0.106)
Works council	0.354	(0.086)**
Plant exports	−0.063	(0.073)
Plant is not part of a larger firm	0.012	(0.073)
Sectoral bargaining agreement	−0.041	(0.067)
Firm-level bargaining agreement	−0.395	(0.156)**
Investment (relative to median)[b]	0.637	(1.677)
5–9 workers	−0.344	(0.096)**
10–19 workers	−0.478	(0.100)**
20–49 workers	−0.551	(0.101)**
50–99 workers	−0.742	(0.126)**
100–199 workers	−1.103	(0.151)**
200–499 workers	−1.259	(0.165)**
500–999 workers	−1.539	(0.249)**
$\geq$ 1000 workers	−2.155	(0.367)**
Plant size bigger than initial size	−0.347	(0.065)**
Profits “good”	−0.083	(0.147)
Profits “satisfactory”	0.207	(0.144)
Profits “just sufficient”	0.544	(0.146)**
Profits “bad”	1.039	(0.148)**
Local unemployment rate	0.028	(0.010)**
Industry concentration index	−1.557	(3.147)
Industry employment growth rate	1.445	(0.615)**
Proportion of non-German workers	0.036	(0.172)
Proportion of females	−0.151	(0.113)
log average wage	−2.843	(0.911)**
(log average wage)^2	0.394	(0.115)**
Average age	0.026	(0.167)
Average age^2/100	−0.096	(0.407)
Average age^3/10000	0.146	(0.323)
Proportion with apprenticeship, no Abitur	−0.185	(0.155)
Proportion with no apprenticeship, with Abitur	−0.816	(0.825)
Proportion with apprenticeship and Abitur	−0.803	(0.335)**
Proportion with technical college degree	−0.498	(0.422)
Proportion with university education	0.037	(0.292)
Proportion with education unknown	−0.145	(0.158)
Average tenure	−0.125	(0.023)**
Average tenure^2/100	0.415	(0.101)**

[a] Regressions also include dummies for location, industry, year and occupation.

[b] Coefficients are $\times 10^{-4}$.

Table 6: Summary of plant-level hazard models

	Coeff.	*Std. Err.*
Raw effect	-0.052	(0.097)
Including $\mathbf{x}_{jk}$ (Table 5)	0.106	(0.106)
Weighted by plant size	0.367	(0.257)
Interaction model:		
Foreign ownership	0.344	(0.138)
Exporting	-0.010	(0.076)
Foreign owned and exporting	-0.485	(0.201)

Table 7: Worker-level separation hazard, base model

	Coeff.	*Std. Err.*
Plant is foreign owned	0.034	(0.004)**
Works council	−0.100	(0.005)**
Plant exports	−0.142	(0.004)**
Plant is not part of a larger firm	−0.045	(0.003)**
Sectoral bargaining agreement	0.011	(0.005)**
Firm-level bargaining agreement	0.006	(0.007)
Investment (relative to median)[b]	−0.246	(0.008)**
5–9 workers	0.056	(0.034)*
10–19 workers	0.111	(0.031)**
20–49 workers	0.078	(0.030)**
50–99 workers	0.154	(0.030)**
100–199 workers	0.169	(0.030)**
200–499 workers	0.163	(0.030)**
500–999 workers	0.088	(0.030)**
$\geq$ 1000 workers	0.046	(0.030)
Profits "good"	−0.010	(0.006)*
Profits "satisfactory"	0.069	(0.006)**
Profits "just sufficient"	0.178	(0.006)**
Profits "bad"	0.193	(0.007)**
Firm age (years)	0.037	(0.001)**
Firm age^2	−0.117	(0.004)**
Non-German worker	0.066	(0.005)**
Female	−0.002	(0.004)
Log wage	−0.151	(0.013)**
Log wage^2	−0.048	(0.002)**
Age	0.228	(0.004)**
Age^2/100	−0.826	(0.011)**
Age^3/10000	0.871	(0.009)**
Apprenticeship, no Abitur	−0.008	(0.006)
No apprenticeship, with Abitur	−0.113	(0.006)**
Apprenticeship, with Abitur	0.220	(0.009)**
Technical college degree	0.017	(0.009)*
University education	0.059	(0.009)**
Education unknown	0.154	(0.009)**

[a] Regressions also include dummies for location, industry, year and occupation.
[b] Coefficients are $\times 10^{-4}$.

Table 8: Summary of worker-level hazard models

	Coeff.	*Std. Err.*
Raw effect (Table 7)	-0.102	(0.004)
Including $\mathbf{x}_{ik}$	0.036	(0.004)
Interaction model:		
Foreign ownership	0.110	(0.007)
Exporting	-0.117	(0.004)
Foreign owned and exporting	-0.117	(0.008)