Foreign-owned plants and job security*

Martyn Andrews
University of Manchester

Lutz Bellmann
Institut für Arbeitsmarkt- und Berufsforschung

Thorsten Schank
Friedrich-Alexander Universität Erlangen-Nürnberg

Richard Upward
University of Nottingham

July 20, 2009

Abstract

We investigate the hypothesis that workers in foreign-owned plants face greater job insecurity than those in domestic-owned plants. Using linked employer-employee data from Germany, we examine whether foreign-owned plants are more likely to close down, and whether workers in foreign-owned plants face higher separation rates. We find that, after controlling for observable and unobservable characteristics of foreign-owned and domestic-owned plants, foreign-owned plants have higher closure rates and their workers have higher separation rates, but the effects are quantitatively small and insignificant. In contrast, foreign-owned plants which do not export have higher closure rates, and foreign-owned plants which are not contracting have lower separation rates. [100 words]

Keywords: Foreign-owned plants, plant closure, worker separations, duration modelling.

JEL codes: F23, J63, C41.

*The authors thank the Economic and Social Research Council (under grant number RES-000-22-1034) and the Leverhulme Centre for Research on Globalisation and Economic Policy (under grant number F/00 114/AM) for financial assistance. We also thank the participants of the 2008 Comparative Analysis of Enterprise Micro Data (CAED) conference in Budapest, the participants of a workshop on Foreign Trade and Globalisation sponsored by the German Council for Social and Economic Data (RatSWD), and the participants of the 2008 Work and Pensions conference in Sheffield for their comments. Stephen Jenkins also helped with clarifying conversations concerning unobserved heterogeneity and left-truncated discrete duration models. The data were kindly supplied by the Institut für Arbeitsmarkt- und Berufsforschung, Nürnberg.

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 suggest that foreign-owned plants are more "footloose" in the sense that they are more likely to exit the market (close down) than similar domestic-owned plants.

There is a small empirical literature (described in Section 2 below) which tests this hypothesis by estimating the probability of plant closure as a function of foreign ownership. This literature is limited to three papers from Ireland, Indonesia and Chile, all of which consider only the manufacturing sector. We use a large, representative sample of plants from all sectors of the economy to test the hypothesis for the German economy. We improve on the existing methodology by estimating the hazard to plant closure using discrete-time duration models which account for unobserved heterogeneity and delayed entry. In contrast to previous studies, our methods allow us to estimate the baseline hazard to plant exit as well as the impact of foreign ownership. The shape of the baseline hazard yields additional insights into the processes which cause plant survival or failure. Our sample of plants can be linked to information on individual workers, and so we are also able to examine the role of workers' characteristics in determining plant survival.

It has also been argued that foreign-owned plants have higher worker turnover rates even if they do not actually shut down. For example, Fabbri, Haskel & Slaughter (2003) argue that multinational firms have more elastic labour demands than domestic-owned firms, which would be consistent with higher worker turnover rates. However, a separate literature has suggested that foreign-owned plants will pay higher wages in order to prevent worker separations, and therefore turnover will actually be lower in foreign-owned plants (Glass & Saggi 2002). The empirical evidence on foreign ownership and worker separation rates is limited to only two papers, only one of which uses micro-level data on workers. We use the same econometric methods to estimate the hazard to worker separation as a function of foreign ownership, conditional on plant survival.

Our results cast doubt on the hypothesis that foreign-owned plants are in general more "footloose", or that jobs in foreign-owned plants are less secure. In the raw data, foreign-owned plants have lower closure rates and lower worker separation rates, but these differences are insignificant. After controlling for different observable and unobservable characteristics, foreign-owned plants do not have significantly higher closure rates and their workers do not have significantly higher separation rates. Our estimates are also quantitatively small.

In Section 2 we discuss the related literature in more detail. In Section 3 we describe the data and how we construct our measures of plant closure 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

Foreign ownership and plant closure

There is a substantial literature on the determinants of firm (or plant) success and failure, where failure is defined in terms of exit from the market (the plant closes). Various theoretical models suggest that larger 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 close down over a five year period as a function of their industry and size. They find that closure 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 or firm closure per period conditional on survival up to that period. This hazard rate is found to be declining in duration (new firms or plants are most likely to fail). Important explanatory variables for closure 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 closure. Görg & Strobl (2003) use Cox Proportional Hazards models, and find that manufacturing plants in Ireland which are owned by foreign multinationals actually have *lower* closure rates. However, foreign-owned plants have characteristics typically associated with lower closure rates. For example, foreign-owned plants tend to be larger. Once these factors are accounted for, Görg & Strobl find that foreign-owned plants have significantly higher hazard rates (Table 2).[1]

Bernard & Sjöholm (2003) find similar results for the census of manufacturing plants in Indonesia. In the raw data, foreign-owned plants are far less likely to shut down than domestic-owned 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-owned plants. The hazard ratio is also large, suggesting that foreign-owned plants are more than 20% more likely to close.

Alvarez & Görg (2009) use data on Chilean manufacturing plants. They estimate Probit models of the probability of exit over the period 1990–2000. They find no significant difference in the probability of exit for foreign-owned plants in the sample as a whole. But they do find a higher exit probability in the second half of the

[1]The estimated coefficient is 0.231 with a standard error of 0.059. However, the ownership variable is also interacted with a number of other characteristics, and some of these interactions are highly significant. The average effect of foreign ownership on the hazard rate may therefore be much smaller than 0.231.

sample period, and for foreign-owned plants which do not export. Foreign-owned plants which export do not have higher closure rates.

One other closely related paper examines the differences in exit rates between plants which belong to multinational firms, as opposed to foreign-owned firms. Bernard & Jensen (2007) use a census of US manufacturing plants and estimate a Probit of plant exit. Plants which are part of multinational enterprises have lower closure rates, but tend to be larger, older and more productive. Once these characteristics are taken into account, plants belonging to multinationals have higher closure rates, with a marginal effect of 0.045 (standard error 0.005).

All of the above papers identify the effect of foreign ownership or multinational status by utilising the cross-section variation in that characteristic, and this is the approach we use in this paper. An alternative approach is to consider the effect of within-plant changes in nationality i.e. foreign acquisition or foreign divestment. Girma & Görg (2004) use plant level data from the UK electronics and food industries. They compare plants which have been taken over with similar plants which have not, and find that foreign takeover increases the hazard rate of exit dramatically.[2]

Foreign ownership and worker turnover

There is also 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).

As noted, it has been suggested that the actions of multinational firms may be

[2]The authors compute hazard ratios of 2.56 for the electronics sector and 10 for the food sector. These estimates are an order of magnitude greater than estimates from cross-sectional comparisons.

associated with greater job turnover, and therefore greater worker turnover. 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 *et al.* (2003) argue that multinational firms have more elastic labour demands than domestic-owned 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 labour 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-owned 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.

Görg & Strobl (2003) also consider job turnover in foreign and domestic plants by examining the persistence of plant-level employment changes. They find that jobs created in foreign-owned firms are actually more persistent than those created in domestic firms.

The only other paper to use linked worker-firm data to investigate this issue is Pesola (2008). She shows that the job separation rate for workers increases after foreign takeover, but that the effects fade after one year, implying a process of restructuring rather than a permanent increase in job insecurity. She also shows that employees in firms that are about to become foreign-owned have higher separation rates before takeover.

In contrast, there is also a recent theoretical literature on human capital spillovers between foreign-owned and domestic-owned 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. 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) will be lower in foreign-owned plants.

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.[3]

Ownership is defined as either Western German, Eastern German, foreign, or public.[4] 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 in the year they enter. In principle, we could compare worker separation rates before and after foreign acquisition, as Pesola (2008) does. However, there are two problems. First, there is a very small number of plants who switch.[5] Second, because we do not observe ownership in 1999 or 2005, we would have to assume it is constant before and after takeover. Instead, therefore, we treat

[3] A detailed description of the IAB Establishment Panel can be found in Fischer, Janik, Müller & Schmucker (2009).

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

[5] For example, there are only 36 plants who were domestic-owned in 2000 and foreign-owned in 2004.

foreign-ownership as a time-invariant characteristic.

The second source of data is the employment statistics register of the German Federal Office of Labour (*Beschäftigtenstatistik*), which covers all workers or trainees registered by the social insurance system. The register covers about 80% of workers in Western Germany and about 85% in Eastern 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).

As noted, we restrict the analysis to the private sector. As almost all workers in the private sector are covered by the social insurance system, the data covers nearly 100% of workers. Furthermore, we only analyse Western German plants. This is because information on the age of plants in the Eastern German sample is more limited, and the process of exit for Eastern German plants seems likely to be completely different. There are almost no Eastern German-owned plants in Western Germany. This leaves us with just two ownership categories: Western German (i.e. domestic) and Foreign.

By using the plant identification number, we link each worker to a plant in the panel. This yields an unbalanced annual panel of workers together with detailed information on the plants in which they work. Our regression sample follows plants and workers over the period 2000–2005. However, the worker-level data (with associated plant identifiers) goes back to 1977. This means that we can accurately date the birth of plants before 2000, and we can accurately date when workers started working in a particular plant. We select all workers in the employment register who were employed by the surveyed plants on June 30th each year. Thus our data comprise an unbalanced panel of plants $j = 1, \ldots, J_t$ observed annually on 30 June for $t = 2000 \ldots 2005$, and a corresponding unbalanced panel of workers $i = 1, \ldots, N_t$.

3.1 Plants

The plant-level data has several important features. First, we are able to identify closure more accurately than is usual in administrative databases. 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. We use the interview question in the IAB panel to confirm whether a plant has genuinely closed.[6]

Second, because the sample period ends in 2005 we have the usual problem of right-censoring. That is, we do not know what happens to any plant which is still observed on 30th June 2005.

Third, the IAB panel suffers from attrition. However, we avoid this problem by making use of the worker-level data. If a plant attrits from the IAB panel, we search the Beschäftigtenstatistik to find out if it is still employing any workers. If it is, the plant is followed until it genuinely exits, or until the end of the sample period in 2005. This is potentially important, because it is likely that attrition from the IAB panel and the outcomes of interest (plant closure and worker separations) are not statistically independent.

Fourth, most plants are born some time before the start of the sample period, and for these plants their age is positive when they are first observed.[7] Thus we have left truncation or delayed entry.[8]

Finally, not all plants are observed at the beginning of the sample period (30th June

[6]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”.

[7]The age of a plant is calculated by finding the year in which the plant first appears in the Beschäftigtenstatistik. However, the earliest year in the Beschäftigtenstatistik is 1977, and so the age of some plants is right-censored.

[8]It is left truncation rather than left censoring because we observe the age of plants as they enter the sample (Wooldridge 2002, Section 20.3.3).

2000). This happens mainly because the IAB expanded the IAB panel considerably over time, but also occurs for plants 'born' after 30th June 2000; for these latter plants, there is no left-truncation.

Table 1 shows the basic movement of plants in and out of the regression 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 4,766 whereas the number of closures is 1,399, and therefore, the sample grows over time. There are 4,209 plants sampled on 30/6/2000; when added to the 4,766 plants who enter, this gives a sample size of 8,975 plants.

Table 1: Plant entry and exit*

Year	*No. of plants* J_t *on 30 June*	*Plants exiting* $O[t, t+1)$	*Plants entering* $I[t, t+1)$
2000	4,209	129	1,547
2001	5,627	238	1,247
2002	6,636	323	1,190
2003	7,503	351	782
2004	7,934	358	—
total	31,909	1,399	4,766

* 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 truncated.

3.2 Workers

For each plant j we observe all workers who are in the Beschäftigtenstatistik. Workers can either exit their plant 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 his plant, it is because his tenure with his employing plant has come to an end. This can happen for one of two reasons. Either the plant closes or the worker separates from the plant. However, with these administrative data we cannot distinguish between separations initiated by the employer (layoffs) or the employee (quits).

Table 2 shows that 516,911 workers exited their plant. Of these, 473,835 workers

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 $i[t, t+1)$, because*	
		plant closes	*worker separates*	*plant joins sample*	*worker hired*
2000	570,413	2,981	70,723	176,273	90,394
2001	763,376	10,660	104,923	93,693	84,096
2002	825,582	10,092	92,032	50,111	79,661
2003	853,230	8,956	92,782	110,412	73,426
2004	935,330	10,387	113,575	—	—
total	3,954,931	43,076	473,835	430,489	327,577

* The table displays the standard stock-flow identity: $N_t = N_{t-1} + i[t-1, t) - o[t-1, t)$.

separated from their plants and 43,076 workers exited because their plant closed. In other words, 8.3% of exits are because of plant closure. 758,066 workers entered the panel. 327,577 of them were hired by plants which already existed in the data, while 430,489 were employed by plants entering the data. When 758,066 workers are added to the 570,413 workers who are observed on 30 June 2000, the sample size is 1,328,479 workers.

4 Methods

We wish to model the probability that a plant closes, and the probability that a worker separates from a plant, as a function of the plant's ownership status. The appropriate econometric framework to use is that of discrete-time duration models, where 'duration' for a plant refers to its age, and 'duration' for a worker refers to his tenure in the plant. The framework is in discrete time because events may occur at any point between 30th June in year t and 29th June in year $t+1$, but we do not observe the precise date on which this happens.

4.1 Plant hazards

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, $h_a(\mathbf{x}_j, u_j)$, 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 $a-1$:

$$h_a(\mathbf{x}_j, u_j) = \Pr(A_j = a | A_j \geq a) = f_a(\mathbf{x}_j, u_j)/S_{a-1}(\mathbf{x}_j, u_j) \quad a = 1, 2, \ldots, a_j,$$

where A_j is the latent age of the plant j, a_j is the completed duration for plant j, $\mathbf{x}_j$ is a vector of observed covariates, u_j is a term capturing all unobserved heterogeneity, $f_a(\mathbf{x}_j, u_j)$ is the probability of observing duration a, and $S_{a-1}(\mathbf{x}_j, u_j)$ is the probability of surviving to duration $a-1$. $\mathbf{x}_j$ includes the foreign ownership dummy. It also includes some worker-level covariates that have been averaged to the plant-level.

For most of the plants in the sample we have delayed entry. Denote the age at which a plant enters the sample as $\underline{a}_j$. Because $\underline{a}_j$ is an integer, and because we round up duration, $\underline{a}_j = 1$ for new entrants and $\underline{a}_j \geq 2$ for late entrants. 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). Finally, we also need to control for the unobserved heterogeneity u_j. Standard references are Wooldridge (2002), Cameron & Trivedi (2005) and Jenkins (2005). It is Jenkins that we use here.

To start, consider the standard case with no delayed entry, but for some of these plants there is right censoring. The log-likelihood function for this sub-sample is given by (Jenkins 2005, Eqn (6.9)):

$$\log L = \sum_j \log \left[\left(\frac{h_{a_j}(\mathbf{x}_j, u_j)}{1 - h_{a_j}(\mathbf{x}_j, u_j)} \right)^{c_j} \prod_{a=1}^{a_j} \log[1 - h_a(\mathbf{x}_j, u_j)] \right]. \tag{1}$$

Here the dummy variable $c_j = 1$ if a plant closes and is 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 $(1-h_{j1})(1-h_{j2})\dots(1-h_{ja_j})$, where h_{ja} is short-hand for $h_a(\mathbf{x}_j, u_j)$.

A standard approach for estimating this model is to expand the data so that each plant contributes a_j rows. Define a dummy variable y_{ja} which takes the value zero unless it is the last year plant j is observed ($a = a_j$) *and* the spell is completed ($c_j = 1$); in this case, $y_{ja} = 1$. We can then write the log-likelihood for this sub-sample as

$$\log L = \sum_j \sum_{a=1}^{a_j} \{y_{ja} \log h_a(\mathbf{x}_j, u_j) + (1 - y_{ja}) \log[1 - h_a(\mathbf{x}_j, u_j)]\}. \tag{2}$$

This is the likelihood for any 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. Then the precise form of the discrete hazard for plant j is given by the complementary log-log link function:

$$h_a(\mathbf{x}_j, u_j) = 1 - \exp(-\exp(\mathbf{x}_j\boldsymbol{\beta} + \gamma_a + u_j)) \quad a = 1, \dots, a_j. \tag{3}$$

The γ_a terms are interpreted as the log of a non-parametric piecewise-linear baseline hazard.

We now deal with the problem of delayed entry or left-truncation. Most of the plants in our sample have $\underline{a}_j > 1$ and so have been at risk of closing for some time. This is a sample selection problem: one is more likely to observe long rather than short durations. For a plant with a left-truncated spell, its contribution to the likelihood is divided by the probability of surviving to the first period of the sample:

$$S_{\underline{a}_j-1}(\mathbf{x}_j, u_j) = \prod_{a=1}^{\underline{a}_j-1} [1 - h_a(\mathbf{x}_j, u_j)].$$

Because the denominator divides into the numerator very neatly, the log-likelihood becomes

$$\log L = \sum_j \log \left[\left(\frac{h_{a_j}(\mathbf{x}_j, u_j)}{1 - h_{a_j}(\mathbf{x}_j, u_j)} \right)^{c_j} \prod_{a=\underline{a}_j}^{a_j} \log[1 - h_a(\mathbf{x}_j, u_j)] \right], \tag{4}$$

and, amending Equation (2), the log-likelihood is also written

$$\log L = \sum_j \sum_{a=\underline{a}_j}^{a_j} \{ y_{ja} \log h_a(\mathbf{x}_j, u_j) + (1 - y_{ja}) \log[1 - h_a(\mathbf{x}_j, u_j)] \}. \tag{5}$$

This convenient cancelling result (Guo 1993, Jenkins 2005) means that Equation (5) is very similar to the standard expression, except that the summation runs from the duration of the plant when it enters the data. As Equations (1) and (2) are special cases of Equations (4) and (5), one can pool the sub-samples with and without late entry.

It is well-known that estimating a model with covariates, but ignoring the unobservable, will bias the estimates of the baseline hazard, even though we assume that u_j and $\mathbf{x}_j$ are (statistically) independent. This means that the heterogeneity needs integrating out:

$$\log L = \sum_j \log \left\{ \int_{-\infty}^{\infty} \left[\prod_{a=\underline{a}_j}^{a_j} h_a(\mathbf{x}_j, u_j)^{y_{ja}} [1 - h_a(\mathbf{x}_j, u_j)]^{1-y_{ja}} \right] f_u(u_j) du_j \right\}, \tag{6}$$

where $f_u(u_j)$ is the density of u_j. We assume that u is Normally distributed; Gaussian quadrature is used to approximate the Normal distribution, and so the unobservable is integrated out numerically. Notice that the left truncation does not cause any further complications (Wooldridge 2002, p.704). However, inference on the variance of the heterogeneity term has to be conditional on the sample drawn, as low draws of u_i are less likely to be observed.

4.2 Worker hazards

The great advantage of linked employer-employee data is that we observe employee separations, 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 a worker separating from plant j to his elapsed time in that plant, namely his tenure. Thus, the econometric model is almost the same as that used for plant closure, except that durations relate to tenure rather than plant age. 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.

A worker's tenure is not zero when a plant enters the data for most workers, because he joined his plant earlier. Thus, delayed entry also occurs in the worker data. In the standard case, the worker is hired by the plant after the plant enters the sample. The likelihood developed earlier for plants applies here. To estimate the hazard to worker separation, the log-likelihood is therefore

$$\log L = \sum_i \log \left\{ \int_{-\infty}^{\infty} \left[\prod_{a=\underline{a}_i}^{a_i} h_a(\mathbf{x}_i, u_i)^{y_{ia}} [1 - h_a(\mathbf{x}_i, u_i)]^{1-y_{ia}} \right] f_u(u_i) du_i \right\}, \qquad (7)$$

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_{ia} is the dummy variable indicating whether the worker separates at elapsed duration a, and

$h_a(\mathbf{x}_i, u_i)$ is the corresponding hazard function:

$$h_a(\mathbf{x}_i, u_i) = 1 - \exp[-\exp(\mathbf{x}_i\boldsymbol{\beta} + \gamma_a + u_i)] \quad a = 1, \ldots, a_i. \tag{8}$$

$\mathbf{x}_i$ denotes worker covariates. This also contains plant-level information, including the foreign ownership dummy. u_i is worker-level heterogeneity, with density $f_u(u_i)$, and is integrated out using Gaussian quadrature.

Note that the unobserved heterogeneity term u_i will be heteroskedastic because it comprises a common component for all workers who are employed by the same plant. We therefore compute a plant-level cluster-robust covariance matrix.

In principle, one could also model the duration until plant closure using the worker-level data. 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 a re-weighting of the plant-level duration model.

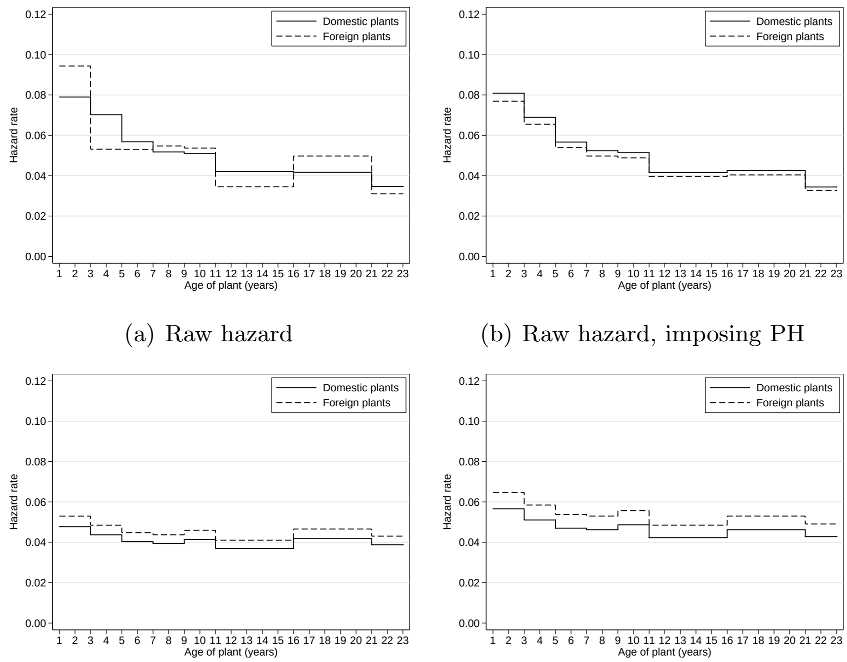
5 Results

5.1 Foreign ownership and plant closure

The raw hazard to plant closure is 0.0438, which means, on average, 4.38% of plants close each year.[9] When this is split between the 756 foreign-owned plants and 8,219 domestic-owned plants, the raw hazards are 0.0429 and 0.0439 respectively, and so the raw difference is -0.001 percentage points in favour of foreign-owned plants; taking the difference in the log of the raw hazards, this raw differential is -0.023 log-points.[10]

[9] This is 1,399 closures divided by 31,909 plant-years at risk. See Table 1. Throughout, all plant-level analyses are based on 31,909 plant-years and 8,975 plants.

[10] It should be emphasised that this difference in log-hazard rates between domestic-owned and foreign-owned plants can be interpreted as an approximate percentage difference in the usual way, measured in log-points.



(a) Raw hazard

(b) Raw hazard, imposing PH

(c) Hazard conditional on $\mathbf{x}_j$, imposing proportional hazards, ...

(d) ...and controlling for unobserved heterogeneity

Figure 1: Plant closure hazards

In Figure 1 we plot estimates of the hazard to plant closure. The raw data is plotted in panel (a), where we have grouped the hazard into age bands of varying widths. The solid line gives the hazard rate for domestic-owned plants. As with the existing literature, surveyed in Section 2, the raw hazard to closing for a domestic-owned plant declines with its age, falling from about 0.08 per year in the first two years to less than 0.04 after 20 years. This is consistent with the raw hazard of 0.0439, because the plant-age distribution is skewed in favour of older plants. The negative duration dependence in the raw hazard either occurs because of selection effects, or because plants' productivities genuinely improve over time. This is explored more fully below. 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-owned and foreign-owned plants.

The estimation methods described in Section 4 rely on the proportional hazards assumption to model the effect of any covariate, such as foreign ownership. As Figure 1 shows, this restriction might be unwarranted between duration groups if the effect of ownership on closure probability varies with plant age. For example, the initially high closure probability of foreign-owned plants might be because they face greater uncertainty about demand conditions compared with a domestic-owned plant. This uncertainty might diminish with plant age, in which case the difference in closure hazards of foreign-owned and domestic-owned plants would not be proportional.

In panel (b) we impose the proportional hazards assumption. This cannot be rejected ($\chi^2(7) = 2.35$, p-value 0.94). The estimated 'foreign ownership effect' in panel (b) is a differential of -0.051 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).

We now consider what happens to the hazard rates for foreign-owned and domestic-owned plants when we control for their observable and unobservable characteristics. The plant-level regressors included in the vector $\mathbf{x}_j$ are summarised in Table 3.

Table 3 shows that foreign-owned plants are more than twice as large, on average, as domestic-owned plants. They are much more likely to operate a works council,[11] are more likely to export their output, are more likely to be part of a larger firm and are 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.[12] They are more likely to be located in the centre of large urban areas, are more likely to be in the producer goods and investment goods industries and are less likely to be in construction and business service industries.

Table 4 compares worker covariates between foreign-owned and domestic-owned plants. It shows that workers in foreign-owned plants are more likely to be male

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

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

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

	Domestic	*Foreign*	*p-value*[a]
Plant age (years)	15.901	14.866	[0.002]
Works council	0.300	0.642	[0.000]
Plant exports	0.292	0.616	[0.000]
Plant not part of larger firm	0.753	0.339	[0.000]
Sectoral bargaining agreement	0.520	0.561	[0.030]
Firm-level bargaining agreement	0.050	0.079	[0.001]
Investment (relative to median)	46.379	136.354	[0.000]
Firm size (number of workers)	127.138	343.294	[0.000]
Profits "very good"	0.054	0.089	[0.000]
Profits "good"	0.276	0.287	[0.524]
Profits "Satisfactory"	0.335	0.320	[0.411]
Profits "Just sufficient"	0.200	0.165	[0.021]
Profits "Bad"	0.134	0.139	[0.732]
Population >500,000 (central)	0.360	0.442	[0.000]
Population >500,000 (outskirts)	0.064	0.056	[0.342]
Population 100,000-500,000 (central)	0.185	0.175	[0.468]
Population 100,000-500,000 (outskirts)	0.116	0.098	[0.131]
Population 50,000-100,000 (central)	0.018	0.015	[0.477]
Population 50,000-100,000 (outskirts)	0.048	0.046	[0.840]
Population 20,000-50,000	0.093	0.086	[0.505]
Population 5,000-20,000	0.078	0.067	[0.304]
Population 2,000-5,000	0.022	0.013	[0.126]
Population <2,000	0.016	0.003	[0.004]
Mining, energy	0.016	0.015	[0.827]
Food	0.042	0.030	[0.125]
Consumer goods	0.058	0.053	[0.554]
Producer goods	0.082	0.179	[0.000]
Investment goods	0.140	0.235	[0.000]
Construction	0.138	0.033	[0.000]
Trade	0.212	0.196	[0.307]
Transport & communications	0.058	0.056	[0.801]
Catering	0.037	0.056	[0.010]
Business services	0.181	0.124	[0.000]
Other services	0.037	0.024	[0.055]
Proportion of sample	0.916	0.084	

[*] 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 8,975 plants (8,219 domestic-owned and 756 foreign-owned).
[a] *P*-value for the *t*-statistic testing the difference in the means.

Table 4: Means for worker-level covariates*

	Domestic	*Foreign*	*p-value*[a]
Tenure (years)	6.921	6.742	[0.527]
Non-German	0.103	0.127	[0.015]
Female	0.279	0.244	[0.020]
Apprentice	0.086	0.059	[0.000]
Part-time worker	0.082	0.049	[0.000]
Home worker	0.001	0.000	[0.033]
Daily wage €	80.911	93.411	[0.000]
Age	36.005	36.336	[0.291]
Without apprenticeship or Abitur	0.215	0.201	[0.271]
Apprenticeship, no Abitur	0.542	0.522	[0.036]
No apprenticeship, with Abitur	0.028	0.031	[0.503]
With apprenticeship and Abitur	0.037	0.046	[0.031]
Technical college degree	0.049	0.065	[0.005]
University education	0.047	0.078	[0.000]
Education unknown	0.084	0.057	[0.001]
Basic manual occupation	0.272	0.341	[0.000]
Qualified manual occupation	0.197	0.152	[0.001]
Engineers and technicians	0.128	0.152	[0.031]
Basic service occupation	0.122	0.074	[0.000]
Qualified service occupation	0.009	0.003	[0.000]
Semi-professional	0.004	0.003	[0.019]
Professional	0.008	0.007	[0.463]
Basic business occupation	0.062	0.060	[0.812]
Qualified business occupation	0.177	0.174	[0.878]
Manager	0.021	0.034	[0.000]
Proportion of sample	0.803	0.197	

* 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 1,328,479 workers (1,066,742 in domestic-owned plants and 261,737 in foreign-owned plants).

[a] P-value for the t-statistic testing the difference in the means, clustered at plant-level.

and non-German, and less likely to be an apprentice. As is well known, they are paid higher wages,[13] they have higher qualification levels and they are more likely to be managers, but also in basic manual occupations. These are averaged to the plant-level and included in $\mathbf{x}_j$.

Given there are observable differences between foreign-owned and domestic-owned plants and their employees, the fact that the raw closure hazards in panels (a) and (b) are very similar could be misleading. Re-estimating the model, now including a detailed set of plant- and worker-characteristics $\mathbf{x}_j$, but not controlling for unobserved heterogeneity, gives us the baseline hazard plotted in panel (c). See Equations (3) and (5). The result is to shift up the hazard of foreign-owned plants relative to domestic-owned plants: the foreign-ownership effect is estimated to be a differential of 0.140 with a standard error of 0.105. When we additionally control for unobserved heterogeneity using Gaussian mixing, see Equations (3), (6) and panel (d), the foreign ownership effect just noted is unaffected, still being 0.140 with a standard error of 0.105. This is because the estimated standard deviation of u_j is very small and insignificant. In both cases, we cannot reject the proportional hazards assumption. Full estimates of panel (d) are reported in Table 5.

The other effect of including $\mathbf{x}_j$ is that the hazard becomes much flatter: compare panel (c) with panel (b). 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

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

hazard shown in panel (b). If our observable characteristics are a good proxy for productivity, then their inclusion will make the hazard flatter. Further controlling for unobserved heterogeneity has no effect on the hazard: compare panels (c) and (d). Indeed, in the preferred 'base' model, one can impose the 7 restrictions that make the two hazards completely flat very easily (p-value 0.57); in other words, plant age is Exponentially distributed.

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 a characteristic is associated with a greater risk of plant closure. 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 plant size and profitability, but is higher for plants with a works council.[14] The effect of plant-size is large. Very small plants (1–4 employees) have more than treble the closure rate of medium-sized plants (100–199 employees) $[100(e^{1.193} - 1) = 230\%]$, who themselves have treble the closure rate of very large plants (1000+ employees) $[100(e^{-2.262+1.193} - 1) = 191\%]$.

Worker characteristics are generally less important, but three results stand out. First, there is a significant relationship between plant closure and average wages. We model average wage in a plant using a dummy for each quintile: plants whose average wage falls in the lowest quintile form the base category. It is clear that plants whose pay is in the highest quintile are significantly more likely to close than the other 80% of plants. This is at odds with the notion that plants with higher unobserved productivity will be less likely to close and more likely to pay higher wages; 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 of the plant's workforce. Plants whose workers have an

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

Table 5: Plant-level closure hazard*

	Est.	*Std. Err.*	*p-value*
Plant-level covariates			
Plant is foreign-owned	0.140	(0.105)	[0.184]
Works council	0.366	(0.086)	[0.000]
Plant exports	−0.052	(0.073)	[0.476]
Plant is not part of a larger firm	0.032	(0.073)	[0.661]
Sectoral bargaining agreement	−0.032	(0.066)	[0.627]
Firm-level bargaining agreement	−0.381	(0.156)	[0.015]
Investment (relative to median)[a]	−0.811	(1.810)	[0.654]
5–9 workers	−0.410	(0.093)	[0.000]
10–19 workers	−0.549	(0.097)	[0.000]
20–49 workers	−0.634	(0.096)	[0.000]
50–99 workers	−0.823	(0.123)	[0.000]
100–199 workers	−1.193	(0.148)	[0.000]
200–499 workers	−1.365	(0.163)	[0.000]
500–999 workers	−1.653	(0.249)	[0.000]
$\geq$ 1000 workers	−2.262	(0.368)	[0.000]
Profits "good"	−0.075	(0.147)	[0.519]
Profits "satisfactory"	0.198	(0.143)	[0.167]
Profits "just sufficient"	0.554	(0.146)	[0.000]
Profits "bad"	1.057	(0.147)	[0.000]
Worker-level covariates			
Proportion non-German workers	−0.009	(0.169)	[0.958]
Proportion females	−0.151	(0.113)	[0.117]
Log average wage: second quintile	−0.104	(0.088)	[0.235]
Log average wage: third quintile	0.122	(0.087)	[0.163]
Log average wage: fourth quintile	0.136	(0.096)	[0.159]
Log average wage: fifth quintile	0.412	(0.105)	[0.000]
Average age: 30–35	−0.080	(0.128)	[0.532]
Average age: 36–40	−0.070	(0.121)	[0.560]
Average age: 41–45	0.063	(0.122)	[0.604]
Average age: 46+	0.378	(0.126)	[0.003]
Proportion with apprenticeship, no Abitur[b]	−0.120	(0.153)	[0.191]
Proportion with no apprenticeship, with Abitur	−0.915	(0.830)	[0.270]
Proportion with apprenticeship and Abitur	−0.782	(0.334)	[0.019]
Proportion with technical college degree	−0.577	(0.423)	[0.173]
Proportion with university education	0.032	(0.290)	[0.911]
Proportion with education unknown	−0.147	(0.156)	[0.311]
Average tenure: 3–5 years	−0.281	(0.089)	[0.002]
Average tenure: 6–8 years	−0.551	(0.116)	[0.000]
Average tenure: 9–11 years	−0.506	(0.128)	[0.000]
Average tenure: 12+ years	−0.519	(0.135)	[0.000]
Standard error unobs het, $\widehat{\text{se}}(u_j)$	0.004	(0.733)	*n/app*

* Proportional hazard, discrete baseline hazard, with Gaussian mixing and delayed entry. Log-likelihood functions given in Equations (3), (6), (8) and (7). The highest log-likelihood was always obtained with Stata's gh optimisation routine (with 12 quadrature points). Regressions also include dummies for location (9), industry (10), and year (4).

[a] Estimates are $\times 10^{-4}$.

[b] In addition to the dummies for the proportion of a plant's workforce with a given qualification, there are similar dummies for occupation.

average tenure of 6 or more years are less likely to close than plants with average tenure of 3–5 years (by a differential of ≈ 0.25) and by another 0.28 compared with plants whose tenure is less than 3 years. This is likely to be a genuine effect as we have already controlled for plant age, with which average tenure is strongly correlated (0.73). This is another reason why there is considerable duration dependence in the raw data: plants who survive longer have more experienced workers. Third, the average age of the workforce, which is likely to be correlated with average tenure, has a *positive* effect on the hazard to closure: plants whose employees have an average age of 45 years or more are more likely to close than the rest by approximately 0.40.

In Table 6 we summarise the estimated coefficient on foreign ownership for a variety of specifications. As discussed above, the raw effect is −0.051 and insignificantly different from zero; this becomes positive, 0.140, but still insignificantly different from zero when controlling for a full set of plant and worker characteristics and unobserved heterogeneity.

Table 6: Summary of foreign-ownership effects in plant-closure and worker-separation models

	Plant level			*Worker level*		
	Est.	*Std.Err.*	*p-value*	*Est.*	*Std.Err.*	*p-value*
Raw effect	−0.051	(0.097)	[0.598]	−0.102	(0.073)	[0.164]
(See Figures 1(b) and 2(b))						
Controlling for observed covariates	0.140	(0.105)	[0.184]	0.023	(0.058)	[0.690]
(See Figures 1(c) and 2(c))						
Controlling for unobserved heterogeneity	0.140	(0.105)	[0.184]	0.024	(0.058)	[0.690]
(See Tables 5 and 7)						
Separate plant-size regressions:						
1–9 employees (2,214 plants; 11,006 emp'ees)	0.603	(0.214)	[0.005]	0.076	(0.114)	[0.505]
10–19 employees (1,439; 21,216)	0.339	(0.339)	[0.318]	0.107	(0.116)	[0.349]
20–99 employees (3,033; 158,916)	−0.157	(0.198)	[0.429]	0.049	(0.053)	[0.357]
100–199 employees (875; 136,839)	−0.359	(0.324)	[0.268]	0.048	(0.076)	[0.527]
200+ (1,414; 1,000,502)	0.026	(0.278)	[0.925]	0.027	(0.069)	[0.700]
Plant Exports:						
Foreign-owned	0.365	(0.138)	[0.008]	0.078	(0.110)	[0.499]
Exporting	−0.001	(0.075)	[0.998]	−0.121	(0.048)	[0.012]
Foreign-owned and exporting	−0.462	(0.201)	[0.021]	−0.092	(0.126)	[0.490]
Plant is contracting:						
Foreign-owned				−0.101	(0.036)	[0.004]
Contracting				0.957	(0.038)	[0.000]
Foreign-owned and contracting				0.188	(0.081)	[0.021]

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. However, Deaton (2000) argues that if the true effect of the covariate (foreign ownership) is heterogenous, then the consistent estimates can only be obtained by modelling the heterogeneity. We do this by running separate regressions for five plant-size categories. Table 6 reveals that very small foreign-owned plants are significantly more likely to close than small domestic-owned plants, while the effect for the three largest size groups is negative, but poorly determined. This suggests that the 'homogeneous' estimate of 0.140 should be viewed as a weighted average of the five effects reported in the table, but, because the standard errors are so large, we cannot be precise about how the foreign ownership effect varies by plant-size.

A key finding of Alvarez & Görg (2005) is that foreign-owned plants are only more 'footloose' if their production is oriented towards domestic markets. We can test this proposition by interacting the foreign ownership dummy with the exporting dummy. The results, also shown in Table 6, confirm Alvarez & Görg's finding, but for Germany rather than Chile. Foreign-owned exporting plants have a significantly lower closure hazard than foreign-owned plants which do not export. To be precise, if a plant is foreign-owned and does not export, the effect is 0.365, whereas if a plant is foreign-owned and exports, the effect is $0.365 - 0.462 = -0.097$. Clearly, the 0.140 estimate above is a weighted average of these two effects. The interaction effect of –0.462 is significant. This large negative interaction effect does not arise because very small firms (1-9 employees) do not export; we have checked that the interaction effect is large and negative for both very small plants (–0.578) and the rest (–0.305). We have also checked that small foreign-owned plants do better than small domestic-owned plants, whether or not they export (0.813 for non-exporters and 0.235 for exporters).

5.2 Foreign ownership and worker separations

Whether or not foreign-owned plants 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 estimate the hazard to separation between workers and their plants, where a separation is defined as when worker i leaves plant j, and the plant does not close down. Note that separations which occur because of plant closure are treated as right censored (see Section 4).

Given 473,835 worker separations, when divided by 3,947,931 worker-years at risk, the raw worker hazard to separating is 0.1200.[15] It is interesting to note that the worker hiring rate is 0.1087 (computed as 328,577/(3,947,931−935,330) from Table 2). This suggests that the sample period is one of contraction amongst existing plants in the Western German labour market.

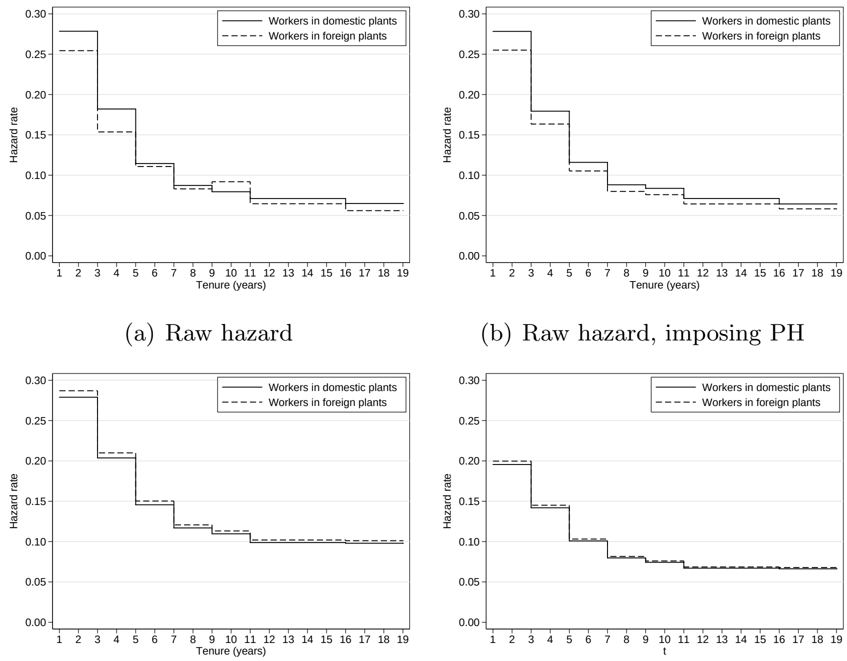
Of the 473,835 worker separations, 89,163 are in foreign-owned plants, and of the 3,947,931 worker-years at risk, 795,867 are in foreign-owned plants. This means that the raw hazard to worker separation is 0.0112 and 0.0122 in foreign-owned and domestic-owned plants respectively, giving a raw differential of –0.091 (0.084) log-points.

Figure 2 draws the estimated hazards for worker separations using the same four specifications used for the plant closure hazards.[16] Panel (a) plots the raw hazard of separating for workers in foreign-owned- and domestic-owned plants. As is well known, the separation hazard exhibits negative duration dependence. A number of

[15] For all our worker-level hazards, our regression sample comprises 3,947,931 worker-years and 1,328,479 workers.

[16] 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.

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.



(a) Raw hazard

(b) Raw hazard, imposing PH

(c) Hazard conditional on $\mathbf{x}_{it}$, imposing proportional hazards, ...

(d) ...and controlling for unobserved heterogeneity

Figure 2: Worker separation hazards

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 quality of 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.

In Figure 2(a), the hazard for workers in foreign-owned plants is below that for workers in domestic-owned plants for every tenure band except for one. As the two hazards are not very far apart, we do not reject the null that the hazard rates are equal ($\chi^2(7) = 12.80$, p-value 0.077). In addition, we cannot reject the null that the difference between the two hazards is a constant proportion ($\chi^2(6) = 9.39$, p-value 0.153). So, the resulting hazards 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 -0.102 log-points with a standard error of 0.073. This is insignificant, and it is a small effect, as shown in panel (b) of Figure 2.

In panel (c) we estimate the hazard after imposing the proportional hazards assumption and controlling for a full set of covariates $\mathbf{x}_i$; these are essentially the same as those used in the plant closure hazards. As with the plant hazard, the proportional hazards assumption is again not rejected: $\chi^2(6) = 9.45$, p-value 0.150. The estimated foreign ownership effect jumps from -0.102 (panel b) to 0.023 (panel c) when we include covariates, but is again insignificant (standard error 0.058). When we additionally control for unobserved heterogeneity, the foreign ownership effect is 0.024 (0.058), that is, does not change at all. So, even though the heterogeneity is significant (the log-likelihood improves by 52.0 log-points), neither the effect nor the shape of the baseline hazard is affected (panel d). Full estimates of panel (d) are reported in Table 7.

Thus, once we control for the differences in observable and unobservable characteristics between workers in foreign-owned and domestic-owned plants, workers in foreign-owned plants actually have *higher* separation rates. This is very similar to the effect of including covariates in the plant closure model. However, in contrast to the plant closure results, controlling for observed and unobserved heterogeneity does not 'flatten' the hazard, as would be expected if the shape of the hazard were drive entirely by selection effects.

A summary of the foreign ownership effect is reported in Table 6. We find that

Table 7: Worker-level separation hazard*

	Est.	*Std. Err.*	*p-value*
Plant-level covariates			
Plant is foreign-owned	0.024	(0.058)	[0.690]
Works council	−0.107	(0.031)	[0.001]
Plant exports	−0.148	(0.051)	[0.004]
Plant is not part of a larger firm	−0.039	(0.042)	[0.355]
Sectoral bargaining agreement	0.016	(0.034)	[0.670]
Plant-level bargaining agreement	0.011	(0.066)	[0.883]
Investment (relative to median)[a]	−0.233	(0.179)	[0.197]
5–9 workers	0.054	(0.046)	[0.254]
10–19 workers	0.098	(0.047)	[0.038]
20–49 workers	0.056	(0.046)	[0.234]
50–99 workers	0.121	(0.051)	[0.020]
100–199 workers	0.136	(0.056)	[0.016]
200–499 workers	0.132	(0.059)	[0.028]
500–999 workers	0.052	(0.069)	[0.454]
$\geq$ 1000 workers	0.017	(0.081)	[0.843]
Profits "good"	−0.019	(0.074)	[0.802]
Profits "satisfactory"	0.068	(0.074)	[0.357]
Profits "just sufficient"	0.177	(0.076)	[0.021]
Profits "bad"	0.197	(0.086)	[0.023]
Plant age: 6–10 years	0.166	(0.067)	[0.019]
Plant age: 11–15 years	0.151	(0.071)	[0.040]
Plant age: 16-20 years	0.134	(0.064)	[0.045]
Plant age: 21+ years	0.142	(0.054)	[0.012]
Worker-level covariates			
Non-German worker	0.065	(0.017)	[0.000]
Female	−0.016	(0.016)	[0.365]
Apprentice	−0.595	(0.038)	[0.000]
Part-time	−0.054	(0.026)	[0.064]
Home worker	−0.080	(0.146)	[0.615]
log wage: second quintile	−0.384	(0.021)	[0.000]
log wage: third quintile	−0.606	(0.030)	[0.000]
log wage: fourth quintile	−0.741	(0.035)	[0.000]
log wage: fifth quintile	−0.639	(0.038)	[0.000]
Age: 21–30	−0.121	(0.018)	[0.000]
Age: 31–40	−0.403	(0.021)	[0.000]
Age: 41–50	−0.505	(0.024)	[0.000]
Age: 51–55	−0.324	(0.042)	[0.000]
Age: 56+	0.616	(0.037)	[0.000]
Apprenticeship, no Abitur	−0.092	(0.022)	[0.000]
No apprenticeship, with Abitur	0.327	(0.032)	[0.000]
Apprenticeship and Abitur	0.039	(0.031)	[0.225]
Technical college degree	0.026	(0.038)	[0.498]
University education	0.096	(0.038)	[0.015]
Education unknown	0.040	(0.030)	[0.218]
Standard error unobs het, $\widehat{\text{se}}(u_i)$	0.266	(0.014)	*n/app*

* Proportional hazard, discrete baseline hazard, with Gaussian mixing and delayed entry. Log-likelihood functions given in Equations (3), (6), (8) and (7). Regressions also include dummies for location (9), industry (10), occupation (10) and year (4). The highest log-likelihood was always obtained with Stata's (default) mvagh optimisation routine (with 12 quadrature points).

[a] Estimates are $\times 10^{-4}$.

exporting plants do have a significantly lower separation rate, but the effect is insignificantly different between foreign-owned plants and domestic-owned plants. Furthermore, the worker separation rate does not vary significantly across plant-size categories, as shown in Table 6.

As already noted, our data do not distinguish between separations that are layoffs, and those that are quits. It is therefore possible that the small overall difference in separation rates between foreign- and domestic-owned plants disguises significantly different quit and layoff rates. Evidence suggests that layoff and quit rates vary systematically with changes in employment at the plant level (Abowd, Corbel & Kramarz 1999). We therefore examine whether the separation rate varies between plants which are contracting their workforce and those that are not.

A plant is defined as 'contracting' if its employment declines by more than 5% in the preceding year.[17] We then interact this dummy with the foreign ownership dummy. If foreign-owned plants have more volatile labour demand, or if they have more aggressive human resource policies (e.g. firing workers in a downturn rather than cutting back on hires), then the exit rate for declining plants will be higher if the plant if foreign-owned. The estimated interaction effect is 0.188 and significant (standard error 0.081). Hence the overall foreign-ownership effect on the hazard to worker separation, 0.024, is a weighted average of 0.087 (0.078) if the worker's plant is contracting and –0.101 (0.036) if it is expanding. Thus, while there is a significant difference between expanding and contracting plants, foreign-owned contracting plants do not have a significantly higher separation rate. Rather, our results are consistent with the idea that the quit rate is lower in foreign-owned plants.

In general, the effects of the covariates in the worker separation model are quite different from those in the plant closure model. The separation hazard is an inverted-U shape in plant size: very small and very large plants have the lowest separation

[17] We experimented with different definitions of a 'contracting' plant, ranging from 0% to 10%; 5% is where the foreign ownership effect is the biggest. We also checked whether expand/contract dummies affects other controls and have interacted it with plant-size. See below.

hazards. Separations are decreasing in profitability: plants with "bad" or "just sufficient" profits have a separation hazard about 0.2 higher than plants with "good" or "very good" profits. Workers in plants who are older than 6 years are more likely to separate than the rest by about 0.14 log-points. Workers in plants which have a works council have a lower separation hazard (-0.107); this contrasts with the positive and significant effect in Table 5, where the plant closure rate was 0.366 (0.086) higher than plants without a works council. Thus, it is a moot point whether it is a better to be in a plant that has a works council, given the sizeable chance of the plant closing down.

There are some interesting effects for worker-level covariates. First, non-German workers are more likely to exit than their German counterparts by 0.065 log-points. Second, as workers get older, they are less likely to separate until their mid-50s, when there is big jump of 0.940 log-points between those aged 51–55 and those aged 56–65. These very large effects arise because of early retirement. A worker aged 41–50 is 0.505 log-points less likely to separate than a worker aged 20 or less. Finally, there are significant effects of the wage on the separation hazard. Workers whose wage is in the middle quintile of the wage distribution are much less likely to separate than workers in the second quintile (by some 0.22 log-points), who themselves are less likely to separate than those in the lowest quintile by 0.38 log-points. The differentials at top end of the distribution are much smaller. So there is a non-linear but negative effect of wages on the separation hazard. Clearly wages are endogenous: plants may pay higher wages to (unobservably) more productive workers, workers who they want to retain. The effects of wages and age on the worker separation hazard are not dissimilar to the same effects on the plant exit hazard.

6 Conclusions

This paper provides the first evidence on the relationship between foreign ownership, plant survival and job security for Germany. We also provide the first evidence based on linked employer-employee data which utilises a consistent econometric method both for plant survival and for job security. Our findings cast doubt on the hypothesis that foreign-owned plants are more "footloose", or that jobs in foreign-owned plants are less secure. In the raw data, foreign-owned plants have lower closure rates and lower worker separation rates, but these differences are insignificant. After controlling for different observable and unobservable characteristics, foreign-owned plants do not have significantly higher closure rates and their workers do not have significantly higher separation rates. Our estimates are also quantitatively small, as clearly illustrated by the hazard rates in Figures 1 and 2.

In fact, our findings on job security are entirely consistent with the small existing literature. Görg & Strobl (2003) find no evidence of greater job turnover in foreign-owned plants; Pesola (2008) finds no permanent effect of takeover on job separation rates.

Our finding that foreign-owned plants are no less likely to survive appears less consistent with some previous findings. But there are a number of reasons why this is unsurprising. First, the literature does not unanimously find significant effects of foreign ownership on plant survival: Alvarez & Görg (2009), for example, find no effect in the sample as a whole. Second, our sample of German plants covers not only manufacturing but the whole economy, and should therefore be regarded as more representative. But even within the manufacturing sector, we find no significant difference in plant survival between foreign-owned and domestic plants.[18] Third, it seems likely that foreign-owned plants in an advanced economy such as Germany are less distinct from domestic plants than in a developing economy, such as Indone-

[18] The estimated foreign-ownership effect in manufacturing from the preferred model is 0.042 with a standard error of 0.172.

sia. It seems entirely plausible that the large effects found by Bernard & Sjöholm (2003) reflect this difference, but it remains to be seen whether the footloose nature of foreign-owned plants is consistently higher in less developed countries.

Our other key finding is that the average effect across the whole sample masks significant differences across different types of plant. First, foreign-owned plants which do not export have significantly higher closure rates. This is consistent with findings for Chile by Alvarez & Görg (2009), who suggest that exporting multinationals are less affected by shocks to the domestic market than multinationals which produce only for the domestic market.

Second, foreign-owned plants which are not contracting their workforce have significantly lower worker separation rates. This finding is consistent with the notion that jobs in foreign-owned plants are better in some other dimension (not in our model), encouraging lower quit-rates.

Our evidence suggests that the presence of foreign-owned plants in Germany is unlikely to be a significant source of greater employment risk, as suggested by the survey results from the UK reported by Scheve & Slaughter (2004). Comparisons of results from several countries, using similar data and methodologies, would be required to discover whether the labour market effects of foreign-owned firms varies in systematic ways between countries.

References

Abowd, J., Corbel, P. & Kramarz, F. (1999), "The entry and exit of workers and the growth of employment: an analysis of French establishments", *The Review of Economics and Statistics* **81**, 170–187.

Addison, J., Bellman, L. & Kölling, A. (2004), "Works councils and plant closings in Germany", *British Journal of Industrial Relations* **42**(1), 125–148.

Addison, J., Schnabel, C. & Wagner, J. (2001), "Works councils in Germany: their effects on establishment performance", *Oxford Economic Papers* **53**, 659–694.

Alvarez, R. & Görg, H. (2005), "Multinationals and plant exit: Evidence from Chile", IZA Discussion Paper 1611.

Alvarez, R. & Görg, H. (2009), "Multinationals and plant exit: Evidence from Chile", *International Review of Economics and Finance* **18**, 45–51.

Anderson, P. & Meyer, B. (1994), "The extent and consequences of job turnover", *Brookings Papers: Microeconomics* .

Andrews, M., Bellmann, L., Schank, T. & Upward, R. (Forthcoming), "The takeover and selection effects of foreign ownership in Germany: An analysis using linked worker-firm data", *Review of World Economics* .

Audretsch, D. & Mahmood, T. (1995), "New firm survival: New results using a hazard function", *The Review of Economics and Statistics* **77**(1), 97–103.

Becker, G. (1962), "Investment in human capital: a theoretical analysis", *Journal of Political Economy* **70 (Sup.)**, 9–49.

Bender, S., Haas, A. & Klose, C. (2000), "IAB employment subsample: opportunities for analysis provided by the anonymised subsample", IZA discussion paper 117. *Forschungsinstitut zur Zukunft der Arbeit*, Bonn.

Bernard, A. & Jensen, J. (2007), "Firm structure, multinationals, and manufacturing plant deaths", *The Review of Economics and Statistics* **89**(2), 193–204.

Bernard, A. & Sjöholm, F. (2003), "Foreign owners and plant survival", NBER working paper 10039.

Cameron, C. & Trivedi, P. (2005), *Microeconometrics: methods and applications*, Cambridge.

Deaton, A. (2000), *The Analysis of Household Surveys: a Microeconometric Approach to Development Policy*, Baltimore: Johns Hopkins.

Disney, R., Haskel, J. & Heden, Y. (2003), "Entry, exit and establishment survival in UK manufacturing", *Journal of Industrial Economics* **LI**(1), 91–112.

Dunne, T., Roberts, M. & Samuelson, L. (1988), "Patterns of firm entry and exit in U.S. manufacturing industries", *RAND Journal of Economics* **19**(4), 495–515.

Dunne, T., Roberts, M. & Samuelson, L. (1989), "The growth and failure of U.S. manufacturing plants", *Quarterly Journal of Economics* **104**(4), 671–698.

Fabbri, F., Haskel, J. & Slaughter, J. (2003), "Does nationality of ownership matter for labor demands?", *Journal of the European Economic Association* **1**(2), 698–707.

Fischer, G., Janik, F., Müller, D. & Schmucker, A. (2009), "The IAB establishment panel — things users should know", *Schmollers Jahrbuch* **129**(1), 133–148.

Girma, S. & Görg, H. (2004), "Blessing or curse? Domestic plants' survival and employment prospects after foreign acquisitions", *Applied Economics Quarterly* **50**(1), 89–110.

Glass, A. & Saggi, K. (2002), "Multinational firms and technology transfer", *Scandinavian Journal of Economics* **104**, 495–513.

Görg, H. & Strobl, E. (2003), "Footloose multinationals", *The Manchester School* **71**(1), 1–19.

Guo, G. (1993), "Event-history analysis for left-truncated data", *in* P. Marsden, ed., *Sociological Methodology*, Cambridge M.A.: Blackwell, pp. 217–243.

Hopenhayn, H. (1992), "Entry, exit, and firm dynamics in long run equilibrium", *Econometrica* **60**(5), 1127–1150.

Jenkins, S. (2005), "Survival analysis", Monograph, University of Essex.

Jovanovic, B. (1979), "Job matching and the theory of turnover", *Journal of Political Economy* **87**, 972–990.

Jovanovic, B. (1982), "Selection and the evolution of industry", *Econometrica* **50**(3), 649–670.

Mata, J. & Portugal, P. (1994), "Life duration of new firms", *Journal of Industrial Economics* **XLII**(3), 227–245.

Navaretti, G., Turrini, A. & Checchi, D. (2003), "Adjusting labor demand: Multinational versus national firms: A cross-European analysis", *Journal of the European Economic Association* **1**(2/3), 698–707.

Pesola, H. (2008), "Labour market transitions following foreign acquisitions", Mimeo, Department of Economics, Helsinki School of Economics.

Scheve, K. & Slaughter, M. (2004), "Economic insecurity and the globalization of production", *American Journal of Political Science* **48**(4), 662–674.

Wagner, J. (1994), "The post-entry performance of new small firms in German manufacturing industries", *Journal of Industrial Economics* **42**(2), 141–154.

Wooldridge, J. (2002), *The econometric analysis of cross section and panel data*, MIT Press, Cambridge MA.