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IT''''S BEEN A LONG TIME: A COMPARATIVE ANALYSIS OF JOB DURATION IN BANKING

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ABSTRACT

Using personnel records from two firms in the banking industry, duration models are estimated to examine separations in the context of Great Britain and Greece. We find that it is sustained, rather than instantaneous, performance that is linked to separations. In common with some earlier studies, we find qualified support for a u-shaped relationship between performance and separations, but only in the case of the British data. Both of the banks under investigation experienced substantial reorganisation activity over the time period considered, and we find that the year following this was characterised by increased separation propensities. While most of our findings are consistent across the firms in the two countries studied, we find that single men are more likely than their female counterparts to quit in Britain, but less likely to quit in Greece. We offer some suggestions about why this should be the case.

Keywords: duration modelling, labour turnover, personnel economics
JEL Classification: C41, J63

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1. Introduction

Labour turnover has, for many years, attracted the attention of analysts interested in improving our understanding of labour markets and business performance. Separation decisions by workers often represent balancing the benefits of job shopping (Johnson, 1978) against those associated with the accumulation of specific human capital (Becker, 1962). The literature on personnel economics emphasises the role played by information, and in this context the asymmetry between current and prospective employers' information about worker performance is of particular interest (Lazear, 1986, 1995). More generally, from the viewpoint of firms, turnover is costly in the presence of hiring (replacement) costs; at the same time, too little turnover can be inefficient in that it perpetuates poor worker-firm matches.

The present paper concerns turnover in a specific industry, namely the banking sector. This is an interesting context in which to analyse turnover because the financial services industry has been much affected by technological and social changes over the last quarter century. The development of information technology has led to substitution of capital for labour; social developments have led to an increased feminisation of the workforce; and at the same time there has been a sustained increase in the demand for banking services.

We have been fortunate to acquire data sets for banks in two countries of the European Union – the United Kingdom and Greece – that provide some interesting points of comparison. Financial services play an important role in both economies, but labour markets are often deemed, by European standards, to be particularly flexible in the UK. In many respects, Greece epitomises a more traditional labour market. An interesting aspect of the present study therefore involves a comparison of how turnover differs across systems.

In analysing the determinants of labour turnover for the two firms in our study, we are able to cast light on an important controversy in the current literature – what is the nature of the relationship between worker performance and turnover? There is debate about whether this relationship is negative or nonmonotonic (u-shaped). We find, for our British data, convincing evidence in favour of the latter; we fail to find any such relationship in the case of our Greek data, and there must therefore be some doubt about how general the u-shaped relationship might be. We are also able to ask and answer some new questions about the dynamics of this relationship. Specifically, is turnover related to instantaneous or sustained performance? A further issue which we are able to address concerns the impact of company reorganisationactivity on turnover.

The structure of the remainder of the paper is as follows. Section 2 briefly reviews some relevant literature. Our data are described in section 3, and the following section reports on the empirical analysis. The paper ends with a conclusion and some suggestions for further research.

2. Literature

Two key precursors of the current research are papers by Trevor *et al*. (1997) and Salamin and Hom (2005). Both use duration modelling as a means of testing the prediction of Jackofsky (1984) that there should be a u-shaped relationship between workers' performance on the job and their propensity to separate from their employer. The rationale for this relationship is that poor performers should be likely to quit (or be fired) as they perceive their prospects within the firm to be poor, while strong performers should perceive strong outside opportunities.

Trevor *et al*. (1997) modelled all quits (over the 1983-90 period) in a proportional hazards framework (Cox, 1972), using as their data base all hires over the 1983-88 period in a large organisation. Salamin and Hom (2005) likewise conduct a Cox analysis using data gathered on new hires over a 5 year period (1994-1999) to a Swiss bank. They analyse quits that occur by the end of this same 5 year period. Amongst their findings are that the quit rate is relatively high for women, younger workers[1], married workers, and workers employed in the lower grades.

The work by Trevor *et al*. (1997), in particular, is widely seen as seminal not least because it overturns the results of an earlier literature which modelled turnover only as a linear function of performance, and which found a negative relationship (see, for example, studies reported in the meta-analysis of Williams and Livingstone, 1994).[2] It is also a key reference because it provides an early example of the use of duration modelling in this context.

An interesting feature of the studies by both Trevor *et al*. (1997) and Salamin and Hom (2005) is the role played by promotion and salary as mechanisms designed to reduce turnover amongst the stronger performers. Indeed it is likely that the finding of a straightforward negative relationship between turnover and performance in earlier studies is partly due to the failure to account for these confounding effects. In the work that follows, we do not include salary and promotion as explicit variables, but we do include as cofactors a vector of variables that includes the common determinants of earnings (such as gender, ethnicity, and polynomial terms in age)[3]; we also include an indicator of the grade at which workers are employed. This being the case, we would expect to observe a u-shaped relationship between performance and separations.

Other work in this area includes a study by Dohmen and Pfann (2004) which estimates a Cox proportional hazard model on data from a large aircraft manufacturer. In this case the firm experienced a period of rapid growth followed by a quick contraction which eventually led to bankruptcy. Using data on separations over a 10 year period, Dohmen and Pfann find that there is a complex relationship between age and quits. As a rule, younger workers are relatively likely to quit. For older workers, notably those of an age suitable for early retirement, the relationship is highly

[1] They do not test for a quadratic relationship between age and quits.
[2] Williams and Livingstone consider also a nonlinear relationship, but base their analysis of this on only eight studies.
[3] We have concerns about the multicollinearity that can result when remuneration is included in models of the type estimated in the sequel.

nonlinear and, inasmuch as it varies over time, clearly affected by the struggles of the firm to manage its own contraction.

Dostie (2005) presents estimates of a duration model using matched employer-employee data for France. The data come from surveys conducted by the Institut National de la Statistique et des Etudes Economiques (INSEE). The duration model in this instance forms part of a larger system in which the impact of seniority on wages is evaluated. Dostie finds that schooling and being male both negatively influence the separation propensity, and that, while both seniority and experience serve to reduce the likelihood of quitting, this effect is strongest for the newest recruits and the least experienced workers respectively.

The studies reported above all adopt a nonparametric form of the hazard function. This is in line with common practice in the literature, but it is worth noting that parametric alternatives exist (Weibull, 1951; Cleves *et al*., 2004). These have the characteristic – at one and the same time an advantage and a disadvantage – that they impose structure on the data. This is an advantage in that it may smooth out noise, but it has the drawback that it could force the relationship predicted by the model between tenure and separation propensity to be simpler than the true relationship. For instance, it might make a nonmonotonic relationship appear to be monotonic.

Another feature of the literature on statistical duration models that has not been captured in most studies reported above concerns unobserved heterogeneity across workers.[4] The panel nature of the data that are characteristically used in studies of this type enable us to allow for differences between workers that can be assumed constant over time, simply by including fixed or random effects (or, more generally, random parameters) in the models. Hence characteristics of workers that are difficult to observe – such as motivation – can be accommodated within models of this type. The unobserved heterogeneity that is addressed by such models is often referred to by statisticians as ‘fragility’. The extent to which fragility can be accommodated within statistical duration models is often data-dependent; such models are computationally demanding, and it is often the case that fragility has to be modelled in a fairly crude fashion. A particularly popular model in this respect, and one that we shall use in the analysis that follows, is that of Heckman and Singer (1984) in which the data are used to group workers into types, each type having a distinct intercept in the hazard function.

In the next section we describe the data upon which our empirical analysis draws. This is followed by the analysis itself.

3. Data

Data for this study come from two sources. The UK data come from the full personnel records of a major bank. These data have been previously analysed by researchers studying remuneration (Audas *et al*., 2004) and the hierarchical structure of organisations (Treble *et al*., 2001). The data cover the period 1989 through 2001. The

[4] Dostie (2005) allows for unobserved heterogeneity in his mixed duration model.

Greek data come from records of leavers from a major commercial bank, drawn over a similar period.

The information contained in these records varies a little across the two banks in question, but for both banks we have data on gender, marital status, age, grade, tenure (in years), region, and a measure of assessed performance.[5] In the case of the British data, we also have information about ethnicity, and limited data on health (a disability indicator) and education (whether or not the respondent is a graduate).

There are other key differences between the two datasets. The British data are collected monthly and refer to all staff employed by the bank. Data on tenure, though, for those who joined the bank before the earliest date for which we have data, are recorded in years.[6] This being the case, we use years as the unit of time in which durations are measured, and employ data only from the January return of each year. Data on gender and ethnicity remain unchanged over the worker's spell of unemployment; age and tenure increase by one year each period; other variables may or may not change from period to period.

In the case of the Greek data, information is collected at only one point in time – the period in which the worker leaves the firm. We are able to infer the way in which age and tenure change over the spell of employment, and we know that gender does not change over time. The remaining variables (marital status, grade, region and performance) could potentially change over the worker's spell of employment. For the purposes of the present exercise, we assume that they do not, since the focus of our analysis is on the separation decision.

The British data cover all workers – leavers and stayers – and so we are able to use these data to analyse leaving decisions up to 2001. In the case of the Greek data, we have information only about leavers, and so it is not immediately obvious how we can analyse these in relation to a control group of stayers. To mitigate this problem, we analyse leavers only up to 1995, and treat workers who have not left by this time as stayers. We recognise that the group of stayers that is identified in this way is a non-random selection of non-leavers; they may, for example, be stayers who are more likely than other stayers to leave over the subsequent few years (for reasons of age or gender, or other observed or unobserved factors). Unfortunately our data do not allow us any obvious alternative approach.[7] While our analyses lead us to think that biases

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[5] In the case of the British data, different performance measures have been employed in various divisions of the company, and so we have collapsed these measures into a simple three point indicator describing below average, average, and above average performance. In the case of Greece, performance is measured on a continuous scale from 60 to 140. However there are very few ratings lower than 110 or higher than 130. This being the case, and for ease of comparison with the British data, we have collapsed the ratings into a three point indicator describing workers whose ratings are in the ranges 60-114, 115-124, and 125-140.

[6] For a small number of observations, data on tenure appear to be incorrectly coded, since they imply that individuals have been with the firm since before they reached school leaving age. These observations are omitted from our sample.

[7] While central records have been kept on leavers, the collection of data on current employees has been conducted in a decentralised manner. For some divisions of the bank, historical data are not available in computerised form. Our decision to use 1995 as a cut-off means that workers who leave after this date are right censored; censoring is discussed in more detail in the next section.

due to sample selection issues in the Greek data are fairly minor, any rigorous assessment of this claim must await the availability of more complete data.

Descriptive statistics for the main variables used in the analysis are reported in Table 1. There are clear differences in the nature of the samples between the two datasets. In large measure these reflect the fact that the Greek sample is drawn from leavers. These workers are, on average, therefore older and (considerably) more senior than their counterparts in the British data – though there are differences in the way in which seniority is measured across the two datasets that exaggerate this gap. Moreover, while the British data comprise mostly women, the Greek data comprise mostly men. This reflects differences in the national structure of employment.

<INSERT TABLE 1 ABOUT HERE>

4. Analysis

In this section we estimate a number of variants of the duration model, using in turn the data drawn from the British and Greek firms. We use a variety of estimation methodologies.

First, we estimate the baseline hazard using a parametric specification, namely the Weibull (1951). The hazard associated with individual j at time t and having characteristics $\mathbf{x}_j$, is given by the equation

$$h(t|x_j) = pt^{p-1}exp(a)exp(x_j\beta_x) \qquad (1)$$

where p, a and the vector are all parameters that are to be estimated. The term $exp(x_j\beta_x)$ is individual-specific, while the term $pt^{p-1}exp(a)$ represents the baseline hazard and is clearly a function of time. If p<1 the baseline hazard is falling over time; if p=1 it is constant, and if p>1 it is rising over time.

A limitation of the Weibull model is that it imposes monotonicity on the hazard (with respect to duration). A popular alternative approach is to assume a nonparametric baseline hazard; such an approach is adopted in the Cox (1972) proportional hazards model. Here the hazard attached to the jth individual at time t, given characteristics x_j, is estimated as

$$h(t|x_j) = h_0(t)exp(x_j\beta_x) \qquad (2)$$

where β_x denotes the vector coefficients to be estimated. In this model, the baseline hazard, h_0, is not estimated parametrically; it is assumed instead that, whatever it looks like, this baseline is the same for every individual. Individuals' hazards differ only as a result of differences in the $exp(x_j\beta_x)$ term. This model therefore allows h_0 to vary with t in a manner that is wholly determined by the data, not by the imposition of any functional form. This is an appealing feature of this model, not least because it allows life-cycle events (such as the tendency for new hires to be job shopping and the tendency for long duration employees to retire) to be accommodated naturally into the model. Since the first term on the right hand side of (2) is nonparametric while the second term is parametric, models of this type are often referred to as semiparametric duration models.

Neither of the approaches discussed above makes full use of the panel nature of the data that are typically available for the analysis of duration. Repeated observations of the same individuals over time enable an allowance to be made for unobserved heterogeneity. The Heckman and Singer (1984) model allows this to be modelled in a simple fashion; this model is particularly appealing because it strikes a balance between the needs, on the one hand, to ensure that heterogeneity is accommodated, and, on the other, for computational feasibility.[8] In essence, the approach is to allow the intercept term, β_0, to take one of two values (say β_0' and β_0''), and then maximise the likelihood associated with the hazard with respect to choice of the parameters β_0', β_0'', the remaining elements of the vector β_x, and information about which individuals comprise the membership of which group (Jenkins, 2005, pp. 85-86).

Two considerations which must be borne in mind for all the models estimated in this section. First, the data that we use are left truncated. In the case of the British bank, the monthly data are available from the start of 1989 onwards, but we include individuals who joined the bank before (and in some cases many years before) this date. There are other individuals who both joined and left the bank before this date, but who we do not observe at all. As a consequence, we have a biased sample of those who entered before the earliest of our monthly censuses. This bias can easily be accommodated in duration models by a modification to the contribution that early entrants make to the likelihood (in the case of parametric models) or by omitting these entrants from the analysis for the truncated period (in the case of semiparametric models). In the case of the Greek data, a similar consideration applies; we have information about leavers only from 1989 onwards, though we know when these leavers joined the bank even if they did so long before our data started to be collected. Again, therefore, there are leavers who we fail to observe because they left before the starting date of our census of leavers.

The second concern is that our data are right censored. For those individuals that leave the firm within the time window for which we have data, our set of relevant information is complete. For others, we know only that they leave the firm at some unspecified time in the future – we do not know when. As is the case for truncation, allowing for censoring involves making minor modifications to the likelihood (for parametric models) or the treatment of individual cases (for semiparametric models). Cleves *et al*. (2004) provide full details of the procedures involved, and these are followed in the estimation of all models reported in the remainder of the present paper.

<INSERT TABLE 2 ABOUT HERE>

[8] In early work we tried to estimate random parameter models that allow for heterogeneity at the level of the individual (that is, without classifying individuals into a limited number of groups). We interrupted the programs after a few days as they had, even with generous convergence criteria, failed to reach a solution within that time.

In Table 2 we present results obtained by estimating Weibull, Cox, and Heckman-Singer models for the data from the British bank. A number of specifications are reported in order to facilitate comparisons both across methods and across the two banks.[9] Results are fairly stable across the Weibull and Cox methods. There is a sign reversal for the marital status variable in some specifications but where this happens the variable is not significantly different from zero. The coefficient on marital status is, however, significantly negative in the models which do not include a quadratic term in age, and this effect is (at a generous level of significance) stronger for men than for women. This result is unsurprising in view of family responsibilities.

In all specifications, we see that men are more likely than women to separate from the firm. This is a somewhat surprising result in view of the well documented gender difference in labour turnover rates. Note, however, that the coefficient on the interaction term denoting men who are married has a negative coefficient, and that this more than offsets the coefficient on the male dummy. Married men are therefore less likely to quit than (married or unmarried) women, and less likely to quit than unmarried men.

The propensity to leave the firm varies nonlinearly with age. Younger and older workers are more likely than prime age workers to leave. The turning point varies somewhat across specifications, but is typically in the late 30s.
Workers in training grades and in the most senior grades are more likely to leave than others. Hence rank exerts a nonlinear impact on separation probability that operates in a similar manner to age.

Graduates are more likely to separate from the firm than are non-graduates. This is likely due to attractive outside employment opportunities that are available for the former group. Part time workers are relatively likely to separate; this is unsurprising in view of the weaker attachment of such workers to the labour force. Ethnicity also exerts a significant effect on the likelihood of separation, with minority workers being more likely than others to separate in any given period.

A full set of dummy variables is used for the standard regions of the UK, though for reasons of space we do not report these coefficients in the table. There is not a great deal of regional variation in separation propensities; an exception is Scotland where separation rates appear to be higher than is the case for workers located elsewhere. Possibly as a result of low migration rates, this effect is not observed in the Heckman-Singer model. It is not clear, therefore, whether the finding of differences between Scotland and the rest of the UK is due to unobserved heterogeneity or a genuine Scottish effect.

[9] Some variables (education, ethnicity, part-time status, and –owing to the fact that data are collected only at one point in time – lagged performance rating) are not available in the case of the Greek data, so we report for the British bank some parsimonious specifications of the model that exclude these variables. Moreover, the Heckman-Singer model failed to converge when age squared is included as an explanatory variable, so to allow straightforward comparison we report some specifications of the Weibull and Cox models where this variable is excluded.

As noted earlier, the bank was involved in a a significant internal reorganisationduring the timeframe covered by this study.[10] We find that separations of workers from the bank were significantly more likely during the year following this reorganisation than at other times. This suggests that, either through natural wastage or redundancy, cuts in staffing were sought following the reorganisation. A curious exception to this result appears in the Heckman-Singer estimates, where the sign on the coefficient becomes negative. We have been unable to pin down the reason for this, but we suspect that it is linked to the fact that the reorganisation had a regional dimension and that we are therefore observing the effects of collinearity between one of the intercept terms in the Heckman-Singer model, the reorganisation term, and the regional dummies. We are therefore inclined to attach more weight to the finding of a positive coefficient on the reorganisation variable in the straightforward Weibull and Cox models.

The impact of the worker's performance rating on separation propensity is interesting, and provides a focal point of the present study. Our measure of rating is an integer variable which is positively related to performance and which ranges from 0 to 2. In all specifications of the model improved performance is associated with a reduced probability of separation. In specifications that include a quadratic term in rating, there is a marked nonlinearity, and a plot of separation propensity against (a continuous measure of) performance is u-shaped with a minimum that lies near 1.6. However if a worker moves from a rating of 1 to a rating of 2, *ceteris paribus*, her separation propensity decreases (though by far less than it would decrease if she were to move from a rating of 0 to a rating of 1). We have, therefore, found a u-shaped relationship of sorts between performance and separations; there is a nonlinearity, but the turning point is too high to allow us to conclude with confidence that the best performers are more likely to separate than are average performers.

A further issue of interest concerns the impact of sustained, as opposed to instantaneous, performance on the probability of separation. In some specifications, we include a lagged rating term. In these equations, both lagged rating and current rating have a significant negative effect on attrition. Workers who have a longstanding problem of poor performance are thus more likely to leave than are others.

<INSERT TABLE 3 ABOUT HERE>

We now turn to consider the Greek data. Table 3 reports results from a variety of models. As in the case of the UK models, the estimated coefficients are broadly robust across specifications, but there are some exceptions. The interaction term between marital status and gender varies across models; while it is insignificant in Weibull and Cox models, it is significantly (and counterintuitively) positive in the Heckman-Singer specification. There is a sign reversal on the age variable between the models that do (on the one hand) and do not (on the other) allow for frailty.

In models that include a quadratic term in age, separations become less likely as a worker ages until she reaches her mid-50s. They then become more likely,

[10] Reorganisations of this kind can result from a number of factors that are not necessarily internal to the firm; these include merger or takeover, new legislation, and step changes in the competitive environment.

presumably as a result of retirement. Note that this turning point occurs at a much later point in the worker's life cycle than is the case in the analysis of UK data.

The separation propensity is significantly lower for men than for women in all specifications, whether respondents are married or not. This provides a contrast with the UK results discussed earlier. In some measure this may reflect the greater flexibility that characterises the UK labour market, with many women switching to part-time work rather than leaving the labour market altogether when they take on family commitments. Marriage makes separation less likely for all workers.

Region is captured using 9 dummies, though these are not reported in the table for reasons of space. There is some evidence, from the Heckman and Singer model, that the propensity to separate is somewhat higher in the northern regions of Thessaly and Macedonia and in the Aegean islands and Crete. Meanwhile, in the Central Greece region (which includes the capital, Athens), the separation rate is somewhat lower than elsewhere, other things being equal.

There is a tendency for increasingly senior workers to be less likely than others to leave the firm. Although the coefficient on training workers is also negative, it is not significant.

As was the case for the British bank, the Greek bank was involved in a reorganisation during the period under consideration. Separation of workers from the company was significantly more likely during the year following this than at other times under consideration.

In the case of the Greek bank, the relationship between performance and separation propensity does not follow a u-shaped relationship. Broadly speaking, improved performance is associated with a reduced probability of separation. The results suggest, however, that the relationship follows an inverse-u shape, with the turning point of the function being at a low value of performance. Taking the estimates from column 1 of the table as an example, the turning point is 0.56. While moving from a performance rating of 0 to 1 would lead to a (slight) increase in separation propensity, a move from x, $0.12<x<1$, to 1 or a move from 1 to 2 would lead to a decrease. We have experimented with specifications of the hazards in which a quadratic term in ratings is not included as an explanatory variable, and these confirm that the relationship between performance and separation in the Greek data is a negative one.

5. Conclusion

The results reported in this paper contribute to what remains at this juncture a very small collection of studies on labour turnover based on company personnel data. In so doing, we have been able to provide some further support for a u-shaped relation between performance and turnover in the case of the British company, but not in the case of the Greek company. We have also contributed new insights about the impact of sustained performance and of company reorganisation on worker-firm separations.

While many of our results point to similar behaviour amongst bank workers in the UK and Greece, there are some differences. With a sample of just one firm in each country, we cannot do more than speculate on the sources of these differences. Yet they are consistent with the view that Britain has developed the more flexible labour market. Further insight on the relationship between overall labour market flexibility and the determinants of turnover within firms must await the availability of personnel records for more companies.

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Table 1 Descriptive statistics

Variable	United Kingdom	Greece
Male	0.3465	0.6230
	(0.4758)	(0.4844)
Married	0.6363	0.6696
	(0.4810)	(0.4704)
Age	34.78	45.3933
	(9.38)	(8.5223)
Training grade	0.2971	0.0078
	(0.4569)	(0.0882)
Middle grade	0.1862	0.2007
	(0.3892)	(0.4005)
Senior grade	0.0480	0.4831
	(0. 2138)	(0.4997)
Performance rating	1.2197	0.9326
	(0.4209)	(0.8145)
White	0.9108	
	(0.2849)	
Part-time	0.1953	
	(0.3964)	
Graduate	0.0470	
	(0.2118)	

Note: Table shows means or proportions and standard deviations in parentheses.

Table 2 Duration analyses for UK bank

Variable	Weibull model 1	Weibull model 2	Weibull model 3	Cox model 1	Cox model 2	Cox model 3	Heckman-Singer model 1	Heckman-Singer model 2
Male	0.2441	0.2509	0.2492	0.2348	0.2383	0.2286	0.2566	0.2951
	(0.0198)	(0.0198)	(0.0200)	(0.0197)	(0.0197)	(0.2003)	(0.0232)	(0.0231)
Married	0.0614	0.1950	-0.0138	0.1011	0.2089	0.01620	0.0250	-0.0829
	(0.0158)	(0.0165)	(0.0176)	(0.0159)	(0.0167)	(0.1770)	(0.0181)	(0.0191)
Male x married	-0.5951	-0.6631	-0.3528	-0.6193	-0.6757	-0.3854	-0.2246	-0.1339
	(0.0261)	(0.0262)	(0.0273)	(0.0262)	(0.0263)	(0.2739)	(0.0291)	(0.0309)
Age	0.0040	-0.1053	-0.1428	0.0075	-0.0771	-0.1141	0.0267	0.0232
	(0.0006)	(0.0036)	(0.0037)	(0.0006)	(0.0037)	(0.0038)	(0.0008)	(0.0008)
Age squared		0.0013	0.0018		0.0011	0.0015		
		(0.0000)	(0.0000)		(0.0000)	(0.0000)		
Training	0.7877	0.6813	0.3260	0.6707	0.59124	0.2392	0.5362	0.4191
	(0.0136)	(0.0141)	(0.0163)	(0.0138)	(0.0142)	(0.0166)	(0.0151)	(0.0169)
Middle	0.0180	0.0870	0.0260	0.0216	0.0749	0.0194	-0.0027	-0.0200
	(0.0186)	(0.0188)	(0.0198)	(0.0187)	(0.0188)	(0.01971)	(0.0203)	(0.0215)
Senior	0.1128	0.1646	0.0945	0.0528	0.0934	0.0379	0.0893	0.0382
	(0.0297)	(0.0298)	(0.0313)	(0.0298)	(0.0299)	(0.0311)	(0.0323)	(0.0358)
Reorganisation	0.4621	0.4640	0.4100	0.4931	0.4971	0.4351	-0.2757	-0.3864
	(.0223)	(0.0224)	(0.0224)	(0.0225)	(0.0225)	(0.0226)	(0.0237)	(0.0223)
Rating	-1.4378	-1.4720	-1.5824	-1.4561	-1.4778	-1.5879	-1.6178	-1.4146
	(0.0930)	(0.0930)	(0.0944)	(0.0931)	(0.0931)	(0.0944)	(0.0971)	(0.1028)
Rating squared	0.4005	0.4217	0.5153	0.4207	0.4342	0.5252	0.4268	0.4004
	(0.0321)	(0.0321)	(0.0323)	(0.0321)	(0.0321)	(0.0323)	(0.0336)	(0.0354)
Lagged rating			-0.1708			-0.1571		-0.2165
			(0.0230)			(0.0235)		(0.0193)
Graduate			0.7263			0.6796		0.3570

			(0.0279)			(0.0278)		(0.0318)
Part-time			0.8596			0.8196		0.4944
			(0.0189)			(0.0189)		(0.0200)
White			-0.1659			-0.1303		-0.1254
			(0.0191)			(0.0191)		(0.0215)
Constant	-1.4958	0.4557903	1.3104				-2.5584	1.5704
	(0.5813)	(0.5848)	(0.5856)				0.0740	(0.0279)
Mass point 2							2.6899	1.5704
							(0.1088)	(0.0279)
Probability of being in type 1							0.9332	0.6132
Number of observations	425629	425629	420405	425629	425629	420405	425629	420405
Log likelihood	-70770.74	-70332.34	-66597.41	-315088.34	-314830.62	-307304.77	-108606.88	-110452.68

Notes: 1.Standard errors in parentheses. 2. The mass point for type 1 observations is normalised to zero so that the intercept for these observations is given by the constant. The intercept for type 2 observations equals the constant plus the coefficient attached to mass point 2.

Table 3 Duration analyses for Greek bank

Variable	Weibull model 1	Weibull model 2	Cox model 1	Cox model 2	Heckman-Singer model
Male	-0.5338	-0.5338	-0.5512	-0.5904	-0.7306
	(0.0499)	(0.0480)	(0.0514)	(0.0507)	(0.0631)
Married	-2.4980	-2.4960	-2.4312	-2.4637	-3.5440
	(0.1118)	(0.1120)	(0.1123)	(0.1124)	(0.1166)
Male x married	0.1258	0.2154	0.0162	0.1485	0.7312
	(0.1318)	(0.1320)	(0.1324)	(0.1327)	(0.1355)
Age	-0.0955	-0.6996	-0.0958	-0.5341	0.0386
	(0.0037)	(0.0235)	(0.0038)	(0.0247)	(0.0033)
Age squared		0.0065		0.0048	
		(0.0002)		(0.0002)	
Training	-0.2053	-0.0784	-0.1186	-0.1472	-0.3151
	(0.2340)	(0.2337)	(0.2342)	(0.2339)	(0.2453)
Middle	-0.9060	-0.8365	-0.9800	-0.8965	-0.2908
	(0.0528)	(0.0542)	(0.0540)	(0.0549)	(0.0608)
Senior	-1.2463	-0.9613	-1.1935	-1.0163	-0.4367
	(0.0464)	(0.0491)	(0.0476)	(0.0497)	(0.0554)
Reorganisation	0.5254	0.5216	0.4076	0.4282	0.3300
	(0.0669)	(0.0668)	(0.0673)	(0.0673)	(0.0683)
Rating	0.1185	0.1888	0.1212	0.1749	0.3404
	(0.0781)	(0.0782)	(0.0780)	(0.0782)	(0.0900)
Rating squared	-0.1061	-0.1357	-0.1070	-0.1282	-0.1983
	(0.0399)	(0.0399)	(0.0399)	(0.0399)	(0.0456)
Number of observations	32260	32260	32260	32260	32260
Constant	-8.1261	3.5353			-6.5325
	(0.1763)	(0.4552)			(0.8162)
Mass point 2					3.9620
					(0.8246)
Probability of being in type 1					0.0757
Log likelihood	-1802.5941	-1511.6845	-19856.4620	-19705.4270	-7295.2244

Notes: See notes to Table 2.